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ESSAYS ON ASYMMETRIC PRICE TRANSMISSION OF INTERNATIONAL SHOCKS IN BRAZIL

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Tese apresentada ao Programa de Pós Graduação em Economia da Universidade Federal de Juiz de Fora como requisito parcial para obtenção do título de Doutor em Economia.

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"Scientific reasoning is useful to anyone in any job because it makes us face the possibility, even the dire reality, that we were mistaken."
Carol Tavris and Elliot Aronson

ABSTRACT

This thesis introduces two essays on the asymmetric transmission of international prices in the Brazilian economy. They are inspired by both society's intuition and academic perception that price adjustment could be higher when they are positive. The first essay investigates the exchange rate pass-through (ERPT) to domestic prices. The work seeks to advance in the methodological field by bringing estimates drawn from different methods. Notably, we investigate if there are significant differences between slope-based parameters in simple regression models and dynamic impulse response functions obtained in system methods. Results indicate that the degree of ERPT and its asymmetry can vary on the methodological choice. Nonetheless, we demonstrate that the ERPT for consumer prices shows positive asymmetry for various models. In its decomposition, the prices said to be "administered" also have a more significant pass-through for depreciations. In the second essay, we study the asymmetric pass-through of international prices to wholesale fuel prices in Brazil. The motivation lies in the allegation of the country's leading refining company that adjustments in fuel prices would follow international prices, like fuel, oil, and currency. The evidence gathered in a bunch of estimates shows signs of positive asymmetry in the transmission of global shocks to gasoline prices from July 2018 to June 2019. On the other hand, the price policy for diesel showed to be symmetric most of the time. When asymmetry was found, it was negative, meaning that decreases in international prices were being passed through more quickly, although this evidence is not constant across time and models. These findings suggest that the company might have more leeway to exert positive price asymmetry in the gasoline market. We hypothesise that the company could utilise its massive market power to keep prices above their equilibrium level.

Keywords: exchange rate pass-through, asymmetry, inflation, fuel market.

RESUMO

Essa tese apresenta dois ensaios sobre transmissões assimétricas de preços internacionais na economia brasileira. Eles se inspiram tanto na intuição popular quanto na percepção dentro da profissão de que ajustes nos preços podem ser maiores quando positivos. No primeiro ensaio, investiga-se a assimetria do *pass-through* da taxa de câmbio para índices domésticos de preços. O trabalho procura avançar no campo metodológico ao trazer estimativas utilizando diferentes métodos. Particularmente, procura-se investigar se há relevante diferença entre estimativas utilizando parâmetros de regressões simples e funções de impulso-resposta em sistemas de equações. Os resultados indicam que tanto o grau de *pass-through* da taxa de câmbio quanto a sua assimetria podem variar a depender da escolha metodológica. Não obstante, demonstra-se que o repasse para o índice de preço ao consumidor apresenta sinais de assimetria para uma gama de modelos. Em sua decomposição, os chamados preços administrados também possuem repasse mais alto para as depreciações do que aqueles tidos como livres. No segundo ensaio, estuda-se a assimetria do repasse de preços internacionais para o preço cobrado nas refinarias da Petrobrás. O trabalho se motiva na alegação por parte da companhia que os reajustes nos preços da gasolina e do diesel seguiriam vis-à-vis os fundamentos internacionais - a saber: preços internacionais dos combustíveis, preço do barril do petróleo e taxa nominal de câmbio. A evidência obtida em uma gama de modelos aponta sinal de assimetria positiva na transmissão de choques internacionais para a gasolina entre julho de 2018 e junho de 2019. Por outro lado, a política de preços para o diesel se mostrou simétrica na maior parte dos casos. Quando assimetrias foram detectadas, elas foram negativas, indicando que decréscimos nos preços internacionais estavam sendo repassados mais facilmente. Tais resultados sugerem que a companhia possui maior liberdade em exercer assimetria positiva de preços no mercado da gasolina. Uma possível hipótese que os justificam parte da utilização de poder de mercado por parte da companhia para manter um preço acima de seu nível de equilíbrio.

Palavras-chave: *pass-throug* cambial, assimetria, inflação, mercado de combustíveis.

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1.1 Introduction

A vital aspect of the exchange rate pass-through (ERPT) literature over the last years is the study of asymmetric and nonlinear adjustments. There are different types of them. Firstly, a sign-asymmetric adjustment regards the possibility that currency depreciations and appreciations affect price changes differently. The motivation comes from recent macroeconomics literature that supports the idea that crucial economic variables have asymmetric adjustment. Enders and Granger (1998) pointed out that firms' greater likelihood of raising instead of shrinking prices is an essential aspect of many of these macroeconomic models. The downward price rigidity is indeed a recurrent theme in recent macroeconomics. Peltzman (2000), for example, argues that prices rise faster than they fall: the response to a positive shock is at least twice the response to a negative shock, and this difference remains from five to eight months. In terms of the adjustment of domestic prices to exchange rates, asymmetric behaviour can arise from a series of micro-founded optimising behaviour, which aggregates to an incomplete and asymmetric ERPT. Its relevance is manifold. Asymmetric ERPT exacerbates the trade-off between inflation targeting and output stability, as the effort to control prices could be higher in depreciation episodes. At the empirical level, if asymmetries are relevant, any inference based on symmetric models could be biased. Secondly, the existence of menu costs underpins size asymmetry in the ERPT - when the adjustment differs depending on the magnitude of the exchange rate change. As such, episodes of sharp changes in the nominal exchange rate can trigger different levels of inflation, which is particularly relevant when global shocks shift emerging countries' currencies to a large extent.

In this paper, we intend to answer questions both from the economic and methodological perspectives. Firstly, one particular issue raised in the new open macroeconomics literature by Engel (2002) and Bacchetta and Wincoop (2003) is the degree of pass-through varies according to the level in the price chain. Import prices tend to absorb exchange rate fluctuations higher than wholesale and consumer prices. Analogously, prices of upstream industries also tend to absorb shocks on international prices and exchange rates than downstream sectors (MCCARTHY, 2007; ITO; SATO, 2008; CHEN; YANG, 2021). This distinction is important since the extent to which pass-though dies out throughout the price chain tells how adjustments to shocks take place in local economies[1]. Moreover, more open and small economies tend to have higher price changes if they have a large share of tradables. No previous works assess asymmetric ERPT in the Brazilian economy at the three stages of the aggregate price distribution chain often defined in the literature - imports, producer, and consumer prices. So far, studies confirmed the intuition that the pass-through tends to alleviate as the exchange rate shocks are transmitted downwards in the price chain (KIM; ROUBINI, 2000; HAHN, 2003; MCCARTHY, 2007; ITO; SATO, 2007; CA'ZORZI; HAHN; SÁNCHEZ, 2007; ITO; SATO, 2008; MIRDALA *et al.*, 2013), but none has answered if and how asymmetries change throughout this chain, which is one of our goals.

Secondly, most studies fail to account for sectoral prices, which may provide a better understanding of how exchange rate shocks affect the aggregate data and are interesting from the policymaker's perspective. Breaking the price indices into sectoral ones allows answering a particular concern raised by Edwards (2006) and Edwards and Cabezas (2021). If the ERPT to nontradables is significant, the country faces a costlier adjustment process, from expenditure switching to expenditure reduction, i.e., a negative wealth effect replaces the substitution effect from imports to domestic goods. To our knowledge, no previous works assessed the effects of exchange rate on sectoral inflation in Brazil.[2]

Finally, we aim to contribute to the methodological debate. The empirical strategies in the ERPT literature have three generally unanswered issues related to the nature of the time series and economic interpretation - and we tackle all of them. Firstly, we attempt to answer the differences between estimating slope coefficients in single-equation regressions and defining asymmetric dynamic responses in system methods. The latter allows for feedback effects on the exchange rate that could affect ERPT in the long run. Thus, the interpretation differs from the slope-based models that represent the majority of the literature (BRUN-AGUERRE; FUERTES; GREENWOOD-NIMMO, 2016; LOPEZ-VILLAVICENCIO; MIGNON *et al.*, 2016; LOURENÇO; VASCONCELOS, 2018). Secondly, but related to the first item, when estimating asymmetries, one should ask how to identify and interpret the results correctly. This issue was first raised on the oil-market literature by Kilian and Vigfusson (2011a, 2011b). Therefore, we formalise a viable way to express dynamic asymmetric responses for the exchange rate pass-through phenomena and compare it to a recent work that uses a similar method but computes asymmetries differently. Lastly, a portion of the single-equation literature relies on

[1] See the expenditure switching mechanism in Obstfeld and Rogoff (1995, 1998, 2000).
[2] There are non-published works with few methodological links with ours.

first-differences of the data to form stationary series. However, if cointegration between long-run level variables is present, it is misspecification that could bias the ERPT estimates. We thus test a range of models that allow for cointegration relationship, and compare them to simpler models to answer if cointegration is a key ingredient in ERPT models.

Our results indicate that the assertion of asymmetry depends on whether the analysis is based upon slope coefficients or dynamic responses. Intriguingly, results can be fully opposite. As such, slope parameters show strong evidence of asymmetric ERPT at the consumer level, loose evidence at the producer level, and no evidence at the import level. On the other hand, when assessing dynamic responses, import and producer prices - which embeds the bulk of tradable goods - seem to have a proper asymmetric dynamic representation in some models, whereas consumer prices do not. Also, our system methods show a higher degree of asymmetry as the appreciation shock triggers what we identify to be wealth effects in the long run, which alleviates the initial drop in the price indices. This wealth effect is one-sided and does not reduce prices after a depreciation shock. Further, we show that size asymmetry is not triggered by a given threshold above which the exchange rate change must lie, but rather conditional on certain episodes of large exchange rate change. Lastly, the decomposition of CPI shows that the ERPT for prices controlled by the public sector (called “administered” in Brazil) embed strong asymmetry, with higher pass-through after a depreciation than “free” prices (those set by supply and demand forces). This indicates that the effects of depreciations on aggregate inflation have the administered prices as a relevant channel.

The remainder of the paper is organised as follows: section 1.2 depicts the theoretical underpinnings for nonlinear and asymmetric ERPT and shows the most recent empirical advances. The third section discusses the methodological approach, motivates the main issue behind the estimation of asymmetric ERPT, and depicts the data sources and manipulation. Section 1.4 discusses some issues with empirical choices. The fifth section shows the results for asymmetric ERPT with a closer look at methodological choices. The sixth section brings some extensions, like system methods, the possibility of cointegration, size asymmetry, and a price decomposition of the consumer-level inflation. Finally, section 1.7 gathers the main findings and contributions and addresses some avenues for future work.

1.2 Literature Overview

The exchange rate pass-through (ERPT) is described as the price response to a 1% change in the nominal exchange rate (NER). In the process that marked the transition from fixed exchange rate regimes to more flexible ones, understanding the pass-through was important in defining alternative currency regimes' main benefits and costs. The first advocates of flexible exchange rate regimes - the Mundell-Fleming model and Friedman (1953) - assume that pass-through is complete because goods prices are unchanged in the producer's currency. This hypothesis, known as *producer currency pricing* (PCP), is present in the eve of the new open macroeconomics literature, pioneered by the works of Obstfeld and Rogoff (1995, 1998, 2000). When PCP is valid, and pass-through is complete, nominal exchange rates are the main adjustment variable. Under this assumption, they impact final goods prices and allow for an important *expenditure-switching* mechanism. It says that whenever there is an imbalance between relative prices of domestic and foreign goods, consumers switch for the cheaper good, and nominal exchange rates move to achieve a new equilibrium. Therefore, an economy with a flexible exchange rate regime achieves a new equilibrium even in a very new-Keynesian setting where prices are sticky. This setting, of course, allows for monetary policy effectiveness under flexible exchange rates.

