Labor Unions and the Electoral Consequences of Trade Liberalization*

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Abstract

We show that the Brazilian trade liberalization in the early 1990s led to a permanent relative decline in the vote share of left-wing presidential candidates in the regions more affected by the tariff cuts. This happened even though the shock, implemented by a right-wing party, induced a contraction in manufacturing and formal employment in the more affected regions, and despite the left's identification with protectionist policies. To rationalize this response, we consider a new institutional channel for the political effects of trade shocks: the weakening of labor unions. We provide support for this mechanism in two steps. First, we show that union presence—proxied by the number of workers directly employed by unions, by union density, and by the number of union establishments—declined in regions that became more exposed to foreign competition. Second, we show that the negative effect of tariff reductions on the votes for the left was driven exclusively by political parties with historical links to unions. Furthermore, the impact of the trade liberalization on the vote share of these parties was significant only in regions that had unions operating before the reform. These findings are consistent with the hypothesis that tariff cuts reduced the vote share of the left partly through the weakening of labor unions. This institutional channel is fundamentally different from the individual-level responses, motivated by economic or identitarian concerns, that have been considered in the literature.

Keywords: Trade Shocks; Elections; Unions; Brazil

JEL classification: F13, D72, J51, F16, F14

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1 Introduction

The political consequences of increased exposure to international trade have become a fiercely debated topic. This issue has gained particular salience in recent years, after Brexit, the election of Donald Trump, and the ongoing trade war between the U.S. and China. Both the economics and political science literature have linked trade shocks, depending on the specific context, to reduced incumbent advantage, to gains for parties identified with the labor movement, and to political radicalization (e.g., Autor et al., 2020; Blanchard et al., 2019; Che et al., 2020; Choi et al., 2021; Dippel et al., 2021; Jensen et al., 2017). These responses are typically interpreted as driven by individual-level reactions, motivated either by economic or identitarian concerns.

In this paper, we show that the Brazilian trade reform of the early 1990s had political implications that do not fit the pattern described above. Specifically, exploiting variation in the initial employment structure across local labor markets, we show that, from 1994 to 2018, regions that faced relatively larger tariff reductions also experienced a persistent relative *reduction* in the support for left-wing presidential candidates.

This result is puzzling for a number of reasons. First, the reform was implemented by a right-wing party, clearly identified as such by the electorate. Since we know from the previous literature that these regions also experienced a deterioration in labor market conditions (Dix-Carneiro and Kovak, 2017, 2019; Kovak, 2013), retrospective voting to punish the incumbent should have led instead to an increase in the vote share of the left. Second, the left in Brazil is historically identified with protectionist stances (Colistete, 2007), so an ideological reaction to the reform should have increased the support for the left. Third, this movement was not associated with increased radicalization. It took place throughout the 1990s and early 2000s, much before the recent wave of extreme-right populism, and was mostly related to a movement towards the center of the political spectrum. And fourth, our sample covers a period when overall voting patterns in Brazil shifted towards the left, with four consecutive presidential elections won by the Labor Party (Partido dos Trabalhadores, PT). Nevertheless, we find that, in relative terms, regions more exposed to increased foreign competition moved towards the center-right.

We argue that this unexpected, long-lasting electoral response was at least partially due to an institutional implication of the trade reform. In several countries, unions tend to be stronger within the manufacturing sector (Visser, 2019). This is particularly true in Brazil, where formal jobs are much more common in the manufacturing sector and formality is intrinsically linked to union strength. Until very recently, the main source of revenue for labor unions in the country was a compulsory annual contribution, equivalent to one working day, for *every* formal worker in the union's regional jurisdiction and category, independently of whether the worker was formally affiliated with the union. Since the trade reform had particularly large impacts on manufacturing formal jobs (Dix-Carneiro and Kovak, 2019; Ponczek and Ulyssea, 2021), one should also expect it to have had a substantial impact on unions' revenues.

As argued by the political science literature, unions can affect election outcomes in various ways: by exerting influence on their members, mobilizing get-out-the-vote efforts, affecting workers' political views, raising campaign contributions, etc. (see, for example, Kim and Margalit, 2017). Their weakening could therefore help to explain changing voting patterns due to the trade reform. In short, since the liberalization in Brazil reduced formal employment in harder-hit regions relative to the country's average, labor organizations in these regions could have lost their financial and organizational capacity to influence elections in favor of their preferred candidates, who are typically from the left.

We investigate this relationship in two steps. First, we show that unions were indeed weakened in regions that became more exposed to foreign competition. We measure "union strength" in three different ways: with the number of union employees, with the number of union establishments, and with a proxy for "union density." Results are similar for all three measures. Second, we present evidence that the relationship between tariff cuts and votes for the left was driven exclusively by candidates from political parties linked to unions. These parties are often, but not always, from the left. Furthermore, the effect of the tariff cuts on the vote share of left-wing parties and parties linked to unions was significant only in regions that had some union presence before the reform started. Altogether, these findings are consistent with the hypothesis that the weakening of labor unions was an important determinant of the negative relationship between regional exposure to the tariff cuts and the vote share of the left.

The magnitude of the relationship we uncover is sizable. A one-standard-deviation increase in the tariff reduction is associated with an average reduction of approximately 4 percentage points in the share of votes for left-wing candidates in the first round of the post-liberalization presidential elections, when compared to the 1989 election. Interestingly, this effect is relatively stable over the entire sample period, indicating that the tariff cuts had a one-time, permanent impact on the electorate. The effect on union strength is also sizable. For example, a one-standard-deviation increase in the tariff reduction would lead to an average reduction in union density in the post-shock years equivalent to 38% of the 1989 median, which corresponded to 19 unionized workers per 1,000 inhabitants (between ages 15 and 64).

As indicated above, we follow the literature on trade and local labor markets and use Brazil's trade reform from the early 1990s as a natural experiment. During the first half of the 1990s, the Brazilian federal government reduced import tariffs of different sectors to varying degrees, and tariffs remained roughly constant afterward. As in Kovak (2013) and several subsequent analyses, we use the heterogeneity in the pre-shock industry mix across regions and the variation in tariff changes across sectors to construct a measure for local labor demand shocks induced by the trade liberalization, which we refer to as *regional tariff reductions*. This literature has shown that regions where the most important industries faced larger tariff cuts experienced relative deterioration in labor market conditions that lasted, in one way or another, for decades (Dix-Carneiro and Kovak 2017, 2019; Dix-Carneiro et al. 2018; Kovak 2013; Ponczek and Ulyssea 2021). The local variation in economic conditions induced by the liberalization represents the negative regional shock that we exploit in this paper.

There are four main reasons why the Brazilian tariff reduction from the 1990s provides a good setting to study the political consequences of trade liberalization. First, the reform was implemented in a centralized way by the federal government, so it can be interpreted as a source of exogenous variation in local economic conditions that, in principle, can be used to identify the causal effects of reductions in trade barriers. Second, as the previous literature has made clear, there is no indication that economic and political-economy factors influenced the design of the reform, which was aimed primarily at reducing, simultaneously, protection levels and the dispersion of tariffs across industries. Third, unlike in many other cases of trade liberalization, the magnitudes of the tariff reductions were substantial: the average import tariff fell from 30.5 percent in 1990 to 12.8 percent in 1995. Fourth, the shock was close to a once-and-for-all event, which makes it possible to identify the cumulative effects of the trade reform on local economies over time. Due to these features, it is not surprising that several authors have used the Brazilian trade liberalization episode to evaluate the impacts of trade shocks on various economic outcomes.[1] Nevertheless, we are the first to investigate the electoral consequences of the shock, as well as its impact on labor unions.

Recent studies, however, have investigated the effects of trade exposure on voting patterns and political outcomes in other contexts. Like us, Che et al. (2020) evaluate the short- and long-run political consequences of a trade shock, in their case the U.S. granting of permanent normal trade relations to China in October 2000. They find that U.S. counties that faced larger exposure to Chinese imports experienced a relative increase in the support for Democratic candidates in American congressional elections until 2010, although not afterwards. Their study supports the idea that more exposure to trade helps parties identified with more protectionist policies—the Democrats during the 2000s, but not afterwards. Autor et al. (2020), on the other hand, find opposite results when analyzing the effects of increased exposure to Chinese imports on U.S. presidential elections. In addition, they find that trade exposure induces more ideological polarization. Dippel et al. (2021) investigate the same relationship, but using German elections, and find that extreme-right candidates benefit from rising trade exposure to China and Eastern Europe. Blanchard et al. (2019) study instead the effects of Trump's trade war, which increased trade barriers, on the American 2018 congressional elections. They find that Republican candidates lost support in the more exposed regions relative to the less affected ones, indicating that, in this case, voters punished the party responsible for implementing the policy. Choi et al. (2021) evaluate the political consequences of NAFTA and reach a similar conclusion: American regions that lost more jobs because of the trade agreement tended to switch from Democrat (who enacted NAFTA) to Republican representatives, as the electorate stopped associating Democrats with trade protection. In turn, Jensen et al. (2017) show that U.S. voters in regions benefiting from increased exports tend to reward the incumbent president's party. Similarly, and despite the very different political context, Campante et al. (2019) show that Chinese prefectures

[1] In addition to the papers already mentioned, see Casagrande and Hidalgo (2019), Costa et al. (2019), Goldberg and Pavcnik (2003), Gonzaga et al. (2006), Hirata and Soares (2020), and Menezes-Filho and Muendler (2011).

hit particularly hard by the 2015-2016 reduction in exports experienced increased political unrest and also increased likelihood of replacement of the local party secretary.

We contribute to this literature in three ways. First, and most importantly, we identify a new institutional channel linking trade exposure to votes: union strength. This mechanism linking trade shocks to political outcomes is still unexplored in the literature and is likely to be relevant in other contexts as well. Second, we provide the first analysis of the relationship between trade shocks and elections in the context of a developing country, where, according to standard trade theories, the effects of trade liberalization can be very different from those observed in developed economies. And third, we use a discrete and well-defined trade policy event, which allows us to observe both its short- and long-run effects on electoral outcomes.

In exploring how the trade shock affected future electoral outcomes through its impact on labor unions, we were inspired by the wider literature on trade and institutions, reviewed by Nunn and Trefler (2014). More specifically, there is an established literature that connects protection levels and the strength of labor unions, but focusing on the reverse effect. A prominent example is Matschke and Sherlund (2006). Starting from the observation that labor unions generally support protectionist stances, they extend Grossman and Helpman (1994)'s "Protection for Sale" model of endogenous trade protection to include lobbying by unions. Testing the predictions of the model using data from U.S. manufacturing, they find that incorporating union lobbying is essential to explain the structure of protection in the U.S. Most reduced-form studies find as well that labor union contributions and lobbying activities are positively correlated with protection levels.[2] Our finding that tariff cuts negatively affect union strength highlights that this line of research needs to pay some attention to reverse causality.[3]

The remainder of the paper is organized as follows. Section 2 describes the institutional context of the Brazilian trade liberalization process and its union organization system. Section 3 summarizes the data and provides descriptive statistics. Section 4 presents the empirical strategy. Section 5 shows the results on election outcomes and associated robustness exercises. Section 6 investigates the weakening of labor unions as a potential mechanism behind the relationship between tariff reductions and electoral outcomes. Section 7 concludes.

[2] For example, Baldwin and Magee (2000) and Conconi et al. (2014) find that union contributions are associated with a higher probability that a U.S. Congress representative votes against trade liberalization bills. An exception is Lake (2015), who, using a detailed dataset linking campaign contributions and lobbying expenditures to specific issues and representatives, does not find support for such a relationship.

[3] There is also a small literature that studies how trade exposure affects labor market outcomes mediated by unions. For example, Gaston and Trefler (1995) find evidence suggesting that union wage premia are sensitive to import competition in American manufacturing, while Baumgarten and Lehwald (2019) find that greater import exposure decreases the probability that German plants participate in industry-wide bargaining agreements.

2 Institutional Context

2.1 The Trade Policy Shock

For almost a hundred years, Brazil pursued a policy of import substitution industrialization, which sought to shield the national economy from foreign competition. This policy ended, rather abruptly, with the implementation of a unilateral trade liberalization program in 1991. Until then, while import duties were very high, the main source of import protection was a wide variety of non-tariff barriers, such as lists of banned products, quantity controls and government procurement restrictions. The protection level effectively faced by a given industry was significantly higher than nominal tariffs alone suggested, leading to a nontransparent structure of protection (see Abreu, 2004, and Kume et al., 2003).

In 1990, the newly elected President, Fernando Collor de Melo, announced a trade reform agenda aimed at reducing the import barriers faced by the Brazilian industry and increasing transparency in trade policy. Trade liberalization started in March 1990 with the abolition of import quotas and the elimination of some administrative import controls (Mérette, 2000). Those non-tariff barriers were replaced by import tariffs that were adjusted to reflect the same level of protection, as measured by the gap between international and local prices (Carvalho, 1992). Hence, tariffs became the principal trade policy instrument and started to accurately reflect the level of protection faced by the national industry in Brazil. The average ad valorem nominal tariff was 30.5%, but varied substantially across industries, reaching 78.7% in the apparel industry (Kume et al., 2003).

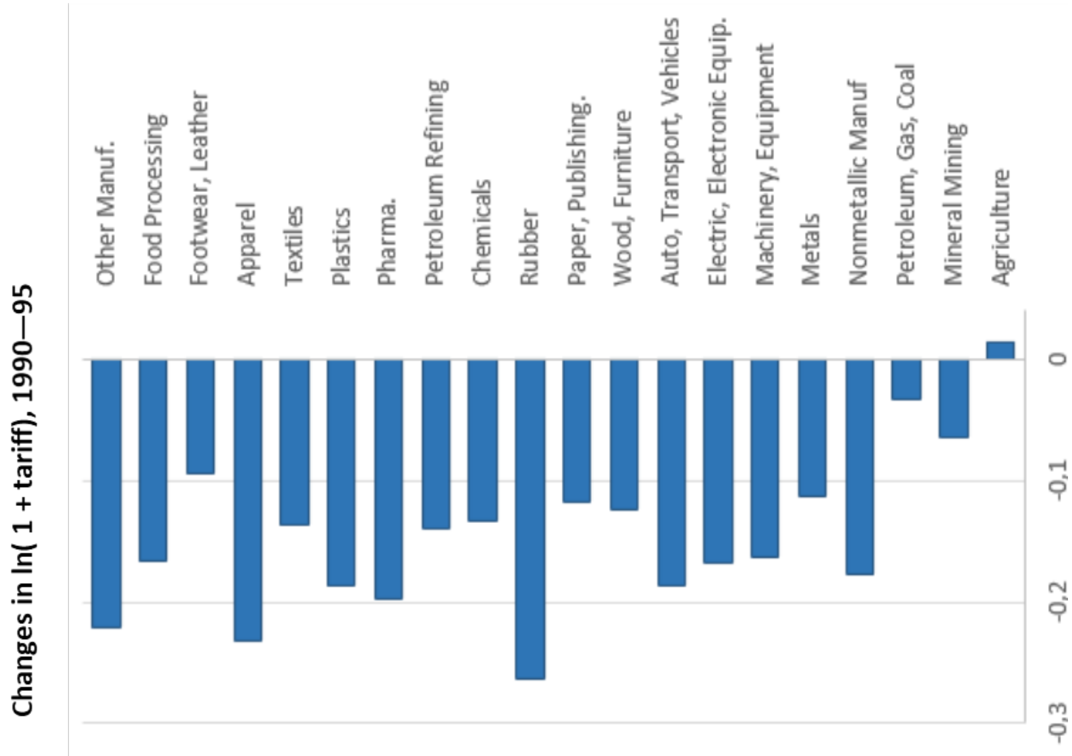
Between 1991 and 1995, Brazil then gradually but significantly reduced its import duties. At the end of this process, the average nominal tariff was 12.8 percent, far below pre-liberalization levels and consistent with the pattern of other developing economies. We measure the extent of liberalization at the industry level using the 1990-1995 difference in the log of one plus the tariff rate. Under the assumption of a small open economy, this would be equal to the change in the local price induced by the trade shock. It is worth noting that there is no novelty in this approach: this is the same measure used in the previous literature on the regional effects of trade liberalization in Brazil.[4]

Figure 1 plots tariff percentage reductions by sector from 1990 to 1995. Clearly, the liberalization affected different sectors by different degrees. For example, the mineral mining and petroleum industries faced small tariff reductions, whereas the rubber and apparel industries experienced large ones; in agriculture, the average tariff actually increased, albeit only slightly. As further discussed in section 4, this cross-sectoral heterogeneity in tariff cuts is central for our identification strategy.

