UNIVERSIDADE FEDERAL DO RIO DE JANEIRO

INSTITUTO DE ECONOMIA

PROGRAMA DE PÓS-GRADUAÇÃO EM ECONOMIA

Miguel Silveira Moreira

STATE TAXATION AND ASYMMETRIC FUEL PRICE ADJUSTMENT: EVIDENCE FROM BRAZIL'S ICMS

Rio de Janeiro

2023

Miguel Silveira Moreira

STATE TAXATION AND ASYMMETRIC FUEL PRICE ADJUSTMENT: EVIDENCE FROM BRAZIL'S ICMS

Dissertação de Mestrado submetida ao Programa de Pós-Graduação em Economia da Indústria e da Tecnologia, Instituto de Economia, Universidade Federal do Rio de Janeiro como requisito parcial à obtenção do título de Mestre em Economia.

Orientador: Prof. Eduardo Pontual Ribeiro

Rio de Janeiro

2023

FICHA CATALOGRÁFICA

M838s Moreira, Miguel Silveira.

State taxation and asymmetric fuel price adjustment: evidence from Brazil's ICMS / Miguel Silveira Moreira. – 2023.

59 f.; 31 cm.

Orientador: Eduardo Pontual Ribeiro.

Tese (doutorado) – Universidade Federal do Rio de Janeiro, Instituto de Economia, Programa de Pós-Graduação em Economia da Indústria e da Tecnologia, 2023.

Bibliografia: f. 53 - 56.

1. Tributos. 2. Assimetria. 3. Imposto sobre circulação de mercadorias e serviços. 4. Gasolina. I. Ribeiro, Eduardo Pontual, orient. II. Universidade Federal do Rio de Janeiro. Instituto de Economia. III. Título.

CDD 336.2

Ficha catalográfica elaborada pela bibliotecária: Luiza Hiromi Arao CRB/7 – 6787
Biblioteca Eugênio Gudin/CCJE/UFRJ

Miguel Silveira Moreira

STATE TAXATION AND ASSYMETRIC FUEL PRICE ADJUSTMENT: EVIDENCE FROM BRAZIL'S ICMS

Dissertação de Mestrado submetida ao Programa de Pós-Graduação em Economia da Indústria e da Tecnologia, Instituto de Economia, Universidade Federal do Rio de Janeiro como requisito parcial à obtenção do título de Mestre em Economia.

Rio de Janeiro, 19 de Setembro de 2023

Prof. Eduardo Pontual Ribeiro
Universidade Federal do Rio de Janeiro (IE-UFRJ)

Prof. Pedro James Frias Hemsley
Universidade Federal do Rio de Janeiro (IE-UFRJ)

Dr. Douglas Sad Silveira
University of Alberta

À Fernanda e Daniele.

AGRADECIMENTOS

Primeiramente gostaria de agradecer à minha família e amigos pelos contínuo suporte e incentivo, sem os quais este trabalho não seria possível. Em especial para minha mãe Fernanda, meu pai Emil e meus irmãos Pedro e José.

Agradeço a Daniele, minha companheira de vida, pelos conselhos e companheirismo. Sua presença foi essencial não somente para este trabalho final, mas durante toda a trajetória.

Gostaria de também de agradecer ao meu orientador, Profº Eduardo Pontual Ribeiro, pelos valiosos aprendizados, pela dedicação e pelo suporte inestimável ao longo deste percurso. Agradeço pela paciência nos momentos de dúvida e pelas horas dedicadas a tornar este trabalho melhor e mais completo.

Aos meus ex-colegas de CADE, especialmente Fernando e Ricardo, que me ajudaram a navegar as nuânces da defesa da concorrência no Brasil. O tempo que passei no DEE foi fundamental para minha formação.

Gostaria de agradecer à UFRJ, pelos anos investidos em minha formação. Este foi o lugar em que encontrei ambiente de aprendizado enriquecedor, e onde conheci professores e colegas incríveis que muito me ensinaram.

Por fim, agradeço ao Conselho Nacional de Desenvolvimento Científico e Tecnológico (CNPq) pelo suporte financeiro que tornou este estudo possível.

RESUMO

Este trabalho estima os efeitos possivelmente assimétricos da incidência do ICMS sobre a gasolina, usando uma nova metodologia de identificação baseada na "*substituição tributária*" dos preços da gasolina no Brasil. As estimativas mostram que não há antecipação nos preços praticados em reposta à mudanças no valor do ICMS cobrado, mesmo que este seja divulgado com antecedência. Os resultados empíricos, frutos da aplicação de um modelo NARDL, indicam repasse possivelmente assimétrico da variação dos impostos aos preços no longo prazo, mas com efeitos de aumento e quedas do imposto muito próximos de um e variáveis no tempo. Mudanças positivas e negativas no tributo parecem ser integralmente repassadas aos consumidores.

Palavras-chave: Repasse asimétrico; NARLD, gasolina.

ABSTRACT

This work estimates both long run and short run possibly asymmetric tax incidence effects of sales tax changes, using a novel identification methodology from ICMS tax withholding or "*substituição tributária*" of gasoline prices in Brazil. The estimates show that there is no anticipation of tax changes, even if new taxes in Reals are publicized in advance. The empirical results, deriving from the application of a NARDL model, indicate possibly asymmetric pass-though of taxes to prices but with the positive and negative effects very close to one and time varying. The tax appears to be fully shifted.

Key-words Asymmetric pass-though, NARDL; gasoline.

List of Figures

List of Tables

LIST OF ABBREVIATIONS AND ACRONYMS

ANP	Agência Nacional do Petróleo, Gás Natural e Biocombustíveis
CADE	Conselho Administrativo de Defesa Econômica
CONFAZ	Conselho Nacional de Política Fazendária
ECM	Error Correction Model
ICMS	Imposto sobre circulação de Mercadorias e Serviços
IBGE	Instituto Brasileiro de Geografia e Estatística
NARDL	Non-linear Autoregressive Distributed Lags
PMPF	Preço Médio Ponderado ao Consumidor Final
SEFAZ	Secretaria de Fazenda
VAR	Vector Autoregressive
VAT	Value added tax

CONTENTS

1 Introduction

The dynamics of gasoline prices and their relationship to collected taxes have long been a topic of interest for both researchers and consumers. Understanding the asymmetric transmission of tax changes to gasoline prices is crucial for policymakers, as it can provide insights into the behavior of fuel markets and guide effective policy interventions. Moreover, consumers are highly sensitive to fluctuations in gasoline prices, making it essential to investigate the factors driving price adjustments and their nature. Lastly, this work may be of interest to policymakers in public finances, as the price's response to ICMS's variations holds significant relevance and implications for fiscal policies and public revenue management and its influence on the final consumer price. Previous studies have explored this topic, but there is still a need for further research, particularly in the Brazilian context.

The supply chain of liquid fuels and petroleum-derived lubricants in Brazil is intricately tied to the Tax on Circulation of Goods and Services (ICMS)[1], the largest non-federal tax in the country. The ICMS is an ad valorem tax that applies variable percentage rates based on the value added at each stage of production. Recent changes in the ICMS tax of fuels (Lei Complementar 19, 2022) raised significant debate but ad earth of studies of the incidence of this tax.

However, the implementation of this tax in Brazil deviates from its traditional form, as tax collection responsibility has been shifted to the refineries to minimize tax evasion. This tax withholding, known as "substituição tributária", streamlines tax collection by reducing the number of tax collectors, and the burden is eventually borne by consumers.

Each state's fiscal authority estimates an expected price to the final consumer to make the tax withholding mechanism possible. This is the so-called "Weighted Average Price to Final Consumers" (PMPF)[2], which is calculated bi-weekly.

The PMPF serves as the tax base for ICMS incidence and is determined through a survey of fuel prices charged at the pump. It is published on the National Board of State Tax Agencies (CONFAZ) website. The PMPF calculation takes into account the prices charged for different fuels and the volumes sold, with larger, high-volume locations having a significant influence on the final estimated

[1] Direct translation from the portuguese "*Impostos sobre Circulação de Mercadoria e Serviços*"
[2] Direct translation from portuguese "*Preço Médio Ponderado a Consumidor Final (PMPF)*"

value.

This work aims to examine the impact of taxes on the retail and distribution prices of Common Gasoline, focusing on changes in the ICMS tax base due to the bi-weekly publication of the PMPF. The present work intends to understand how these variations in costs are transmitted to the final consumer price, considering the unique characteristics of the PMPF calculation as a volume-weighted average. Additionally, the paper explores an alternative causal chain where changes in tax due at the refinery level may impact retail prices directly or indirectly through anticipation by retailers or distributors.

This work aims to fill this gap by examining the possibly asymmetric pass-through of changes in the ICMS tax, a significant component of gasoline prices in Brazil, to both retail and distribution prices of Common Gasoline.

The incidence of sales tax is a topic with a long literature (Atkinson and Stiglitz (1972), e.g.). More recent studies have considered both asymmetric effects (Benzarti et al., 2020; Asplund et al., 2000 and Schmerer and Hansen, 2023) and anticipation effects (Coglianese et al., 2017). The main concern is that tax increases may be fully shifted to consumers, and that tax decreases may be cashed in by retailers at the expense of consumers. While asymmetric effects of cost increases in fuel have been studied in Brazil (Melo et al., 2021 on LPG, Canêdo-Pinheiro, 2012 on diesel), there are no models that account for the possibility of asymmetric responses both in the short run and long run.

By employing a NARDL model we estimate the impact of changes in the ICMS due to the retail and distribution prices of Common Gasoline. The models account for pass-through asymmetries in both the short and long run. The methodology proposed presents a flexible framework, allowing the construction of restricted models of interest, as imposing symmetry conditions for the short and long runs alternatively.

The results indicate that short-run pass-through asymmetries are statistically insignificant in the majority of the periods analyzed. However, significant long-run asymmetries are observed. For the retail price, positive shocks in the ICMS collection lead to a higher average pass-through compared to negative shocks in the first three periods. In the fourth period, the opposite trend is observed, with negative changes in costs being slightly more passed on to the final price. Similar patterns are observed for the distribution price, with statistically significant long-run asymmetries.

It's important to note that, even though the long-run asymmetries are statistically significant, they are numerically close, and both negative and positive shocks tend to integral passthrough.

Additionally, the models that consider asymmetries only in the short run also fail to find statistically significant asymmetries, except for the first period in the retail price model. The overall pass-through estimates remain similar to the unrestricted models.

Finally, the models that impose symmetry in both the short and long run yield similar cumulative pass-through effects, with values close to unity for most periods, indicating integral pass-through from ICMS variations.

These findings are in line with previous studies that have reported short run asymmetries in fuel markets. The results suggest that, in the Brazilian context, pass-through asymmetries are more pronounced in the long run compared to the short run. However, the overall pass-through remains close to unity, indicating a significant transmission of changes in the ICMS to gasoline prices.

The subsequent sections of the study will delve deeper into the institutional context, data sources, and existing literature on the topic to further analyze and understand the relationship between taxes and fuel prices in Brazil.

2 Institutional Context

The supply chain of liquid fuels and petroleum-derived lubricants is interwoven with a multitude of taxes, among which the ICMS (Tax on Circulation of Goods and Services) represents the largest non-federal tax. Governed by individual federal entities, the ICMS functions as an ad valorem tax - VAT or *vlaue added tax* - that applies variable percentage rates based on the value added, Δ in Figure 1, at each stage of production. Interestingly, this tax implementation in the Brazilian context deviates from its traditional form; administrative strategies have shifted the tax collection responsibility to the refineries in an effort to minimize tax evasion.

These refineries are burdened with tax withholding, or "substituição tributária" in Portuguese, which streamlines tax collection by narrowing down the pool of tax collectors. The tax is eventually borne by consumers, after being transferred through the fuel supply chain. Such withholding requires an accurate estimate of prices or profit margins across the supply chain, which is achieved through a bi-weekly calculation of the Average Consumer Price (PMPF). This PMPF, which sets the tax base for the ICMS, is determined by a survey of prices charged for different fuels at the pump and is published on the website of CONFAZ.

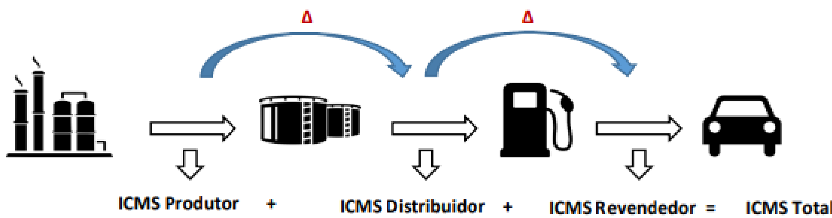
The tax withholding procedure was further formalized into a statutory consumption tax, which essentially translated the VAT tax rate into a fixed-value consumption tax. In this scenario, the tax due becomes a fixed amount (in Reals), determined by the product of the PMPF (estimated sale price) and the tax rate relative to the refinery price.