However, empirical evidence throughout the nineties shows limited pass-through to consumer prices, i.e., the approach that assumes that nominal prices are set in producers' currencies was at odds with empirical evidence (ENGEL, 2002). This was a justification for models with *local currency pricing* (LCP), as they assume that exporters set prices in foreign currency. In models like Devereux and Engel (1998, 2003), such hypothesis leaves little or no space for expenditure-switching mechanisms, and fixed-exchange rate regimes outperform more flexible ones in terms of monetary policy. Such models also help explain why exchange rates are volatile and disconnected from the real economy. If they have little to no power to impact relative prices, they must change by a large extent to perform adjustments (KRUGMAN, 1989; DEVEREUX; ENGEL, 2002; CORSETTI; DEDOLA; LEDUC, 2008). Nonetheless, it did not take long for claims that both LCP and PCP assumptions are too extreme and not consistent with evidence (BACCHETTA; WINCOOP, 2003). Even though LCP models captured the zero pass-through to final retail prices, they did not reflect the partial pass-through to wholesale prices implied by some micro studies on pass-through (OBSTFELD, 2001). Additionally, the LCP hypothesis predicts that when a country's currency depreciates, its terms of trade improve, which is at odds with the data (OBSTFELD; ROGOFF, 2000).

Rogoff (1996) points out that changes in the exchange rate may not be totally-passed through due to several adjustment costs. Indeed, empirical results supported incomplete ERPT, refuting both "pure" LCP and PCP. Aron, Macdonald and Muellbauer (2014) gathered the main causes for the occurrence and magnitude of incomplete ERPT to import prices. It may be incomplete due to three channels depending on (i) the degree of markup, (ii) the marginal cost, and (iii) nominal rigidities. The two first channels depend crucially on the market structure. The variation in the *markups* allows firms to alter their prices, leading to incomplete ERPT. The presence of variable markup depends on the functional form attributed to the demand curve, which consists in the extent of competition in the domestic market, the ease of substitutability between domestic and foreign goods, and the degree of market segmentation. Exporters' *marginal costs* could explain incomplete ERPT if they vary with the exchange rate, offsetting gains of marginal revenue, which happens when the exporter also needs to import goods as inputs. Finally, *nominal rigidities* are less explored in the literature and come from price stickiness micro-studies. Rogoff (1996) argues that it is insufficient to account for the sluggishness of the adjustment of aggregate prices. Corsetti, Dedola and Leduc (2008) suggest that nominal rigidities might play a role in determining a high portion of LCP. However, they are not strongly required in models that include nontradability, distribution, and price discrimination.

1.2.1 Pass-through at different stages

As seen, neither the LCP nor the PCP assumption is consistent with pieces of evidence. As Bacchetta and Wincoop (2003) observe, the heart of the matter is that the degree of ERPT to consumer prices is much lower than the ERPT to import prices. Engel (2002) also shares the importance of this distinction.

When imported products reach lower stages of the distribution chain, their influence on aggregate domestic prices diminishes. Two related theoretical approaches try to explain the causes of this disconnect between import- and consumer-prices pass-through. First, imported goods must go through a distribution sector to reach consumers. Suppose the imported goods incorporate a significant share of local value added[3]. In that case, consumer prices will not be so sensitive to exchange rate changes (MCCALLUM; NELSON, 2000). In this approach, the pass-through coefficient will be determined by

[3] This local value added embeds costs in the destination market like transport, taxes, tariffs, storage, marketing, advertising, finance, insurance, and rents.

the share of foreign firms selling final goods in the domestic market relative to domestic firms (DORNBUSCH, 1985). Burstein, Eichenbaum and Rebelo (2002) find two channels that muted pass-through from large devaluation episodes. They show that distribution costs and substitution away from imports to lower-quality goods (*"flight from quality"*) can account for the difference between the expected rate of inflation under the PPP hypothesis and the actual rate observed in the data.[4]

In the second approach, imports are intermediate goods, often mixed with domestically produced goods to produce a final good sold to consumers (BACCHETTA; WINCOOP, 2003). The degree of ERPT depends on the share of imported inputs in the technology varieties of the domestic industries.[5] The framework by Obstfeld (2001) has PCP (and therefore full pass-through) at the level of intermediate goods but zero pass-through to consumer prices. A consequence of this setting is that his model embeds a substantial expenditure-switching effect of exchange rate changes, which operates at the firm rather than at the consumer level. As such, he shows that zero pass-through at the consumer level does not suppress the effects of exchange rates on price adjustments. In such context, full pass-through occurs at some upstream level but is muted in final retail prices as producers practice expenditure switching. The model by citeonlinebacchetta2003consumer reaches a similar conclusion but with a different mechanism. They assume that monopolistically competitive exporters sell intermediate inputs to monopolistically competitive final good producers. Firms optimally choose in which currency to set their price. The result indicates that this choice depends on the competitive pressure in the domestic market. The justification goes as follows: traded and nontradable goods compete in their model. As the consumer substitute between them, the larger the nontradable sector, the more producers prefer not to pass through their variation in costs. On the other hand, foreign exporters only compete with other intermediate goods producers. Thus, there is LCP in the sector of final consumption goods and PCP in the intermediate sector. The size of the nontradable sector also mutes the PCP in the intermediate sector.

[4] See also Burstein, Neves and Rebelo (2003) for the role of distribution costs as one of the reasons why purchasing power parity fails.

[5] In Bacchetta and Wincoop (2005), the steady-state share of imports enters the equation of the log output. This defines a route by which exchange rate changes affect the price of domestic goods.

1.2.2 Asymmetry

Some theoretical mechanisms scattered in the literature help explain why we could expect asymmetric behaviour of prices after a shock in the exchange rate. They consist in five main assumptions.[6] These assumptions regard how exporters may react - in terms of choices of prices and output - when faced with an exchange rate change. For the remainder of this section, consider that a *Home* country (H) imports goods from a *Foreign* country (F). P^X is the exporter price represented in monetary units of F, and P^M is the importer price in monetary units of H. The nominal exchange rate (E) is defined as units of F's currency per unit of H's currency (a decrease is an appreciation in H and a depreciation in F). Pass-through to import prices here is the effect on P^M after a change in E.

The first theoretical channel is the market constraints hypothesis (FOSTER; BALDWIN, 1986; WEBBER, 2000). Consider the effect of an appreciation (depreciation) in the Home (Foreign) currency over the price of imported goods. After the devaluation in their currency, foreign exporters gained price competitiveness: if they keep their previous prices (P^X), it will be possible to increase the quantity sold, as P^M will be lower. However, if the production capacity is already on its maximum short-run level or if the costs of adjustment are high, foreign producers are unlikely to expand their supply (or at least it will take some time until they do it).[7]. Thus, higher demand will lead to a hike in the prices, offsetting, totally or partially, the initial effect of the appreciation in H's currency. This hypothesis yields a higher ERPT for depreciations than for appreciations.

The second hypothesis is the presence of market share objectives (FROOT; KLEMPERER, 1988; MARSTON, 1990; KRUGMAN, 1986), and it is derived from monopolistic competitive models. Consider now a depreciation (appreciation) in the Home (Foreign) currency. An appreciation in the exporters' currency causes a loss of price competitiveness and, therefore, a loss of market share overseas. To offset the appreciation, the exporters can lower the prices in their currency (i.e., lower P^X).[8] For H, the ERPT after a deprecation is compensated and thus lower than after an appreciation.

However, there may exist a threshold from which the producers do not consider optimal a decrease in their prices. One can thus define the third hypothesis specifying

[6] For a more complete exposition, see Pollard and Coughlin (2004) and Aron, Macdonald and Muellbauer (2014).

[7] The same mechanism may occur in the case of a trade restriction. In this case, the constraint is not a limit in the industry capacity but a restriction imposed in the destination market.

[8] Notice that if the exporter sets a single price for his good regardless the destination market (LCP hypothesis), the appreciation also tends to reduce the prices at the domestic level in F.

downward price rigidities (KNETTER, 1994). If the appreciation in F is too large, demanding a large reduction in prices and profit margins, the firm could eventually reach a negative mark-up. The exchange rate fluctuation is thus better offset in an episode of depreciation in F, as prices are easily reset upwards than downwards (PELTZMAN, 2000). Turning to the Home currency standpoint, this hypothesis leads to a large ERPT after a devaluation.

So far, the hypothesis assumed P^X as the exporters' choice variable. In the production switching, hypothesis (WARE; WINTER, 1988), the exporter chooses quantity (Q). Suppose that the global market for the tradable good is competitive - so it has an international equilibrium price. The exporters' firms in F are price takers. Consider also that this firm can alter its production between technologies intensive in imported or domestic inputs. In the former, an exchange rate fluctuation will affect both marginal costs and revenue equally[9] and will keep the production decision unaltered. Therefore, being $P^M(Q)$ the inverse demand function defined in Hs currency, the price does not change. In the latter, the exchange rate affects only the exporters' marginal revenue, leading to a shift in the supply and the following price adjustment, leading to some degree of ERPT. The core of this hypothesis is that when the composition of the exporters' inputs varies between domestic and imported intermediate goods, their production incentives also vary when facing an exchange rate fluctuation. To simplify, it is assumed that whenever F's currency depreciates, the exporter changes his inputs to domestic sources and the other way around after an appreciation. This hypothesis results in a larger ERPT for appreciations in H.

Lastly, the fifth hypothesis regards firms facing menu costs[10]. The implication of non-linear pass-through is that domestic firms in F will adjust prices infrequently, even facing changes in the costs of their imported inputs. They are thus prone to absorb small exchange rate changes in their margins and only transmit to prices those changes exceeding a high threshold (ARON; MACDONALD; MUELLBAUER, 2014). As this behaviour may occur in both directions, it is often described as a size asymmetry (the previous hypothesis defines sign asymmetry).

[9] The effect is equal by assumption. However, the responsiveness of marginal costs and revenues to the exchange rate can be different and elicit different net outcomes. Shortly, marginal revenue response to the NER depends on the share of sales overseas, and marginal cost response depends on the import-intensity of inputs in the production. So there might be adjustments in the output, which complicates the analysis without a proper micro model.

[10] The menu costs hypothesis can be extended to a broader class of adjustment and transaction costs with similar reasoning (DELGADO, 1991; SERCU; UPPAL; HULLE, 1995; KLEMPERER, 1995; GIOVANNINI, 1988).

All these assumptions can be summarised as the exporters' capability to offset a shift in the exchange rate. Consider that the export price changes according to $\Delta P^X = \alpha \Delta e$. The exporter offsets a shift in the exchange rate if he can choose a value for α. For the reasons mentioned above, α is different whether the exchange rate appreciates or depreciates or if it changes to a small or large extent. These hypotheses encompass the traditional ERPT literature, which studies the relationship between exchange rate and import prices. When studying prices at the domestic level, all these mechanisms can function. Still, there are also more general models that not only look at import prices but at the general price-setting behaviour of firms (see, for example, Ball and Mankiw, 1994).

1.2.3 The degree of ERPT in Brazil

Lastly, we introduce a few contextualisation on how the ERPT literature has progressed in the Brazilian case and when studies assessed asymmetric effects. In methodological terms, one can divide the literature into three groups. The first embeds macroeconometric models; the second has structural or semi-structural approaches, and the last with micro evidence based on firm- or industry-level data. In terms of coverage, most of them studied the Brazilian case particularly, and only a few included Brazil in a batch of countries, and performed multi-country analysis (BRUN-AGUERRE; FUERTES; GREENWOOD-NIMMO, 2013). Lastly, in the price decomposition scope, the majority focus on either the domestic inflation (CPI) or the price of imports, but none has focused on two or multiple stages of the price chain like ours. As such, our paper fits the macroeconometric models and seeks to investigate methodological issues better and provide a better understanding of the asymmetries throughout the price chain and how the prices of nontradable goods behave.

In appendix A, we show a summary of what we consider to be the main econometric works. Despite different methods and possible nonlinear/asymmetric effects, the ERPT is computed with slope coefficients in all but two cases. Therefore, the few works that study asymmetric responses do it by slope coefficients, usually dummies. The exception is Pimentel, Luporini and Modenesi (2016). As we argue throughout this text, assessing the dynamic responses could indicate different results. We will come back to this point later.

Overall, the ERPT for aggregate CPI inflation in Brazil is low and between 5% and 12%. A few studies reported in appendix A allow for nonlinearities, asymmetries, and price disaggregation, yielding different ERPT estimates than the most simple and linear cases. Accordingly, Carneiro, Monteiro and Wu (2002) show that nonlinearities depending on the level of the real exchange rate and the unemployment rate play a role. When they are not considered, the linear model could either over or underestimate the pass-through coefficient. They also show that prices under government control have lower ERPT than "free" prices, which we shall revisit later in this text. On the other hand, Maciel (2006) obtains fairly high ERPT coefficients for all sub-groups (tradable, non-tradable, and administered). Lourenço and Vasconcelos (2018) studied asymmetries in consumer prices and found asymmetric long-run pass-through (around 25% for depreciations and 12% for appreciations). Finally, Correa and Minella (2010) finds three scopes of nonlinearities in the ERPT: it is higher when the economy is booming when the exchange rate depreciates above a certain threshold, and when exchange rate volatility is lower.