A key characteristic of the Brazilian trade liberalization process of the 1990s is that

[4] See, for example, Costa et al. (2019), Dix-Carneiro and Kovak (2017), Dix-Carneiro et al. (2018) and Kovak (2013).

Figure 1: Changes in ln(1 + tariff), 1990-95



Notes: Changes in the log of one plus the nominal tariffs between 1990 and 1995, by sector.

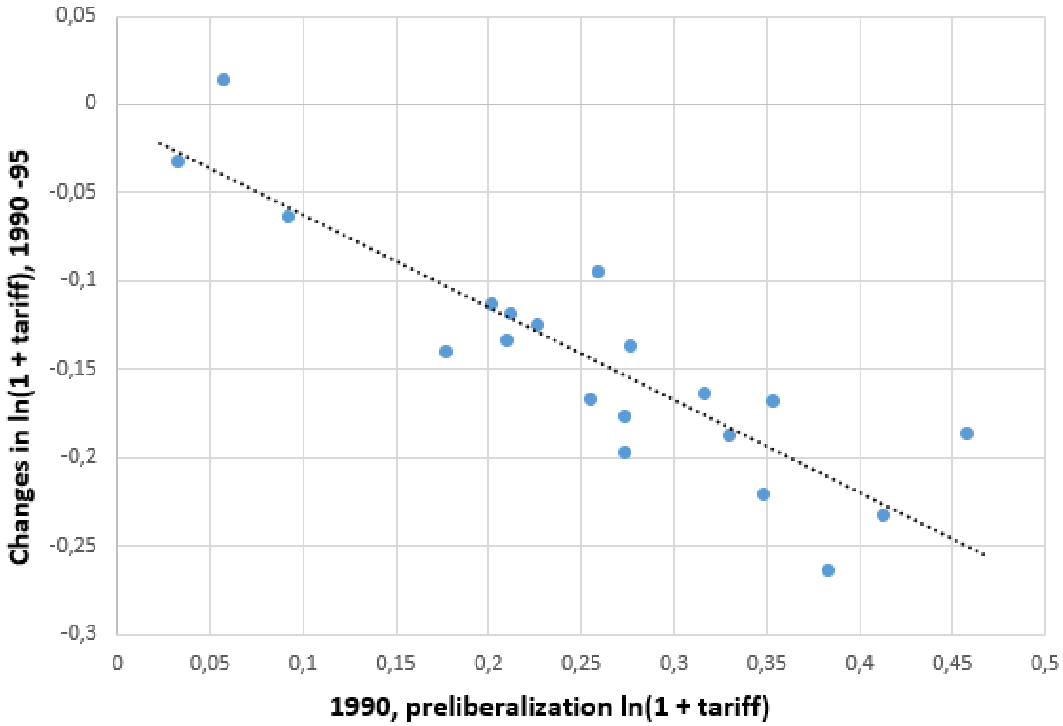
tariff cuts were strongly correlated with the initial tariff level. As Figure 2 shows, the sectors with high tariffs before liberalization experienced the greatest cuts. Since pre-liberalization tariff levels reflected a protection structure imposed more than 30 years before the tariff cuts started (Kume et al., 2003), the trade reform does not appear to have been influenced by previous industry performance or other contemporaneous political economy issues. Rather, it had the purpose of simultaneously reducing protection levels and the dispersion of tariffs.[5]

Another characteristic of the Brazilian trade shock, which is particularly useful for the identification of long-run effects, is that trade policy in the country after 1995 has remained largely unchanged.[6] Multilateral import tariffs are roughly at the level they were set in 1995. Participation in free trade agreements beyond Mercosur (discussed in subsection 5.1) has been very limited. Use of antidumping and other countervailing measures has fluctuated over time, but affected only a small share of the country's trade volume. Hence, at least until 2018, one can reasonably consider the trade policy shock from 1991 to 1995 as "permanent."

[5] This argument was first proposed by Goldberg and Pavcnik (2005) to show that Colombian tariff reductions in the period of 1983–1998 were exogenously predetermined by its initial levels, and has been widely used for the Brazilian liberalization episode. See Costa et al. (2019), Dix-Carneiro and Kovak (2017), Dix-Carneiro and Kovak (2019), Kovak (2013), and Pavcnik et al. (2004).

[6] See Ferraz et al. (2020) for a detailed account of the trade policy landscape in Brazil since the 1990s.

Figure 2: Tariff Changes and Pre-liberalization Tariffs



Notes: Correlation between the changes in the log of one plus the nominal tariff between 1990 and 1995 and the log of one plus the nominal tariff in 1990. Correlation : -0.881.

2.2 Labor Unions in Brazil

Labor unions in Brazil were institutionalized in 1931 and have since played an important role in domestic politics. Historically, they have been associated with left-wing parties. This relationship became more salient in the end of the military regime in the 1980s, when many union leaders started political careers in left-wing parties (Coradini, 2007). During this same period, the main national union confederations were created by union leaders affiliated to left-wing parties.[7] Nowadays, they are still led by politicians connected to the left. In 2013, more than 75% of the labor unions in the country were affiliated to these union confederations (Cardoso, 2014), indicating proximity to left-wing parties.

Importantly, labor unions in Brazil represent not only the workers formally affiliated with them, but all workers in their occupational category and geographic jurisdiction. A consequence of this structure is that, until 2017, whenever an occupational category was formally represented by a union in a municipality, all workers in that category, unionized or

[7]Union confederations are national associations of unions that help unions in bargaining processes and coordinate political actions across unions. The most traditional and politically strongest confederations are CUT ("Central Única dos Trabalhadores"), created by union leaders affiliated to PT (the Labor Party) in 1983, and UGT ("União Geral dos Trabalhadores"), formerly CGT ("Central Geral dos Trabalhadores"), founded by union leaders from PCB (the Brazilian Communist Party) in 1986.

not, had to contribute to the union with a compulsory annual payment equivalent to one working day. This compulsory contribution—called the "union contribution"—has been a key source of revenue for labor unions in the country.

Despite this compulsory payment, formal affiliation was still fundamental for political action, since unionized workers influence mobilization, strikes, and collective agreements led by unions. For this reason, affiliation to unions is considered an indicator of proximity between workers and unions (Campos, 2016).

During our period of analysis, the number of labor union establishments in Brazil increased substantially, from less than 7,500 in 1995 to around 12,000 in 2017. Despite this increase, the share of workers affiliated with unions fell from 26% in 1995 to 20% in 2015 (Instituto Brasileiro de Geografia e Estatísitca, 1985-2014).

3 Data

Following the previous literature on the regional effects of trade liberalization in Brazil, we carry out the analysis at the microregional level, defined as a group of economically integrated contiguous municipalities with similar geographic and productive characteristics (Instituto Brasileiro de Geografia e Estatísitca, 2002). Since some boundaries have changed during the period of analysis, we aggregate minimum comparable areas, as described in Reis et al. (2011), to construct identifiable microregions over time.[8] This process yields 484 microregions.[9] In the next subsections, we discuss the classification of parties in the right-left political spectrum and describe the different sources of data.

3.1 Party Classification in the Left-Right Political Spectrum

We classify all political parties that ran in at least one presidential election between 1989 and 2018 into two groups: left wing and non left wing. A party is classified as left wing if it is defined as communist, socialist, or left wing in its official sources. When the classification is not clear and official sources do not contain enough information, we rely on the party's characterization in the mainstream media. The resulting classification is described in Table A1 (Appendix A).

[8] From 1991 to 2015, more than one thousand municipalities were founded in Brazil, representing an increase of 24%. Following other papers that use the same unit of analysis, we omit the microregions of Manaus, which is a Free Trade Zone, and Fernando de Noronha, which is a tiny archipelago representing less than 0.0014% of the Brazilian population.

[9] We also estimate our results using a more aggregated definition that is consistent over a longer period. Using this more aggregated definition, we are left with 411 microregions. The main results are similar when using this alternative aggregation.

Our classification is consistent with others from the political science and economics literature. Using the methodology proposed by Rodrigues (2002), which classifies political parties based on the characterization in the media and the view of political scientist, Sakurai and Menezes-Filho (2008) present a classification very similar to ours, except that we consider a larger number of parties. Power and Zucco Jr (2009) estimate the position of some Brazilian political parties on the left-right spectrum using survey responses from more than 850 federal legislators on their political ideology and on the ideological position of parties, from 1990 to 2005. Their ordering is also consistent with our classification. Figueiredo and Limongi (1999) evaluate the pattern of votes in the Brazilian congress between 1989 and 1994 and classify parties in three groups—right, center, and left—based on the similarity in their voting patterns. Their classification is also consistent with ours.[10] Nevertheless, when discussing the robustness of our main results (subsection 5.1), we consider some plausible variations in this definition.

An important observation about the proposed classification is that the Party of the Brazilian Social Democracy (Partido da Social Democracia Brasileira, PSDB) is considered center-right in our main definition. In the literature, there is some debate on how PSDB should be classified in its initial years. In the 1989 presidential election, soon after its foundation, PSDB supported PT in the election runoff, which indicates certain proximity in ideological terms. However, after 1993, PSDB took a turn to the right and became more clearly identified with the center-right (Power and Zucco Jr, 2009). For methodological consistency, and to be conservative, we classify it as center-right throughout the entire period. Yet we note that, if we classify PSDB as left wing between 1989 and 1993 and center-right after 1994, our qualitative results do not change and point estimates become larger (we present results using this alternative classification in subsection 5.1).

Finally, notice that this discussion is relevant only for the first round of elections. In the runoffs, there was always a clearly defined party on the left (PT) running against a non-left party.

3.2 Elections Data

We use data on presidential elections spanning the period from 1989 to 2018. We focus on presidential elections for two main reasons. First, all trade policy decisions in Brazil are responsibility of the federal executive branch.[11] Therefore, it is more appropriate to study the relationship between policies that lower barriers to trade and voters' behavior in presidential elections. Second, there is no consistent data available for state or local elections before the Brazilian liberalization. Hence, it is only possible to conduct this type of analysis using data from the presidential elections.

[10] It is important to emphasize that we classify a larger number of parties than all of these other studies.

[11] In the 1980s, the Foreign Trade Department of the Bank of Brazil, which was subordinate to the federal government, was the public body that formulated and implemented trade policies. In the 1990s, this function was transferred to the Foreign Trade Chamber, which was also subordinate to the federal government.

Table 1: Descriptive Statistics – Elections

Variables	1989	1994	1998	2002	2006	2010	2014	2018
Eligible Voters (Thousand)	81,468.64	94,034.36	105,210.51	114,233.28	124,708.91	134,333.57	141,080.13	145,311.32
V.Sh Left-Wing Parties (First Round)	35.25%	30.27%	43.64%	76.70%	57.80%	66.98%	65.14%	42.50%
V.Sh Left-Wing Parties (Runoff Elections)	47.03%	N/A	N/A	61.15%	60.58%	55.84%	51.61%	44.98%
Turnout (First Round)	88.12%	82.29%	78.53%	82.26%	83.24%	81.86%	80.66%	79.74%
Turnout (Runoff Elections)	85.66%	N/A	N/A	79.56%	81.01%	78.51%	78.96%	78.77%

Notes: This table displays the number of eligible voters, the vote share for the left-wing parties in the first round of voting and runoff elections, and the turnout rate in both rounds of voting in each presidential election in Brazil between 1989 and 2018.

Electoral data come from the Brazilian Superior Electoral Court (TSE), which reports municipality level data for a wide variety of electoral outcomes for the first and second rounds of voting for every presidential election after the transition to democracy, starting in 1989. The TSE electoral data for the 1989 and 1994 Brazilian elections are available for download at the Institute of Applied Economic Research website.[12] For the elections after 1994, data are available from the TSE website.[13] The electoral outcome used in this paper is based on the number of votes cast for each candidate in each election. We aggregate them over all left and non-left political parties to obtain the vote share of the left in each election. We focus on the first round of voting in each election because there were no runoff elections in 1994 and 1998, when the PSDB candidate, Fernando Henrique Cardoso, received more than 50% of the valid votes in the first round. But we also report results using the runoff data for the other election years.

Table 1 shows the number of eligible voters, the vote share for left-wing parties, and the turnout rate for each presidential election between 1989 and 2018. The descriptive statistics indicate that the vote share for left-wing candidates increased substantially in the 2002 election and, although falling a bit in subsequent elections, remained above 50% up to 2014. The table also shows that turnout fluctuates slightly over time, with no clear trend.[14]

[12] Elections data can be found at `http://www.ipeadata.gov.br/Default.aspx`. We do not have data on the votes cast for PMB in the 1989 elections. However, PMB received only 4,363 votes in the whole country, corresponding to less than 0.007% of the valid votes in 1989.

[13] Election data for elections after 1994 can be found at `TSE.jus.br`.

[14] Turnout may look exceedingly high, but notice that voting in Brazil is compulsory for literate citizens aged between 18 and 70, and is optional only for illiterates and citizens aged between 16 and 18 or above 70. Voters who fail to vote must justify their absence or pay a fine. Otherwise, they become ineligible to register for civil service entrance examinations, to have passports or ID cards issued, to renew registration at federal educational establishments, to borrow from public banks, and to participate in public or administrative bidding.

3.3 Tariff Data

Tariff data come from Kume et al. (2003), who report effective rates of protection and nominal tariffs at the *Nível 50* Brazilian industry classification level (similar to the two-digit SIC) from 1987 to 1998. We use nominal tariffs to measure the protection level in each industry, but the results are very similar using effective rates of protection.[15] As in Dix-Carneiro and Kovak (2017), we aggregate tariffs to have an industry classification consistent with the 1991 Demographic Census.[16]

3.4 Other Variables

We use individual-level data from the 1991 Brazilian demographic census (Instituto Brasileiro de Geografia e Estatísitca, 1991) to compute several variables of interest. First, we calculate the share of employees by sector of economic activity. We also calculate per capita household inequality (Gini index) and the share of the population aged over 60, the share of females, the share of whites, the share living in urban areas, the share of employment in manufacturing, the employment rate, and the share of adult population with complete high school.

4 Empirical Strategy

The empirical analysis follows the literature on the regional labor market effects of trade, which considers that regions whose most important industries faced larger tariff reductions experienced larger reductions in labor demand. Thus, although the tariff changes are the same across all regions within the country, the heterogeneity in the regional industry mix, together with the fact that tariff shocks affected industries in different degrees, allows us to measure regional trade shocks with the combination of sector-specific trade policy shocks and sectoral composition.

This idea was formalized within the framework of the specific-factors model for regional economies by Kovak (2013). We follow his approach and measure the trade shock as the "regional tariff reduction" (RTR) in region r. The RTR_r is calculated as the weighted average of import tariff reductions faced by the industries in region r, where the weights are

[15] Effective rates of protection take into account the input tariffs, measuring the difference in the value-added per unit of output with the protection structure relative to the free-trade scenario.

[16] Table A2 (Appendix A) provides the classification and matching between tariffs and census industries. For more information, see Dix-Carneiro and Kovak (2017).

given by the relevance of each industry in region r.[17] Formally:

$$RTR_r = -\sum_{i\in I} \beta_{ri}\, \Delta ln(1 + \text{tariff}_i),$$

where

$$\beta_{ri} \equiv \frac{\lambda_{ri}\frac{1}{\phi_i}}{\sum_{j\in I}\lambda_{rj}\frac{1}{\phi_j}},$$

the operator Δ represents the long difference from 1990 to 1995, λ_{ri} is the initial share of region r workers employed in industry i, and ϕ_i is the cost share of nonlabor factors, calculated as one minus the wage bill share of industry i using 1990 national accounts data from the Brazilian Institute of Geography and Statistics (IBGE).

Figure 3, Panel A, shows the distribution of RTR_r across regions and some descriptive statistics. Darker regions are those facing larger tariff reductions, while the less affected regions are shown as lighter. The image shows a large degree of heterogeneity in the RTR_r across regions, even within the same state. The heterogeneity within states is important for our identification, since we estimate our model in differences and include state fixed effects in the analysis (which is equivalent to including state-specific time trends in a panel setting).

Figure 3, Panel B, shows the changes from 1989 to 2018 in the vote share for left-wing candidates. Regions facing larger declines (or more commonly, smaller increases) in the vote share of the left are presented as darker. Comparing panels A and B, one notices that, in general, the South, Southeast and Midwest regions faced both larger tariff cuts and reductions (or smaller increases) in the vote share for the parties from the left. This suggests a negative relationship between changes in trade exposure and in the political support for the left.