Building on this dynamic, this study seeks to examine the impact of taxes on the retail and distribution prices of Common Gasoline. The focus lies on the changes in the ICMS tax base due to the bi-weekly publication of the PMPF, which subsequently leads to shifts in the tax due. We aim to understand the transmission of these cost variations towards the final consumer price, given the unique characteristics of the PMPF calculation, such as the weightage of volume sold. We delve into an alternative causal chain where tax changes due at the refinery level impact retail prices either directly or indirectly through anticipation by retailers or distributors.

2.1 The value added tax (ICMS) in the fuel supply chain

Among all the taxes imposed on the sale of liquid fuels and petroleum-derived lubricants, the largest that's not under federal jurisdiction is the ICMS (Tax on Circulation of Goods and Services). ICMS is an ad valorem tax with rates de facto defined independently by each federal entity.

Figure 1: The ICMS incidence in the liquid fuels production chain.



Source: Reproduction of EPE (2020)

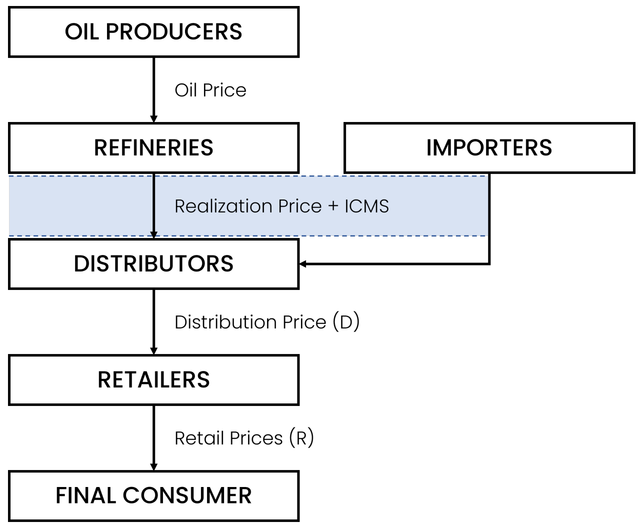
Figure 1 presents a diagram of the tax base on which the ICMS is imposed, and the base on which the tax due is calculated: the value added (Δ) throughout the supply chain, i.e. the difference between the cost of inputs bought and the price of goods sold. The tax rate is a percentage, not a fixed, rate. As a value-added tax, it should be collected at every stage of the production chain (Ebrill et al., 2001, e.g.). For administrative reasons, the ICMS on liquid fuels in Brazil is subject to a different implementation, where the refinery is responsible for withholding the tax due on the next steps of the fuel chain (distributors and retail gas stations).

Tax withholding or "*substituição tributária*" in Portuguese, aims to reduce the pool of taxpayers responsible for collecting tax in order to reduce tax evasion. The taxpayer chosen to withhold the tax usually have complex organizational structures with detailed commercial and tax records that can afford the additional burden.

In the case of petroleum-derived fuels and ethanol, the refinery and the importers are usually responsible for collecting and withholding the tax on subsequent operations.The burden generated by the anticipation of ICMS is successively transferred along the chain until it reaches the final consumer. (Cavalcanti, 2006)

The tax substitution regime requires an estimate of the price or profit margin throughout the

Figure 2: Gasoline production chain and the ICMS collection.



Source: Adaptation from Canêdo-Pinheiro (2012).

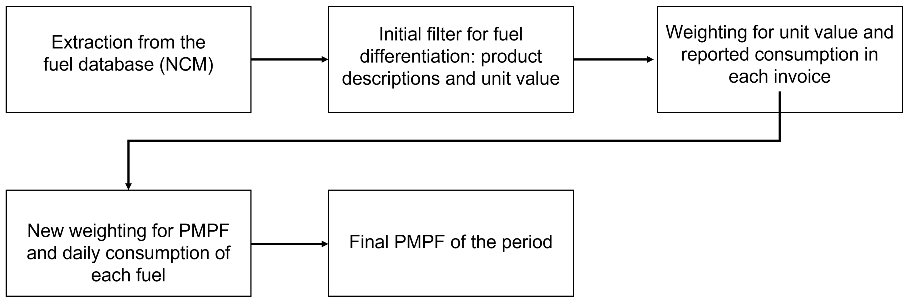
chain. The main methodology for calculating ICMS on fuels, currently adopted by all states and the Federal District, is based on an average price for the final consumer (PMPF). The PMPF is calculated and published every two weeks on the CONFAZ website.

A survey based on prices charged at the pump for Regular Gasoline, Premium Gasoline, Ethanol, Diesel S-500, and Diesel S-10 determines the PMPF base price. PMPF values for the first half of a given month are determined by the retail prices charged in the first half of the previous month. This same dynamics is used for the second half of each month.

The PMPF calculation is carried out by the Treasury Department (SEFAZ) of each state, using data from all municipalities in the state. The SEFAZ of each state uses the prices and volumes registered with each invoice at the tax agencies. The consumed amount of each fuel determines the proportion used in the final PMPF calculation. Figure 3 shows a flowchart of the calculation methodology used by each state's SEFAZ.

Note that the ICMS tax withholding effectively translates the VAT tax rate into a fixed value (in Reals) consumption tax. When posting the PMPF, the tax due is now fixed in Reals, from the calculation of the PMPF estimate of the sale price with respect to the refinery price times the tax rate. In fact, in 2023, a constitutional law changed altered the exceptional ICMS tax withholding procedure into a statutory consumption tax, where the tax due is defined in Reals.

Figure 3: Methodology for calculating PMPF



Source: Translation and reproduction of the Technical Note ‘Metodologia de Cálculo dos PMPF dos Combustíveis”. (SEFAZ/ES, 2020)

Recently, the ICMS incidence attracted great political attention. In late 2021 the state governments froze adjustments in the PMPF, as a measure to control the liquid fuel consumer price. (Rodrigues, 2022, Pupo, 2022). In addition, the federal government unilaterally, and temporarily reduced federal taxes. Presidential and state government elections were held in October 2022.

In 2023, the state’s transition (back) to a variable ICMS is marked by a simultaneous change in fuel ICMS. The Complementary Law nº192 of 2022 established a change from an *ad valorem* to an *ad rem* tax, with a uniform value per liter to all states (Verdélio, 2023). Additionally, the ICMS will vary in longer intervals than the previous fortnight. The Results section discusses how this may affect Regular Gasoline prices in Brazil.

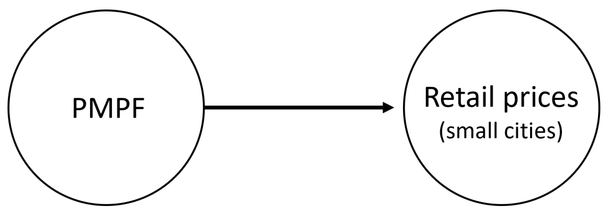
2.2 PMPF and retail prices.

The proposed causal effect identification mechanism uses the PMPF calculation methodology to analyze the effect of taxes on the retail and distribution prices of liquid fuels. The tax that makes up the final price of fuels in each state may change from two effects: the current tax rate and the PMPF published by the state treasury department. Due to the withholding tax regime, the value of ICMS depends on the ”presumed” value of the retail price, the PMPF, which is based on prices observed in the retail market in the previous month.

As mentioned above, the PMPF is published biweekly by CONFAZ (National Council of Fiscal Policy) through the ICMS agreement signed with the federal entities. The published PMPFs alter the ”tax base” of the ad valorem tax. This generates exogenous changes - with respect to the retail price - in the tax due. Our interest is to investigate how the transmission of these costs changes toward the final consumer price.

An important characteristic of the calculation of the PMPF carried out by the treasury departments is explored: the weighting of the price practiced by the volume sold. Given volumes sold across municipalities, the PMPF - a weighted average of gas station prices and volumes, is predominantly determined in large, high-volume locations. Therefore, in smaller cities, changes in tax due would be exogenous in relation to market dynamics, if local demand shocks are independent of large market shocks. Nevertheless, aggregate, large market shocks can be controlled for in a panel analysis. The general path proposed can be observed in Figure 4.

Figure 4: Identification I



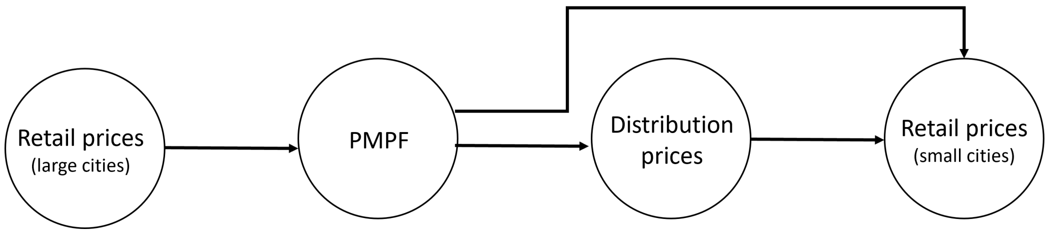
Source: Self elaboration.

The tax increase at the refinery may reach the retail gas station either because the input bought at the distributor has a higher price that is transmitted to the retail. Or the retailer itself anticipates an increase in the distribution price and passes on the expected tax increase directly, before purchasing fuel at the higher tax price from the distributor.

In either case, the changes in fuel costs to retailers are exogenous to local demand conditions in smaller localities, isolated from large markets, as the PMPF calculation basis accounts for the volume of consumption. The additional causal chain may change the timing of the effect, from the regular distributor-retailer supply cycle to an immediate change when the new tax is due. The second possible causal chain points to the need to control for anticipation of the tax increase by retailers (or distributors). Figure 5 presents this alternative causal chain.

The proposed methodology aims to contribute to the discussion on how taxes affect the retail and distribution price, and if gas stations and distributors anticipate a tax variation. This methodology also makes it possible to assess the existence of asymmetries between tax increases and decreases,

Figure 5: Identification II



Fonte: Elaboração própria.

both in the short and long run, and to estimate the cumulative effect of these changes in the investigated price. Thus, this work aims to analyze both the short-run and long-run relationship between the ICMS and the price.

3 Data

This section will present the data sources used in this study. The main sources used are the weekly price data for the retail and distribution of gasoline, from the Brazilian National Agency of Petroleum, Natural Gas and Biofuels (ANP). The PMPF prices are published every 15 days by the National Council for Fiscal Policy (CONFAZ). And population (municipality size) obtained from the Brazilian Institute of Geography and Statistics (IBGE). A similar data set is used by Melo et al. (2021) to estimate pass-through in the fuel chain given an exogenous change in a tax increase.

3.1 Gasoline prices

The weekly prices for the retail and distribution of gasoline were obtained from ANP, the Brazilian regulatory agency of the fuel sector.[3] This is the main data source for the proposed methodology, and its developed in the context of the Fuel Price and Commercialization Margins Survey (LPMCC), a survey held by ANP that collects the price for several fuels in gas stations across Brazil, using a seemingly random selection of gas stations.

The ANP has been conducting a version of the LPMCC since 1996. This survey aims to inform the general public about fuel prices and margins. Citizens could compare local prices or prices from a gas station with regional and national prices at the ANP website. This was expected to increase competition from increased transparency.

However, the survey has changed since its inception in the early 2000's. The changes may be significant and create non-representative or large changes in average prices. We document the most significant changes, that will influence our data time and regional coverage.

Up to October 30, 2004, the resale and distribution prices were calculated using a simple arithmetic average at the state, region, and national level. From November 2004 these average prices were weighted averages based on the sales information provided by distributors to ANP. For this reason only periods later than this date were chosen for the present work.

Later, in August 2015, due to budgetary issues, the number of surveyed localities was reduced

[3] The used data can be found at https://www.gov.br/anp/pt-br/assuntos/precos-e-defesa-da-concorrencia .

from 555 to 501. The exclusion of localities was limited to municipalities with fewer than nine operating automotive fuel retail stations in March 2015.

Between July 30, 2017, and December 30, 2017, the sample was further reduced: the geographical localities decreased from 501 to 459. More importantly, the survey frequency changed: State Capitals (almost always the largest city in the state) and the Federal District were surveyed weekly, while another 432 municipalities were surveyed every other week on alternating weeks. On December 31, 2017, the weekly frequency was resumed for all 459 municipalities included in the LPMCC (Fuel Price and Commercialization Margins Survey).