Pimentel, Luporini and Modenesi (2016) is the only one to apply impulse response analysis to a model with an asymmetric decomposition of the exchange rate in Brazil. They compare the symmetric and asymmetric models to show a strong positive asymmetry in all specifications for the CPI response. After 12 months, the effect of a depreciation is 11.38%, and the impact of an appreciation is 2.84%. Their analysis has possible improvements that we seek to contribute upon. Firstly, they do not discuss how their impulse-response functions are obtained, leading us to conclude that they compute it the standard way, i.e., they construct impulse responses exactly as in linear structural VAR models. Hence, it seems that the Kilian and Vigfusson (2011a) (KV) critique of asymmetric VAR models from the oil shock literature applies to their work. Accordingly, the computation of impulse response functions in asymmetric systems should embody the history of the variables and the shock realisation, as will be more explicit in the next section. A second issue within Pimentel, Luporini and Modenesi (2016), also related to impulse responses, is the absence of confidence intervals or error bands in their analysis. Without any assessment of how wide the confidence intervals are at each step of the IRFs, all their discussion on asymmetric effects is imprecise. Also, they do not explicitly test for the significance of asymmetries as they report only the point estimates of their responses. Thirdly, their SVAR model has explicit equations for the asymmetric decomposition of the exchange rate, i.e., they set equations for x^+ and x^- as dependent variables. This incurs inconsistencies, as also pointed out by Kilian and Vigfusson (2011a, p. 433). Lastly, they study only the aggregate CPI inflation. To complement their analysis, we look at the whole price chain from import to consumer prices and seek to decompose the latter into groups to verify where asymmetry is stronger.

1.3 Methodology and Data

1.3.1 Analytical Framework

Our methodological approach relied on the simple theoretical structure standard in the ERPT literature and was popularised by Campa and Goldberg (2005). This structure tries to build with aggregate data the main relations between the prices chosen by the exporters and the domestic and international fundamentals. Consider lower-case letters as natural logs, p_i^m as the import price in the currency of the destination market i; p_j^x as the export price in the currency of the exporting market j. The import price of country i is given by an identity that transforms the export price of j to the importers' currency using the nominal bilateral exchange rate, represented in terms of i's currency for a unit of j's currency:

$$p_i^m = p_j^x + e_{ij} \tag{1.1}$$

The estimation of expression in 1.1 with aggregate price indices as proxies for p_j^x and p_i^m could yield biased ERPT estimates, as these prices are contemporaneously defined in an international general equilibrium context. Moreover, aggregate import/export indices are highly correlated, even for countries with different commodity compositions. That makes the log-linear form of 1.1 oversimplified and not empirically feasible. As for small economies, p_j^x also depends on e_{ij}, the literature works with an alternative and now usual approach that consists of understanding how the exporter firms can set export prices when the exchange rate swifts. The main takeaway is that the degree of pass-through depends on the firms' choice of inputs and the structure of competition in the industry at the international level. Hence, the exporting firms in j try to optimise their price decision by choosing a fraction α_j, $0 \leq \alpha_j \leq 1$ that can offset changes in the exchange rate (BUSSIERE, 2013):

$$\Delta p_j^x = -\alpha_j \Delta e_{ij} \tag{1.2}$$

Bussiere (2013) shows that the degree of pass-through in the destination market is 1 - α_j. To understand what is behind α_j, we follow a general class of models of monopolistic competition based on Amiti, Itskhoki and Konings (2019, 2020) and assume that the exporters have a desired price in their currency $\tilde{p}_i^x$ and that it follows a price identity[11]:

$$\widetilde{p}_i^{\,x} = \widetilde{mk}_j + mc_j, \tag{1.3}$$

where mc_j is the marginal cost of an exporting firm in country j, and $\widetilde{mk}_j$ is the desired markup, which follows a reaction function that depends on the firms' own price ($\widetilde{p}_i^{\,x}$) and the price of its competitors in the exporters' currency ($p_i^* - e_{ij}$), i.e., $\widetilde{mk}_j = \mathfrak{F}(\widetilde{p}_i^{\,x}, p_i^* - e_{ij})$. Amiti, Itskhoki and Konings (2020) works with the following decomposition, which is the full differential of 1.3:

$$\widetilde{p}_i^{\,x} = \kappa\,(p_i^* - e_{ij}) + (1-\kappa)\,mc_j + \varepsilon_j \tag{1.4}$$

where p_i^* is the competitor price in the destination currency, ε_j is a demand shock, and κ is a term that embodies the elasticity of the desired markup concerning realised prices in the exporter's market. Therefore, κ is the own cost pass-through elasticity - it is the elasticity of the firm's price concerning its marginal cost, and $1-\kappa$ reflects the degree of strategic complementarities, i.e., the degree of price competitiveness in the international market. When $\kappa = 0$, the firm does not have strategic complementarities in price setting. When the nominal exchange rate change, the adjustment is given by:

$$\frac{\partial \widetilde{p}_j^{\,x}}{\partial e_{ij}} = \alpha_j = (1 - \varphi_j - \gamma_j) \tag{1.5}$$

where $\varphi_j = {}^{\partial mc_j}/_{\partial e_{ij}}$ captures the sensitiveness of the firm's marginal cost to the foreign currency and $\gamma_j = -\kappa\,\frac{\partial[p_i^* - e_{ij} - mc_i]}{\partial e_{ij}}$ captures the exposure of the firm's desired markup to the foreign currency via the gap between competitor price and the firm's markup.[12] This result implies that when the nominal exchange rate changes: i) the marginal cost changes, especially via the use of intermediate inputs, and ii) the markup changes because, for example, everything else constant, it is more difficult to get a large markup if domestic prices in the destination market are too low compared to the prices in the exporter market. Notice that an aggregation of the term $p_i^* - e_{ij} - mc_i$ can be seen as the real exchange rate deviation following a change in the nominal exchange rate. By 1.5 if the firm does not have substantial amounts of imported inputs on its technology, then its marginal costs remain stable after a change in the exchange rate ($\varphi_j = 0$). Also, if the

[11] Corsetti, Dedola and Leduc (2008) also derives a log-linear expression for the price of exports very similar to this one.

[12] This result follows from 1.3 and 1.4 by observing that $\widetilde{mk}_j = -\kappa\,(p_i^* - e_{ij} - mc_i) + \varepsilon_j$.

firm practices PCP and do not pursue a price strategy for foreign markets, it does not have strategic complementarities in its price setting ($\gamma_j = 0$). When both scenarios are true, α approximates zero, and the degree of ERPT approximates 1.

The structure in 1.5 imposes some simplifications. Firstly, the markup is assumed to vary only with the real exchange rate as we hold constant the demand elasticity in the destination market, the degree of product differentiation[13], and other features of the market structure.[14] Analogously, the marginal costs are assumed to vary only via the input costs channel as the costs of labour and capital that typically enter the production technology do not vary with the bilateral exchange rate.

In this framework, asymmetry implies that there must be a certain degree of convexity/concavity in the function g: $de_{ij} \rightarrow dp^x_j$. This requires nonlinearities in the marginal costs and desired markup functions. For example, according to the market share hypothesis, after an appreciation of j's currency, p^x_j decreases more than it increases after a depreciation. The theoretical consequences of each possible price-setting behaviour are not the aim but the empirical functional forms that address asymmetries in a regression framework.

A typical empirical specification of 1.5 assumes some usual aggregation that allows writing prices at the firm level as price index series at the country level. Usually, the open macroeconomics literature achieves aggregate demand curves for imported goods and price indexes by Dixit-Stiglitz and Armington aggregators.[15] As such, an aggregate import price is a composition of a weighted average of industry-specific import price indices. The typical consequence of such aggregation is that our estimation of pass-through effects depends on changes in the weights of different products in the overall import bundle (CAMPA; GOLDBERG, 2005). One limitation is that aggregation bias may arise and blur the estimates if cross-sectoral differences in microeconomic behaviour cancel out in the aggregate (BUSSIERE, 2013). Still, we can consistently estimate the elasticities by choosing proper functional forms and identifying adequate proxies for theoretical variables.

Following the survey by Aron, Macdonald and Muellbauer (2014), the empirical

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[13] See Barde *et al.* (2008) for an argument on why product differentiation is important in modelling imperfectly competitive industries. They show that including variable markups that depend on strategic behaviour allows for certain pricing behaviour, such as the pricing-to-market hypothesis.

[14] Campa and Goldberg (2005) offer a structure where markups depend on macroeconomic conditions, expressed only as function of the exchange rate, and on industry-specific fixed effects.

[15] This literature is too wide to be covered here. See Chari, Kehoe and McGrattan (2002), Barde *et al.* (2008), and Gust, Leduc and Vigfusson (2010) for a few examples.

specification requires prices in the destination market (w^m) as a proxy for competitors' prices. We also include input costs in the exporters' market (w^x) to account for marginal costs and international costs - such as an index for raw commodities - faced by the exporter's economy (p^{comm}) to account for the costs of imported inputs faced by the exporter. This leads us to:

$$p^m = F(e_i^*, w^x, w^m, p^{comm}) \tag{1.6}$$

where e_i^* is now an effective nominal exchange rate that accounts for the main trading partners of the importing economy, i.e., $e_i^* = \sum_{j=0}^{N} \bar{w}_{ij}\, e_{ij}$, being $\bar{w}_{ij}$ the weight of exports coming from country j to country i. Notice that, despite p^{comm} and w^x both representing price variables of world economies and being thus correlated, they embed different theoretical mechanisms. p^{comm} implies mainly costs of production; whereas w^x besides *also* representing domestic costs (inputs and materials, e.g., energy), it enters the real exchange rate expression. Higher domestic prices can change the exporting decisions toward domestic sales and lower the supply for exports, eventually raising p_j^x. These mechanisms can operate at different rates of adjustment. Whereas higher international prices p^{comm} could imply higher production costs immediately, higher domestic prices causing a supply adjustment via real exchange rates can take longer. For most of our empirical methodology, we use them interchangeably as a measure of international prices (from here, we call just p^w). As we will see next, we test models with both p^{comm} and w^x.

Lastly, Aron, Macdonald and Muellbauer (2014), Aron *et al.* (2014) builds on a more general relationship that also includes the real activity in both destination and exporter's market (y^m and y^x, respectively). Accordingly, omitting control variables correlated with exchange rates could result in biased estimates of the ERPT coefficient. Despite these works using this more general approach in single equation frameworks, the inclusion of real variables is often in the structural general equilibrium approach for the ERPT, which model explicitly the sources in the exchange rate changes (SHAMBAUGH, 2008; BORENSZTEIN; QUEIJO, 2016; FORBES; HJORTSOE; NENOVA, 2017; FORBES; HJORTSOE; NENOVA, 2018; HA; STOCKER; YILMAZKUDAY, 2019). To account for these mechanisms, we also consider the following relation:

$$p^m = F(e_i^*, w^x, w^m, p^{comm}, y^x, y^m) \tag{1.7}$$

We will build log-linear relations based on the structures 1.6 and 1.7 for all import, producer, and consumer prices. This is possible because of the price chain approach

(MCCARTHY, 2007; ITO; SATO, 2007), which sets downstream prices as a function of upstream prices and other controls. Therefore, substituting the determinants of import prices in 1.6 and 1.7 into an expression for producer prices and then further substituting these determinants into an equation for consumer prices would allow the elimination of both import prices and producer prices. Thus, 1.6 and 1.7 are generally valid for prices at different stages of the production chain.