To compare the evolution of election outcomes in regions facing larger regional tariff cuts to those in regions facing smaller tariff cuts, we estimate, for each presidential election year t between 1994 and 2018, the following regression:

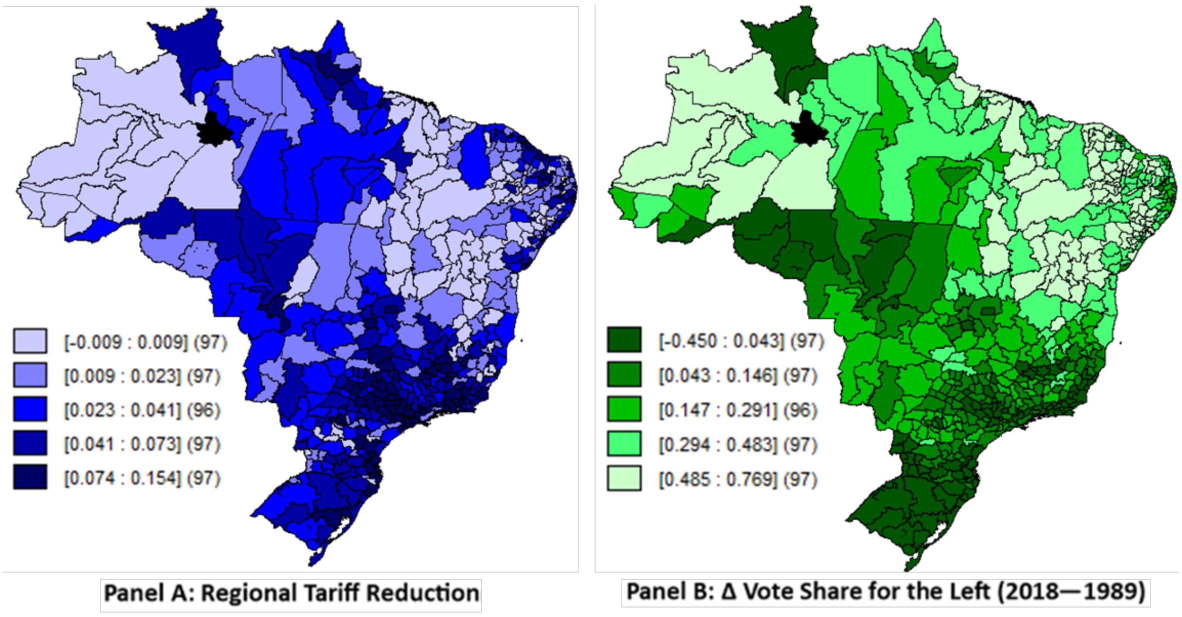
$$y_{rt} - y_{r,1989} = \theta_t RTR_r + \alpha_{st} + \psi_t X_r + \epsilon_{rt}, \tag{1}$$

where y_{rt} is an outcome in region r, such as the vote share for left-wing parties; θ_t is the cumulative effect of the liberalization on the outcome variable; α_{st} is a state fixed effect (which is allowed to vary over the years); X_r is a vector of pre-liberalization socio-demographic controls discussed in further detail below; and ϵ_{rt} is an error term. To account for potential spatial correlation in outcomes across neighboring regions, we cluster standard errors at the mesoregion level.[18]

[17] As in Kovak (2013), we drop the nontradable sector based on the result that its price in a particular region follows the price of locally produced tradable goods. Dix-Carneiro and Kovak (2017) provide empirical support for this result in the context of the Brazilian trade shock of the 1990s.

[18] Mesoregions are groups of microregions with similar characteristics defined by IBGE, numbering 114. Adao et al. (2019) propose an alternative approach to estimate the standard errors in shift-share regression designs. However, their method assumes that the number of industries is large, which is not true in our case (we have 20 tradable industries).

Figure 3: Regional Tariff Reduction and Changes in the Vote Share of the Left



Mean	SD	P10	P90
0.043	0.039	0.002	0.106

Mean	SD	P10	P90
0.230	0.252	-0.101	0.558

Notes: Panel A displays the distribution of regional tariff reductions across microregions. Panel B shows the distribution of the changes between 1989 and 2018 in the vote share for the left-wing parties. Both panels also display the mean, standard deviations and the 10^{th}, and 90^{th} percentiles for each variable.

State fixed effects are included for two reasons. First, Brazilian states differ in cultural and economic characteristics.[19] Second, since Brazil is a federal republic, many public security, health and educational policies are delegated to state governments. As both cultural and economic characteristics influence elections, state fixed effects provide a better treatment-control comparison and a more transparent analysis. We also control for pre-liberalization (1991) socio-demographic characteristics that are found to be important in the economics and political science literature on voting. These variables capture socioeconomic conditions that are correlated with the initial employment structure (used to construct RTR_r) and that could also potentially influence the evolution of political and institutional outcomes: shares of whites, females, high-school graduates, and individuals aged over 60 in the population, share of workers employed in manufacturing, employment and urbanization rates, and the Gini inequality index.[20] Notice that, since we are running our regressions in differences, this is equivalent to including interactions of the initial values of these variables with time dum-

[19]For example, in 2018, Brazil's most affluent state, São Paulo, accounted for more than a third of the national GDP and had a per capita income more than three times higher than that of the poorest federal unit, Maranhão.

[20]These variables are calculated with the 1991 census. Many other studies in the literature that explore the link between trade shocks and electoral outcomes include similar variables (Autor et al., 2020; Che et al., 2020; Dippel et al., 2021).

mies in a panel data setting. That is, baseline covariates control for initial characteristics of regions that might be correlated with the *evolution* of voters' behavior over time. Similarly, our state fixed effects can be interpreted as capturing state-specific time trends.

It is important to emphasize that the variable RTR_r does not vary over time, since it reflects the regional tariff reductions that occurred during the Brazilian trade liberalization process. For this reason, one can interpret the coefficient θ_t as the cumulative effect of tariff changes on electoral outcomes over time. This characterization is only possible because the liberalization was a one-time, discrete, and permanent shock, with tariffs being reduced between 1990 and 1995 and remaining stable afterwards.

Specifically, our identifying assumption is that, conditional on our set of controls, RTR_r is orthogonal to the local political and institutional dynamics across microregions. The discussion on the institutional background of the Brazilian trade reform in section 2.1 makes the case that the reductions in tariffs seemed to be orthogonal to local institutional and political dynamics. The pre-intervention placebo tests, which we run whenever data availability allows, provide further evidence supporting this assumption.[21] As for the potential inference issues discussed by Adao et al. (2019) in shift-share designs in general, Ferman (2021) shows that, in settings like ours, their proposed inference method leads to large over-rejection due to the relatively small number of industries. The pre-intervention placebo tests that we present in section 6 suggest that clustering standard errors at the mesoregion level is enough to capture the correlation across error terms. In other words, over-rejection of the null hypothesis does not seem to be a problem in our specification.

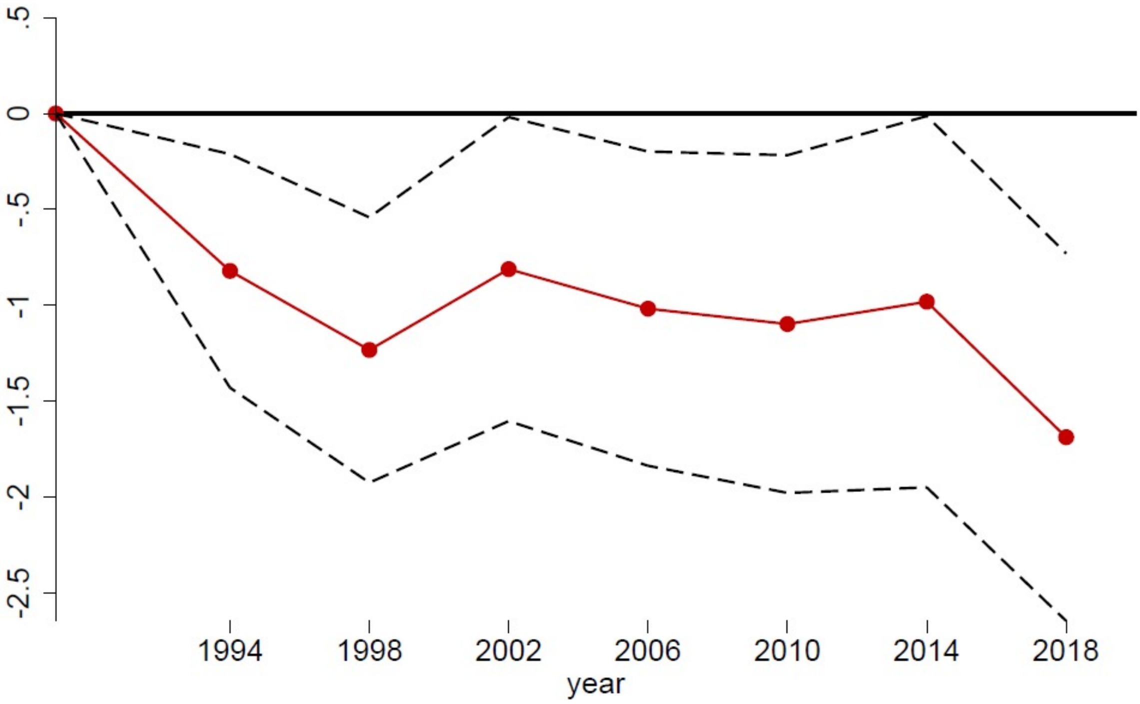
5 Trade Liberalization and the Vote Share for the Left

Figure 4 plots the estimates of $\hat{\theta}_t$ from our main specification in equation (1), where the dependent variable is the vote share for left-wing parties, together with 95% confidence intervals. We find negative and statistically significant estimates for the tariff cuts in every election after the policy shock. The results indicate that, after the trade liberalization, candidates from left-wing parties experienced larger reductions (or smaller increases) in their vote share in the most affected regions, when compared to the less affected ones. Moreover, the coefficients are fairly stable over time, implying that the tariff cuts affected voting patterns permanently.

These estimates are reported in Panel A of Table 2. The results reveal a considerable impact of the regional tariff reduction on the vote share for left-wing candidates. According to the estimates in Panel A, a one-standard-deviation increase in RTR_r is associated with an average reduction of approximately 4 percentage points in the share of votes for the

[21] Since we consider a linear shift-share design in which the exposure shares sum to one in all microregions, the identifications issues raised by Borusyak and Hull (2021) do not apply to our setting (see also Borusyak et al., 2021).

Figure 4: Impact of the Tariff Cuts on the Vote Share for Left-wing Candidates



Notes: Each dot represents the estimated coefficient $\hat{\theta}_t$ reported in Panel A of Table 2. Negative estimates imply fewer votes for left-wing candidates in regions facing larger tariff reductions. All regressions include state fixed effects and pre-liberalization socio-demographic controls. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 114 mesoregion clusters.

left, relative to the 1989 election. Moving a region from the 10^{th} to the 90^{th} percentile of the distribution, in turn, is associated with an average reduction of just over 11 percentage points. To get the average effect over the entire period, in column (8) we estimate a pooled OLS regression using observations for all years and controlling for year fixed effects.

In Panel B, we use an alternative measure for the local labor demand shock induced by the trade liberalization. This alternative measure is similar to the regional tariff reduction, RTR_r, but is constructed using changes in effective rates of protection instead of nominal tariffs. All estimates from this specification are once again negative and statistically significant, indicating that the main results are robust to this alternative definition. The magnitudes of the coefficients are smaller than the ones from Panel A, but the reason is that the scale of the treatment variable is different. Quantitatively, the estimated average effects are similar to the ones in the first panel. As before, moving a region from the 10^{th} to the 90^{th} percentile of the alternative measure of regional tariff reduction would also imply a reduction of about 11 percentage points in the vote share for the left. This was expected, since the magnitudes of the changes in effective rates of protection during the trade reform were larger than for nominal tariffs, but reflected similar information.

In Panel C, we use the same specification as in Panel A. However, the dependent

Table 2: Regional Tariff Reduction and the Vote Share of the Left

Dep var: Δ Vote share for left-wing parties	(1) $t = 1994$	(2) $t = 1998$	(3) $t = 2002$	(4) $t = 2006$	(5) $t = 2010$	(6) $t = 2014$	(7) $t = 2018$	(8) Pooled
A. Main Specification								
Regional Tariff Reduction (RTR_r)	-0.829***	-1.243***	-0.811**	-1.032**	-1.103**	-0.986**	-1.733***	-1.105***
	(0.307)	(0.347)	(0.398)	(0.411)	(0.441)	(0.487)	(0.485)	(0.359)
B. Effective Protection								
Regional Effective Tariff Reduction ($RETR_r$)	-0.480***	-0.662***	-0.494**	-0.675***	-0.756***	-0.686**	-1.072***	-0.689***
	(0.181)	(0.205)	(0.248)	(0.253)	(0.265)	(0.270)	(0.276)	(0.217)
C. Runoff Elections								
Regional Tariff Reduction (RTR_r)	N/A	N/A	-1.227***	-1.280***	-1.815***	-1.785***	-1.949***	-1.611***
			(0.390)	(0.406)	(0.490)	(0.491)	(0.495)	(0.412)
Observations	484	484	484	484	484	484	484	3,388 or 2,420

Notes: Coefficients obtained from OLS regressions of the changes in vote share for left-wing parties on regional tariff reductions. All regressions control for state x time fixed effects and the 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the 1991 Gini index. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Panel A presents our benchmark specification. Panel B uses the alternative measures for labor demand shock using the changes in effective rates of protection instead of the changes in nominal tariffs. Panel C uses runoff elections to construct the dependent variable. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

variable is constructed using the second round of voting instead of the first round. The coefficients for 1994 and 1998 are not reported because there were no runoff elections in those years. The estimates using this specification are negative and significant at the 1% level, and are larger (in absolute value) than those from the main specification (almost 50% larger in the case of the pooled regression), indicating that the negative effects of the tariff reductions on the vote share for the left are stronger in runoff elections.

Overall, our findings show that the trade exposure generated by the Brazilian trade reform of the 1990s reduced the support for left-wing candidates in harder-hit regions relative to the less affected ones. Moreover, the estimates are stable over time, indicating that the effect on the vote share for the left was permanent.

Ideally, we would also be able to provide evidence that these same trends were not present before the trade reform started in the late 1980s. Or, in terms of the difference-in-differences interpretation of our empirical strategy, we would be able to show that pre-existing trends are not a concern in our setting. Unfortunately, this exercise is impossible because 1989 marks the first Brazilian direct presidential election after the end of the military dictatorship in 1985. We are able to provide evidence of the absence of pre-existing trends when looking at the variables related to union strength in section 6 but, in terms of political outcomes, we are constrained by the fact that there were no direct presidential elections between 1960 and 1989.

The only pre-1989 election data available at the microregion level for the entire country refers to the 1982 elections for state governors. One possibility would be to classify the votes for state governors in 1982 as left vs. non left and use the 1982 state elections data, together with the data from the 1989 presidential election, to test for pre-existing trends in electoral outcomes.[22] We do not think this is a sensible approach, since state elections in Brazil have very different dynamics—in terms of political alliances and ideological identification—when compared to presidential elections, and also because trade policy is under the auspices of the federal government. Moreover, there were other differences between the 1982 and 1989 elections in addition to the office in question. In the 1982 governor elections, some states did not participate at all (because of their official status during the dictatorship) and some others did not have a left-wing candidate running. Still, for the sake of completeness, we estimate this regression (change in vote share for the left between 1982 and 1989 on the RTR_r).[23] We find a negative and non-significant coefficient—point estimate equal to -0.611 and standard error equal to 0.457—indicating that microregions that would be subject to larger tariff reductions in the early 1990s were observing a small and non-significant relative *increase* in the vote share for the left before the trade reform started. If anything, bearing in mind its limitations, this exercise indicates that the left was not losing ground in the regions that were to experience larger reductions in tariffs before the trade liberalization process actually started.