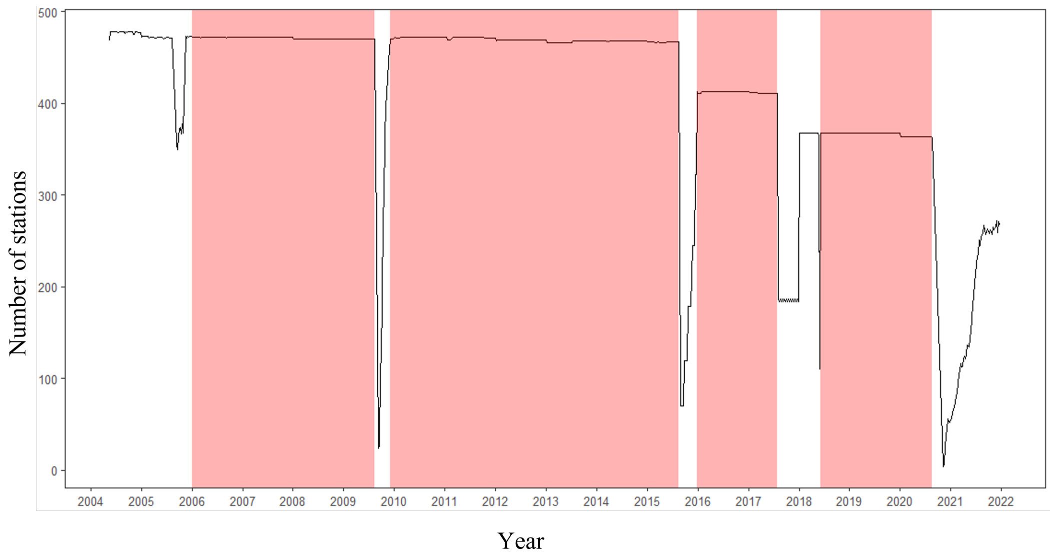
The most significant and relevant discontinuity in the survey occurred with the renewal of the survey contract, in late 2021, as can be seen in Figure 6. Similar to what happened between July 30, 2017, and December 30, 2017, the sample was reduced, and a staggered implementation of fieldwork of the survey was created. The reduction in the number of interviewed stations was more significant than in the previous case. In this more recent case, the groups were not periodically surveyed, and a progressive implementation schedule was created, where the locations to be added were mostly at the discretion of the contracted company. Additionally, in the absence of more detailed explanations, the stations interviewed within a location or municipality are also at the discretion of the contracted company.

To surpass the above problems of continuity in the data set, the panel data used is composed of four periods that maximize the sample size across time, while mitigating the effect of survey changes. The data was selected as follows:

- 1º period: from 01/01/2006 to 15/08/2009, or 189 weeks;
- 2º period: from 05/12/2009 to 15/08/2017, or 298 weeks;
- 3º period: from 23/12/2015 to 29/07/2017, or 84 weeks; and
- 4º period: from 09/06/2019 to 22/08/2020, or 116 weeks;

This division can be visualized in Figure 6, and it’s clear where the interruptions and discontinuities in the survey held by ANP are localized. Four periods are selected based on the criteria of creating continuous periods where it's possible to avoid survey discontinuities and create the biggest possible balanced panels.

Figure 6: Gas stations interviewed every week



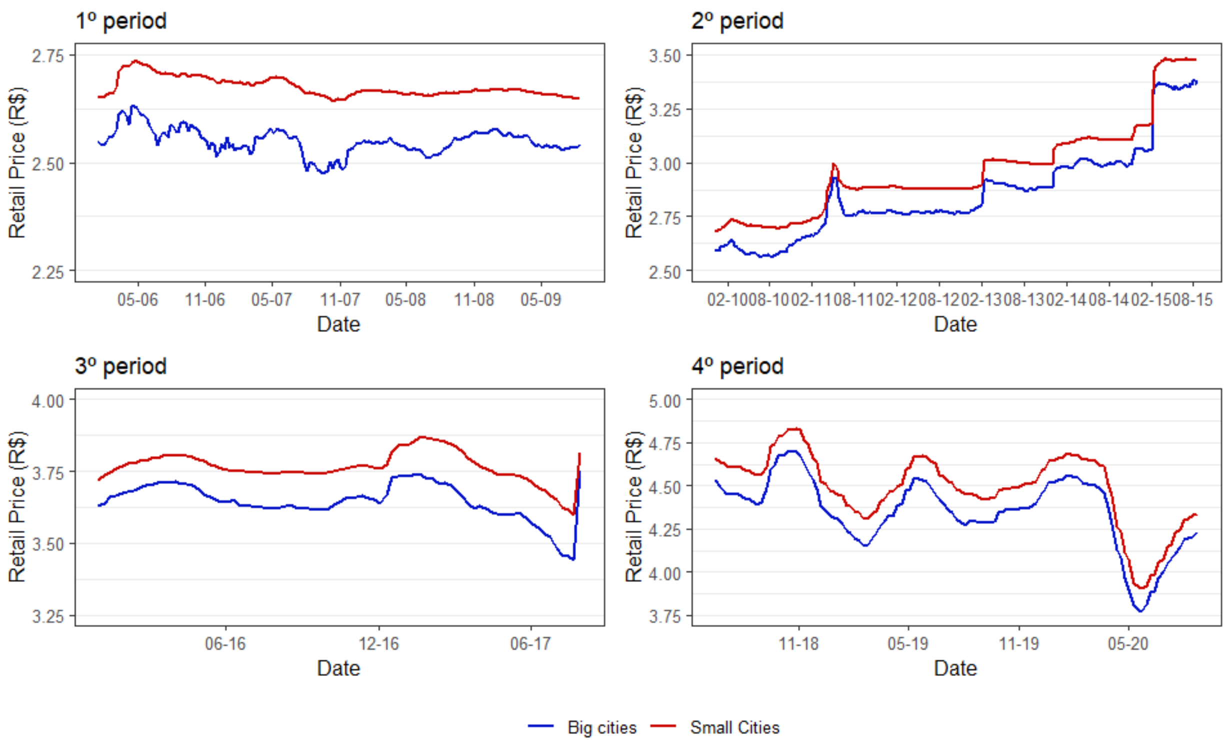
Source: self-elaboration based on data retrieved from ANP's website. The periods coloured in red signalize the periods chosen to carry the estimation of the proposed methodology.

Figures 7 and 8 display the mean price difference between large cities, those with at least 300.000 residents, and smaller cities, those with less than 300.000 residents for all four periods considered. A common trend emerges from the close observation of the time series plotted - the means of the retail and distribution price. The retail price appears to be significantly bigger in smaller cities, compared to cities with more than 300.000 residents. Also, this does not appear to be the case regarding the distribution price, which presents a much smaller apparent difference between the mean price observed in smaller and bigger cities.

3.2 CONFAZ

As stated in Section 2, the ICMS is due at the refinery level of the supply chain, and distribution, and retail levels of the product chain is withheld at the refinery or the importer. To levy the tax based on the cumulative value added across the chain, an estimate of the final price to the consumer will is required. The PMPF is calculated by each state's tax agency (SEFAZ) and published biweekly

Figure 7: Retail price difference between small and large cities



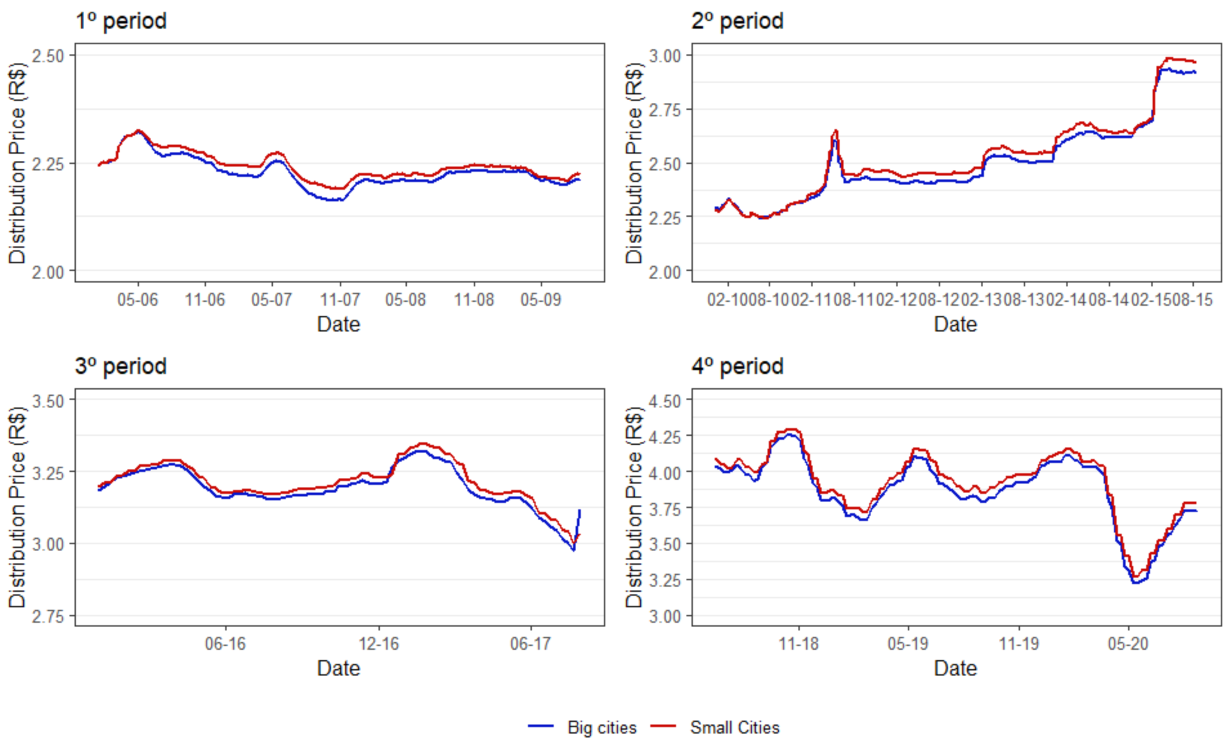
Source: Self-elaboration, based on data from ANP and IBGE. The vertical lines present the mean price in R$ for the liter of Common Gasoline, and the horizontal axis represents the date. Additionally, the blue line represents big cities and the red line represents small cities, and the panels are arranged in order of the periods

by the CONFAZ, as described in section 3.

The PMPF data is published biweekly on the CONFAZ's website, however, no aggregate information or series is available and has to be manually retrieved for every period. Considering the period length proposed in the presented work a *webscrapping* solution was employed to retrieve those prices and, in addition to the statutory percentual tax rate defined by each state's legislation[4], to calculate the unit tax charged in each liter of fuel, in our case Common Gasoline.

[4]The tax rate was obtained from the FECOMERCIO's (Trade Federation) annual reports available at their website.

Figure 8: Distribution price difference between small and large cities



Source: Self-elaboration, based on data from ANP and IBGE. The vertical lines present the mean price in R$ for the liter of Common Gasoline, and the horizontal axis represents the date. Additionally, the blue line represents big cities and the red line represents small cities, and the panels are arranged in order of the periods

3.3 IBGE

Finally, from IBGE it was retrieved the populational data necessary to carry out the division between small and big cities is presented in the previous sections. We use both the census data, as well as the estimated population to determine the number of residents in each municipality.

4 Literature Review

Acknowledging the non-linearity inherent in many macroeconomic variables and processes has been well-established over time. Keynes (1936, p. 314) famously remarked that substituting a downward trend with an upward one often occurs suddenly and forcefully, while transitioning from an upward trend to a downward one typically lacks a sharp turning point. [5]

The correlation between fluctuations in oil prices and their impact on gasoline prices holds considerable significance for both consumers, who display heightened sensitivity to the financial implications of fuel expenditures for their vehicles, and researchers, who frequently encounter requests to offer viable justifications for the observed temporal dynamics in the relationship between oil and gasoline prices.

The general consensus among consumers that gasoline prices exhibit swift responses to increases in oil prices but slower adjustments to decreases in oil prices has been widely investigated. Additionally, recent price levels observed in both oil and gasoline markets, coupled with uncertainties surrounding supply and reserve availability, as well as the escalating global energy demand, have revitalized interest in exploring the asymmetric transmission of oil price changes to gasoline prices.

The literature regarding the asymmetric price response is extensive, primarily related to the fuel market. The question of whether there exists a difference in response to price increases and decreases, both in the speed of adjustment and absolute value was first considered by Bacon (1991). Using biweekly data from the United Kingdom's gasoline market the author found that the increases in the cost of refined gasoline were usually integrally transmitted to the final consumers in a period of two months, while decreases in the cost took an extra two weeks to be fully transmitted.

The use of Error Correction Models (ECM's) has a long history in the investigation of asymmetric relationships in the oil and fuel industry. According to Frey and Manera (2007), one of the first papers to propose a rigorous ECM model for assessing asymmetric price adjustments was Kirchgässner and Kübler (1992). Using monthly price data for Western Germany in a 17-year period, comprising from 1972 to 1989, the authors tested for cointegration and, when it cannot be

[5] As pointed by Shin et al. (2014, p. 281)

rejected, fitted both symmetric and asymmetric ECM's models. Although not allowing asymmetry to exist in the long run, the author found considerable asymmetry in the short run.

Shin (1994), however, investigating the relationship between the price of products and the price of oil for the US in the period spanning from 1982 to 1990 found no evidence, through applying dynamic models, of asymmetric pass-trough between the "cost" of oil and the final price of oil products. In fact, previous studies, such as Karrenbrock et al. (1991) studied the transmission of cost to the prices of several types of gasoline - mainly leaded, unleaded, and premium.