1.4 Issues with empirical choices

Our methodological framework lies in the empirical, applied field that attempts to estimate the price transmission of international shocks into domestic price compositions. This literature embodies works not only on exchange rate pass-through, but also on commodity price pass-through[16] and the effects of oil shocks[17]. To this point, several empirical strategies have been used in this literature. Aron, Macdonald and Muellbauer (2014) provide a guideline and show the main issues on applied modelling: whether to use single or system of equations and whether to use first-differenced variables or cointegration techniques. One further issue is how to model and interpret asymmetry. This last topic is the central subject of the debate between James Hamilton's and Lutz Kilian's works on the effects of oil shocks on the U.S. economy.[18] As Hamilton (2018) introduces and motivates this issue, a regression of y on x yields reasonable estimates of the population linear projection of y on x, which is a sensible answer to the following conditional forecasting question — historically, when x went up, on average by how much did y change? Accordingly, as long as the residuals are conditionally normal, OLS regression will yield asymptotically efficient estimates of the true conditional expectation of consumer price inflation given historical values of the explanatory variables.

However, much of the researchers' and policymakers' interest is in the economy's response over time to an unexpected increase in a given international price. According to Kilian and Vigfusson (2011a), introducing asymmetries in the price transmittal is tricky

[16] Helbling *et al.* (2008), International Monetary Fund (2008), Ferrucci, Jiménez-Rodríguez and Onorante (2010), Gelos and Ustyugova (2017).

[17] Hamilton (2003, 2009), Bernanke, Gertler and Watson (2004), Kilian and Vigfusson (2011a), Kilian and Vigfusson (2011b).

[18] This literature is not our focus, but the contextualisation is that several works, including the ones by Hamilton (1996, 2003, 2009, 2011), conclude that increases in oil prices can trigger recessions in the United States. On the other hand, Kilian and Vigfusson (2011a, 2011b, 2017) criticised these works and among their arguments is the idea that asymmetry was wrongly computed and interpreted. Later, Hamilton (2018) in a non-empirical text argues in favour of simpler OLS-based and local projection methods.

in system models, as slope-based inference and impulse responses can vary dramatically. As much of the asymmetric ERPT literature relies on simple regression models following a pioneer work by Campa and Goldberg (2005), we explore if and to what extent dynamic models differ from traditional estimates. In this process, we compute dynamic asymmetric shocks following the steps of Kilian and Vigfusson (2011a), which was not the case in applied works of the Brazilian case like Pimentel, Luporini and Modenesi (2016). Apart from the main distinction between coefficients from regression models and impulse response computations, we also briefly investigate if cointegration plays a role in the dynamic relationship between domestic prices, exchange rates, and other covariates. We do so because if cointegration between long-run level variables is present, it could lead to bias in the ERPT estimates, especially in the long-run adjustment. We introduce this discussion in section 1.6.4. Before turning to our model, we introduce a brief characterisation of each issue.

1.4.1 Single equation and system of equations

So far, many studies have applied single equations within a framework where the exchange rate is assumed to be exogenous. The data property that underpins this assumption is the stylised statement that free-floating exchange rates follow a random walk process, so improving its modelling with a richer structural approach is challenging. The advantages of such an approach are that single equations can more easily handle structural breaks and asymmetries. Moreover, single equation models in first differences will be more robust to shifts in the mean due to structural breaks. However, this kind of assumption does not allow the variables within the system to exert feedback on the path of exchange rates, being implicit a partial equilibrium environment. An alternative constitutes system methods with a more general equilibrium intuition, where exchange rates are determined by domestic and international shocks and feedback that affects macroeconomic variables' temporal evolution. This design allows for studying the source of the change in the exchange rate. They also allow for assessing the role played by the economic policy in the aftermath of exogenous shocks. If, for example, the interest lies on the central bank's reaction function, the ERPT should be modelled in a system since the exchange rate is endogenous to policy with sometimes non-negligible feedback effects from the monetary policy. As argued by Aron, Macdonald and Muellbauer (2014), impulse response functions drawn from systems will differ from those of single equation estimates that ignore potential offsetting feedbacks. In what follows, we will work with both approaches to assess if their results show consistency and, if they diverge, the possible reasons behind the divergence.

1.4.2 Modelling asymmetry

An unanswered question in the asymmetric pass-through literature is whether the modelling choice affects the underlying asymmetry. The main concern regards how to treat variables that are transformations from the original data. This concern comes from the fact that the most usual way to study asymmetry is by employing the following asymmetric decomposition:

$$\begin{aligned} \Delta e_t^{(+)} &= max\left(\Delta e_t, 0\right) \\ \Delta e_t^{(-)} &= min\left(\Delta e_t, 0\right) \end{aligned} \tag{1.8}$$

From this decomposition, a simple specification would thus be:

$$\begin{aligned} \Delta p_t = c + \alpha_1 \Delta p_{t-1} + \ldots + \alpha_p \Delta p_{t-p} + \beta_1 \Delta e_{t-1}^{+} + \ldots + \beta_p \Delta e_{t-p}^{+} + \\ + \gamma_1 \Delta e_{t-1}^{-} + \ldots + \gamma_p^{-} \Delta e_{t-p}^{-} + \varepsilon_t, \end{aligned} \tag{1.9}$$

As shown in Hamilton (2018), there are three main approaches for assessing the exchange rate pass-through asymmetry. The first, widely used in the literature, is estimating 1.9 by single-equation techniques, like a simple OLS estimation. From such model, one can test for the null of symmetry by building hypothesis of equality between $\beta_1, \ldots \beta_p$ and $\gamma_1, \ldots \gamma_p$. He argues that the linear projection may not differ too much from the true forecast, which is the validity of using OLS to estimate dynamic relations like this. If the residuals are conditionally normal, then this method yields asymptotically efficient estimates of the true conditional expectation of the price inflation given the past values of the explanatory variables. The second method is by Kilian and Vigfusson (2011a), and it regards two further concerns related to the specification and interpretation of results often seen in models like 1.9 in the literature. The concern related to the specification is estimating VAR models where the censored variables calculated in 1.8 are included as dependent variables. Accordingly, they argue there is no way to construct valid structural impulse response functions from such models. The inconsistency arises because the DGP cannot be represented as a VAR model where the asymmetric regressors Δe_t^{+} and Δe_t^{-} substitute the original variable (Δe_t). The concern related to the interpretation of the asymmetry regards how different the estimates of slopes and impulse response functions can be in asymmetric systems. A crucial result in KV is that it is possible to obtain asymmetry in the slope coefficients and symmetry in the dynamic responses and the other way round – symmetry in the slope coefficients and asymmetry in the dynamic responses.[19] This indicates that the econometrician should be aware that asymmetries drawn from slope parameters and dynamic responses could bear different interpretations. Lastly, the third method comes from predictive regressors as in Jordà (2005) and is the choice by Caselli and Roitman (2016). The method consists in simply estimating a separate forecasting equation for each horizon (s) of interest:

$$\begin{aligned}\Delta p_{t+s} = c_s + \alpha_{1,s}\Delta p_{t-1} + \ldots + \alpha_{p,s}\Delta p_{t-p} + \beta_{1,s}\Delta e^{+}_{t-1} + \ldots + \beta_{p,s}\Delta e^{+}_{t-p} + \\ + \gamma_{1,s}\Delta e^{-}_{t-1} + \ldots + \gamma^{-}_{p,s}\Delta e^{-}_{t-p} + \varepsilon_{t+s,s},\end{aligned} \tag{1.10}$$

The OLS for each s gives the optimal forecast of Δp_{t+s} with the information available up to $t-1$. Such a local projection (LP) technique is flexible and easy to implement. LP methods do not involve any non-linear transformation of the estimated slope coefficients to obtain the impulse response functions, as they depend only on the quality of the local approximation (KILIAN; KIM, 2011). In the next section, we cover two out of these three methods.

1.4.3 The empirical Model

We start with a single equation setting, following most of the literature since Campa and Goldberg (2005). The equation models the first-difference of a given price index in domestic currency as a function of the current and lagged asymmetric decomposition of exchange rates ($\Delta e_t^{(+)}$ and $\Delta e_t^{(-)}$) and a set of stationary covariates (x_t) as defined in either functions 1.6 or 1.7:

$$\Delta p_t = \beta_0 + \sum_{i=1}^{p} \beta_{1,i}\Delta p_{t-i} + \sum_{i=0}^{p} \left[\beta_{2,i}^{(+)}\Delta e_{t-i}^{(+)} + \beta_{2,i}^{(-)}\Delta e_{t-i}^{(-)} + \beta_{3,i}x_{t-i}\right] + u_t, \tag{1.11}$$

$$\Delta e_t^{(+)} = max\,(\Delta e_t, 0)$$
$$\Delta e_t^{(-)} = min\,(\Delta e_t, 0)$$

We can compute the slope-based asymmetric exchange rate pass-through in three different horizons. We obtain the short run coefficients by setting $i = 0$; the cumulative

[19] The appendix therein formulates the reasons why this is possible. Shortly, the uncertainty of the realisation of the shocks can drive the dynamic response functions to very different values compared to the slope parameters.

short-run effect up to lag p by the summation from $i = 0$ to p; and the the long run pass-through, given by the expressions $[\sum_0^p \beta_{1,i}]/\left[1 - \sum_0^p \beta_{2,i}^+\right]$ and $[\sum_0^p \beta_{1,i}]/\left[1 - \sum_0^p \beta_{2,i}^-\right]$. Following Brun-Aguerre, Fuertes and Greenwood-Nimmo (2013), we then formulate a set of hypotheses that account for i) the significance of the pass-through; ii) its completeness, and iii) its symmetry:

Zero ERPT

$$\begin{aligned} H_0^1 &: \beta_{2,0}^+ = 0 \quad \text{and} \quad \beta_{2,0}^- = 0; \\ H_0^2 &: \frac{\sum_0^p \beta_{1,i}}{1 - \sum_0^p \beta_{2,i}^+} = 0 \quad \text{and} \quad \frac{\sum_0^p \beta_{1,i}}{1 - \sum_0^p \beta_{2,i}^-} = 0. \end{aligned} \tag{1.12}$$

Complete ERPT

$$\begin{aligned} H_0^3 &: \beta_{2,0}^+ = 1 \quad \text{and} \quad \beta_{2,0}^- = 1; \\ H_0^4 &: \frac{\sum_0^p \beta_{1,i}}{1 - \sum_0^p \beta_{2,i}^+} = 1 \quad \text{and} \quad \frac{\sum_0^p \beta_{1,i}}{1 - \sum_0^p \beta_{2,i}^-} = 1. \end{aligned} \tag{1.13}$$

Symmetric ERPT

$$\begin{aligned} H_0^5 &: \beta_{2,0}^+ = \beta_{2,0}^- \\ H_0^6 &: \frac{\sum_0^p \beta_{1,i}}{1 - \sum_0^p \beta_{2,i}^+} = \frac{\sum_0^p \beta_{1,i}}{1 - \sum_0^p \beta_{2,i}^-} \end{aligned} \tag{1.14}$$

Lastly, we need a choice for the lag length p. Notice from 1.11 that we assume the same lag length for all variables. A more flexible alternative would be choosing information criteria and allowing p to change across the covariates. Campa and Goldberg (2005) use four lags in quarterly data, which covers a year. Caselli and Roitman (2016) use Akaike and Bayesian information criteria to reach three lags with monthly data, although they work with panel data. Bussiere (2013) and Brun-Aguerre, Fuertes and Greenwood-Nimmo (2013) use one and two lags respectively with quarterly data. We estimate models with fixed 3, 6, and 12 lags to allow for more flexibility. Then we turn to information criteria and allow lags to vary for each dependent variable. Section 1.5 further discusses the results.