A potential question related to our previous results is whether the reduction in votes for the left was associated with increased political radicalization, as documented by Autor et al. (2020) and Dippel et al. (2021) in other contexts. In Table A3 (Appendix A), we show that this does not seem to have been the case. We calculate the vote shares for center and traditional right-wing parties and, separately, for far-right parties and re-estimate our benchmark specification using these two vote shares as dependent variables.[24] The results show that the reduction in votes for the left was accompanied by an almost one-to-one increase in votes for center and traditional right-wing parties. The impact on votes for far-right parties was much smaller in magnitude and non-robust in terms of statistical significance and estimated sign. So, in general, our previous results do not reflect increased political radicalization.[25]

[22]Microregion data for state governors elections for 1986 and 1990 are not available. Thus, analyses based entirely on elections for state governors, in addition to their shortcomings related to the different nature of presidential and state elections, are also not possible due to data limitations.

[23]Since we use 1982 data to estimate the regression with state elections, we use a more aggregate classification of minimum comparable areas (in order to keep the consistency over time). In addition, we use fewer regions in this analysis because some states did not hold elections in 1982 and some others did not have a left candidate running. To make the results comparable, we also re-estimate our main specification with the same restricted sample used in this pre-trend analysis. The results using this alternative sample are all negative and similar in magnitude to the ones with the full sample, and with exception of the 2002 coefficient, they are also statistically significant.

[24]We classify PRONA, PSC, PSDC, PSL, NOVO and PRTB as far-right parties. Before PSL, which was electorally irrelevant prior to gaining prominence in (and winning) the 2018 elections, among all parties that obtained at least 2.5% of the votes in presidential elections, only PRONA in 1994 can be considered far right. In particular, none of the country's traditional political movements has been associated with the radical right. Possibly for this reason, to the best of our knowledge, there is no systematic classification of far-right parties in Brazil. Nevertheless, our classification is consistent with the general view of political scientists in the country (https://www.gazetadopovo.com.br/rodrigo-constantino/artigos/existem-partidos-de-extrema-direita-no-brasil, https://politica.estadao.com.br/noticias/eleicoes,o-que-significam-direita-esquerda-e-centro-na-politica,70002314116).

[25]We also estimate the same regression using voter turnout as dependent variable and find no significant effect for any of the election years. This should be expected, given that voting is compulsory in Brazil.

5.1 Robustness

To estimate the causal effect of the trade liberalization on electoral outcomes consistently, the error term ϵ_{rt} in equation (1) must be uncorrelated with RTR_r conditional on the other covariates. This identification assumption would not hold if there were an omitted variable correlated with RTR_r that impacted election outcomes and were not captured by the controls. In this section, we confirm that the results in Table 2 are robust to possible confounding effects due to other shocks that happened after and during the trade reform. We also show that they are present in some restricted samples and are robust to alternative classifications of political parties.

Table 3 shows that the results are robust to controlling for other economic shocks and to restricting the sample to a more homogeneous set of observations. For comparison purposes, Panel A replicates the estimates from the main specification. Since we analyze the effects of tariff reductions that occurred in the period from 1990 to 1995, it is essential to guarantee that post-liberalization tariff changes are not interfering with the results. For this reason, in Panel B we carry out the same robustness test proposed by Dix-Carneiro and Kovak (2017), including a variable that captures the changes in import tariffs for the period after liberalization. This variable is similar to the RTR_r, but uses tariff changes for the period between 1995 and each year of analysis $t > 1995$. Since Kume et al. (2003) do not report tariff data after 1998, we use UNCTAD TRAINS data to construct post-liberalization tariffs. The coefficients on these other tariff changes vary over time and are generally indistinguishable from zero. Most importantly, the estimates for θ_t presented in Panel B are very similar to the ones in Panel A, indicating that the results from the main specification are robust to the inclusion of tariff changes after 1995.

In Panel C, we analyze whether the results are driven mostly by the contrast between regions concentrated in agriculture vs. manufacturing, or whether the variation in tariffs across manufacturing sectors is indeed providing some identification. With this objective, we estimate equation (1) excluding microregions with more than 50% of the workforce in agriculture. This is similar to excluding the top 25% microregions in terms of share of employment in agriculture. The table shows that, when we exclude these regions, the coefficients remain fairly similar to those in the main specification, and remain statistically significant despite the reduced sample size. In addition, it is worth noting that the results are still negative for all years and statistically significant for the pooled specification even when we exclude all regions above the median of the distribution of workers in agriculture, which cuts the sample in half. This shows that our results are not driven exclusively by the differences across agricultural and non-agricultural regions.

During the sample period, Brazil privatized some state-owned firms. The privatization process began in 1991 during the Collor administration but accelerated substantially during Cardoso's administration (1995-2002). Since that process affected different industries to different degrees, it was potentially correlated with the RTR_r. Therefore, in Panels D and E we include controls to address the impact of privatization on elections. Panel D adds

Table 3: Regional Tariff Reduction and the Vote Share of the Left: Robustness

Dep.Var: Δ Vote share for left-wing parties	(1) $t = 1994$	(2) $t = 1998$	(3) $t = 2002$	(4) $t = 2006$	(5) $t = 2010$	(6) $t = 2014$	(7) $t = 2018$	(8) Pooled
A. Main Specification								
Regional Tariff Reduction (RTR_r)	-0.829***	-1.243***	-0.811**	-1.032**	-1.103**	-0.986**	-1.733***	-1.105***
	(0.307)	(0.347)	(0.398)	(0.411)	(0.441)	(0.487)	(0.485)	(0.359)
B. Post-Liberalization Tariff Reduction								
Regional Tariff Reduction (RTR_r)	N/A	-1.244***	-0.726*	-0.912**	-1.176**	-1.169**	-1.730***	-1.141***
		(0.347)	(0.404)	(0.414)	(0.485)	(0.546)	(0.486)	(0.375)
C. Less than 50% of the work force working in Agriculture								
Regional Tariff Reduction (RTR_r)	-0.880***	-1.254***	-0.681*	-0.804*	-1.030**	-0.932*	-1.655***	-1.034***
	(0.333)	(0.383)	(0.403)	(0.420)	(0.458)	(0.522)	(0.500)	(0.380)
Observations:	338	338	338	338	338	338	338	2,366
D. Privatization - Initial state-owned employment share								
Regional Tariff Reduction (RTR_r)	N/A	-1.232***	-0.804**	-0.993**	-1.063**	-0.936*	-1.690***	-1.120***
		(0.349)	(0.397)	(0.426)	(0.446)	(0.499)	(0.494)	(0.384)
E. Privatization - Change in state-owned employment share								
Regional Tariff Reduction (RTR_r)	N/A	-1.243***	-0.808**	-1.028**	-1.099**	-0.973**	-1.714***	-1.147***
		(0.347)	(0.399)	(0.413)	(0.442)	(0.490)	(0.487)	(0.378)
F. Real Exchange Rates								
Regional Tariff Reduction (RTR_r)	-1.008***	-1.369***	-0.882**	-1.085***	-1.119**	-1.115**	-1.804***	-1.197***
	(0.306)	(0.354)	(0.405)	(0.407)	(0.442)	(0.511)	(0.469)	(0.356)
G. Tariff Changes in Mercosur								
Regional Tariff Reduction (RTR_r)	-1.000***	-1.291***	-0.990**	-1.484***	-1.753***	-1.555***	-2.212***	-1.469***
	(0.351)	(0.419)	(0.466)	(0.476)	(0.515)	(0.583)	(0.531)	(0.410)
Observations	484	484	484	484	484	484	484	3388 or 2904

Notes: Coefficients obtained from OLS regressions of the changes in vote share for left-wing parties on regional tariff reductions. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Panel A presents our benchmark specification. Panel B includes controls for tariff changes after 1995. Panel C estimates the main equations using only regions with less than 50% of the workforce allocated in agriculture. Panel D controls the main specification for quartile indicators for the 1995 share of regional employment in state-owned firms. Panel E controls for the change in the share of regional workforce employed in state-owned firms from 1995 to subsequent election year. Estimates for the first column in panels D and E are not provided, since data on the number of workers in state-owned firms are only available from 1995 onwards. Panel F includes controls for real exchange rates changes to the main specification. Panel G controls the main specification for the preferential tariff reductions implemented to Brazil due to Mercosur and for the tariff changes imposed by the Mercosur partners. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

quartile indicators for the 1995 share of regional employment in state-owned firms, while Panel E controls for the change in employment share in state-owned companies from 1995 to subsequent election years.[26] In both cases, the privatization controls hardly affect the estimates for θ_t, indicating that the baseline results presented are robust to this policy shock.

In Panel F, we verify whether the results are robust to swings in the Brazilian currency, which may also have differential effects across industries. To do that, we construct two variables to control for the changes in regional real exchange rates. In Appendix B.1, we describe in detail how these exchange rate controls are constructed, but the key point is that they are analogous to the RTR_r variable, built from industry specific exchange rates and weighted by initial employment shares. Once again, the estimates for θ_t are similar to the ones in Panel A, so we conclude that exchange rates are not relevant drivers of the results discussed earlier.

[26]Data on employment in state-owned firms come from Brazilian Ministry of Labor's RAIS dataset, which started to provide this information in 1995. RAIS - *Relações Anuais de Informações Sociais* (Ministério do Trabalho, 1985-2014) is an administrative data set, reported by employers, covering all workers formally employed in the country. See Appendix C.1 for additional information about RAIS.

Finally, there was another important change in trade policy during the period of Brazil's unilateral trade liberalization, which could be potentially correlated with the tariff cuts. In 1991, Argentina, Brazil, Paraguay and Uruguay formed Mercosur, a preferential trading bloc. The bloc entailed a process of reciprocal tariff reductions that also took place mostly from 1991 to 1995. Hence, regions that were larger importers from the other members of the bloc may have experienced larger effective tariff cuts than the ones captured by the unilateral trade liberalization. On the other hand, regions that were larger exporters to the other bloc members may have experienced an opposite effect stemming from the reduction of tariffs by Mercosur partners.

To account for these possibilities, in Panel G we add two variables. One captures the preferential tariff reductions implemented by Brazil due to Mercosur. The other captures the preferential tariff reductions in the other Mercosur members. These variables are calculated for each microregion analogously to the way we constructed RTR_r (these calculations are described in detail in Appendix B.2). Panel G shows that their introduction as controls increases the magnitude of the estimates of the RTR_r coefficient, but the sign and statistical significance of these coefficients are unaffected.[27]

We also verify whether our results are robust to alternative classifications of political parties in the right-left spectrum. These robustness exercises are reported in Table 4. Panel A replicates the estimates of the main specification. The rationale for the different classifications in Panels B and C are as follows. In Brazil, in all presidential elections since 1989, there have been more than a dozen candidates, and sometimes there were over twenty. While the main political parties are perennial, the fringe parties often merge, split, dissolve and change names. This makes the definition of their political orientation less precise. To verify whether potential errors in the classification of these fringe parties in the left-right spectrum are driving our results, in Panel B we show the coefficient of RTR_r for a classification that considers all "less relevant parties" as right wing. Panel C does the same but assumes that all such parties are left wing. We use a rather lax definition for "less relevant parties": all political parties that did not reach at least 5% of the total votes in any election since 1989. This includes some traditional parties, such as PMDB and PL (which were important parties from a legislative perspective). Despite the apparent arbitrariness of this reclassification, the results remain similar to those in Panel A, indicating that they are not driven by the votes for small parites. If we use cutoffs lower than 5%, the results become even more similar to our benchmark specification.

Finally, as mentioned before, in Panel D we present results using a classification that considers PSDB as a left-wing party in the 1989 election. The coefficients remain negative

[27]The coefficient for the variable that captures the preferential tariff reductions implemented by Brazil due to Mercosur is negative and statistically significant at the 10% level in the pooled regression. This result is in line with the impact found from the unilateral liberalization. However, the magnitude of the impact from the Mercosur tariff changes is almost an order of magnitude smaller than that of the unilateral tariff reduction. Moving a region from the 10^{th} percentile to the 90^{th} percentile of the distribution of the variable capturing Brazil's tariff reductions to the Mercosur partners would be associated with an average reduction of 1.38 p.p. in the vote share for the left in presidential elections.

Table 4: Robustness: Alternative Definitions of "Left"

Dep.Var: Δ Vote share for left-wing parties	(1) $t = 1994$	(2) $t = 1998$	(3) $t = 2002$	(4) $t = 2006$	(5) $t = 2010$	(6) $t = 2014$	(7) $t = 2018$	(8) Pooled
A. Main Specification								
Regional Tariff Reduction (RTR_r)	-0.829***	-1.243***	-0.811**	-1.032**	-1.103**	-0.986**	-1.733***	-1.105***
	(0.307)	(0.347)	(0.398)	(0.411)	(0.441)	(0.487)	(0.485)	(0.359)
B. All Less Relevant Parties Right-Wing								
Regional Tariff Reduction (RTR_r)	-0.728**	-1.122***	-0.739*	-0.931**	-1.009**	-0.893*	-1.637***	-1.008***
	(0.300)	(0.344)	(0.393)	(0.408)	(0.439)	(0.484)	(0.482)	(0.354)
C. All Less Relevant Parties Left-Wing								
Regional Tariff Reduction (RTR_r)	-0.849***	-0.981**	-0.529	-0.742	-0.814*	-0.608	-1.334**	-0.837**
	(0.317)	(0.400)	(0.420)	(0.455)	(0.485)	(0.525)	(0.522)	(0.392)
D. PSDB left-wing party in 1989								
Regional Tariff Reduction (RTR_r)	-1.361***	-1.776***	-1.344***	-1.564***	-1.636***	-1.519***	-2.265***	-1.638***
	(0.316)	(0.344)	(0.364)	(0.401)	(0.434)	(0.490)	(0.470)	(0.349)
Observations	484	484	484	484	484	484	484	3.388

Notes: Coefficients obtained from OLS regressions of the changes in vote share for left-wing parties on regional tariff reductions. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Panel A presents our benchmark specification. Panel B constructs the dependent variable considering that all lees relevant parties are right-wing parties. Panel C constructs the dependent variable considering that all lees relevant parties are left-wing parties. Less relevant parties are defined as the parties that did not reach at least 5% of the total votes in any election since 1989. Panel D constructs the dependent variable considering that PSDB was a left-wing party in the 1989 presidential elections. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

and statistically significant in every year, and the estimates (in absolute value) are somewhat larger relative to the ones from the main specification.

Altogether, the results in this section demonstrate that the estimates presented in Table 2 are robust to alternative measurements of the trade reform and classification of parties, and to a wide variety of economic shocks as possible confounders.

6 Trade Liberalization, Labor Unions, and Elections

The shift of voting patterns in Brazil to the center-right in regions more affected by the reduction in tariffs seems, at first sight, puzzling. As discussed in the introduction, both punishment of the incumbent party or an ideological reaction to the trade reform in this context should have led to an increase in the vote share of the left. These responses, which have been documented in other settings (e.g., Autor et al., 2020; Blanchard et al., 2019; Che et al., 2020; Dippel et al., 2021; Jensen et al., 2017), are absent in the case of the Brazilian trade reform: the patterns documented in the previous section reflect a movement away from the left and towards the center-right.

We propose here an institutional mechanism that works through the weakening of labor unions—and therefore of the organized labor movement in general—in regions most impacted by the reduction in tariffs. This mechanism has not yet been considered by the literature. Still, it is likely to be present in other contexts as well, since labor unions tend to be particularly strong in the manufacturing sector (Visser, 2019) and the manufacturing sector tends to be particularly affected by trade shocks.

In the case of Brazil, this potential mechanism is likely to be further enhanced by the intrinsic relationship between formal employment and union strength and by the much higher incidence of labor formality in the manufacturing sector. We know from the previous literature that the tariff cuts of the Brazilian trade liberalization reduced relative formal employment rates in the most affected regions (Dix-Carneiro and Kovak, 2017; Ponczek and Ulyssea, 2021). As a result, because of the compulsory contribution that formal workers had to make to unions during our sample period, the tariff cuts reduced unions' main source of revenue.[28]

This potential effect on labor unions may end up influencing electoral outcomes. The relationship between labor unions and elections has been explored by several authors in the political science literature. Burns et al. (2000), Leighley and Nagler (2007) and Radcliff and Davis (2000) discuss unions' ability to influence electoral outcomes via campaign contributions and turnout mobilization, while Kim and Margalit (2017) show that unions are capable of influencing the political preferences of their members. These channels may be particularly true for Brazil, where labor unions and left-wing parties are historically connected; as Colistete (2007) points out, most of the main left-wing parties were founded by former union leaders. Therefore, the results in Figure 4 may be partly driven by the effect of the trade liberalization on unions' capacity to influence elections.