One hallmark work on the use o ECM's to measure asymmetries in the fuel market was presented by Borenstein et al. (1997). The authors use weekly data from the US to detect asymmetries in different levels of the distribution chain, for that the authors use four different price sources: crude spot prices, gasoline spot prices, wholesale prices, and retail prices. Additionally Borenstein et al. (1997) use an instrumental variable approach by applying a TSLS where the crude oil spots are used as instruments for the change in the cost of gasoline. It found strong and persistent asymmetries throughout the fuel chain, where the negative decreases in prices were not fully passed to the final consumer.

As noted by Grasso and Manera (2007) the authors of Borenstein et al. (1997) propose three theoretical interpretations for the reasons for the asymmetries found. First, relay the existence of a focal point for collusion of the oligopolistic sellers following a drop in the price. Second, conjectures that production lags and inventory management allow for a faster accommodation of negative shocks. Third, proposes that, due to the volatility in crude oil prices, there's a lower payoff from consumer search that makes outlets less competitive, this hypothesis is supported in later work at Borenstein and Shepard (2002).

An interesting theoretical proposition comes from Radchenko (2005). The authors use a time-series approach, through a VAR model, to support the construction that the degree of asymmetry in gasoline prices declines with an increase in oil price volatility. Using GARCH estimates, the author found results that support the thesis that oligopolistic coordination is a possible explanation for the asymmetries observed in the fuel market. As stated by the authors:

"This theory is based on the assumption that the observed asymmetry in the response

of gasoline prices is evidence of imperfect competition among retailers.4 In the retail gasoline market, firms have imperfect knowledge about the price charged by other competitors and retailers may charge a price above the competitive level if their sales remain above a threshold level. In this case, price reduction occurs only if there is a significant drop in sales indicating price cutting by other retailers." (Radchenko, 2005, p. 711)

Building upon the work of Borenstein et al. (1997), Balke et al. (1998) uses the distributed lag model and an ECM model that allows for both short-run and long-run asymmetries in price transmission. This is probably the first work that robustly utilizes both forms of asymmetries, although it does not find compelling evidence of its existence.

The Brazilian literature on the theme is also noteworthy. Canêdo-Pinheiro (2012) investigates the asymmetry in price transmission of wholesale diesel to final consumers in Brazil. The results show that price adjustments are asymmetric both in the short and long run, with a complete response to wholesale price reductions taking up to ten months. The author arrives at these findings by employing an error correction model to explain the mechanism of diesel price transmission, and it tests various types of asymmetry, both short-term and long-term. The paper also innovates by investigating the impact of the presence of structural breaks in long-term relationships on asymmetry tests, as suggested in Meyer and von Cramon-Taubadel (2004).

Additionally, Uchôa (2008), in previous work, has looked at another link of the gasoline supply chain. The author looks at the asymmetry between oil prices and retail prices and employs a Threshold Autoregressive (TAR) model and a Momentum-threshold Autoregressive (M-TAR) model to estimate the speed of adjustment for positive and negative discrepancies, after testing for cointegration on the relevant variables. Using data from the nominal exchange rate BRL/USD, the resale price of gasoline, and the price of oil on the international market the results show that gasoline prices recover 90% of negative variations from one period to another, but only 5% of positive differences are adjusted. Also, it is found that positive variations have a faster speed of adjustment, compared to negative variations.

Until this moment, the economic literature faced two main obstacles when modeling asymmetries: first, the lack of a flexible and reliable framework that allows for customizing both the long-run and the short-run between the investigated variables. Second, and more frequent in the case of the fuel supply chain, is the possible endogeneity between the related variable of cost and price. Even within a distributed lags framework[6], a few authors, like Borenstein et al. (1997), have employed a two-stages approach to dealing with the endogenous nature of costs and price.

The next relevant advance in the literature related to the estimation of asymmetries, proposed by Shin et al. (2014), aims at answering these problems. The methodology described by the authors is the foundation of the work developed in this study and will be explained in detail in the Methodology section. The motivation behind the development of this method by the authors was to construct a straightforward and adaptable nonlinear dynamic framework that can effectively model asymmetries in both the underlying long-term relationship and the dynamic adjustment patterns. The authors establish the dynamic error correction representation associated with the asymmetric long-term cointegrating regression, thereby introducing the nonlinear autoregressive distributed lag (NARDL) model.

Additionally, Shin et al. (2014) calculate asymmetric cumulative dynamic multipliers that enable us to analyze the asymmetric adjustment patterns following positive and negative shocks to the explanatory variables. This feature holds significant theoretical appeal as it allows for an intuitive depiction of the transition to a new equilibrium after a disturbance to the system. The flexibility of our framework is noteworthy as it readily accommodates various specifications and can effectively capture the complex dynamics of the underlying process.

Since its development, it has been successfully utilized to further expand the ongoing research on asymmetric passthrough. Pal and Mitra (2015) uses a multiple threshold NARDL model to investigate the asymmetric relationship between crude oil and petroleum products. The authors employ the model described in Shin et al. (2014) and use weekly time series for the United States comprising of prices for common gasoline and diesel for different locations as well as the price and volume of crude oil. The authors conclude for the existence of relatively high asymmetries in the investigated passthrough. On the same line, Atil et al. (2014) uses the same methodology to investigate pass-through of crude oil prices into gasoline and natural gas prices. The authors find

[6] According to Pesaran et al. (2001) the ARDL model is applicable even when the explanatory variables are endogenous

significant asymmetries both in the short and long run. [7]

[7] More applications of NARDL models in similar context can be found at Apergis and Vouzavalis (2018), Greenwood-Nimmo and Shin (2013) and Bagnai and Ospina (2015)

5 Methodology

This section outlines the proposed methodology and presents its uses in the proposed identification strategy. As previously mentioned in Section 4 although the intuition behind the model proposed by Shin et al. (2014) has been used since Borenstein et al. (1997) by applying partial sum decomposition to the shock or cost variable. This section presents the basic functioning of the NARDL (Non-linear Autoregressive distributed lag) model. The framework proposed by Shin et al. (2014) provides a flexible framework that allows for modeling for restrictions in the short-run and long-run asymmetries.

The equations used to obtain the results, for all four periods, are presented and will be the basis for interpreting the results presented in Section 6.

5.1 Base NARDL specification

Consider the following long-run asymmetric regression:

$$y_t = \beta^+ x_t^+ + \beta^- x_t^- + u_t, \tag{1}$$

$$\Delta x_t = v_t \tag{2}$$

where $x_t = x_0 + x_t^+ + x_t^-$ and where x_t^+ and x_t^- are partial cumulative sum process of positive and negatives changes in x_t. It's described as follows:

$$\begin{cases} \boldsymbol{x}_t^+ = \sum_{j=1}^{t} \Delta \boldsymbol{x}_j^+ = \sum_{j=1}^{t} \max\left(\Delta \boldsymbol{x}_j, 0\right) \\ \boldsymbol{x}_t^- = \sum_{j=1}^{t} \boldsymbol{\Delta} x_j^- = \sum_{j=1}^{t} \min\left(\Delta \boldsymbol{x}_j, 0\right) \end{cases} \tag{3}$$

Schorderet et al. (2003) generalizes this concept, defines this system of equations, and investigates its proprieties. First, assume that $\{x_t^+\}_{t=1}^T$ and $\{x_t^-\}_{t=1}^T$ are the time-series generated by the partial sum process of the positive and negative shocks defining the level of the original time-series at time t, as described in equation 3.

Given the usual dynamics of time series, in order to have a dynamically complete model we should consider an autoregressive, distributed lag model (ARDL). The time series properties of the variables may influence the specification. Schorderet et al. (2003) explored in detail the properties of the variables in the asymmetric regression, assuming that $\{x_t\}$ is a random walk process, with a possible linear time trend in the mean.

Additionally, assuming that $\{x_t\}$ is a random walk process, $\{x_t^+\}$ and $\{x_t^-\}$ are integrated of order 1. And $Cov(\Delta\{x_t^+\}, \Delta\{x_t^-\}) \neq 0$ emphasizes that $\{x_t^+\}$ and $\{x_t^-\}$ are not independent.

Schorderet et al. (2003) expands on this idea, and builds the following stationary linear combination of the partial sum components:

$$z_t = \beta_0^+ y_t^+ + \beta_0^- y_t^- + \beta_1^+ x_t^{++} + \beta_1^- x_t^-$$

The long-run asymmetric regression can be incorporated into the ARDL approach that was popularized by Pesaran et al. (1995) and Pesaran et al. (2001). Consider the following nonlinear $ARDL(p, q)$ model:

$$y_t = \sum_{j=1}^{p} \rho_j y_{t-j} + \sum_{j=0}^{q} \left(\boldsymbol{\theta}_j^{+\prime} \boldsymbol{x}_{t-j}^+ + \boldsymbol{\theta}_j^{-\prime} \boldsymbol{x}_{t-j}^- \right) + \varepsilon_t \tag{4}$$

where $\boldsymbol{x}_t$ is a $k \times 1$ vector of regressors such that $boldsymbol x_t = \boldsymbol{x}_0 + \boldsymbol{x}_t^+ + \boldsymbol{x}_t^-$, ρ_j is the autoregressive parameter, θ_j^- and θ_j^+ are the asymmetric distributed lag parameters, and ε_t is a stochastic error term.

As pointed out by Shin et al. (2014), it is possible to rewrite equation (4) in the error correction form. The final form of interest is represented below:

$$\begin{aligned} \Delta y_t &= \rho y_{t-1} + \boldsymbol{\theta}^{+\prime} \boldsymbol{x}_{t-1}^+ + \boldsymbol{\theta}^{-\prime} \boldsymbol{x}_{t-1}^- + \sum_{j=1}^{p-1} \gamma_j \Delta y_{t-j} \\ &\quad + \sum_{j=0}^{q-1} \left(\varphi_j^{+\prime} \Delta \boldsymbol{x}_{t-j}^+ + \varphi_j^{-\prime} \Delta \boldsymbol{x}_{t-j}^- \right) + \varepsilon_t \\ &= \rho \xi_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta y_{t-j} + \sum_{j=0}^{q-1} \left(\varphi_j^{+\prime} \Delta \boldsymbol{x}_{t-j}^+ + \varphi_j^{-\prime} \Delta \boldsymbol{x}_{t-j}^- \right) + \varepsilon_t \end{aligned} \tag{5}$$

where $\phi = \sum_{j=1}^{p} \rho_j - 1$, $\gamma_j = -\sum_{i=j+1}^{p} \rho_i$ for $j = 1, \ldots, p-1$, $\theta^+ = \sum_{j=0}^{q} \boldsymbol{\theta}_j^+$ and $\boldsymbol{\theta}^- = \sum_{j=0}^{q} \boldsymbol{\theta}_j^-$, $\varphi_0^+ = \boldsymbol{\theta}_0^+$ and $\boldsymbol{\varphi}_0^- = \boldsymbol{\theta}_0^-$, $\boldsymbol{\varphi}_j^+ = -\sum_{i=j+1}^{q} \theta_j^+$ for $j = 1, \ldots, q-1$, $\boldsymbol{\varphi}_j^- = -\sum_{i=j+1}^{q} \theta_j^-$ for $j = 1, \ldots, q-1$, $\xi_t = y_t - \boldsymbol{\beta}^{+\prime}\boldsymbol{x}_t^+ - \boldsymbol{\beta}^{-\prime}\boldsymbol{x}_t^-$ is the nonlinear error correction term where $\boldsymbol{\beta}^+ = -\boldsymbol{\theta}^+/\phi$ and $\boldsymbol{\beta}^- = -\boldsymbol{\theta}^-/\phi$ are the associated asymmetric long-run parameters.