1.4.4 The Kilian-Vigfusson approach to asymmetry

In the last section, we depart from a simple single-equation method frequently used in the empirical literature to assess asymmetry. In doing so, we perform standard

Wald tests given by the null hypothesis H_0^5 and H_0^6. However, Kilian and Vigfusson (2011a) argues that "[...] slope-based tests are useful in assessing the symmetry of the slope parameters of single-equation regression models. Still, they are not informative about the degree of symmetry of the impulse response function obtained from a fully specified dynamic structural model."(KILIAN; VIGFUSSON, 2011a, p.436). From the same model, one can also compute the impulse response function (IRF), which is the dynamic response of Δp_t at various horizons following a one-off shock in the exchange rate. Unlike previous studies, though, we allow the exchange rate to adjust to a shock on itself by assuming that Δe_t follows an AR(1) and by computing the IRF via recursive forecasts. The dynamic responses depend on the full characterisation of the model, its parameters, past history, and residual disturbances. First, consider the following system model:[20]

$$\begin{aligned} &\Delta e_t = \alpha_1 \Delta e_{t-1} + u_{1t}, \\ &\Delta p_t \;=\; \beta_0 \;+\; \sum_{i=1}^{p} \beta_{1,i} \Delta p_{t-i} \;+\; \sum_{i=0}^{p} \left[\beta_{2,i}^{(+)} \Delta e_{t-i}^{(+)} \;+\; \beta_{2,i}^{(-)} \Delta e_{t-i}^{(-)} \;+\; \beta_{3,i} x_{t-i} \right] \;+\; u_{2t}. \end{aligned} \tag{1.15}$$

Exogeneity in macroeconomics is generally hard to achieve. However, two considerations underpin the structure in 1.15. First, as mentioned, the AR(1) is a good representation of nominal exchange rates, firstly because of its statistical properties, and second, because it is rather difficult to build relations from theory with strong empirical predictive power for the exchange rate (KILIAN; TAYLOR, 2003). Secondly, a parsimonious specification leaves us with few parameters to estimate. Compared to some of the monthly VAR models recurrent in the literature, we have nearly four times fewer coefficients to estimate. That also affects confidence intervals.

Notice that, despite having only one endogenous variable (Δp_t), this setting is similar to a bivariate VARX model[21] with lag restrictions. By assuming that the first difference of the exchange follows an AR(1) (from which shocks are drawn), we are imposing a VAR model where the first equation is exogenous to the model, i.e., there is no contemporaneous structural relation from the price changes to the exchange rate equation. To be a complete VARX model, we would have to define also equations for each covariate of the vector x_t and restrict them to depend only on their past values. We could, for example, define an AR(1) for each covariate, but this would not change anything in our interpretation as the interest is only on the exchange rate shock.

[20] A similar class of asymmetric VAR models became popular in the oil price shock literature, and equation-by-equation robust OLS can efficiently estimate them (KILIAN; VIGFUSSON, 2011a; KILIAN; VIGFUSSON, 2011b; KILIAN; LÜTKEPOHL, 2017).

[21] **V**ector **A**uto**R**egressive model with e**X**ogenous variables.

The main feature of equation 1.15 is that the first variable (Δe_t) does not appear in the second equation, but its asymmetric decomposition - $\Delta e_t^{(-)}$ and $\Delta e_t^{(+)}$ - does.[22] Thus, at each step of the recursive forecast, one must define the identities as in 1.11 before estimating the response of Δp_t in the second equation of 1.15. The AR(1) and the asymmetric identities are thus predetermined to the Δp_t equation.

Finally, to make the pass-through coefficients readily interpretable regardless of the magnitude of the shocks, we follow Belaisch (2003) and compute the cumulative pass-through from the asymmetric impulse-response functions as:

$$
\begin{aligned}
PT_{t+j}^{(+)} &= \sum_{j=1}^{H} \Delta p_{t+j} \Big/ \sum_{j=1}^{H} \Delta e_{t+j}^{(+)} \\
PT_{t+j}^{(-)} &= \sum_{j=1}^{H} \Delta p_{t+j} \Big/ \sum_{j=1}^{H} \Delta e_{t+j}^{(-)},
\end{aligned}
\tag{1.16}
$$

where $\sum_{j=1}^{H} \Delta p_{t+j}$ and $\sum_{j=1}^{H} \Delta e_{t+j}$ are the cumulative effects on the price level and on the exchange rate at j months after the shock, respectively.

To obtain valid impulse response functions in systems with asymmetric regressors, we resort to a bootstrap procedure using recursive forecasts, following Kilian and Vigfusson (2011a, 2011b). In such systems, one can distinguish two asymmetric effects - one regarding the slope coefficients and the other regarding the dynamic response in a given horizon. The latter type of asymmetry is highly nonlinear functions of *all* the parameters, the uncertainty of the innovations, and the history of the process (KILIAN; VIGFUSSON, 2011a). The reason is that the realisation of the variables Δe_t^{+} and Δe_t^{-} is a stochastic process and varies with the previous paths of the variables and the distribution of the innovations.[23] Put differently, their paths could vary significantly even conditional on the same structural parameters and distribution of innovations, but with different histories. Moreover, the uncertainty brought forth by the innovations can lead to symmetry in the slopes and asymmetry in the dynamic responses and the other way round - asymmetry in the slopes and symmetry in the dynamic responses.[24] These properties imply that

[22] Forero, Vega *et al.* (2016) estimate two systems separately: first one with Δe_t and $\Delta e_t^{(+)}$, then one with Δe_t and $\Delta e_t^{(-)}$. The main advantage of our way is the possibility to interpret the slopes directly as depreciation and appreciation. For example, the effect of an appreciation in a system with Δe_t and $\Delta e_t^{(+)}$ is the sum between their coefficients. Both ways are valid. The construction is only problematic when negative changes of e_t are totally omitted from the second equation, for example, a bivariate system between $e_t^{(+)}$ and p_t.

[23] Appendix B illustrates it.

[24] We show this with a simulation exercise in appendix B.

the computation of the traditional impulse response function that treats the innovations as equal to zero under the counterfactual path and places initial conditions equal to zero (or the mean of the process) is incorrect. The computation of impulse responses in a nonlinear dynamic model is not straightforward (FORERO; VEGA *et al.*, 2016). In this sense, the work by Pimentel, Luporini and Modenesi (2016) seems to compute shocks without taking these factors into account, although no further details are discussed therein. The algorithm follows Kilian and Vigfusson (2011b):

Algorithm 1

---

1. Estimate each equation of the model 1.15 by OLS using the whole sample period and store the vector of residuals $\hat{\varepsilon}_t$;

2. Set the numbers:
H: *the horizon for IRFs;*
L: *number of draws for different initial forecasting points;*
M: *number of draws for different shock realisations;*

3) For each $l = 1, ..., L$:

A) Take a block of p consecutive values of the first differences of p_t, $e_t^{(+)}$, $e_t^{(-)}$, and x_t. This defines a history Ω^i.

B) For each $m = 1, ..., M$:

a) Simulate three realisations for ε_{t+h}, for $h = 0, ..., H$, by drawing with replacement from the empirical distribution of the residuals. The realisations are identical, except for that $\varepsilon_{j,t}$ (i.e., $h = 0$) equal to δ and $-\delta$ in two of them, where j is the variable we want to shock. The other one is the baseline.

b) Use the three bootstrapped realisations from a) and the history Ω^i to simulate the pass-through paths $\left\{\hat{PT}_t\right\}_T^{T+h}$, $\left\{\hat{PT}_t^+\right\}_T^{T+h}$, and $\left\{\hat{PT}_t^-\right\}_T^{T+h}$ based on equation equation 1.16.

c) Calculate:

. $\left\{\hat{PT}_t^+\right\}_T^{T+h} - \left\{\hat{PT}_t\right\}_T^{T+h}$, and call it IRF_m^+;

. $\left\{\hat{PT}_t^-\right\}_T^{T+h} - \left\{\hat{PT}_t\right\}_T^{T+h}$, and call it IRF_m^-;

C) Average IRF_m^+ and IRF_m^- to get IRF_l^+ and IRF_l^-.

4) Compute the median and the 68% percentiles from the distribution of IRF_l^+ and IRF_l^- and generate plots.

The model yields a different dynamic response conditional on each possible previous history (Ω^i). These conditional responses are relevant for forecasting and policy purposes. The unconditional response throughout all Ω^i (i.e., the average obtained on step C) is our interest as it is the statistic that tells us the overall importance of exchange rate shocks on the inflation dynamics.

A few details are worth mentioning. Firstly, in the single equation models, the shocks are drawn from an AR(1) for the first-difference of the exchange rate. The remaining exogenous variables do not have equations wherein they are determined throughout the forecasting path. Therefore, throughout these paths, they follow a random non-overlapping block bootstrap of size p. Secondly, we draw residuals directly from the empirical distribution. In this matter, wild bootstrap approaches - which uses an auxiliary distribution to draw random sequences - can also be tested. Thirdly, at step *c)*, we accumulate the IRFs to obtain the cumulative responses:

$$CIRF_t = \sum_{h=0}^{t} IRF_h \tag{1.17}$$

Finally, we observe that the literature on ERPT rarely relates the empirical findings obtained through diverse methodologies. The issue is the lack of analysis between slope-based asymmetry and asymmetric dynamic responses. The asymmetric oil price shock literature has provided some valuable interpretations within this discussion. To borrow the analysis embedded in Kilian and Vigfusson (2011b), one should interpret the slope-based evidence of asymmetry with caution for two reasons. The first is that asymmetric slopes are not necessary nor sufficient for the existence of dynamic long-run asymmetry. This is due to the observation that, as already acknowledged, symmetry (asymmetry) in the slopes does not imply symmetry (asymmetry) in the dynamic responses. The second is that they are likely to be less informative on the dynamics studied by economists. The rejection of symmetry in the slopes of a given model (regardless it is a single equation or a system) only tells us that the asymmetric relationship has better predictive power than the symmetric case. Nonetheless, this is not the same as establishing asymmetric causal links or dynamic relations between the variables. Of course, the literature mentioned above studies the real effects of oil price shocks, while ours is the interaction between nominal variables. Still, even without a full structural model, the ERPT phenomenon is more complex than the average slope response because we cannot ensure exchange rate exogeneity towards domestic variables[25]. Therefore, regardless of following an AR(1), a partial equilibrium VAR model, or a full structural general equilibrium design, the pattern of adjustment of the exchange rates matters if we are to understand long-run exchange rate pass-through. Analysis based on slopes ignores such patterns.

1.4.5 Data

We use monthly data from June 1999 to December 2018. We chose the beginning of the sample to be a semester after Brazil shifted from a fixed currency regime to a floating one to avoid the adjustment process marked by a strong depreciation in that period. It was also the first month of the inflation target regime. Table 1 shows all the data used in the estimations, the corresponding theoretical counterpart, their sources and brief comments. Our dependent variables are three domestic price indices at the import (IPI), producer (PPI) and consumer (CPI) levels. The PPI and CPI series were gathered from the International Monetary Fund, whereas the IPI is from a domestic source (Funcex) because the IMF series is too short.

We use two measures of foreign costs (w^x). Oil prices are a good approximation for international costs in industrialised countries - as they are intertwined with several inputs and transportation costs. In contrast, foreign producer prices and foreign wages account for aggregate industry-level costs faced by foreign producers. We select two commodity price variable (p^{comm}). The Brazilian commodity index merely weights international commodity prices based on their ability to predict inflation, whereas the trade-weighted commodity import price index accounts for commodity prices that affect the costs of foreign exporters relevant to Brazilian trade. Both are in U.S. Dollars, and all the time series representing w^x and p^{comm} are correlated to some extent (see appendix C). The proxies for foreign demand (y^x) are two: the composite leading indicator by the OECD is an index that captures fluctuations in the global business cycle, whereas the Kilian index is also a leading indicator based on the price of dry bulk cargo freights. Both domestic

[25] Comparing to the oil price literature, domestic variables were considered exogenous in some theoretic models in the eighties (BERNANKE, 1983), but the modern approach models the price of oil with important endogenous components (KILIAN, 2008). On the other hand, the literature on exchange rate determination has different ingredients (portfolio and balance of payment adjustments, to name a few) - but all of them are related to the purchasing power parity hypothesis. One of them (MUSSA, 1984) builds, for example, a general price level channel, through which a change in either the expected or unexpected component of this general price will have effects on the demand for money and result in nominal exchange rate adjustments (see MacDonald (2007) for a thorough presentation of such models). Therefore, despite appealing at the empirical level, the case for an exogenous nominal exchange rate has loose theoretical ground.

output variables (y^m) are based on seasonally adjusted industrial production. The first is the first-differenced series (PROD), whereas the second is the gap between the data in level and its long-run trend captured by the HP filter (GAP). Both are interpreted as a measure of slackness in domestic demand. The policy rate is the open inter-banking market rate. Finally, the exchange rate is the nominal effective, measured in foreign currencies per unit of the domestic currency.