We establish this point in two steps. In the next subsection, we present evidence that unions were relatively weakened in the microregions more affected by the tariff cuts. In subsection 6.2, we then show that the negative relationship between votes for the left and tariff cuts was driven predominantly by: (i) political parties associated with labor unions; and (ii) microregions where unions were present in the pre-shock period.

6.1 Trade Liberalization and Union Strength

There is no obvious measure to define how strong unions are in a certain region and period. Hence, using all available data, we construct three different proxies for unions' strength to verify whether labor unions were affected by the trade liberalization process. The first is the number of people who work for labor unions. The assumption behind this proxy is that unions that employ more people are larger and, therefore, more likely to be politically powerful. If this assumption holds, then it is possible to measure the "political strength" of a union by looking at the number of workers it employs. To construct this measure, we use data between 1985 and 2018 for the universe of formal workers in the country from RAIS, and identify workers who were employed by labor union establishments. Unfortunately, prior

[28] As discussed in section 2, until 2017 the main source of revenue for labor unions in Brazil was a compulsory annual contribution, equivalent to one working day, for every formal worker in the union's regional jurisdiction and category, independently of whether the worker was formally affiliated with the union.

to 1995, we cannot distinguish labor union establishments from professional associations or employers' associations. For this reason, to be consistent over time, we construct our first proxy as the number of workers employed by establishments in either of these three categories. By looking at post-1995 data, however, we know that the vast majority of workers employed by these types of establishments were actually working for labor unions. In 1995, for example, this number was above 75%.

The second proxy is the number of labor unions in each region. If trade liberalization affected unions' strength, then we should observe a relative reduction in the number of these organizations in the harder-hit regions. We construct this variable by using RAIS data between 1985 and 2018 and counting the number of establishments classified as labor unions, professional associations, or employers' associations with at least one worker employed in December of each year.

The third proxy is an estimate of "union density." The use of this measure as a proxy for union strength assumes that the more members a union has in relation to the number of workers in the economy, the more influential it is. The influence stems both from more revenue collected with membership fees and from a greater capacity to persuade the electorate. For both reasons, unions with more unionized workers can be associated with greater political strength.

Unfortunately, information on union membership is only available at the state level. To impute this number for microregions, we use a weighted sum of the union membership by industry from the state-level data, where the weights are given by each microregion's share in state-level formal employment by industry. Formally, we construct the unionized workers variable by microregion (UW_{rt}) as follows:

$$UW_{rt} = \sum_{i=1}^{I} \frac{\text{Formal Workers}_{irt}}{\text{Formal Workers}_{ist}} \text{Unionized Workers}_{ist},$$

where Formal Workers_{irt} denotes the number of formal workers in industry i, region r and time t, Formal Workers_{ist} denotes the number of formal workers in industry i, state s and time t, and Unionized Workers_{ist} indicates the number of formal unionized workers in industry i, state s and year t. The state-level number of workers affiliated with unions by industry from 1986 to 2015 comes from the *Pesquisa Nacional por Amostra de Domicílio* (Instituto Brasileiro de Geografia e Estatísitca, 1985-2014), which is the Brazilian National Household Survey, conducted by IBGE. See Appendix C.2 for additional information about PNAD.[29]

Because our imputation of union membership generates some outliers, we use median regressions instead of ordinary least squares in this section. Specifically, we estimate a median regression version of equation (1), using the difference between each year $t \in 1985,..., 2018$ and 1989, for each of our proxies for union strength. Since there are some municipalities that do not appear in RAIS in some of the years, we have to drop 10 microregions to get a balanced

[29]Table A4 displays descriptive statistics for the variables used as proxy for union strength in the baseline year.

sample for the entire period.[30] Importantly, because we have data pre-1989 for all the union strength variables, we can use the pre-shock years to estimate "placebo" coefficients and test for pre-existing trends. In Table A5 (Appendix A), we report all the estimates discussed in this section using OLS regressions (instead of median regressions). The table shows that qualitative results are unchanged by the estimation method employed. For completeness, in Tables A6 and A7 (Appendix A) we also report the estimates from Table 2 using median regressions and the sample with 474 microregions, respectively, showing that the election results are also not sensitive to the specific estimation method or the sample used.

We first analyze the relationship between tariff reductions and the number of union employees relative to the population of the microregion. More precisely, our dependent variable is the change in the number of union employees per thousand people aged between 15 and 64. The estimates for the impact of the regional tariff reduction are displayed in Figure 5. The coefficients for the years between 1990 and 2018 are reported in red diamonds. The figure also shows 95% confidence intervals (dashed lines). The effect of the tariff cuts on unions' workforce is negative in every year after 1989. In most years during the transition years of 1991-1995 (in gray), the estimates are not statistically different from zero, but after 1997 they become stable and statistically significant at the 5% level. The coefficients' negative sign reveals that regions that faced deeper tariff cuts experienced a reduction in labor unions' workforce relative to regions facing smaller tariff cuts. This indicates that labor unions were affected by the trade liberalization negatively and permanently.

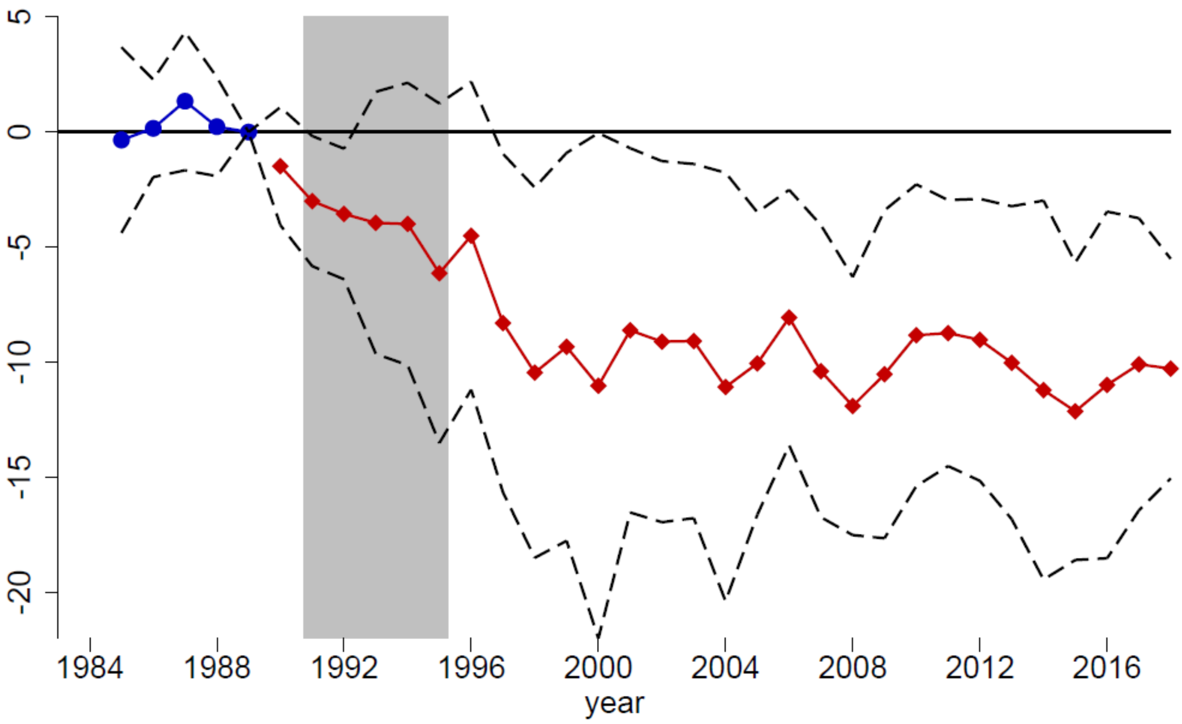
As mentioned above, we carry out placebo tests using as dependent variables the pre-intervention differences in the number of union employees per thousand people aged between 15 and 64 years old. These coefficients are represented in blue circles in Figure 5. In addition to being very small in absolute value, none is statistically significant, indicating that pre-existing trends in unions' employment levels were uncorrelated with the tariff cuts. Thus, the distribution of sectoral tariff reductions seems indeed to have been orthogonal to local institutional dynamics before the trade liberalization process.

Next, we assess the relationship between the liberalization and the number of unions in each region (number of active union establishments per thousand inhabitants aged between 15 and 64). The results are presented in Figure 6, following the same patterns used in Figure 5. Estimated coefficients are small and not statistically significant during the transition years. After 1995, they increase in absolute value and become statistically significant, or nearly so, in most years, indicating that the number of unions in the more affected regions fell relative to that in the less affected ones. The pre-shock coefficients are again small and not statistically significant, showing no sign of pre-trends interfering with the results.

Finally, the relationship between union density (number of unionized workers per thousand people aged between 15 and 64) and tariff reductions is displayed in Figure 7, following

[30] We drop the 7 microregions in Tocantins state, which was only founded in 1988, and 3 other regions omitted in RAIS in at least one of the years in our sample.

Figure 5: Impact of the Tariff Cuts on Labor Unions' Employment



Notes: Each point reflects an individual median regression coefficient $\hat{\theta}_t$ following regression (1), where the dependent variable is the change in the number of union employees per thousand people between 15 and 64 years old and the independent variable is regional tariff reductions (RTR_r). Red diamonds indicate the coefficients from regressions considering changes from 1989 to the year in the x-axis. Blue circles indicate analogous coefficients for changes in years pre-1989. All regressions include state fixed effects and pre-liberalization demographic controls. The shaded area indicates that the liberalization process began in 1991 and ended in 1995. Dashed lines indicate 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

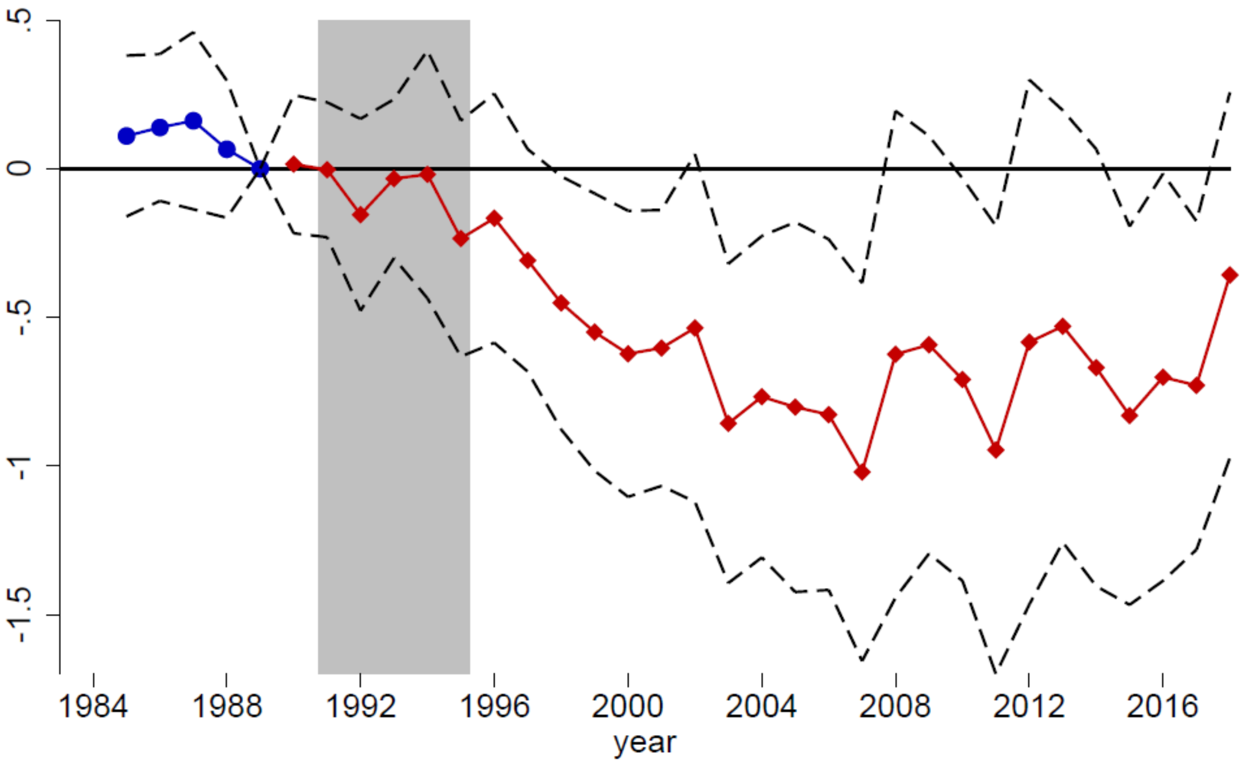
once more the same pattern from previous figures.[31] Once again, we observe a negative relationship between increased trade exposure and union density. As before, pre-trends do not seem to be a problem.

In Table 5, we report the coefficients displayed in the previous figures, together with additional estimates using alternative definitions of the dependent variables. In addition to normalizing the variables by the population between 15 and 64, we also present results for dependent variables normalized by the total population. To avoid putting too much information in the table, we report only the coefficients for election years and for the pooled regressions with all years.[32] The numbers in the table confirm the results already discussed in the previous figures. Column (8) of Panel A.1, for example, indicates that a one-standard-

[31]PNAD data on the number of workers affiliated with labor unions is not available for 1985, 1987, 1989, 1990, 1991, 1994, 2000, 2010 and 2016 to 2018. Therefore, we do not estimate the regression for these years. This is the reason why we have fewer point estimates in Figure 7 when compared to the previous two figures.

[32]We estimate the pooled regression using all years between 1991 and 2018. In this specification, we also include an interaction with a dummy variable indicating transition years (1991 - 1995)

Figure 6: Impact of the Tariff Cuts on the Number of Union Establishments

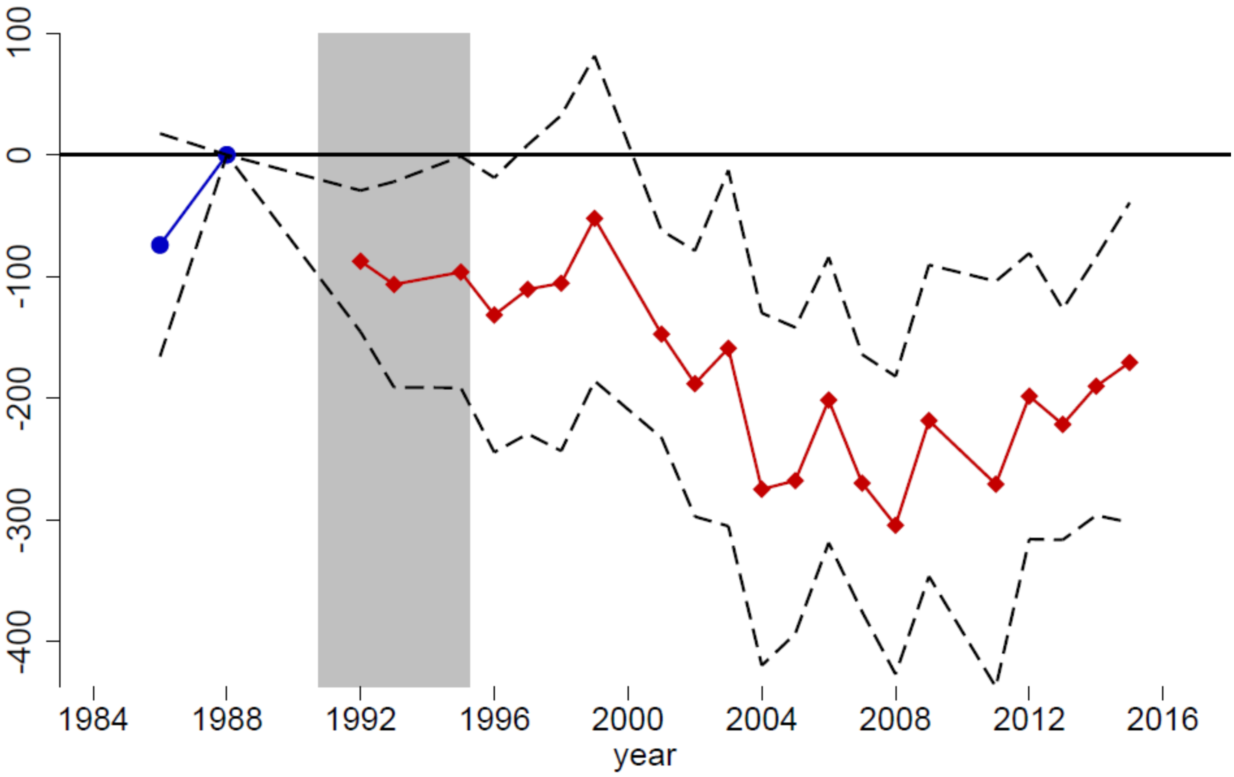


Notes: Each point reflects an individual regression coefficient $\hat{\theta}_t$ following regression (1), where the dependent variable is the change in the number of labor unions per thousand people between 15 and 64 years old, and the independent variable is regional tariff reductions (RTR_r). Red diamonds indicate the coefficients from regressions considering changes from 1989 to the year in the x-axis. Blue circles indicate analogous coefficients for changes in years pre-1989. All regressions include state fixed effects and pre-liberalization demographic controls. The shaded area indicates that the liberalization process began in 1991 and ended in 1995. Dashed lines indicate 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

deviation increase in the RTR_r would lead to an average reduction in union employees in the post-shock years equivalent to 78% of the 1989 median, which corresponded to 0.48 unionized worker per 1,000 inhabitants (between ages 15 and 64). Moving a region from the 10^{th} to the 90^{th} percentile of the RTR_r distribution corresponds to a relative reduction of 0.99 in the number of workers employed by labor unions. Using the other measures of union strength would yield less stark but still large results (for example, a one-standard-deviation increase in the RTR_r would lead to a 38% reduction in union density). To put this result in perspective, the median number of unions operating in a region during our sample period is 7. The extreme movement from the 10^{th} to the 90^{th} percentile of tariff reductions would lead to the virtual shutdown of union activities in the region with the median level of union presence.