In the context of the study conducted here, equation (5) will be estimated for the impact of the unit cost in *Reais/(R$)* of ICMS tax on the distribution price (D_t) and the retail price (R_t) price for every city considered in the sample. In terms of the variable of equation (5) we have that

$$
\begin{aligned}
\boldsymbol{x}_{t-1}^+ &= \sum \mathrm{ICMS}_{t-1}^+ = \mathbf{ICMS}_{t-1}^+ \\
\boldsymbol{x}_{t-1}^- &= \sum \mathrm{ICMS}_{t-1}^- = \mathbf{ICMS}_{t-1}^-.
\end{aligned}
$$

It's then possible to define $\boldsymbol{x}_{t-j}^+ = \Delta\mathrm{ICMS}_{t-j}^+$ and $\boldsymbol{x}_{t-j}^- = \Delta\mathrm{ICMS}_{t-j}^-$. The outcome variable ($y_t$) will either the distribution price (D_t) or the retail price (R_t). We can then rewrite equation (5) in the terms of the panel to be estimated:

$$
\begin{aligned}
\Delta R_{t,i} = {} & \rho R_{t-1,i} + \boldsymbol{\theta}^{+\prime}\mathbf{ICMS}_{t-1,i}^+ + \boldsymbol{\theta}^{-\prime}\mathbf{ICMS}_{t-1,i}^- + \sum_{j=1}^{p-1} \gamma_j \Delta R_{t-j,i} \\
& + \sum_{j=0}^{q-1} \left(\varphi_j^{+\prime}\Delta\mathbf{ICMS}_{t-j,i}^+ + \varphi_j^{-\prime}\Delta\mathbf{ICMS}_{t-j,i}^- \right) + X_{t,i} + \varepsilon_{t,i}
\end{aligned}
\tag{6}
$$

$$
\begin{aligned}
\Delta D_{t,i} = {} & \rho D_{t-1,i} + \boldsymbol{\theta}^{+\prime}\mathbf{ICMS}_{t-1,i}^+ + \boldsymbol{\theta}^{-\prime}\mathbf{ICMS}_{t-1,i}^- + \sum_{j=1}^{p-1} \gamma_j \Delta D_{t-j,i} \\
& + \sum_{j=0}^{q-1} \left(\varphi_j^{+\prime}\Delta\mathbf{ICMS}_{t-j,i}^+ + \varphi_j^{-\prime}\Delta\mathbf{ICMS}_{t-j,i}^- \right) + X_{t,i} + \varepsilon_t
\end{aligned}
\tag{7}
$$

where $\Delta R_{t,i}$ and $\Delta D_{t,i}$ are respectively the first difference of the retail and distribution price in a week t for the city i, and $X_{t,i}$ is a fixed effects vector.

As pointed out by Bertsatos et al. (2022), panel dynamic fixed effects have been broadly used in empirical works to derive and analyze panel datasets and obtain long-run dynamic multipliers, using a distributed lags approach. However, it is known that a model with lagged variables and fixed effects in the form of *dummy* variables can be biased when the time dimension fo the panel, in our case weeks, is small. On the other hand, it's important to note that: "*When the time dimension, T, is large then, it is legit to use this estimator since the familiar downward lagged dependent bias (...) is eliminated given a large number of cross sections, N.*" (Bertsatos et al., 2022, p. 2)

That is precisely our case, as the smallest period considered (the third period) consists of a large T (84 weeks) and a large N (375 municipalities), and our biggest period, the second one, accounts for 298 consecutive weeks. Those numbers are well within the proposed by Judson and Owen (1999) for balanced panels, as it also is the case of the dataset employed in the estimations developed.

Furthermore, the fixed effects specification used includes time (week) and municipality. In regard to the recent literature about "two-way" althou effects, our model is not a difference-in-differences, as we do not use a control group. The continuous treatment is simultaneous for all observations, as the monthly change in PMPF goes into effect on the same date (week) for all observations. This data and institutional environment do not fit in the recent critique of asynchronous binary treatment two-way models of Callaway and Sant'Anna (2021).

5.2 Asymmetric Dynamic Multipliers

As proposed by Shin et al. (2014) for deriving the cumulative effect, or cumulative dynamic multiplier, of a positive and negative change in ICMS, we first consider the ARDL-in-levels representation of equation (5):

$$\phi(L)y_t = \boldsymbol{\theta}^+(L)\boldsymbol{x}_t^+ + \boldsymbol{\theta}^-(L)\boldsymbol{x}_t^- + e_t \tag{8}$$

Here, $\phi(L) = 1 - \sum_{i=1}^{p-1} \phi_i L^i, \theta^+(L) = \sum_{i=0}^{q} \theta_i^+ L^i$, and $\theta^-(L) = \sum_{i=0}^{q} \theta_i^{-1} L^i$. By premultiplying the previous equation by the inverse of $\phi(L)$, we derive:

$$y_t = \lambda^+(L)x_t^+ + \lambda^-(L)x_{t-i}^- + [\phi(L)]^{-1}e_t, \tag{9}$$

where $\lambda^+(L)\left(= \sum_{j=0}^{\infty} \lambda_j^+\right) = \phi(L)^{-1}\theta^+(L)$ and $\lambda^-(L)\left(= \sum_{j=0}^{\infty} \lambda_j^-\right) = \phi(L)^{-1}\boldsymbol{\theta}^-(L)$. The cumulative dynamic multiplier impacts of $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$ on y_t can be assessed as follows:

$$
\begin{aligned}
\boldsymbol{L}^{+}_{\text{ICMS,h}} &= \sum_{j=0}^{h} \frac{\partial y_{t+j}}{\partial \textbf{ICMS}^{+}_{t}} = \sum_{j=0}^{h} \lambda^{+}_{j}, \\
\boldsymbol{L}^{-}_{\text{ICMS,h}} &= \sum_{j=0}^{h} \frac{\partial y_{t+j}}{\partial \textbf{ICMS}^{-}_{t}} = \sum_{j=0}^{h} \lambda^{-}_{j}, \text{ for } h = 0, 1, 2 \dots
\end{aligned}
\tag{10}
$$

From equation (10) we can derive that as $h \to \infty$ then $\beta^{+} \to \theta^{+}/\rho$ and $\beta^{-} \to \theta^{-}/\rho$ that are the asymmetric long-run coefficients defined in equation (5).

To obtain the dynamic multiplier in a model that only considers the existence of short-run pass-through asymmetries, such as the model developed by Borenstein et al. (1997), we first have to add a restriction to equation (5) to impose symmetry in the long-run. That is, if $\boldsymbol{\theta}^{+} = \boldsymbol{\theta}^{-} = \boldsymbol{\theta}$ equation (5) simplifies to:

$$
\Delta y_t = \rho y_{t-1} + \boldsymbol{\theta} \boldsymbol{x}_{t-1} + \sum_{i=1}^{p-1} \gamma_i \Delta y_{t-i} + \sum_{i=0}^{q-1} \left(\boldsymbol{\varphi}^{+} \Delta \boldsymbol{x}^{+}_{t-i} + \varphi^{-} \Delta \boldsymbol{x}^{-}_{t-i} \right) + e_t \tag{11}
$$

In this model, in line with the derivation presented in the previous section, the dynamic multiplier will be defined as the ratio $-\boldsymbol{\theta}/\rho$, the same will be true for the more restrictive model defined below.

Analogously, it's possible to define a model that accounts only for long-run asymmetries imposing that $\varphi^{+} = \varphi^{-} = \varphi$. In the same manner as was done for equation (11) we can change equation (5) to be represented under this restriction:

$$
\Delta y_t = \rho y_{t-1} + \boldsymbol{\theta}^{+} \boldsymbol{x}^{+}_{t-1} + \boldsymbol{\theta}^{-} \boldsymbol{x}^{-}_{t-1} + \sum_{i=1}^{p-1} \gamma \Delta y_{t-i} + \sum_{i=0}^{q-1} \varphi \Delta \boldsymbol{x}_{t-i} + e_t \tag{12}
$$

Finally, the most restrictive model that can be estimated using the NARDL framework proposed by Shin et al. (2014) is considering both the previous symmetric impositions, for the long-run and the short-run, such that the equation is reduced to the form on equation (12).

5.3 Bounds-Testing

The ARDL model is suited for cointegration testing, using the bounds-testing analysis to investigate the long-run asymmetric relationship between the variables selected for the study. It is hypothesized that asymmetries exist in the short-run dynamics of these variables. Shin et al. (2014) provides a detailed operational procedure for testing the existence of asymmetric cointegration in the long-run dynamics of the base model, specified in equation (5) if the variables are weakly exogenous.

As in Shin et al. (2014), one begins by using a t-statistic to test whether $\rho = 0$ against $\rho < 0$ in (5). We will call that statistic t_{BDM}. Following Pesaran et al. (2001) an F-test on the joint null hypothesis: $\rho = \theta^{+} = \theta^{-}$ in (5) should be used. This will be called F_{PSS} statistics.

The actual critical values of the statistics depend on the integration level of the variables.[8] This may be subject to pre-test bias as unit root tests are well known for their distortions and lack of power. To overcome these difficulties, Shin et al. (2014) proposes undertaking the pragmatic "bounds-testing" proposed by Pesaran et al. (2001). That is:

> *Two extreme cases can be identified, one in which the level regressors $\boldsymbol{x}_t^{+}$and $\boldsymbol{x}_t^{-}$ in (9.10) are all $I(1)$, and the other in which they are all $I(0)$. It follows that critical values tabulated for these two scenarios provide critical value bounds for all classifications, irrespective of whether the regressors are $I(0), I(1)$, or mutually cointegrated. This is an important property in the current context due to the various dependence structures (including cointegration) that may exist between $\boldsymbol{x}_t^{+}$and $\boldsymbol{x}_t^{-}$.*

Additionally, we test for the significance of the cumulative effect of positive and negative changes in the retail and distribution price. We begin by obtaining the significance interval of the cumulative dynamic multiplier defined in section 5.2 through a block bootstrap method considering 500 repetitions. Then, a T-test is performed in which the null hypothesis is $\boldsymbol{L}_{\text{ICMS}}^{+or-} = 0$ for the mean of the bootstrap estimation. This test will be referred to as the "Asymmetry Test".

[8] constant terms and their role in the cointegrating equation may affect the distribution as well, as in the Johansen cointegration test.

5.4 Tax Anticipation and Model Selection

In the study conducted by Coglianese et al. (2017) compelling evidence is presented that demonstrates anticipatory behavior among gasoline buyers in response to tax changes. The researchers found that consumers strategically adjust their purchasing behavior in anticipation of tax fluctuations. Specifically, an increase in gasoline purchases was observed prior to tax hikes, and a delay in purchases was noted preceding tax reductions.

Moreover, the research in Coglianese et al. (2017) also reveals that this anticipatory behavior is not exclusive to retail consumers. The authors provide evidence that gasoline station operators and gasoline distributors also adjust their purchase and storage choices in the days leading up to a gasoline tax change. This finding suggests that the observed spike in the quantity of gasoline sold prior to gasoline tax increases likely reflects not only shifts in the purchases of final consumers but also stockpiling by gasoline distributors and gasoline stations. The researchers also provide evidence that large elasticity estimates are an artifact of not having accounted for shifts in gasoline purchases in anticipation of gasoline tax changes. They argue that these large elasticity estimates are a result of the failure to account for this anticipatory behavior.

Finally, the pattern of anticipatory behavior appears to be approximately symmetric in tax increases and decreases. The study found that gas purchases decrease during the month leading up to gasoline tax decreases. However, this effect was not statistically significant due to fewer tax decreases in the data Coglianese et al. (2017). This finding suggests that both tax increases and decreases trigger anticipatory behavior among gasoline buyers, although the evidence is stronger for tax increases.

A similar result comes from Dieler et al. (2015) that shows significant anticipatory behavior using European data. Additionally, they propose the following mechanism: the anticipation effect comes into play when consumers, aware of an upcoming fuel tax increase, choose to fill their cars' fuel tanks shortly before its implementation. As a result, these consumers end up refueling less in the following month. The reasons proposed for the reduction in consumption following a tax hike suggested by the authors are two. First, based on a demand reaction, the increase in the retail price of motor fuel because of the tax increase makes fuel more expensive, leading to less consumption.

Second, Due to anticipatory behavior, fuel tanks are fuller at the beginning of the tax month than in months without a tax increase. The aspect of a fuller fuel tank at the start of the tax month is identified as an anticipation effect or an inter-temporal shifting of fuel purchases. This shift does not change overall fuel consumption, it merely alters the timing of when the fuel is purchased.

For that reason when selecting for the optimal lags (p, q) in the model in equation (5) it was allowed for the possibility of anticipation, that is, for the existence of negative lags in the considered models. The *benchmark* used for selecting the optimal model was information criteria, more specifically AIC and SIC following the selection method used in Atil et al. (2014). The criteria are defined as follows:

$$
\begin{aligned}
\text{AIC} &= T \ln(\text{ sum of squared residuals}) + 2n \\
\text{SBC} &= T \ln(\text{ sum of squared residuals}) + n \ln(T)
\end{aligned}
\tag{13}
$$

where $n =$ number of parameters estimated ($p + q+$ possible constant term) $T =$ number of usable observations.

To select the number of lags to be considered in our model we allow p, in equation (5) to assume every integer value in the interval $(0, 20)$ and q to assume the value of all integers in the interval $(-20, 20)$. For all the four periods considered the number of lags that are consistently chosen is $(p, q) = (3, 2)$. Additionally, we observe no benefits of including anticipation *lags*, as their introduction does not appear to have any downward effect on the information criteria applied.

6 Results

This section presents the results of the proposed methodology described in subsection 5.1, as well as the bounds tests proposed in subsection 5.3. The models will be estimated for the four proposed periods presented in the Data section, more specifically in subsection 3.1, which presents the gas prices data. All models presented in this section, as specified in Section 5, are modeled in "Reais" (R$).