Two series used (FPPI and CEPI) are trade-weighted indices. We do so because the weights capture more precisely the relevance of international prices in determining specifically the export prices chosen by the most important Brazilian trade partners[26]. This difference tends to grow larger if some country-specific shock on domestic costs plays a big part in determining FOB export prices destined to Brazil. Nonetheless, the data does not show this mechanism as the weighted trade series hold a very high correlation with the broad indices, as the ones representing only the United States or aggregating some industrialised economies. Their interchangeability does not seem to alter our results, although we did not text for it.

Lastly, all but two variables first-differenced data to reach stationarity. The exceptions are the series that measure slackness of the domestic demand (y^m), as both the de-trended series (PROD) and the gap from a long-run trend (GAP) are already stationary.

[26] The weights used are displayed in the appendix D

Table 1 – Variables and data sources

Theoretical definition		Code	Variable Description	Source
w^x	Foreign costs	OIL	**Oil price** Average between Crude Oil and WTI.	FRED
		FPPI	**Trade weighted PPI** Composite of foreign PPIs weighted by the share in Brazilian imports (yearly rolling weights).	IMF (IFS); UNC; ONW
p^{comm}	Commodity prices	ICBR	**Brazilian Commodity Index** Composite of raw commodities weighted by the degree of pass-through to the inflation.	BCB
		CEPI	**Commodity Export Price Index** Trade weighted index of export prices of 45 commodities.	IMF(CTT); UNC; ONW
p^m	Domestic prices and costs	IPI	**Import Price Index**	Funcex
p^p		PPI	**Producer Price Index**	IMF (IFS)
p^c		CPI	**Consumer Price Index**	IMF (IFS)
w^m		W	**Wages** Real minimum wages.	Ipea
y^x	Foreign demand	WGDP	**Global production** Composite Leading Indicator for real economic activity.	OECD
		KILIAN	**Kilian Index** Index based on ocean dry bulk cargo freight rates.	Kilian (2009)
y^m	Domestic demand	PROD	**Industrial Production** Seasonally adjusted and detrended series.	IMF (IFS)
		GAP	**Output gap** HP filter on the industrial production series.	ONW
r	Policy rate	R	**SELIC interest rate** Policy rate.	BCB
E	Exchange Rate	NEER	**Nominal Effective Exchange Rate** Trade weighted basket of currencies.	BIS

Notes:
BCB - Banco Central do Brasil;
BIS - Bank for International Settlements;
Funcex - Foreign Trade Studies Foundation, Brazil;
FRED - Federal Reserve of St.' Louis Data;
IMF - International Monetary Fund:
(CTT) - Commodity Terms of Trade;
(IFS) - International Financial Statistics;
Ipea - Applied Institute for Economic Research, Brazil;
OECD - Organisation for Economic Co-operation and Development;
ONW - Onw calculations;
UN-C - United Nations Comtrade.
Source: own.

1.5 Results

1.5.1 Slope-based coefficients

We initially analyse slope-based ERPT coefficients that are comparable to the majority of applied works. With monthly data from June 1999 to December 2018, we estimate equation 1.11 by OLS with all the variables as the first-differences of natural log transformation to form stationary time series. For the baseline specification, the vector of covariates x_t has international prices p^w, proxied by the international price of oil, and a measure for domestic slackness w^m, proxied by the output gap.[27] As put in section 2.4.3, we estimate models with 4, 8, and 12 lags and show the results in Table 02.

First, one assumption is needed for the import price model. There is no import price index measured in domestic currency for Brazil. Both the IMF (IFS) and Funcex series are in U.S. Dollars. One way to deal with it is to transform it with the bilateral nominal exchange rate between BRL and USD. However, by doing so, we incur a high contemporaneous correlation with the NEER, which is one of our explanatory variables. This tends to overestimate pass-through because the transformed series will vary almost linearly with the exchange rate. To partially deal with it, we do not estimate import price equations with Δe_0, but only with its lagged values. Doing so, we assume that exporters set new prices in month t based on the exchange rates of months $t-1$ to $t-p$. This way, however, we could miss the instantaneous pass-through of raw goods and commodities, especially from exporters practising PCP. We estimate import price equations with both assumptions - with and without a contemporary link between the NEER and p^m.

In Table 2, panels a.1 and a.2 show the results with both assumptions stated in the last paragraph. Both the degree of ERPT and the direction of asymmetry vary between them. When a contemporaneous relationship is allowed, ERPT is complete and symmetric, given the low values for the Wald statistics that test H_0^5 and H_0^6. When contemporaneity is assumed away (panel a.2), the degree of ERPT falls. Moreover, asymmetry becomes more pronounced, although we still fail to reject H_0^5 and H_0^6.

In the short run ("impact" column), pass-through is higher for depreciations in the import price equation, although we fail to reject hypothesis H_0^5 and H_0^6. In the long run ("cumulative over p months" column), the cumulative coefficients are non-significant. The larger short-run ERPT contrasts with most of the results found in Campa and Goldberg

[27] We test several other specifications in Appendix E.

(2005), where the adjustment for developed countries tends to be bigger in the long run (for example, for countries like Japan, Portugal, Switzerland, and France, pass-through tend to be considerably higher in the long run). The pattern found here is only similar at best to Austria and Ireland, where short-run ERPT is larger, and long-run ERPT is non-significant, which indicates speedy adjustment. The difference, however, is that Brazil has a higher pass-through, especially considering the effects of depreciations ($\beta^{(-)}$ in Table 02), which are up to 0.57. Moreover, the result for depreciation surpasses the average effect found for 27 emerging economies in Caselli and Roitman (2016). Therein, the median response of import prices 12 months after a depreciation shock is 35%, and the upper bound is 50%. Our price response after an appreciation (0.26 to 0.37) is also superior than the upper bound for appreciations in that work, which is around 0.21. This is evidence of a comparably large ERPT for import prices in Brazil.

Panel *b* of Table 2 shows the response of domestic prices at the producer level (p^p). These prices embed the bulk of industrial inputs but also energy and capital costs, the costs of labour, transportation and other distribution services. ERPT is significant for the depreciation variable. In the short run, it is 8%, whereas in the long run, it is between 53% and 60%. Despite the pronounced asymmetry in the point estimate, the chi-squared Wald test again fails to reject H_0^5 and H_0^6 for all three models. Lastly, panel c depicts pass-through at the consumer level (p^c). The impact effect is null. After four months, CPI increases by 15% for depreciations and reaches 24% over a year. The Wald test indicates positive asymmetry at the consumer level. Observe that for all lags, an exchange rate appreciation has a null effect over CPI.

1.5.2 Impulse Response Functions

We now turn to impulse response functions based on recursive forecasts, as depicted in algorithm 1. Our baseline system is composed of an AR(1) for the log changes of the nominal exchange rate and a pass-through equation, as in 1.15. We chose $L = 250$, which approximates the sample size. Thus, all possible blocks for initial conditions have a nearly equal probability of being drawn eventually, $M = 1000$, which suffices to reach normality according to Forero, Vega *et al.* (2016). This yields 250,000 simulations.[28]

Firstly, notice that the expected sign of the pass-through coefficient - defined in 1.16 and using the exchange rate in terms of dollars per domestic currency - is negative. To ease the visualisation, the IRFs are normalised (multiplied by -1). Therefore, whenever

[28] Results achieve convergence and remain consistent with fewer iterations, but we decided to keep those values to follow previous empirical works.

Table 2 – Slope-based asymmetric coefficients

Panel a.1 - Import Prices (assuming $\sum_{i=0}^{p}\beta_{2,i}$)						
	Impact			**Cumulative over *p* months**		
Lags	$\beta_0^{(+)}$	$\beta_0^{(-)}$	Chi-Sq.	$\sum\beta^{(+)}$	$\sum\beta^{(-)}$	Chi-Sq.
4	-1.09$^{\ddagger *}$	-1.03$^{\ddagger *}$	0.40	-1.00$^{\ddagger *}$	-0.94$^{\ddagger *}$	0.15
8	-1.01$^{\ddagger *}$	-1.08$^{\ddagger *}$	0.40	-1.05$^{\ddagger *}$	-0.87$^{\ddagger *}$	1.52
12	-1.05$^{\ddagger *}$	-1.11$^{\ddagger *}$	0.23	-0.97$^{\ddagger *}$	-0.88$^{\ddagger *}$	0.57
Panel a.2 - Import Prices (assuming $\sum_{i=1}^{p}\beta_{2,i}$)						
Lags	$\beta_0^{(+)}$	$\beta_0^{(-)}$	Chi-Sq.	$\sum\beta^{(+)}$	$\sum\beta^{(-)}$	Chi-Sq.
4	-0.26	-0.43	0.33	-0.17*	-0.32*	0.10
8	-0.37	-0.57*	0.47	-0.24	-0.30	0.02
12	-0.32	-0.53*	0.67	-0.05	-0.09	0.01
Panel b - Producer Prices						
Lags	$\beta_0^{(+)}$	$\beta_0^{(-)}$	Chi-Sq.	$\sum\beta^{(+)}$	$\sum\beta^{(-)}$	Chi-Sq.
4	-0.04	-0.07*	0.34	-0.33*	-0.55*	1.42
8	-0.04	-0.08*	0.90	-0.28	-0.60*	2.31
12	-0.01	-0.08*	2.47	-0.10	-0.53*	3.24
Panel c - Consumer Prices						
Lags	$\beta_0^{(+)}$	$\beta_0^{(-)}$	Chi-Sq.	$\sum\beta^{(+)}$	$\sum\beta^{(-)}$	Chi-Sq.
4	0.02*	0.01	0.95	0.00	-0.15*	*3.61*
8	0.02	0.00	1.81	0.01	-0.22*	**6.87**
12	0.02	0.00	1.52	0.01	-0.24*	**5.62**

The Wald tests follow a chi-squared distribution and regard the hypothesis H_0^5 and H_0^6 of symmetry for the short and long run, respectively. **Bold** values indicate p-value < 0.05 and *Italic* values indicate p-value < 0.10;
* indicates ERPT $\neq = 0$ at 95% (hypothesis H_0^1 and H_0^2);
‡ indicates ERPT $= 1$ at 95% (hypothesis H_0^3 and H_0^4).
Source: own calculation.

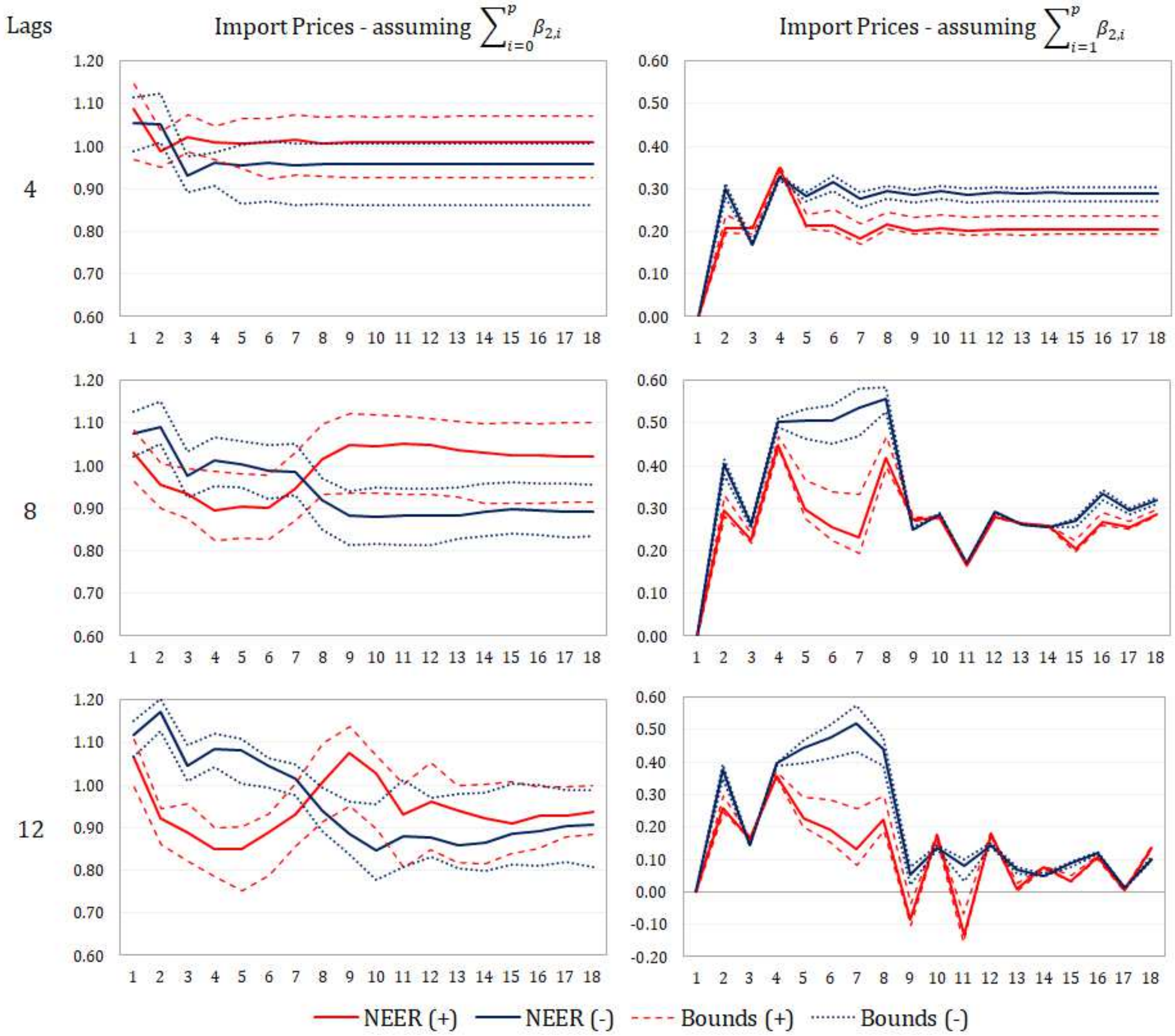
the response is positive, it has the expected sign: positive responses in the plots mean a fall (increase) in the inflation of that particular domestic price after an appreciation (depreciation) shock. Lastly, we plot bands based on the percentiles of the distribution of the 250,000 responses obtained in algorithm 1. As it is unclear in the literature, we leave the discussion of confidence intervals and symmetry tests for section 1.6.2.