Altogether, these results indicate that labor unions shrank in microregions that experienced relatively larger impacts from the trade liberalization process. If this loss of strength

Figure 7: Impact of the Tariff Cuts on Union Density



Notes: Each point reflects an individual regression coefficient $\hat{\theta}_t$ following regression (1), where the dependent variable is the change in the estimated number of unionized workers per thousand people between 15 and 64 years old, and the independent variable is regional tariff reductions (RTR_r). Red diamonds indicate the coefficients from regressions considering changes from 1989 to the year in the x-axis. Blue circles indicate analogous coefficients for changes in years pre-1989. All regressions include state fixed effects and pre-liberalization demographic controls. The shaded area indicates that the liberalization process began in 1991 and ended in 1995. Dashed lines indicate 95 percent confidence intervals. Standard errors adjusted for 112 mesoregion clusters.

impacted the capacity of these organizations to influence elections, then it is plausible that the weakening of labor unions may be an important driver of the results depicted in Figure 4. To verify this possibility, in the next subsection we analyze whether the effects of the tariff reduction on the vote share for left-wing candidates stemmed mostly from parties associated with unions, and whether they were heterogeneous across regions with and without the presence of unions in the pre-liberalization period.

6.2 Parties Identified with Labor Unions

Given the results from the previous sections, one may be tempted to adopt an instrumental variable strategy to estimate the impact of union strength on votes for the left, with RTR_r being used as an instrument for union strength. But the trade liberalization process

Table 5: Regional Tariff Reduction and Unions Strength

	(1) 1994	(2) 1998	(3) 2002	(4) 2006	(5) 2010	(6) 2014	(7) 2018	(8) Pooled
A.1 - Union Employees Per Thousand People (Between 15 and 64 Years)								
Regional Tariff Reduction (RTR_r)	-3.993	-10.46**	-9.108**	-8.062***	-8.830***	-11.21***	-10.28***	-9.549***
	(3.121)	(4.093)	(3.992)	(2.824)	(3.335)	(4.190)	(2.427)	(2.937)
A.2 - Union Employees Per Capita (Thousand People)								
Regional Tariff Reduction (RTR_r)	-2.450	-6.236**	-6.026***	-5.022***	-5.837**	-7.081***	-6.258***	-6.166***
	(1.913)	(2.989)	(1.675)	(1.355)	(2.820)	(2.176)	(1.612)	(1.686)
B.1 - Unions per Thousand People (Between 15 and 64 Years)								
Regional Tariff Reduction (RTR_r)	-0.0190	-0.451**	-0.536*	-0.827***	-0.709**	-0.669*	-0.357	-0.604**
	(0.212)	(0.217)	(0.298)	(0.300)	(0.344)	(0.374)	(0.313)	(0.242)
B.2 - Unions per Capita (Thousand People)								
Regional Tariff Reduction (RTR_r)	0.00444	-0.305***	-0.325	-0.536**	-0.484**	-0.607***	-0.368*	-0.402**
	(0.125)	(0.109)	(0.198)	(0.224)	(0.235)	(0.223)	(0.189)	(0.196)
C.1 - Unionized Workers per Thousand People (Between 15 and 64 Years)								
Regional Tariff Reduction (RTR_r)	N/A	-105.4	-188.0***	-201.5***	N/A	-190.1***	N/A	-186.0***
		(70.14)	(55.57)	(59.66)		(53.99)		(39.33)
C.2 - Unionized Workers per Capita (Thousand People)								
Regional Tariff Reduction (RTR_r)	N/A	-46.16	-137.8***	-127.5***	N/A	-133.1***	N/A	-114.8***
		(36.07)	(35.09)	(41.68)		(40.46)		(26.60)
Number of Microregions	474	474	474	474	474	474	474	474
Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Coefficients obtained from median regressions of the changes in the variables described in Panels titles on regional tariff reductions. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The pooled regression (column 8) controls for a interaction between RTR_r and a dummy indicating the transition years and interactions between control variables and year dummy variables. Each panel estimates the regression using a different dependent variable. Panel A.1 uses the changes in the number of workers employed in labor unions per thousand people between 15 and 64 years old at each region as the dependent variable. Panel A.2 uses the changes in the number of union employees per thousand people. In Panel B.1, the dependent variable is the change in the number of formal union establishments per thousand people between 15 and 64 years old. Panel B.2 uses the changes in the number of unions per thousand people as the dependent variable. The dependent variable in Panel C.1 is the changes in the number of unionized workers per thousand people between 15 and 64 years old. Panel C.2 uses the changes in the number of unionized workers per thousand people as the dependent. Standard errors in parentheses, clustered at the mesoregion level (112 clusters).
Significance at the *10%, **5%, *** 1% levels.

also impacted other dimensions of local labor markets and public good provision in ways that are bound to have electoral consequences, so that the exclusion restriction required by this estimation is unlikely to hold (see, for example, the evidence presented in Dix-Carneiro et al., 2018).

Hence, to provide evidence that part of the electoral effects estimated in section 5 are driven by the weakening of labor unions, we dig deeper into some institutional features of the party system in Brazil. Several Brazilian parties have a quasi-official connection with labor unions. If the weakening of unions were indeed an important mechanism driving the impact of the liberalization on electoral outcomes, we should expect the effect to be particularly evident in the support for parties more closely identified with the labor movement. To verify whether this is the case, we re-estimate equation (1) using the change in the vote share of parties identified with labor unions as the dependent variable; for brevity, we call these parties "union parties." Following Queiroz (2017), we classify PT, PDT, PMDB, PCB, PSB and PSD as union parties. The classification takes into account the historical relationship between political parties and the national union confederations discussed before.

The results are shown in Panel A of Table 6. The estimates are negative in every election year, indicating that these parties lost votes in the regions more affected by the tariff cuts, relative to the less affected ones. Indeed, the results are very similar to the main specification in Table 2, with the exception of 2002 and 2014, the only years when the impact of the tariff cuts on the vote share of union parties is not statistically significant. For example, the estimate of the pooled specification in Panel A corresponds to 93 percent of the magnitude of the respective coefficient in our main specification and is statistically significant at the 1% level.

Table 6: Regional Tariff Reduction and the Vote Share of Parties Identified with Unions

Dep var: Δ_{1989-t} Vote share for:	(1) $t = 1994$	(2) $t = 1998$	(3) $t = 2002$	(4) $t = 2006$	(5) $t = 2010$	(6) $t = 2014$	(7) $t = 2018$	(8) Pooled
A: Union Parties								
Regional Tariff Reduction (RTR_r)	-0.901***	-0.797**	-0.355	-1.090**	-1.749***	-0.802	-1.453***	-1.021***
	(0.308)	(0.344)	(0.408)	(0.449)	(0.545)	(0.516)	(0.525)	(0.384)
B: Left Without Union Parties								
Regional Tariff Reduction (RTR_r)	N/A	-0.164	-0.174	0.340***	0.928***	0.099***	0.024**	0.149***
		(0.191)	(0.208)	(0.054)	(0.173)	(0.0174)	(0.010)	(0.050)
C: Left Without Union Parties - Excluding Union Parties								
Regional Tariff Reduction (RTR_r)	N/A	-0.033	0.288	0.577***	0.931**	0.273***	0.009	0.290**
		(0.291)	(0.314)	(0.127)	(0.390)	(0.082)	(0.020)	(0.127)
Observations	484	484	484	484	484	484	484	3388

Notes: Panel A displays the coefficients from OLS regression of the changes in the vote share for parties identified with unions on regional tariff reductions. Panel B shows the coefficients of the OLS regression of the changes in the vote share for left-wing parties that are not identified with unions on regional tariff reductions. Panel C displays the coefficients from OLS regressions of the vote share for left-wing parties that are not identified with unions, but excluding the number of votes cast for union parties to calculate the share of votes. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the 1991 Gini index. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

To verify how other left-wing parties were impacted by the liberalization, we re-estimate the main specification but using the vote share of the other left-wing parties (i.e., those not identified with labor unions). We show these estimates in Panel B of Table 6.[33] We find that, in most years, the share of votes for left-wing parties not identified with labor unions *increased* in the most affected regions, relative to less affected ones. As discussed before, this is what one should expect when considering individual-led reactions or an ideological response from the electorate. In 1998 and 2002, the effect was indistinguishable from zero. To ensure that the results in Panel B are not mechanically driven by the reduction of votes for the left-wing parties identified with labor unions, we exclude them to construct the vote share for the left without considering union parties. We show the results in Panel C. They are very similar to those in Panel B. These results make clear that the negative impact of the tariff reduction holds exclusively for the parties that were connected with unions, providing support to the idea that the weakening of these labor organizations is indeed key to explain the effect of tariff reductions on electoral outcomes.

[33]There is no estimation for 1994 because there was no candidate in the 1994 presidential election from a left-wing party not identified with labor unions.

Now, if the weakening of labor unions were an important mechanism to explain the negative effect of the tariff reductions on the vote share of left-wing parties or "union parties," then one would expect to observe a stronger negative relationship between tariff cuts and the vote share for those parties in regions where unions were present before the reform. Accordingly, we assess whether there is a heterogeneous effect of the tariff reduction on election results depending on the presence of unions before the liberalization.

To carry out this analysis, we re-estimate equation (1) including interactions between the regional tariff reduction and a dummy variable indicating whether region r had any union operating in 1989. Specifically, we estimate the following equation for every presidential election in year t after 1989:

$$y_{rt} - y_{r,1989} = \theta_t^0[RTR_r \times (1 - I_r)] + \theta_t^1(RTR_r \times I_r) + \alpha_{st} + \psi X_r + \epsilon_{rt}, \tag{2}$$

where y_{rt} represents the vote share for left-wing or union parties in the presidential elections in year t after 1989 and I_r is an indicator of whether region r had at least one union establishment operating in the baseline year.[34] We run this specification considering our three main dependent variables: vote share for the left in the first rounds, vote share for the left in runoffs, and vote share for union parties in the first rounds.

Table 7 displays the results for the heterogeneity analysis. Panel A displays the results using the difference in the vote share for left-wing parties in the first round of voting. Panel B uses the same difference but in runoff elections. Panel C shows the results when considering the changes in the vote share for union parties as the dependent variable. The first row of each panel presents the estimates for θ_t^0, which indicate the effect for regions without labor unions in the baseline, while the second row of each panel presents the estimates for θ_t^1, which indicate the effect for regions with some presence of unions in the baseline.

The coefficients on the regional tariff reductions in regions without unions are not statistically significant in any election year and, in most cases, are small in magnitude (in some cases, they even become positive). The coefficients in regions with labor unions basically reproduce the results discussed extensively in previous sections. The pooled regressions in column 8, for example, show that the point estimate for regions with active unions in the baseline is always at least three times larger than that in regions without unions. In the case of union parties, which is probably the more precise test of our hypothesis, the coefficient in the pooled regression for regions without unions is a precisely estimated zero, more than two orders of magnitude smaller than the estimated coefficient for regions with operating unions in the baseline.

The results indicate that the impact of the trade liberalization on the vote share for left-wing and union parties only occurred in regions with some presence of unions prior to the shock. This pattern gives support to the hypothesis that the weakening of unions is an important mechanism through which the tariff reductions from the 1990s affected electoral

[34] In this analysis, we drop the microregion of Chorozinho, since we do not have information on unions for this region in 1989.

Table 7: Regional Tariff Reduction and Elections: Heterogeneous Effects

Dep. Var: Vote share for left-wing or union parties	(1) t = 1994	(2) t = 1998	(3) t = 2002	(4) t = 2006	(5) t = 2010	(6) t = 2014	(7) t = 2018	(8) Pooled
A. Vote Share for Left Wing Parties								
RTR_r in Regions with Unions in 1989	-0.853***	-1.339***	-0.902**	-1.095**	-1.148**	-1.034**	-1.759***	-1.161***
	(0.314)	(0.357)	(0.415)	(0.422)	(0.449)	(0.500)	(0.492)	(0.368)
RTR_r in Regions without Unions in 1989	-0.514	-0.346	0.166	-0.236	-0.463	-0.348	-1.098	-0.406
	(0.404)	(0.918)	(0.573)	(0.562)	(0.580)	(0.626)	(0.705)	(0.515)
B. Vote Share for Left Wing Parties in Runoff Elections								
RTR_r in Regions with Unions in 1989	N/A	N/A	-1.370***	-1.383***	-1.903***	-1.871***	-1.977***	-1.701***
			(0.415)	(0.420)	(0.498)	(0.496)	(0.502)	(0.420)
RTR_r in Regions without Unions in 1989	N/A	N/A	0.241	-0.161	-0.742	-0.750	-1.393	-0.561
			(0.834)	(0.573)	(0.628)	(0.701)	(0.918)	(0.641)
C. Vote Share for Union Parties								
RTR_r in Regions with Unions in 1989	-0.969***	-0.871**	-0.427	-1.203**	-1.873***	-0.890*	-1.517***	-1.107***
	(0.306)	(0.347)	(0.423)	(0.460)	(0.542)	(0.528)	(0.529)	(0.389)
RTR_r in Regions without Unions in 1989	-0.145	-0.120	0.453	0.224	-0.234	0.225	-0.450	-0.00664
	(0.413)	(0.670)	(0.483)	(0.559)	(0.599)	(0.644)	(0.710)	(0.496)
Observations	483	483	483	483	483	483	483	483

Notes: Coefficients obtained from OLS regressions of the changes in vote share for left-wing parties (Panel A), vote share for the left in the runoff elections (Panel B), and vote share for union parties (Panel C) on the interactions between the regional tariff reduction and dummy variables indicating whether region had at least one union operating in 1989. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Panel A uses the changes in the vote share for left parties in the first round of voting as the dependent variable. Panel B uses the changes in the votes hare for left parties in runoff elections as the dependent variable. Panel C uses the vote share for union parties to construct the dependent variable. The first row of each panel displays the estimates for the interaction of regional tariff reduction and the indicator which is equal to one when the region had no union operating in 1989. The second row of each panel displays the estimates for the interaction of the regional tariff reduction and the indicator variable that is equal to one if the region had at least one union operating in 1989. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

outcomes afterwards. It is also interesting to note that the pooled coefficient for regions with labor unions in Panel C of Table 7 is virtually identical to that in Panel A, which, in turn, is virtually identical to that from our benchmark specification in Table 2. This means that the entire average effect of tariff reductions on voting patterns can be statistically accounted for by the effect on votes for parties linked to unions in regions where unions operated before 1989.

A potential concern in this heterogeneity analysis is that regions without unions in 1989 may also be different in many other dimensions. In fact, there are only 54 microregions—just over 11% of the sample—without presence of unions in 1989 and they tend to be smaller and less urbanized, so this concern is indeed legitimate. To evaluate whether other microregion characteristics are partly behind the results from Table 7, we re-estimate a similar specification allowing for heterogeneity along the most relevant dimensions. First, since we have roughly 10% of the sample without labor unions in 1989, we create three dummy variables equal to 1 for the bottom decile of microregions in the distributions of population, urbanization rate, and share of workers in manufacturing. We then repeat the same estimation procedure, with the same dependent variables, considering the heterogeneity in the impact of RTR_r along five groups of microregions: (i) microregions with unions and outside of the bottom decile in all categories considered; (ii) microregions without unions in 1989; (iii) microregions in the bottom decile of population in 1989; (iv) microregions in the bottom decile of urbanization in 1991; and (v) microregions in the bottom decile of share of workers in manufacturing in 1991.