The results are derived from the application of the model proposed in Section 5, in the functional forms described in equations (6) and (7) for respectively the retail and distribution prices. The equations (11) and (12) will be used to estimate the models under symmetric restrictions for the long and short runs, as previously described. Further, a symmetric model is estimated, to provide a basis for comparison.

Table 1: Kao (1999) test for cointegration in the retail price model.

Period	Modified Dickey–Fuller t	Dickey–Fuller t	Augmented Dickey–Fuller t	Unadjusted modified Dickey–Fuller t	Unadjusted Dickey–Fuller t
1	-1.0e+03	-2.5e+02	-1.7e+02	-1.0e+03	-2.5e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)
2	-1.6e+03	-3.2e+02	-2.1e+02	-1.6e+03	-3.2e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)
3	-5.4e+02	-2.0e+02	-1.2e+02	-5.5e+02	-2.0e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)
4	-7.8e+02	-2.5e+02	-1.6e+02	-8.1e+02	-2.5e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)

This table presents the results for the Kao (1999) cointegration test, for all four periods. The proposed tests are structures as the null hypothesis (H_0) is "No cointegration" and the alternative hypothesis (H_A) is "All panels are cointegrated". P-values are displayed in parentheses.

As previously stated in section 5.3, the described bounds test (F_{PSS}) is displayed in the "*Fit Statistics*" portion of the results tables below in this section. Additionally, the long-run symmetry test, labeled "*Simmetry Test*" is also displayed, the test is obtained by performing a T-Test on the null for the equality of the long-run multipliers, obtained through a block-bootstrap procedure with 500 repetitions. Also, a test for the presence of short-run asymmetries is presented in the tables - and labeled "*SR Asymmetry Test*" - that tests for the joint hypothesis the sum of both positive and negative short-run coefficients are equal.

While the necessary bounds test is presented, we further estimate supplementary cointegration evidence for added robustness. Given the nature of the panel data, it's pertinent to conduct a cointegration test that assesses all panels, thereby enhancing the validity of our estimations. Consequently, we adopt the test outlined in Kao (1999) to examine the cointegration relations investigated, for both the retail and distribution prices. The test performed is rigorous in the sense that the alternative hypothesis is that all panels are cointegrated. The results are presented in Tables 1 and 2.

Table 2: Kao (1999) test for cointegration in the distribution price model.

Period	Modified Dickey–Fuller t	Dickey–Fuller t	Augmented Dickey–Fuller t	Unadjusted modified Dickey–Fuller t	Unadjusted Dickey–Fuller t
1	-7.5e+02	-2.2e+02	-1.2e+02	-8.3e+02	-2.2e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)
2	-4.7e+02	-93.1532	-56.9769	-4.8e+02	-93.0536
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)
3	-3.1e+02	-1.3e+02	-91.2551	-3.7e+02	-1.3e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)
4	-4.6e+02	-1.2e+02	-61.3622	-4.0e+02	-1.2e+02
	(0,0001)	(0,0001)	(0,0001)	(0,0001)	(0,0001)

This table presents the results for the Kao (1999) cointegration test, for all four periods. The proposed tests are structures as the null hypothesis (H_0) is "No cointegration" and the alternative hypothesis (H_A) is "All panels are cointegrated". P-values are displayed in parentheses.

6.1 Asymmetries in the short and long run

This section presents the results of the model specified in equations (7) and (6) that accounts for pass-through asymmetries in both the short and long run for the retail and distribution prices of Common Gasoline. The tables 3 and 4 display the results for the retail and distribution price, respectively. In each table columns (1)-(4) represents the four different periods outlined in section 3.1.

Additionally, the results tables, such as Table 3, present tests that allow us to better interpret the measured asymmetries. The field named "Asymmetry T-Test" presents the p-value of a two Sample t-test of the bootstrap samples where the Null Hypothesis (H_0) is that the long-run dynamic multipliers (either L^{+}_{ICMS} or L^{-}_{ICMS}) are equal to zero. Further, the field named "SR Asymmetry Test" tests for the equality between the sum of the positive and negative short-run coefficients (H_0), and the p-values are displayed.

Table 3: Main Results: Asymmetric response on the retail price

Dependent Variable:	ΔR_t			
Model:	(1)	(2)	(3)	(4)
Variables				
R_{t-1}	-0.0890***	-0.0712***	-0.1311***	-0.1399***
	(0.0107)	(0.0039)	(0.0131)	(0.0057)
ICMS^+_{t-1}	0.0650***	0.0711***	0.1193***	0.1549***
	(0.0177)	(0.0078)	(0.0205)	(0.0104)
ICMS^-_{t-1}	0.0472***	0.0672***	0.1036***	0.1564***
	(0.0168)	(0.0086)	(0.0203)	(0.0094)
ΔR_{t-1}	-0.1949***	-0.2621***	-0.1231*	-0.0076
	(0.0297)	(0.0213)	(0.0645)	(0.0087)
ΔR_{t-2}	-0.0272	0.0133	0.0939	-0.0731***
	(0.0299)	(0.0175)	(0.0686)	(0.0111)
ΔICMS^+_t	0.0255	0.2397***	0.2210***	0.1346***
	(0.0464)	(0.0305)	(0.0396)	(0.0273)
$\Delta\text{ICMS}^+_{t-1}$	0.0879**	0.0873***	0.0614	0.1143***
	(0.0414)	(0.0288)	(0.0398)	(0.0243)
ΔICMS^-_t	0.0322	0.1772***	-0.0011	0.0924***
	(0.0438)	(0.0495)	(0.0679)	(0.0309)
$\Delta\text{ICMS}^-_{t-1}$	-0.0596	0.1800***	0.0296	0.0310
	(0.0487)	(0.0452)	(0.1166)	(0.0365)
L^+_{ICMS}	0.721***	1***	0.91***	1.108***
L^-_{ICMS}	0.534***	0.949***	0.797***	1.119***
Fixed-effects				
Municipality	Yes	Yes	Yes	Yes
Date	Yes	Yes	Yes	Yes
Fit statistics				
Asymmetry T-Test	2.18×10^{-52}	7.71×10^{-17}	4.76×10^{-48}	0.0036
R^2	0.11679	0.33439	0.39727	0.47636
Observations	47,988	72,570	30,375	40,906
BIC	-227,717.7	-326,837.3	-113,400.4	-131,219.4
F_{PSS}	4.73×10^{-15}	3.02×10^{-74}	2.28×10^{-22}	4.87×10^{-131}
SR Asymmetry Test	0.15354	0.69816	0.13417	0.05198

Clustered (cidade_estado) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*
L^+_{ICMS} and L^-_{ICMS} represents the long-term cumulative effect of the ICMS's incidence over the prices. Their reported *p-values* were obtained through a bootstrap with panel block resampling considering 500 replications.

The results presented in Table 3 indicate that we find symmetric short-run pass-trough in the first three periods, as we fail to reject the null hypothesis of the "SR Asymmetry Test", the short-run asymmetry of the for period is significant at 90% confidence level. However, it's noteworthy that, for the first and third periods - columns (1) and (3) - the negative short-run coefficients are not statistically significant, despite their sum being equal to their positive counterparts.

The long-run coefficients show a different pattern, as in all four periods we find statistically significant asymmetry. For the first three periods, we observe that positive shocks are, on average, more intensely passed on to retail prices. For example, in period (3) we see that, on average, for every R$1,00 increase in the ICMS collection, R$0,91 is passed on to the final price, and this value reduces to approximately R$0,80 when we observe a R$1,00 reduction in the ICMS collection. A similar pass-trough is observed in the second period, model in column (2), as for every R$1,00 in the positive difference we observe an R$ 1,00 increase in the final consumer price, a 1:1 ratio. This transmission drop to R$ 0,95 when observing negative fluctuation in ICMS.

The fourth period observes a reverse of this trend where the negative changes in costs (measured in the ICMS) are slightly more passed on to the final costs, R$1,12 for every R$1,00 variation, as opposed to R$1,11 for the positive variations. So, with the exception of the first period, we observe integral pass-trough with limited asymmetries in both the short-run, where they are statistically nonexistent and in the long-run, where they are numerically proximate. This integral passthrough, even for negative changes in cost through a decrease in VAT is not unheard of, as Schmerer and Hansen (2023) finds similar results when analyzing a recent fuel tax rebate made by the German government.

A similar conclusion to the ones drawn from the retail prices model arises from the analysis of the distribution prices. Once again, to the exemption of the first period, we found no statistical evidence of asymmetry in the short-run, through an analysis of the “SR Asymmetry Test”.

Although there exists a statistically relevant asymmetry in the long-run dynamic multipliers, L^{+}_{ICMS} and L^{-}_{ICMS}, the estimated coefficients are numerically proximate, with the biggest difference being perceived in the third period where a unit change in ICMS caused an R$0,98 increase in the distribution price when the variation is positive, and a decrease of R$0,81 when the variation is negative.

6.2 Asymmetries only in the short run

As proposed by Borenstein et al. (1997), and later presented by Shin et al. (2014) we can repeat the same model of Tables 3 and 4 imposing symmetrical long-run pass-through, as previously

Table 4: Main Results: Asymmetric response on the distribution price

Dependent Variable:	$\Delta D_{i,t}$			
Model:	(1)	(2)	(3)	(4)
Variables				
D_{t-1}	-0.1699***	-0.1666***	-0.2058***	-0.2770***
	(0.0192)	(0.0174)	(0.0137)	(0.0130)
ICMS^+_{t-1}	0.1704***	0.1652***	0.2017***	0.3044***
	(0.0201)	(0.0261)	(0.0210)	(0.0246)
ICMS^-_{t-1}	0.1563***	0.1648***	0.1683***	0.2728***
	(0.0186)	(0.0257)	(0.0234)	(0.0179)
ΔD_{t-1}	-0.3717***	-0.1626***	-0.0506***	0.0365***
	(0.0253)	(0.0275)	(0.0188)	(0.0130)
ΔD_{t-2}	-0.0699***	-0.1220***	-0.1716***	-0.2001***
	(0.0255)	(0.0242)	(0.0210)	(0.0176)
ΔICMS^+_t	0.0649*	0.1942**	0.1044	-0.0276
	(0.0368)	(0.0716)	(0.0674)	(0.0426)
$\Delta\text{ICMS}^+_{t-1}$	0.2184***	0.1881**	-0.1223***	-0.0296
	(0.0451)	(0.0885)	(0.0400)	(0.0495)
ΔICMS^-_t	-0.0273	0.0674	0.1476**	0.0592
	(0.0490)	(0.1020)	(0.0696)	(0.0503)
$\Delta\text{ICMS}^-_{t-1}$	-0.0243	0.3431**	0.0270	-0.0499
	(0.0418)	(0.1529)	(0.1045)	(0.0514)
L^+_{ICMS}	1***	0.995***	0.979***	1.098***
L^-_{ICMS}	0.918***	1.001***	0.811***	0.982***
Fixed-effects				
Municipality	Yes	Yes	Yes	Yes
Date	Yes	Yes	Yes	Yes
Fit statistics				
Asymmetry T-Test	1.24×10^{-39}	0.2802	4.44×10^{-128}	2.66×10^{-110}
R^2	0.29339	0.44331	0.31747	0.67143
Observations	21,204	7,670	16,767	16,950
BIC	-121,111.2	-37,367.7	-67,465.0	-56,780.8
F_{PSS}	3.01×10^{-20}	6.19×10^{-23}	9.83×10^{-52}	1.06×10^{-105}
SR Asymmetry Test	0.00014	0.90077	0.22852	0.55067

Clustered (cidade_estado) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*
L^+_{ICMS} and L^-_{ICMS} represents the long-term cumulative effect of the ICMS's incidente over the prices. Their reported *p-values* were obtained through a bootstrap with panel block resampling considering 500 replications.

presented in equation 11. Tables 5 and 6 present the results of such estimations for the retail and distribution prices respectively, both measured in "reais" (R$). By and large, we observe the same general results pattern as displayed in the unrestricted model repeats.

For all periods, with the exemption of the last one for the retail price model, the short-run asymmetry test fails to reject the null hypothesis, indicating we do not observe short-sun asymmetries in

Table 5: Asymmetric response on the distribution price. Asymmetries are considered only for the short run.