Figure 1 shows the impulse response functions after 18 months for a 2 s.d. asymmetric shock in the nominal effective exchange rate. The instantaneous response (at $t = 1$) to shocks follow the assumptions made in the pass-through equation. When a contemporaneous link is allowed, ERPT is complete with slightly positive asymmetry. When it is not, ERPT is zero and import prices have lagged response. In the first case, the asymmetry changes sign in the long run, although we cannot state that they are significant due to the wide bands. In the second case, asymmetry keeps positive throughout most of the path and is symmetric in the models with 8 and 12 lags.

Table 3 compares the results of the two previous sections. Although they are generally similar, especially for import prices, it might be interesting to put such numbers into perspective in terms of actual inflation. That is, how important is the divergence between these results in terms of expected inflation after a shock in the currency? The disturbance applied to compute the impulse responses is a 2-standard-deviation-shock in the residuals of the AR(1). Our sample yields a 6.6% change in the nominal effective exchange rate. A simple calculation can obtain the disturbance in percentage points in each of these inflation measures after such a shock. For example, for the 12-lag model for producer prices, the long-run effect of an appreciation shock would reduce inflation by 0.33% if the regression coefficients are a good fit of the DGP. In contrast, it would reduce inflation by 0.86% if we are to rely on the dynamic responses. The same exercise for the 12-lag-model of CPI shows that a depreciation shock would cause an increase in inflation of around 1.58% using regression coefficients and of 1.12% using IRFs.[29]. Taking the average monthly inflation of such price indexes (0.7% for PPI and 0.5% for CPI) as a reference, these differences do not seem negligible. Notwithstanding, only 16 out of 251 monthly observations of the NEER lie above the ±6.6% change used as a disturbance in the exchange rate. These differences tend to be minor with mote typical-sized shocks.

[29] Notice that these results are conditional on the shock size (2 s.d.) used to compute IRFs. Different intensity of shocks can disturb the model as Kilian and Vigfusson (2011a) put and we show in section 1.4

Figure 1 – Response of import prices to asymmetric shock in the nominal effective exchange rate
Equation 1.15; $x_t = [p_t^w, y_t^m]$



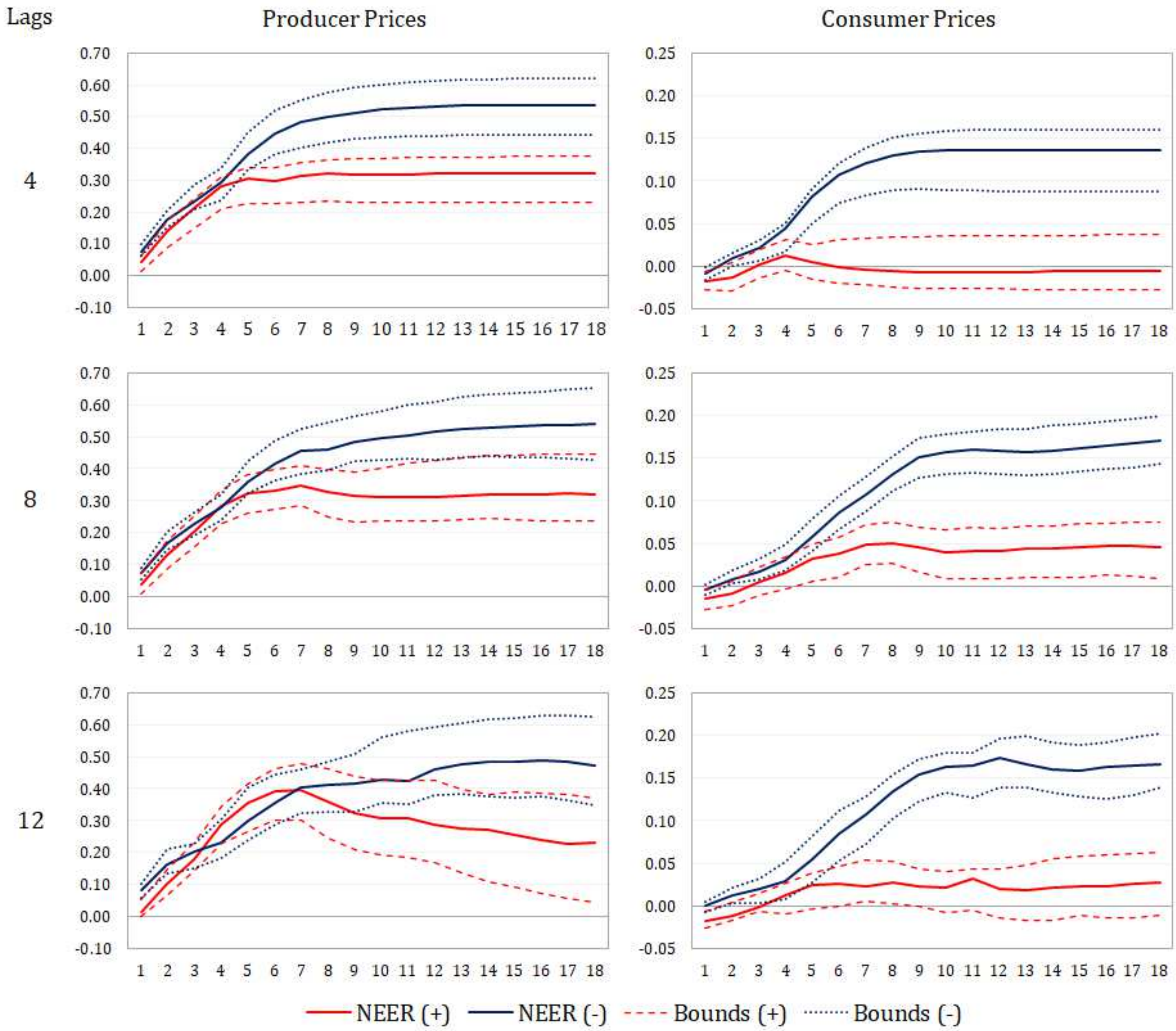
i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

1.6 Extensions

1.6.1 VAR model

The previous discussion departs from a simple structure where shocks are drawn from an AR(1) model, and the exchange rate does not have feedback from other variables. In this setting, the KV method and simple coefficients of OLS regressions show similar patterns, although with some non-negligible differences. This helps answer the first part

Figure 2 – Response of producer and consumer prices to asymmetric shock in the nominal effective exchange rate
Equation 1.15; $x_t = [p_t^w, y_t^m]$



i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

of their argumentation, which is about correctly computing IRFs so that results could potentially differ from slope coefficients. The second part of the argumentation says that building a full structural model avoids the ambiguities of defining a shock in nonlinear reduced-form models, like the ones of the previous session, which are common in the ERPT literature.

Here we formalise a k-variable asymmetric VAR model following the intuition of Kilian and Vigfusson's oil price shock model to assess if something is missing in terms

Table 3 – Comparative: ERPT computed with regression's slopes and dynamic responses

		Import Prices $\sum_{i=0}^{p}\beta_{2,i}$			Import Prices $\sum_{i=1}^{p}\beta_{2,i}$			Producer Prices			Consumer Prices		
Variable	Lags	4	8	12	4	8	12	4	8	12	4	8	12
β^{+}	**Slope**	1.00	1.05	0.97	0.17	0.24	0.05	0.33	0.28	0.10	0.00	-0.01	-0.01
	IRF	1.01	1.02	0.94	0.20	0.29	0.13	0.32	0.32	0.23	0.00	0.05	0.03
β^{-}	**Slope**	0.94	0.87	0.88	0.32	0.30	0.09	0.55	0.60	0.53	0.15	0.22	0.24
	IRF	0.96	0.89	0.91	0.29	0.32	0.10	0.54	0.54	0.47	0.14	0.17	0.17

Notes: IRFs are cumulative responses after 18 months and the slopes are the cumulative over p months, where p is the number of lags in the model; we omit the negative signal for conciseness.
Source: own calculation.

of asymmetric responses obtained in simpler models. Answering this is relevant since simpler models could suffice to give policymakers plausible forecasts on the relationship between prices and the exchange rate. First, from Schorderet *et al.* (2003), notice that a time series can be written as the sum of its positive and negative decomposition. In first differences, this yields:

$$\begin{aligned} &\Delta Y_t = \Delta Y_t^{(+)} + \Delta Y_t^{(-)} \\ &\text{where:} \\ &\Delta Y_t^{(+)} = max\,(Y_t, 0) \\ &\Delta Y_t^{(-)} = min\,(Y_t, 0) \end{aligned} \tag{1.18}$$

For simplicity, the exposition of a k-variable asymmetric system follows the structure of a VAR(1) (it can be easily extended to a p-lag system though):

$$A_0 Y_t = A_0 + A_1 Y_{t-1} + B_0^{+} Y_t^{+} + B_0^{-} Y_t^{-} + B_1^{+} Y_{t-1}^{+} + B_1^{-} Y_{t-1}^{-} + u_t, \tag{1.19}$$

where Y_t is a k-vector of endogenous variables and Y_t^{+} and Y_t^{-} are k-vectors with the asymmetric decomposition of the endogenous variables as in 1.18. A_i, B_i^{+} and B_i^{-}, $i = 0, 1$ are vectors of contemporaneous and lagged coefficients.

A few points are worth mentioning. First, the system uses asymmetric variables only when they are covariates, i.e., it does not define specific equations for Y_t^{+} and Y_t^{-}. Apart from the complexity that emerges when identifying theoretical relations for a variable if whether it increases or decreases (would covariates and lags be the same in both equations?), even if we are interested only in exchange rate depreciations, according to Kilian and Vigfusson (2011a), working with censored explanatory variables renders inconsistent estimators. A proper structural model between y and a censored variable x^+ is not possible.

Secondly, an intuitive consequence of the relation shown in 1.18 is that whenever a given row j of the vectors B_i^+ and B_i^-, $i = [0, 1]$ are different from zero, the respective row in vector A_i is zero, i.e., for one equation j in the system, we whether represent the original variable or its sign decomposition 1.18. This follows the logic of dummy variables. As shown in Kilian and Vigfusson (2011a), after imposing both censored variables accounting for positive and negative changes, one need not include the original variable in level, i.e., the matrix A_1 has zeros in the equations with asymmetric structure and matrices B^+ and B^- have zeros in the equations without symmetric structure.