The results from this exercise are presented in Table A8 in Appendix A. The table shows that there is enough variation along the different dimensions considered and that the results from Table 7 seem indeed to reflect the presence of unions, not other sources of heterogeneity. The coefficients for microregions without unions in 1989 remain small and are not statistically significant, as before. In addition, among the 9 coefficients estimated for the alternative dimensions of heterogeneity, 8 are quite close in magnitude to those for municipalities with unions and outside the bottom deciles (some of these are not statistically significant due to lack of precision, but point estimates remain very similar to those from the first row in Table A8). This is true for all the coefficients indicating less populated and less urbanized regions, and is also true for 2 out of the 3 coefficients indicating regions with less employment in manufacturing. Hence, the patterns in Table 7 definitely do not reflect simply smaller, less urbanized, and more rural regions. Maybe surprisingly, given the high correlation between union presence and all of these dimensions, there is enough variation in the data to show that the heterogeneity seems to be particularly related to the presence of unions.

In sum, the results presented in this section provide strong support to the idea that the effects of the tariff cuts on the vote share for the left were driven, at least partly, by the weakening of unions' influence. First, we show that the tariff reductions weakened labor unions (as measured by different proxies). Second, we present evidence that the effects of the tariff cuts on the votes for the left were driven by the parties identified with labor unions, which are predominantly, but not exclusively, from the left. Finally, we show that the impact of the liberalization on the vote share for left wing and union parties was observed only in regions where unions were present before the shock.

7 Conclusion

We study the effects of the economic shock induced by the Brazilian trade liberalization of the 1990s on subsequent presidential elections in the country. We find that the vote share for left-wing parties in regions that faced larger tariff cuts fell significantly in all presidential elections from 1994 to 2018, when compared to their vote share in other regions. In light of the existing literature, this result was largely unexpected.

We present evidence suggesting that the weakening of labor unions played an important role in explaining this effect. Previous research has documented that the regions facing the largest tariff cuts in Brazil experienced a relative deterioration in labor market conditions (Dix-Carneiro and Kovak 2017, 2019; Kovak 2013). This affected the main source of income and membership of labor unions, potentially disrupting their mobilizing capacity. We show that, indeed, regions facing larger tariff reductions suffered a reduction in union membership, in the number of operating union establishments, and in the number of employees working directly for unions. This suggests that the trade policy shock weakened unions and, presumably, also their capacity to mobilize voters for the left, as they have historically done in Brazil (and elsewhere). Furthermore, the electoral loss of the left was driven exclusively by parties associated with unions, and happened only in regions where unions were present prior to the shock.

Our paper contributes to a recent but growing literature that investigates how trade shocks shape political outcomes, both in the short and long runs. We highlight a mechanism that so far has been largely overlooked in the trade policy literature. Trade liberalization typically has larger impacts in areas whose main activities are concentrated in manufacturing. But these are also areas where organized labor is strongest. Since organized labor can have profound political and social influences—including on voting behavior—the political consequences of trade liberalization may be wider than usually believed. Costa et al. (2019), for example, show that the labor shocks from the Brazilian trade liberalization led to an accelerated expansion in Pentecostal Evangelicalism in the country. A possibility raised by our results, and that deserves further investigation, is whether Pentecostal churches moved in to fill the social gap left by shrinking labor unions.

In addition, our findings may help to explain why there was no policy reversal in Brazil, even when left-leaning parties took power afterwards. An important line of research, best exemplified by the analyses of Maggi and Rodriguez-Clare (1998, 2007), shows that governments can sign trade agreements as a way to tie the hands of future governments and prevent a quid pro quo with protectionist agents. Although it is difficult to ascertain the ultimate goal of the Brazilian government that implemented the reform, it may have worked in the way envisioned by the Maggi and Rodriguez-Clare models, even though the liberalization was unilateral, rather than reciprocal. The reason in this case would be that the reform weakened a key institutional player that lobbies for protection and supports political parties with similar policy views. Therefore, it also undermined the possibility of a future policy reversal backed by this same institutional player.

More generally, our findings reinforce the growing view that large processes of trade liberalization can have consequences that go far beyond economic outcomes. Although the mechanisms through which these consequences play out may be subtle, their long-term effects can be significant.

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Appendix A - Additional Tables

Table A1: Left-Right Orientation of Brazil's Political Parties

Left-Wing Parties		Non-Left-Wing Parties		
PT*	PDT*	PRN	PSDB	PSD*
PV	PSB*	PSL	PRONA	PDCdoB
PPS	PSOL	PDS	PMDB*	PRTB
PCO	PSTU	PL	PTN	PFL
PCB*	PMN	PPR	PCN	PSC
PPL		PP	PSDC	PPB
		PTB	PN	PSN
		PSP	PTdoB	PRP
		PLP	REDE	PATRI
		NOVO	PODE	DC

Notes: This table displays the classification of Brazil's political parties in the left-right political spectrum. Parties with strong links with labor unions (Union Parties) are identified with *.

	Industry	Industry Name	Nivel 50	1970, 1980, 1991 Census (*atividade*)	2000, 2010 Census (*CNAE-Dom*)
Tradable	1	Agriculture	1	011-037, 041, 042, 581	1101-1118, 1201-1209, 1300, 1401, 1402, 2001, 2002, 5001, 5002
	2	Mineral Mining (except combustibles)	2	050, 053-059	12000, 13001, 13002, 14001-14004
	3	Petroleum and Gas Extraction and Coal Mining	3	051-052	10000, 11000
	4	Nonmetallic Mineral Goods Manufacturing	4	100	26010, 26091, 26092
	5	Iron and Steel, Nonferrous, and Other Metal Production and Processing	5-7	110	27001-27003, 28001, 28002
	8	Machinery, Equipment, Commercial Installation Manufacturing, and Tractor Manufacturing	8	120	29001
	10	Electrical, Electronic, and Communication Equipment and Components Manufacturing	10-11	130	29002, 30000, 31001, 31002, 32000, 33003
	12	Automobile, Transportation, and Vehicle Parts Manufacturing	12-13	140	34001-34003, 35010, 35020, 35030, 35090
	14	Wood Products, Furniture Manufacturing, and Peat Production	14	150, 151, 160	20000, 36010
	15	Paper Manufacturing, Publishing, and Printing	15	170, 290	21001, 21002, 22000
	16	Rubber Product Manufacturing	16	180	25010
	17	Chemical Product Manufacturing	17,19	200	23010, 23030, 23400, 24010, 24090
	18	Petroleum Refining and Petrochemical Manufacturing	18	201, 202, 352, 477	23020
	20	Pharmaceutical Products, Perfumes and Detergents Manufacturing	20	210, 220	24020, 24030
	21	Plastics Products Manufacturing	21	230	25020
	22	Textiles Manufacturing	22	240, 241	17001, 17002
	23	Apparel and Apparel Accessories Manufacturing	23	250,532	18001, 18002
	24	Footwear and Leather and Hide Products Manufacturing	24	190, 251	19011, 19012, 19020
	25	Food Processing (Coffee, Plant Products, Meat, Dairy, Sugar, Oils, Beverages, and Other)	25-31	260, 261, 270, 280	15010, 15021, 15022, 15030, 15041-15043, 15050, 16000
	32	Miscellaneous Other Products Manufacturing	32	300	33001, 33002, 33004, 33005, 36090, 37000
Nontradable	91	Utilities	33	351, 353	40010, 40020, 41000
	92	Construction	34	340, 524	45001-45005
	93	Wholesale and Retail Trade	35	410-424, 582, 583	50010, 50030, 50040, 50050, 53010 ,53020, 53030, 53041, 53042, 53050, 53061-53068, 53070, 53080, 53090, 53101, 53102, 55020
	94	Financial Institutions	38	451-453, 585, 612	65000, 66000, 67010, 67020
	95	Real Estate and Corporate Services	40, 41	461-464, 543, 552, 571-578, 584, 589	63022, 70001, 71020, 72010, 74011, 74012, 74021, 74022, 74030, 74040, 74050, 74090, 92013, 92014, 92015, 92020
	96	Transportation and Communications	36, 37	471-476, 481, 482, 588	60010, 60020, 60031, 60032, 60040, 60091, 60092, 61000, 62000, 63010, 63021 ,64010 ,64020, 91010
	97	Private Services	39, 43	511, 512, 521-523, 525, 531, 533, 541, 542. 544, 545, 551, 577, 586, 587, 613-619, 622-624, 632, 901, 902	1500, 50020, 53111, 53112, 53113, 55010, 55030, 63030, 70002, 71010, 71030, 72020, 73000, 74060, 80011, 80012, 80090, 85011, 85012, 85013, 85020, 85030, 90000, 91020, 91091, 91092, 92011, 92012, 92030, 92040, 93010, 93020, 93030, 93091, 93092, 95000
	98	Public Administration	42	354, 610, 611, 621, 631, 711-717, 721-727	75011-75017, 75020

Table A2: Consistent Industry Classification Across Censuses and Tariff Data
Source: Dix-Carneiro and Kovak (2017)

Table A3: Vote Share for Far-Right Parties

Dep var: Δ_{1989-t} Vote share:	$t = 1994$	$t = 1998$	$t = 2002$	$t = 2006$	$t = 2010$	$t = 2014$	$t = 2018$	Pooled
Panel A. Far-Right Parties								
Regional Tariff Reduction (RTR_r)	0.266***	-0.035	N/A	-0.021**	-0.023***	0.065***	0.304	0.093*
	(0.077)	(0.021)		(0.008)	(0.008)	(0.014)	(0.315)	(0.049)
Panel B. Other Center and Right Parties								
Regional Tariff Reduction (RTR_r)	0.562*	1.278***	0.842**	1.053**	1.127**	0.921*	1.428***	1.030***
	(0.336)	(0.354)	(0.398)	(0.411)	(0.441)	(0.491)	(0.434)	(0.352)
Observations	484	484	484	484	484	484	484	3,388 or 2,904

Notes: Coefficients obtained from OLS regressions of the changes in vote share for far-right parties on the regional tariff reductions in Panel A and for other non-left parties on the regional tariff reductions in Panel B. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Standard errors in parentheses, clustered at the mesoregion level (114 clusters). Significance at the *10%, **5%, *** 1% levels.

Table A4: Descriptive Statistics – Labor Unions

Variable:	Mean	SD	Min	Max	P10	P25	P50	P75	P90
Number of Union Employees	436.63	2013.84	0.00	27,855	0	6	36	174	688
Number of Union Employees per Thousand People (Between 15 and 64 Years Old)	1.11	1.72	0.00	13.73	0.00	0.11	0.48	1.41	2.92
Number of Union Employees Per Capita (Thousand People)	0.68	1.08	0.00	8.89	0.00	0.06	0.28	0.81	1.83
Number of Unions	25.98	85.54	0.00	1,283	0	2	7	21	51
Number of Unions Per Thousand People (Between 15 and 64 Years Old)	0.11	0.10	0.00	0.53	0.00	0.03	0.09	0.18	0.24
Number of Unions Per Capita (Thousand People)	0.07	0.06	0.00	0.32	0.00	0.02	0.05	0.11	0.15
Number of Unionized Workers	10088.72	45383.58	1.39	675,000	118	407	1,574	5,121	15,601
Number of Unionized Workers Per Thousand People (Between 15 and 64 Years Old)	31.06	34.50	0.11	351.23	2.86	6.96	18.87	42.82	75.35
Number of Unionized Workers Per Capita (Thousand People)	18.64	21.09	0.06	199.02	1.41	3.60	10.77	26.06	46.26

Notes: This table displays descriptive statistics for the measures used as a proxy for the strength of unions in 1989. For the measures related to the number of unionized workers - the last four variables - we display the statistics for 1988. For each measure, we display the average value and standard deviations across microregions. We also display the minimum value, maximum value, and the 10, 25, 50, 75, and 90 percentiles.

Table A5: Unions Strength and RTR – OLS regressions

	(1) 1994	(2) 1998	(3) 2002	(4) 2006	(5) 2010	(6) 2014	(7) 2018	(8) Pooled
A.1 - Union Employees Per Thousand People (Between 15 and 64 Years)								
Regional Tariff Reduction (RTR_r)	2.971	-20.55***	-22.09***	-17.06**	-23.22***	-24.65***	-25.19***	-20.78***
	(12.93)	(6.538)	(6.443)	(7.878)	(5.417)	(4.946)	(5.593)	(4.972)
A.2 - Union Employees Per Capita (Thousand People)								
Regional Tariff Reduction (RTR_r)	1.507	-12.53***	-13.74***	-11.13**	-14.35***	-15.18***	-15.53***	-12.83***
	(7.838)	(3.965)	(4.041)	(4.450)	(3.405)	(3.084)	(3.357)	(2.980)
B.1 - Unions per Thousand People (Between 15 and 64 Years)								
Regional Tariff Reduction (RTR_r)	-0.580	-1.212**	-1.087*	-1.181**	-1.134**	-1.108**	-0.935	-1.120**
	(0.559)	(0.559)	(0.563)	(0.560)	(0.530)	(0.527)	(0.578)	(0.545)
B.2 - Unions per Capita (Thousand People)								
Regional Tariff Reduction (RTR_r)	-0.360	-0.752**	-0.701**	-0.762**	-0.745**	-0.723**	-0.606*	-0.723**
	(0.339)	(0.340)	(0.343)	(0.342)	(0.322)	(0.317)	(0.351)	(0.330)
C.1 - Unionized Workers per Thousand People (Between 15 and 64 Years)								
Regional Tariff Reduction (RTR_r)	N/A	-142.8	-296.2***	-345.3***	N/A	-225.6**	N/A	-251.0***
		(92.45)	(92.21)	(100.5)		(91.56)		(84.94)
C.2 - Unionized Workers per Capita (Thousand People)								
Regional Tariff Reduction (RTR_r)	N/A	-85.52	-187.4***	-221.9***	N/A	-141.5**	N/A	-159.5***
		(54.38)	(53.75)	(60.72)		(55.48)		(49.22)
Number of Microregions	474	474	474	474	474	474	474	474
Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Coefficients obtained from OLS regressions of the changes in the variables described in Panels titles on regional tariff reductions. All regressions control for state x time fixed effects and 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The pooled regression (column 8) controls for a interaction between the RTR_r and a dummy indicating the transition years and interactions between control variables and year dummy variables. Each panel estimates the regression using a different dependent variable. Panel A.1 uses the changes in the number of workers employed in labor unions per thousand people between 15 and 64 years old at each region as the dependent variable. Panel A.2 uses the changes in the number of union employees per thousand people. In Panel B.1, the dependent variable is the change in the number of formal union establishments per thousand people between 15 and 64 years old. Panel B.2 uses the changes in the number of unions per thousand people as the dependent variable. The dependent variable in Panel C.1 is the changes in the number of unionized workers per thousand people between 15 and 64 years old. Panel C.2 uses the changes in the number of unionized workers per thousand people as the dependent. Standard errors in parentheses, clustered at the mesoregion level (112 clusters).
Significance at the *10%, **5%, *** 1% levels.