Dependent Variable:			ΔR_t	
Model:	(1)	(2)	(3)	(4)
Variables				
R_{t-1}	-0.0884***	-0.0711***	-0.1307***	-0.1399***
	(0.0107)	(0.0039)	(0.0130)	(0.0057)
ICMS_{t-1}	0.0534***	0.0708***	0.1123***	0.1560***
	(0.0168)	(0.0077)	(0.0191)	(0.0091)
ΔR_{t-1}	-0.1951***	-0.2621***	-0.1234*	-0.0076
	(0.0297)	(0.0213)	(0.0645)	(0.0087)
ΔR_{t-2}	-0.0272	0.0132	0.0937	-0.0731***
	(0.0299)	(0.0175)	(0.0686)	(0.0111)
ΔICMS_t^+	0.0196	0.2389***	0.2201***	0.1350***
	(0.0466)	(0.0306)	(0.0396)	(0.0273)
$\Delta\text{ICMS}_{t-1}^+$	0.0935**	0.0868***	0.0675*	0.1137***
	(0.0414)	(0.0288)	(0.0392)	(0.0239)
ΔICMS_t^-	0.0396	0.1816***	-0.0033	0.0921***
	(0.0446)	(0.0490)	(0.0680)	(0.0311)
$\Delta\text{ICMS}_{t-1}^-$	-0.0585	0.1807***	0.0187	0.0311
	(0.0486)	(0.0450)	(0.1172)	(0.0364)
L_{ICMS}	0.595***	1***	0.855***	1.119***
Fixed-effects				
Municipality	Yes	Yes	Yes	Yes
Date	Yes	Yes	Yes	Yes
Fit statistics				
R^2	0.11653	0.33438	0.39724	0.47636
Observations	47,988	72,570	30,375	40,906
BIC	-227,714.5	-326,846.9	-113,409.1	-131,229.9
SR Asymmetry Test	0.18245	0.63360	0.11090	0.05199

Clustered (cidade_estado) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*
L_{ICMS} represents the long-term cumulative effect of the ICMS's incidence over the prices. Their reported *p-values* were obtained through a bootstrap with panel block resampling considering 500 replications.

the ICMS passthrough to both the retail and distribution prices. Once again, a few periods, for both the retail and distribution prices, have short-term negative coefficients statistically equal to zero, even though we observe symmetry through the performed asymmetry t-test.

For this restricted model L_{ICMS} represents the cumulative effect of a unit change, in Reais (R$), in ICMS on the retail price. The overall values confirm the results of the first table, in terms of cumulative effects. We observe an almost integral pass-through of changes in the collected ICMS, with the exception being again the first period for the retail model, where a relatively low dynamic multiplier is found, of just R$0,59 for every R$1,00 change in the ICMS. The remainder of the periods presents a cumulative effect between approximately R$0,85 to R$1,12.

Table 6: Asymmetric response on the distribution price. Asymmetries are considered only for the short run.

Dependent Variable:		$\Delta D_{i,t}$		
Model:	(1)	(2)	(3)	(4)
Variables				
D_{t-1}	-0.1694***	-0.1666***	-0.2055***	-0.2757***
	(0.0187)	(0.0174)	(0.0138)	(0.0135)
ICMS_{t-1}	0.1588***	0.1651***	0.1879***	0.2782***
	(0.0187)	(0.0256)	(0.0195)	(0.0173)
ΔD_{t-1}	-0.3717***	-0.1626***	-0.0508***	0.0359***
	(0.0253)	(0.0275)	(0.0188)	(0.0130)
ΔD_{t-2}	-0.0697***	-0.1220***	-0.1717***	-0.2007***
	(0.0255)	(0.0242)	(0.0210)	(0.0176)
ΔICMS_t^+	0.0610	0.1940**	0.1045	-0.0382
	(0.0372)	(0.0724)	(0.0670)	(0.0439)
$\Delta\text{ICMS}_{t-1}^+$	0.2259***	0.1880**	-0.1084***	-0.0163
	(0.0448)	(0.0887)	(0.0406)	(0.0464)
ΔICMS_t^-	-0.0230	0.0685	0.1380**	0.0644
	(0.0493)	(0.1036)	(0.0684)	(0.0510)
$\Delta\text{ICMS}_{t-1}^-$	-0.0224	0.3439**	-0.0023	-0.0498
	(0.0419)	(0.1456)	(0.1019)	(0.0511)
L_{ICMS}	0.933***	0.992***	0.916***	1.007***
Fixed-effects				
Municipality	Yes	Yes	Yes	Yes
Date	Yes	Yes	Yes	Yes
Fit statistics				
R^2	0.29303	0.44331	0.31726	0.67121
Observations	21,204	7,670	16,767	16,950
BIC	-121,110.3	-37,376.6	-67,469.6	-56,779.4
SR Asymmetry Test	0.00016	0.89114	0.38178	0.53922

Clustered (cidade_estado) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*
L_{ICMS} represents the long-term cumulative effect of the ICMS's incidence over the prices. Their reported *p-values* were obtained through a bootstrap with panel block resampling considering 500 replications.

A very similar set of results is found when looking at the distribution price in Table 6, with a few differences. In this specification, only the first period presents short-run asymmetry, identical to the main results for the distribution price (Table 4). The symmetric long-run dynamic multiplier ranges between 0.93 and 1.00 according to the periods, these results are in line with the ones previously presented in the unrestricted model.

6.3 Symmetric response

Finally, the results for the model that imposes symmetry in the long-run and the short-run are presented in Tables 7 and 8 for the retail and distribution prices. Similar results from the previous models are found in this version of the restricted model in relation to the cumulative multiplier estimated.

For the retail price model, we find cumulative effects similar to the previous model, with the first period being once again with the lowest ICMS cumulative pas-trough, and the remainder of the periods with cumulative values close to unit. For the distribution price once again cumulative values close to the unit are found, indicating integral (1:1) passthrough from variations in the ICMS collected.

Table 7: Symmetric model

Dependent Variable:	ΔR_t			
Model:	(1)	(2)	(3)	(4)
Variables				
R_{t-1}	-0.0884***	-0.0711***	-0.1303***	-0.1398***
	(0.0107)	(0.0039)	(0.0130)	(0.0057)
ICMS_{t-1}	0.0530***	0.0707***	0.1073***	0.1544***
	(0.0168)	(0.0077)	(0.0188)	(0.0090)
ΔR_{t-1}	-0.1951***	-0.2623***	-0.1239*	-0.0077
	(0.0297)	(0.0213)	(0.0644)	(0.0087)
ΔR_{t-2}	-0.0271	0.0132	0.0935	-0.0733***
	(0.0299)	(0.0175)	(0.0686)	(0.0110)
ΔICMS_t	0.0299	0.2287***	0.1262***	0.1145***
	(0.0295)	(0.0272)	(0.0333)	(0.0206)
ΔICMS_{t-1}	0.0207	0.1030***	0.0514	0.0760***
	(0.0341)	(0.0249)	(0.0540)	(0.0217)
L_{ICMS}	0.593***	0.999***	0.858***	1.115***
Fixed-effects				
Municipality	Yes	Yes	Yes	Yes
Date	Yes	Yes	Yes	Yes
Fit statistics				
R^2	0.11640	0.33433	0.39703	0.47628
Observations	47,988	72,570	30,375	40,906
BIC	-227,728.6	-326,864.0	-113,419.6	-131,245.1

Clustered (cidade_estado) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*
L_{ICMS} represents the long-term cumulative effect of the ICMS's incidence over the prices. Their reported *p-values* were obtained through a bootstrap with panel block resampling considering 500 replications.

There are a few possible explanations for the observed behavior. As mentioned in Section 4 previous literature found regular frequency passthrough asymmetry in the fuel market, in both the chains investigated in the current work: retail and distribution.

Table 8: Symmetric model

Dependent Variable:	$\Delta D_{i,t}$			
Model:	(1)	(2)	(3)	(4)
Variables				
D_{t-1}	-0.1691***	-0.1666***	-0.2055***	-0.2759***
	(0.0187)	(0.0174)	(0.0138)	(0.0134)
ICMS_{t-1}	0.1558***	0.1652***	0.1906***	0.2793***
	(0.0188)	(0.0253)	(0.0194)	(0.0170)
ΔD_{t-1}	-0.3717***	-0.1630***	-0.0507***	0.0359***
	(0.0253)	(0.0275)	(0.0188)	(0.0130)
ΔD_{t-2}	-0.0695***	-0.1219***	-0.1716***	-0.2007***
	(0.0255)	(0.0242)	(0.0210)	(0.0176)
ΔICMS_t	0.0211	0.1790***	0.1173**	0.0160
	(0.0310)	(0.0636)	(0.0478)	(0.0321)
ΔICMS_{t-1}	0.1055***	0.2076**	-0.0696	-0.0348
	(0.0315)	(0.0779)	(0.0491)	(0.0323)
L_{ICMS}	0.945***	0.987***	0.912***	1.011***
Fixed-effects				
Municipality	Yes	Yes	Yes	Yes
Date	Yes	Yes	Yes	Yes
Fit statistics				
R^2	0.29223	0.44314	0.31721	0.67115
Observations	21,204	7,670	16,767	16,950
BIC	-121,106.5	-37,392.1	-67,487.8	-56,795.7

Clustered (cidade_estado) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*
L_{ICMS} represents the long-term cumulative effect of the ICMS's incidence over the prices. Their reported *p-values* were obtained through a bootstrap with panel block resampling considering 500 replications.

Although not included in the estimation sample, as stated in the institutional context and data sections, there were recent changes in the liquid fuel ICMS. More specifically, a unified rate across states regarding common gasoline was introduced, effective July 1^{st} of 2023. With the end of the ICMS's "freezing" and the transition to a fixed and common *ad rem* collection, the chosen value for the ICMS over the liter of Common Gasoline, in a joint decision between all state's tax board and the federal government, was R$1,22. This value represents an increase in overall taxation for almost all states, as can be seen in Figure 9, as it ranged from R$0.85 In the Goias state to R$1.15 in the Ceara state, i.e., an increase ranging from R$0.37 to R$0.07.In light of the presented results, we should expect a more than proportional transmission to prices.

Using the long-term cumulative multiplier from the last period presented in Table 3 we can estimate an expected increase in the retail price of approximately R$0,41 for the state of Goiás (GO). On the other hand, the only state that can expect a reduction in the retail price is Piauí (PI), the state with the biggest ICMS collection, which will have an effective reduction of R$0,02 on the ICMS charge, which can be transmitted to prices as an R$0,023 reduction in retail prices price.

We hypothesize that the main results observe, i.e. the lack of symmetry in the short-run and close cumulative passthrough in the long-run, with close to integral cost transmission, is in part due to the ICMS mechanism. The published PMPF may be acting as a focal point for determining the retail and distribution prices in the market.

The publicity of prices has been known to affect market price dynamics. For example, Luco (2019), when analyzing a policy of mandatory price publicity led by the regulation authority in Chile suggest that price-disclosure policies may have important distributional effects. [9]

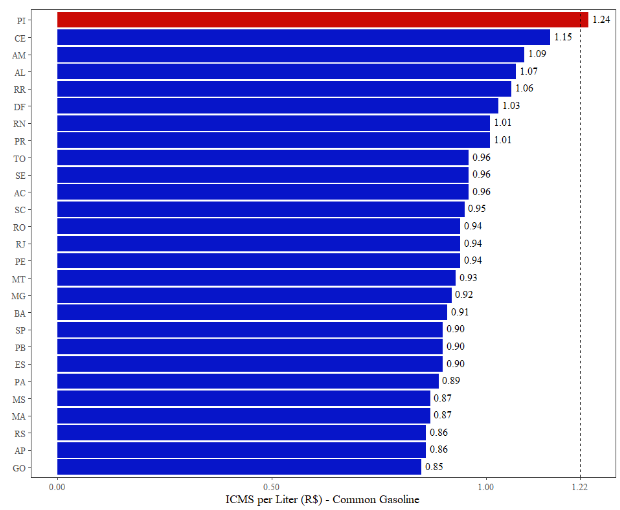
In a similar manner Dewenter et al. (2017)shows a significant impact of the publicity of retail prices in the gasoline market. The authors present the problem as two orthogonal forces: while consumers accelerate their decisions as a consequence of the information gains, firms also have additional information. The authors relate the expected results to Kühn and Vives (1995) who identify two effects of uncertain demand and information sharing among firms. First, the "output adjustment effect", decreases output adjustment to demand due to increased information. This effect's magnitude depends on market competition; price setting leads to more output adjustment with more information, while strategic price usage reduces output adjustment, increasing deadweight loss. Second, the 'preference for variety' effect influences output uniformity based on shared demand information. If demand shocks are common, outputs become uniform; if firm-specific, outputs diverge.

In the case analyzed in the current work, each state's fiscal authority releases, bi-weekly, the weighted mean of prices practiced in the state in the previous period. When analyzing tax changes, there's a natural focal point, since such changes are known in advance. While examining how prices

[9] Luco (2019) in his works suggests that the publicity of retail fuel prices led to a lessening of competition, characterized by an increase in margins and a reduction in price dispersion.

adjust for changes in tax, Asplund et al. (2000) also found that a tax change produces an integral and fast price adjustment. The study also finds long-run price asymmetry while looking at the spot market price of premium leaded gasoline, with its model being more explanatory for price increases than price cuts.

Figure 9: New *ad rem* value for the ICMS over Common Gasoline.



Source: Adaptation from Pamplona (2023).

The overall results found in Asplund et al. (2000) approximate the results presented in this section, and some explanations can be drawn, in accordance with the previous work of Luco (2019) and Dewenter et al. (2017). The disclosure of a mean weighted price in each state, every two weeks, may create a focal point for price determination, both in the retail as well in the distribution link of the supply chain, a theory that has been partially supported by the Brazilian competition authority (CADE) in CADE (2018). If the presented holds true, the gas stations and distributors would have incentives to follow the adjustment path of the PMPF's variations. An increase in the PMPF would prompt the market agents to pass on the trend to the prices, as it would be seen as a natural market trend. On the other hand, negative variations would also be passed on to prices, even if not integrally, as deviating from a market trend could implicate a negative demand shock, as the entire market reduces prices and only one retailer/distributor holds prices equal.

7 Conclusion

This study explores the potential asymmetrical relationships between ICMS (a value-added tax) variations and pricing dynamics within Brazil's liquid fuels supply chain, an area that has been relatively under-explored despite the broader context being a well-studied domain. A significant distinction of our research lies in our model's emphasis on ICMS, a crucial component in gasoline pricing. Utilizing the computational method of ICMS, we employ a panel NARDL model to examine the effects of cost alterations on price structuring, specifically on retail and distribution

This study, through the application of a NARDL model, examines both short-term and long-term asymmetries and quantifies the cumulative long-term influence on prices. This aids in refining our comprehension of the asymmetries within the liquid fuels production chain. Furthermore, we explore the potential anticipatory behavior of retailers and distributors in response to an expected tax hike.

The empirical results reveal that while long-term asymmetry is evident, short-term dynamics are symmetrical. The cumulative impact of ICMS fluctuations underscores that any modifications in the tax — either a rise or reduction — are integrally transmitted to both retail and distribution prices. Additionally, we do not find evidence that the agents - distributors and retailers - anticipate tax changes.

Overall, our data do not substantiate the existence of short-term passthrough asymmetries. Nonetheless, discernible long-term asymmetries emerge, with positive ICMS adjustments been transmitted to a greater degree than positive ones, with some periods characterised by a more than proportional passtrough. Our restricted models further confirm this interpretation, indicating a nearly complete (1:1) passthrough for both distribution and retail, irrespective of the nature of the ICMS change. Also, the restricted models point to a symmetrical passthrough in the short-run

To summarise, passthrough asymmetries only become evident upon examining the cumulative long-term effect. For the vast majority of the observed spans, the short-term asymmetries do not hold statistical significance.

These insights cohere with extant empirical studies regarding the existence of passthrough asymmetries, but the ability to gauge the enduring long-term effects brings a fresh perspective to the

Brazilian market.

8 Bibliography

References

Apergis, N., Vouzavalis, G., 2018. Asymmetric pass through of oil prices to gasoline prices: Evidence from a new country sample. Energy policy 114, 519–528.

Asplund, M., Eriksson, R., Friberg, R., 2000. Price adjustments by a gasoline retail chain. Scandinavian Journal of Economics 102, 101–121.

Atil, A., Lahiani, A., Nguyen, D.K., 2014. Asymmetric and nonlinear pass-through of crude oil prices to gasoline and natural gas prices. Energy Policy 65, 567–573.

Atkinson, A.B., Stiglitz, J.E., 1972. The structure of indirect taxation and economic efficiency. Journal of Public economics 1, 97–119.

Bacon, R., 1991. Rockets and feathers: the asymmetric speed of adjustment of uk retail gasoline prices to cost changes. Energy economics 13, 211–218.

Bagnai, A., Ospina, C.A.M., 2015. Long-and short-run price asymmetries and hysteresis in the italian gasoline market. Energy Policy 78, 41–50.

Balke, N.S., Brown, S.P., Yucel, M.K., 1998. Crude oil and gasoline prices: an asymmetric relationship? Economic Review-Federal Reserve Bank of Dallas , 2.

Benzarti, Y., Carloni, D., Harju, J., Kosonen, T., 2020. What goes up may not come down: asymmetric incidence of value-added taxes. Journal of Political Economy 128, 4438–4474.

Bertsatos, G., Sakellaris, P., Tsionas, M., 2022. A panel bounds testing procedure. Available at SSRN 4064916 .

Borenstein, S., Cameron, A., Gilbert, R., 1997. Do gasoline prices respond asymmetrically to crude oil price changes? The Quarterly journal of economics 112, 305–339.

Borenstein, S., Shepard, A., 2002. Sticky prices, inventories, and market power in wholesale gasoline markets. RAND Journal of Economics , 116–139.

CADE, 2018. Repensando o setor de combustíveis medidas pró-concorrência. Contribuições do Cade , 1–36.

Callaway, B., Sant'Anna, P.H., 2021. Difference-in-differences with multiple time periods. Journal of econometrics 225, 200–230.

Canêdo-Pinheiro, M., 2012. Assimetrias na transmissão dos preços dos combustíveis: o caso do óleo diesel no brasil. Revista Brasileira de Economia 66, 469–490.

Cavalcanti, M.C.B., 2006. Análise dos tributos incidentes sobre os combustíveis automotivos no brasil. Universidade Federal do Rio de Janeiro, COPPE. Rio de Janeiro .

Coglianese, J., Davis, L., Kilian, L., Stock, J., 2017. Anticipation, tax avoidance, and the price elasticity of gasoline demand. Journal of Applied Econometrics 32, 1–15.

Dewenter, R., Heimeshoff, U., Lüth, H., 2017. The impact of the market transparency unit for fuels on gasoline prices in germany. Applied Economics Letters 24, 302–305.

Dieler, J., Jus, D., Zimmer, M., 2015. Fill'er up: The effect of fuel taxes on carbon emissions. Working Paper, CESifo, Munich .

Ebrill, M.L.P., Keen, M.M., Perry, M.V.P., 2001. The modern VAT. International Monetary Fund.

EPE, 2020. Carga tributária incidente sobre a comercialização de combustíveis no brasil: 2020. Série Formação de Preços de Combustíveis .

Frey, G., Manera, M., 2007. Econometric models of asymmetric price transmission. Journal of Economic surveys 21, 349–415.

Grasso, M., Manera, M., 2007. Asymmetric error correction models for the oil–gasoline price relationship. Energy Policy 35, 156–177.

Greenwood-Nimmo, M., Shin, Y., 2013. Taxation and the asymmetric adjustment of selected retail energy prices in the uk. Economics Letters 121, 411–416.

Judson, R.A., Owen, A.L., 1999. Estimating dynamic panel data models: a guide for macroeconomists. Economics letters 65, 9–15.

Kao, C., 1999. Spurious regression and residual-based tests for cointegration in panel data. Journal of econometrics 90, 1–44.

Karrenbrock, J.D., et al., 1991. The behavior of retail gasoline prices: symmetric or not? Federal Reserve Bank of St. Louis Review 73, 19–29.

Keynes, J.M., 1936. The general theory of interest, employment and money.

Kirchgässner, G., Kübler, K., 1992. Symmetric or asymmetric price adjustments in the oil market: an empirical analysis of the relations between international and domestic prices in the federal republic of germany, 1972–1989. Energy Economics 14, 171–185.

Luco, F., 2019. Who benefits from information disclosure? the case of retail gasoline. American Economic Journal: Microeconomics 11, 277–305.

Melo, C., Moita, R., Sunao, S., 2021. Passing through the supply chain: Implications for market power. International Journal of Industrial Organization 79, 102796.

Meyer, J., von Cramon-Taubadel, S., 2004. Asymmetric price transmission: a survey. Journal of agricultural economics 55, 581–611.

Pal, D., Mitra, S.K., 2015. Asymmetric impact of crude price on oil product pricing in the united states: An application of multiple threshold nonlinear autoregressive distributed lag model. Economic Modelling 51, 436–443.

Pamplona, N., 2023. Estados definem icms único de r$ 1,45 por litro e pressionam preço da gasolina. Folha de São Paulo URL: https://www1.folha.uol.com.br/mercado/2023/03/estados-definem-icms-unico-de-r-145-por-litro-e-pressionam-preco-da-gasolina.shtml.

Pesaran, M.H., Shin, Y., Smith, R.J., 2001. Bounds testing approaches to the analysis of level relationships. Journal of applied econometrics 16, 289–326.

Pesaran, M.H., Shin, Y., et al., 1995. An autoregressive distributed lag modelling approach to cointegration analysis .

Pupo, F., 2022. Governadores prorrogam congelamento de icms da gasolina até fim de junho. Folha de São PAulo URL: https://www1.folha.uol.com.br/mercado/2022/03/governadores-prorrogam-congelamento-de-icms-da-gasolina-ate-fim-de-junho.shtml.

Radchenko, S., 2005. Oil price volatility and the asymmetric response of gasoline prices to oil price increases and decreases. Energy economics 27, 708–730.

Rodrigues, E., 2022. Secretários aprovam congelamento do icms sobre combustíveis até fim de março. Estado de São PAulo URL: https://www.estadao.com.br/economia/secretarios-aprovam-congelamento-do-icms-sobre-combustiveis-ate-fim-de-marco/.

Schmerer, H.J., Hansen, J., 2023. Pass-through effects of a temporary tax rebate on german fuel prices. Economics Letters 227, 111104.

Schorderet, Y., et al., 2003. Asymmetric cointegration. Université de Genève/Faculté des sciences économiques et sociales Genève .

SEFAZ/ES, 2020. Nota técnica: Metodologia de cálculo do pmpf dos combustíveis .

Shin, D., 1994. Do product prices respond symmetrically to changes in crude prices? OPEC review 18, 137–157.

Shin, Y., Yu, B., Greenwood-Nimmo, M., 2014. Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ardl framework. Festschrift in honor of Peter Schmidt: Econometric methods and applications , 281–314.

Uchôa, C.F.A., 2008. Testando a assimetria nos preços da gasolina brasileira. Revista Brasileira de Economia 62, 103–117.

Verdélio, A., 2023. Alíquota fixa e única do icms nos combustíveis entra hoje em vigor preços nas bombas devem aumentar para o consumidor final. Agência Brasil URL: https://agenciabrasil.ebc.com.br/economia/noticia/2023-06/aliquota-fixa-e-unica-do-icms-nos-combustiveis-entra-hoje-em-vigor.

9 Appendix

This appendix presents the statistics of each block bootstrap performed for the NARDL model specifications in the Results section. Each statistics table in this section has the reference of the reference model and all accounts for 500 repetitions.

9.1 Bootstrap report tables

Table 9: Bootstrap long-run dynamic multiplier: Retail price model from Table 3

Model	L^{+}_{ICMS}	CI^{+}	L^{-}_{ICMS}	CI^{-}	T Test
(1)	0.721***	[0.36,1.08]	0.534***	[0.18,0.89]	2.185E-52
(2)	1***	[0.83,1.17]	0.949***	[0.74,1.15]	7.713E-17
(3)	0.91***	[0.7,1.12]	0.797***	[0.55,1.04]	4.755E-48
(4)	1.108***	[0.98,1.24]	1.119***	[1.01,1.23]	3.614E-03

Table 10: Bootstrap long-run dynamic multiplier: Distribution price model - Table 4

Model	L^{+}_{ICMS}	CI^{+}	L^{-}_{ICMS}	CI^{-}	T Test
(1)	1***	[0.83,1.15]	0.918***	[0.73,1.11]	4.749E-32
(2)	0.995***	[0.83,1.17]	1.001***	[0.83,1.15]	4.790E-02
(3)	0.979***	[0.8,1.16]	0.811***	[0.62,1.01]	1.206E-127
(4)	1.098***	[0.92,1.27]	0.982***	[0.89,1.09]	9.808E-95

Table 11: Bootstrap long-run dynamic multiplier: Retail and Distribution price model from Tables 5 and 6 for the restricted model with only short-run asymmetries

Model ($\Delta R_{i,t}$)	L_{ICMS}	CI^{+}	Model ($\Delta D_{i,t}$)	L_{ICMS}	CI^{+}
(1)	0.604***	[0.25,0.96]	(1)	0.938***	[0.75,1.13]
(2)	1.003***	[0.83,1.18]	(2)	0.997***	[0.84,1.15]
(3)	0.846***	[0.65,1.05]	(3)	0.912***	[0.76,1.06]
(4)	1.113***	[1,1.22]	(4)	1.005***	[0.9,1.12]

Table 12: ootstrap long-run dynamic multiplier: Retail and Distribution price model from Tables 7 and 8 for the restricted model without asymmetries

Model ($\Delta R_{i,t}$)	L_{ICMS}	CI^+	Model ($\Delta D_{i,t}$)	L_{ICMS}	CI^+
(1)	0.603***	[0.25,0.95]	(1)	0.942***	[0.76,1.12]
(2)	0.997***	[0.82,1.18]	(2)	0.99***	[0.83,1.15]
(3)	0.853***	[0.65,1.06]	(3)	0.909***	[0.75,1.06]
(4)	1.116***	[1.02,1.21]	(4)	1.009***	[0.9,1.12]

...