Table A6: Vote Share for the Left and Regional Tariff Reduction – Median Regressions

Dep var: Δ_{1989-t} Vote share for left-wing parties	(1) $t = 1994$	(2) $t = 1998$	(3) $t = 2002$	(4) $t = 2006$	(5) $t = 2010$	(6) $t = 2014$	(7) $t = 2018$	(8) Pooled
Panel A: Main Specification								
Regional Tariff Reduction (RTR_r)	-0.952***	-0.947**	-1.060**	-0.873*	-1.296**	-1.114*	-2.256***	-1.145***
	(0.310)	(0.352)	(0.495)	(0.459)	(0.578)	(0.615)	(0.622)	(0.365)
Panel B: Effective Protection								
Regional Effective Tariff Reduction ($RETR_r$)	-0.619***	-0.592***	-0.515**	-0.495**	-0.691**	-0.627**	-1.356***	-0.634***
	(0.165)	(0.209)	(0.242)	(0.231)	(0.304)	(0.300)	(0.428)	(0.176)
Panel C: Second-Round of Voting								
Regional Tariff Reduction (RTR_r)	N/A	N/A	-1.369**	-1.659**	-1.779***	-2.019***	-1.933***	-1.721***
	N/A	N/A	(0.556)	(0.772)	(0.592)	(0.469)	(0.698)	(0.411)
Observations	484	484	484	484	484	484	484	3,388 or 2,420

Notes: Coefficients obtained from median regressions of the changes in vote share for left-wing parties on regional tariff reductions. All regressions control for state x time fixed effects and the 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the 1991 Gini index. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Panel A presents our benchmark specification. Panel B uses the alternative measures for labor demand shock using the changes in effective rates of protection instead of the changes in nominal tariffs. Panel C uses runoff elections to construct the dependent variable. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

Table A7: Vote Share for the Left and Regional Tariff Reduction – Alternative Sample

Dep. var: δ Vote share for left-wing parties	(1) t = 1994	(2) t = 1998	(3) t = 2002	(4) t = 2006	(5) t = 2010	(6) t = 2014	(7) t = 2018	(8) Pooled
A. Main Specification								
Regional Tariff Reduction (RTR_r)	-0.849***	-1.285***	-0.828**	-1.063**	-1.133**	-1.004**	-1.732***	-1.128***
	(0.311)	(0.350)	(0.402)	(0.414)	(0.447)	(0.493)	(0.489)	(0.363)
B. Effective Protection								
Regional Effective Tariff Reduction ($RETR_r$)	-0.485***	-0.677***	-0.492*	-0.676***	-0.759***	-0.690**	-1.067***	-0.692***
	(0.183)	(0.208)	(0.250)	(0.255)	(0.269)	(0.273)	(0.276)	(0.219)
C. Runoff Elections								
Regional Tariff Reduction (RTR_r)	N/A	N/A	-1.229***	-1.289***	-1.820***	-1.776***	-1.912***	-1.605***
			(0.395)	(0.411)	(0.499)	(0.500)	(0.501)	(0.418)
Observations	474	474	474	474	474	474	474	3,318 or 2,370

Notes: Coefficients obtained from an OLS regressions of the changes in vote share for left-wing parties on regional tariff reductions. All regressions control for state x time fixed effects and the 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the 1991 Gini index. The pooled regression (column 8) controls for interactions between control variables and year dummy variables. Panel A presents our benchmark specification. Panel B uses the alternative measures for labor demand shock using the changes in effective rates of protection instead of the changes in nominal tariffs. Panel C uses runoff elections to construct the dependent variable. Standard errors in parentheses, clustered at the mesoregion level (112 clusters).
Significance at the *10%, **5%, *** 1% levels.

Table A8: Regional Tariff Reduction and Elections: Heterogeneous Effects

Dep var:	(1) V.sh. for the Left	(2) V.sh. for the Left (Runoff Elections)	(3) V.sh. for Union Parties
RTR_r in regions with unions in 1989 and above the first decile of the 3 dimensions considered	-1.175**	-1.700***	-1.113**
	(0.372)	(0.424)	(0.389)
RTR_r in regions without unions in 1989	-0.623	-0.497	-0.023
	(0.557)	(0.755)	(0.569)
RTR_r in the 10% less populated regions in 1989	-1.217	-1.933***	-1.221**
	(0.485)	(0.551)	(0.554)
RTR_r in the 10% less urbanized regions in 1991	-1.312	-1.657**	-0.955
	(0.877)	(0.803)	(0.776)
RTR_r in regions with fewer workers in manufacturing in 1991	0.134	-1.567*	-0.857
	(0.829)	(0.856)	(0.680)
Observations	3,381	2,415	3,381

Notes:Coefficients obtained from a pooled OLS regressions of the dependent variable in column title on regional tariff reduction and the interactions of regional tariff reduction and the indicators that the region had no presence of unions in 1989, the indicator that regions were one of the 10% less populated in 1989, the indicator that the region was one of the 10% less urbanized regions in 1991, and the indicator that the region was one of 10% with few workers in manufacturing sector. All regressions control for state x time fixed effects and the interaction between year dummy variables and the following controls: 1991 shares of whites, females, high-school graduates, and individuals aged over 60 in the population, 1991 share of workers employed in manufacturing, employment and urbanization rates, and the Gini index in 1991. The first row represents the effect of regional tariff reduction in regions with the presence of unions in 1989, the 90% most populated, the 90% most urbanized, and the 90% with more workers in manufacturing. the second row displays the effect of regional tariff reduction for the regions without unions but in the 90% most populated, 90% most urbanized, and 90% with more workers in manufacturing. The interpretation of rows 3, 4, and 5 are similar to the one from row 2, but changing the presence of unions for being one of the 10% less populated, 10% less urbanized or 10% with fewer workers in manufacturing, respectively. Standard errors in parentheses, clustered at the mesoregion level (114 clusters).
Significance at the *10%, **5%, *** 1% levels.

Table A9: Consistent Industry Classification Across RAIS and Tariff Data

	Industry	Industry Name	Nivel 50	1985-1993 RAIS (Sub-atividade)	RAIS CNAE 1.0
Tradable	1	Agriculture	1	0110-0360; 5810	01112-05126
	2	Mineral Mining (Except Combustibles)	2	0510; 0520; 0540	13102-14214; 14290; 14293
	3	Petroleum and Gas Extraction and Coal Mining	3	0530	10006-11207
	4	Nonmetallic Mineral Goods Manufacturing	4	1010-1090	26115-26999
	5	Iron and Steel, Nonferous, and Other Metal Production and Processing	5-7	1100-1190	27111-28126; 28312-28436; 28916-28991
	8	Machinery, Equipment, Commercial Installation Manufacturing, and Tractor Manufacturing	8	1210-1250; 1270; 1280	28134-28223; 29114-29726
	10	Electrical, Electronic, and Communication Equipment and Components Manufacturing	10-11	1310-1330; 1350; 1370-1390	29815; 29890; 30112-31410; 31518; 31526; 31917-32301; 33308
	12	Automobile, Transportation, and Vehicle Parts Manufacturing	12-13	1340; 1410-1490	31429; 31607; 34100-35998
	14	Wood Products, Furniture Manufacturing, and Peat Production	14	1510-1570; 1610-1650; 1690	20109-20290; 36110-36145
	15	Paper Manufacturing, Publishing, and Printing	15	1710-1750; 1790; 2910; 2920; 2980; 2990; 3060	21105-22349
	16	Rubber Product Manufacturing	16	1810-1850; 1890	25119-25194
	17	Chemical Product Manufacturing	17,19	2000; 2030; 2050-2090; 2750	23108; 23302-24198; 24295; 24619-24694; 24724; 24813-24996
	18	Petroleum Refining and Petrochemical Manufacturing	18	2010-2020	23205-23299; 24210; 24228; 24317-24422
	20	Pharmaceutical Products, Perfume, and Detergents Manufacturing	20	2110; 2210; 2220	24511-24546; 24716;24732
	21	Plastics Products Manufacturing	21	2310-2370; 2390	25216-25291
	22	Textiles Manufacturing	22	2410-2460; 2490; 2550	17116-17710
	23	Apparel and Apparel Accessories Manufacturing	23	2510; 2520; 2540; 5320	17728-18228
	24	Footwear and Leather and Hide Products Manufacturing	24	1910-1930; 1990; 2530	19100-19399
	25	Food Processing (Coffe, Plant Products, Meat, Dairy, Sugar, Oils, Beverages, and Others)	25-31	2040; 2600-2690; 2710-2740; 2810-2830	14222; 15113-16004
	32	Miscellaneous Other Products Manufacturing	32	1260; 2230; 3000-3050; 3070-3090	33103; 33200; 33405; 33502; 36919-37206; 92118
Nontradable (99)	91	Utilities	33	3510-3530	40100-41009; 90000
	92	Construction	34	3410-3490; 5270	45110-45608
	93	Wholesale and Retail Trade	35	4110-4195; 4301-4307; 4310-4395; 5830	50105; 50300; 50415; 50504-52698; 63118; 63126
	94	Financial Institutions	38	4510-4540; 4640-4660; 5840	65102-67202
	95	Real Estate and Corporate Services	40; 41	4610-4630; 5460; 5520; 5710-5770; 5790; 5890	70106-72907; 74110-74500; 74918-74993; 92215; 92223; 92401
	96	Transportation and Communications	36; 37	4711-4760; 4810-4830; 5820	60100-62308; 63215-64203
	97	Private Services	39; 43	5110;5120; 5210-5260; 5290; 5310; 5390; 5410-5450; 5490; 5510; 5780; 6130-6190; 6210-6230; 6310; 6320; 9000	28819; 28827; 29912-29963; 31810-31895; 32905; 33910-33944; 50202; 50423; 52710-55298; 73105; 73202; 74608; 74705; 80110-85200; 91111-91995; 92126; 92134; 92312-92398; 92517-95001
	98	Public Administration	42	3540; 6110; 6120; 7011-7029; 8010; 8020	75116-75302; 85316; 85324; 99007

Appendix B - Control Variables

B.1 Real Exchange Rate Controls

We construct regional exchange rates controls as in Dix-Carneiro and Kovak (2017). We first calculate industry-specific real exchange rates (IER) as a weighted average of country-specific real exchange rates between Brazil and its trade partners,[35] weighted either by country's 1989 import share or export share. Formally:

$$IER_{it}^{K} = \sum^{P} S_{(ip,1989)}^{K} er_{pt},$$

where $S_{(ip,1989)}^{K}$ is country p's share of k ={imports, exports} in industry i in 1989,[36] and er_{pt} is the country-specific real exchange rate between Brazil and country p at time t.[37] Then, we use IER to construct regional real exchange rate shocks as the weighted average of the changes in log IER from 1990 to each subsequent electoral year, weighted by the 1991 industry distribution of employment. That is:

$$RER_{rt}^{K} = \sum_{i=1}^{I} \lambda_{ri} \Delta_{(t-90)} ln(IER_{it}^{K}),$$

where λ_{ri} is the 1991 share of region r workers employed in industry i.

[35] Exchange rates are expressed as the number of Brazilian reais needed to buy a dollar.

[36] Each country's 1989 shares of Brazil's imports and exports in each industry are constructed using UN-Comtrade data.

[37] Exchange rates data come from revision 10.0 of the Penn World Table. Real exchange rates can differ from other studies, such as Dix-Carneiro and Kovak (2017), since they use revision 7.1 of the Penn World Table, which uses ICP prices for 2005 as the benchmark instead of 2011's.

B.2 Tariff Changes due to Mercosur

We construct two variables to control for tariff changes due to Mercosur. The first variable (MT_r^{exp}) controls for the import tariff changes imposed by other members of Mercosur to the Brazilian goods. The second variable (MT_r^{imp}) controls for the changes in the Brazilian tariffs to imported goods from the Mercosur members.

To construct these controls, we first calculate 1991 and 1995 industry-specific import tariffs from Mercosur members (IT_{it}^{exp}) and to Mercosur partners (IT_{it}^{imp}). The industry-specific import tariffs from Mercosur members are constructed as the weighted average of tariffs to Brazilian products in each partner p, using the share of Brazil's exports to each country as weights. Similarly, the industry-specific import tariffs to Mercosur partners are calculated as the weighted average of Brazilian import tariffs to Mercosur partners, using the share of Brazil's import from each partner p as weights. Formally,

$$IT_{it}^{exp} = \sum^{P} S_{(ip,1991)}^{exp} \tau_{ipt}^{exp}, \tag{3}$$

$$IT_{it}^{imp} = \sum^{P} S_{(ip,1991)}^{imp} \tau_{ipt}^{imp}, \tag{4}$$

where $S_{(ip,1991)}^{exp}$ is the 1991 share of Brazilian exports to country p in industry i, $S_{(ip,1991)}^{imp}$ is the 1991 share of Brazilian imports from country p in industry i, τ_{ipt}^{exp} is the import tariff to Brazilian goods imposed by Mercosur members p in industry i and year $t = \{1991, 1995\}$, and τ_{ipt}^{imp} is the Brazilian import tariff to country p goods in industry i and year $t = \{1991, 1995\}$.

Using the differences in the industry-specific import tariffs from 1991 to 1995, we construct MT_r^{exp} and MT_r^{imp} similarly to the regional tariff reduction. That is:

$$MRWT_r^{exp} = \sum_{i=1}^{I} \lambda_{r,i} \Delta_{95-91} ln(1 + IT_{it}^{exp}), \tag{5}$$

$$MRWT_r^{imp} = \sum_{i=1}^{I} \lambda_{r,i} \Delta_{95-91} ln(1 + IT_{it}^{imp}), \tag{6}$$

where λ_{ri} is region r share of workers employed in industry i in 1991.

The share of Brazilian exports and imports are constructed using data from UN-Comtrade Database, and the Mercosur preferential tariffs at 4-digit ISIC level come from individual country sources, which has been compiled by Estevadeordal et al. (2008).[38] To calculate the industry-specific tariffs, we first aggregate the 4-digit ISIC-level tariffs to the industry level used in the whole paper using simple averages. We had to drop 4 ISIC industries, because there was no consistent correspondence with our industry level.

[38] See Estevadeordal et al. (2008) for more details about the data on Mercosur tariffs.

Appendix C - Data

C.1 RAIS

The Annual Relation of Social Information (RAIS) is an administrative data set reported by the Brazilian Ministry of Labor that provides high-quality data of Brazilian formal labor market. RAIS was instituted by the decree nº 76.900 on December 23 of 1975 to (i) Provide information regarding the formal labor market in Brazil, (ii) monitor the entry of foreign workers in the Brazilian labor market. (iii) provides statistical information for government decisions (iv) generate data for different governmental benefit programs as FGTS, unemployment insurances, PIS, PASEP, and "abono salárial".

Since report the information on every formal employee is a costly task for companies, the federal government created a mechanism to guarantee that the information reported are complete and accurate. First, companies that delay sending data or send false or incomplete information face fines until they send the complete information. Second, some governmental benefits paid to workers are conditional on having their job information correctly declared in RAIS. Hence, workers have an incentive to require the employer to send the correct information. Therefore, both workers and employers have an incentive to report accurate and complete information to the Ministry of labor, ensuring the quality of the data.

We collect the following information of RAIS:

1. The number of formal employees in each industry for the years between 1986 and 2015: To construct this variable, we considered only individuals between 15 and 64 years old, employed on the last day of December, and with a positive earning in that month. In cases in which the same worker appeared twice in the sample, we keep only the highest paying job in December. To get a consistent industry classification across RAIS data and the industry classification used in the paper, we classified the CNAE activities in RAIS using the correspondence displayed in table A9. This correspondence is mainly based on the classification from CNAE codes to Nivel 50 available at `https://concla.ibge.gov.br/classificacoes/correspondencias/atividades-economicas.html`

2. The number of employees working in state-owned firms: Starting in 1995, RAIS indicates whether the company is state-owned or not. Hence, from 1995 to 2014, we have the number of workers in state organizations. Again, we only consider the individuals between 15 and 64 years old, employed on the last day of December, with a positive earning in that month, and we keep only the highest paying job in December.

3. The number of people that work in labor unions: As commented before, we cannot identify whether the company is either a labor union, a professional association, or an employers' association. For this reason, we construct the variable of union employees as the number of people who works for one of this three group of companies. As before,

we only consider the individuals with age between 15 and 64 years old, employed on the last day of December, and with a positive earning in that month. To avoid counting the same worker twice in each region, we count each worker only once per microregion. We also conduct a robustness test counting each union worker only once, considering its highest salary, and the results are very similar.

4. The number of labor unions in each year: This variable is constructed as the number of establishments with at least one formal employee registered in December classified in one of the three activity groups mentioned in the previous bullet point.

C.2 PNAD

The National Household Sample Survey (PNAD) is an administrative data set reported annually by IBGE, with exception for 1994 and the census years (1980, 1991, 2000 and 2010). This data set provides general characteristics of the population, education, labor, income and housing at the household level, but only representative at the federal unit level. During the 49 years of existence, PNAD has been used by many researchers as an instrument to investigate the effects of policies and economic shocks on socioeconomic and living condition in Brazil.

In this paper we use the PNAD data set to access the percentage of workers that are associated to labor unions in each of the 27 federal units in Brazil. Since we do not have data by municipality or microregion, we use the RAIS data on the number of formal worker in each microregion and calculate an estimate for union density as a weighted average of the proportion of formal workers affiliated to labor unions in each industry, weighted by the fraction of formal workers in region r in that industry.