The third restriction assumes asymmetric shocks departing from only one variable. Apart from the exchange rate shock, we are not interested in configurations with other asymmetric effects within the system. However, even assuming asymmetry departing from only one variable and at one equation, the whole set of k impulse responses might be asymmetric, not only the variable whose equation has explicit asymmetric coefficients. Being i the equation with asymmetric effects and j the positioning of Δe_t in the vector Y_t, we have:

$$b_i^+ = [0, \ldots, b_{ij}^+, \ldots, 0], \quad b_i^- = [0, \ldots, b_{ij}^-, \ldots, 0], \tag{1.20}$$

It is a simple structure because the main interest consists of an unexpected shock at the exchange rate equation, so the other shocks within the system need not be fully identified.

The vector of variables in the VAR model is the same as defined in the previous exercise. At the international level, we use only a price variable (p^w) that could be either foreign producer costs (w^x) or the global prices of commodities (p^{comm}). Both series can capture lagged effects of world supply and demand shocks, although the latter is noisier due to financial markets. As before, p^w is proxied by either the price of oil, the trade-weighted foreign producer prices or the trade-weighted commodity import price index. At the domestic level, we use the domestic demand (y^m), the central bank policy rate (r), and three price indexes at different stages - p^m, p^p, p^c.

Our identification follows the standard textbook recursive ordering of variables. We first depart from the same set of variables we have defined previously. As such, we build a simple 4-variable VAR model with international prices (p^w), the NEER (e), domestic demand (y^m), and a price index (either p^m, p^p, or p^c). Now, the exchange rate equation is not an AR(1), but an equation that allows adjustment with economic rationale. International prices are the first variable in the ordering[30]. Basher, Haug and Sadorsky (2012) shows that, in emerging markets, exchange rates respond in the short run to oil prices. Thus, we define the nominal effective exchange rate as the second equation.[31] Doing so, however, lead us to relax the assumption that import prices do not have a contemporary effect on the exchange rate. Indeed, the position of the exchange rate is not consensual. Ito and Sato (2008) order it after the output gap and the monetary policy and before domestic prices, Kim and Roubini (2000) order it last, while the majority (ROWLAND, 2003; HAHN, 2003; CA'ZORZI; HAHN; SÁNCHEZ, 2007; MCCARTHY, 2007; MIRDALA *et al.*, 2013) order it before all domestic variables. Regardless, the rule of thumb consists in testing different orderings. To make the results of the VAR model readily comparable to those in the previous section, we also consider positioning the exchange rate last so that the import price equation is identical.

4-variable VAR model

Ordering a) p^w, e, y^m, p;

Ordering b) p^w, y^m, p, e,

where p is either p^m, p^p, or p^c - imports, producer and consumer prices, respectively.

In what follows, the estimation of IRFs is the same as in algorithm 1, excluding the block bootstrap for variables not defined in the model, as now all of them are endogenous. Figure 3 shows the results for ordering *a)* and indicates lower ERPT for appreciations (red lines) for all price indexes and lags. Asymmetry becomes more noticeable in the

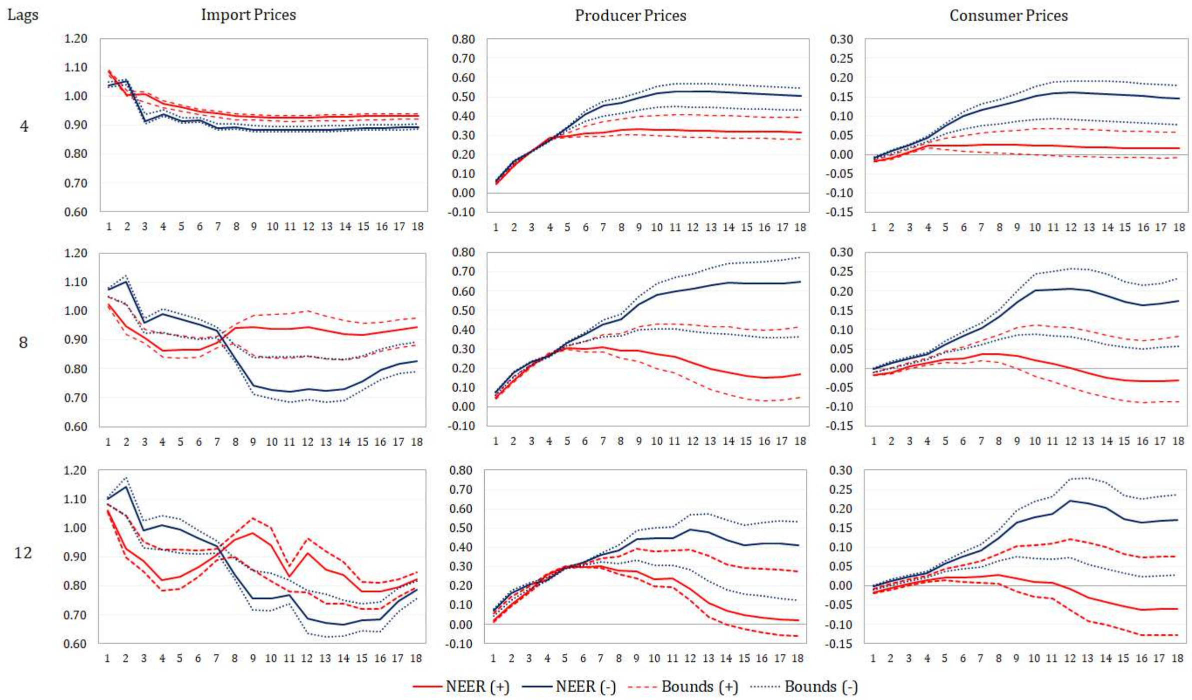
[30] Our identification is not structural in modelling demand and supply forces. Therefore, international price inflation may be affected contemporaneously by both aggregate and supply shocks (MCCARTHY, 2007). The inclusion of a real lagged variable at the international level (y^x) does not change our results.

[31] There is extensive literature on the links between nominal and real exchange rates and nominal and real oil prices. Beckmann, Czudaj and Arora (2017) gathers a big portion of this literature. Initially, they show that there are many theoretical causal links between them and the period under analysis matter. Most studies show that oil price affects the exchange rate in the long run, but not vice versa. In the short run, results vary in both directions. Beckmann and Czudaj (2013) show that an increase in oil prices is associated with a nominal appreciation of the Brazilian currency. Moreover, we use the Dollar price of oil, which is likely to not suffer from exchange rate shocks if we rely on the usual small country assumption. Notice that our ordering choice tests for the sensitivity of positioning the NER elsewhere in the system.

producer and consumer price models with 8 and 12 lags. For ordering *b)* (Figure 4), results show more instability. With four lags, ERPT to import prices has opposing signs to what one would expect in the long run, indicating that a depreciation (appreciation) could decrease (increase) prices. On the other hand, when the VAR model has eight lags, results become more aligned with the ones shown in Table 5, with symmetric long-run ERPT at the ballpark of 0.24. The IRFs of the 12-lag model have high dispersion, with percentiles reaching an implausible range, and we do not show them for the simplicity of our results. Regarding PPI and CPI, results are more well-behaved and similar to Figure 3. When we put contemporaneous relationships aside, ERPT tends to be lower. The impact effect is zero by construction, and long-run ERPT is lower than in ordering *a)*. As such, the ordering matters because impact-ERPT matters even for CPI inflation, despite its tendency to respond slowly to an exchange rate shock.

Figure 3 – Response of domestic price indices to asymmetric shock in the nominal effective exchange rate

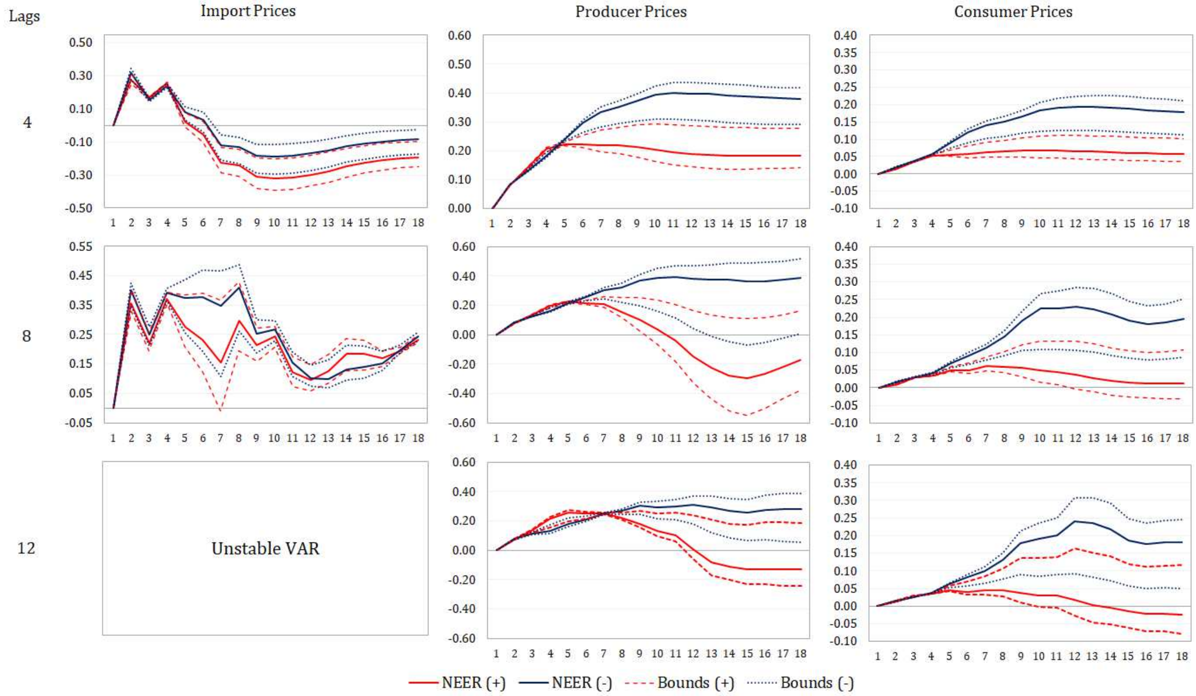
Equation 1.19 (VAR model); Ordering a) p^w, e, y^m, p.



i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

Figure 4 – Response of domestic price indices to asymmetric shock in the nominal effective exchange rate

Equation 1.19 (VAR model); Ordering b) p^w, y^m, p, e.



i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

We now follow Ito and Sato (2007, 2008) and build a 5-variable system, with the addition of the policy interest rate to account for monetary policy reaction. Usually, the interest rate is ordered last (MCCARTHY, 2007; CA'ZORZI; HAHN; SÁNCHEZ, 2007), allowing monetary policy to react contemporaneously to all variables in the model. However, the adjustments to the policy rate in Brazil follow a pre-specified schedule, and it has an instrument target, i.e., the policymaker uses open market operations to pursue a given target for the interest rate. Therefore, for any shock to have contemporaneous effect on the interest rate on a monthly basis would require an automatic reaction from the policymaker, which could not be realistic for a central banker that operates under rules. The baseline ordering places the interest rate after the domestic demand. As before, we allow the exchange rate to have either contemporaneous effects over domestic variables (c) or lagged effects only (d).

5-variable VAR model

Ordering c) p^w, e, y^m, r, p;

Ordering d) p^w, y^m, r, p, e,

where p is either p^m, p^p, or p^c - imports, producer and consumer prices, respectively.

Results reveal IRFs that are akin to the patterns seen in the 4-variable VAR model and we leave them in appendix E.

Observe that in the foregoing specification, there is a different VAR model for each price index. A frequently used system method builds on a price distribution chain in a differenced VAR model, ordered in a way to "identify" the transmission from the "most exogenous" international prices down to the "most endogenous" domestic price at the consumer level (MCCARTHY, 2007; HAHN, 2003; ITO; SATO, 2007). The price chain approach places import, producer and consumer prices in that order. Its main contribution is to account for import penetration and distribution costs, as the extent of consumer-level inflation after a depreciation depends on the degree of imported inputs being used in domestic activities and the presence of distribution costs (BURSTEIN; EICHENBAUM; REBELO, 2002; BURSTEIN; EICHENBAUM; REBELO, 2005). Depending on how distributors adjust their margins throughout the stages of the distribution chain, the effect on domestic inflation can be magnified or diluted. Following the identical orderings as before, we define a system with the price chain structure, which now has seven variables:

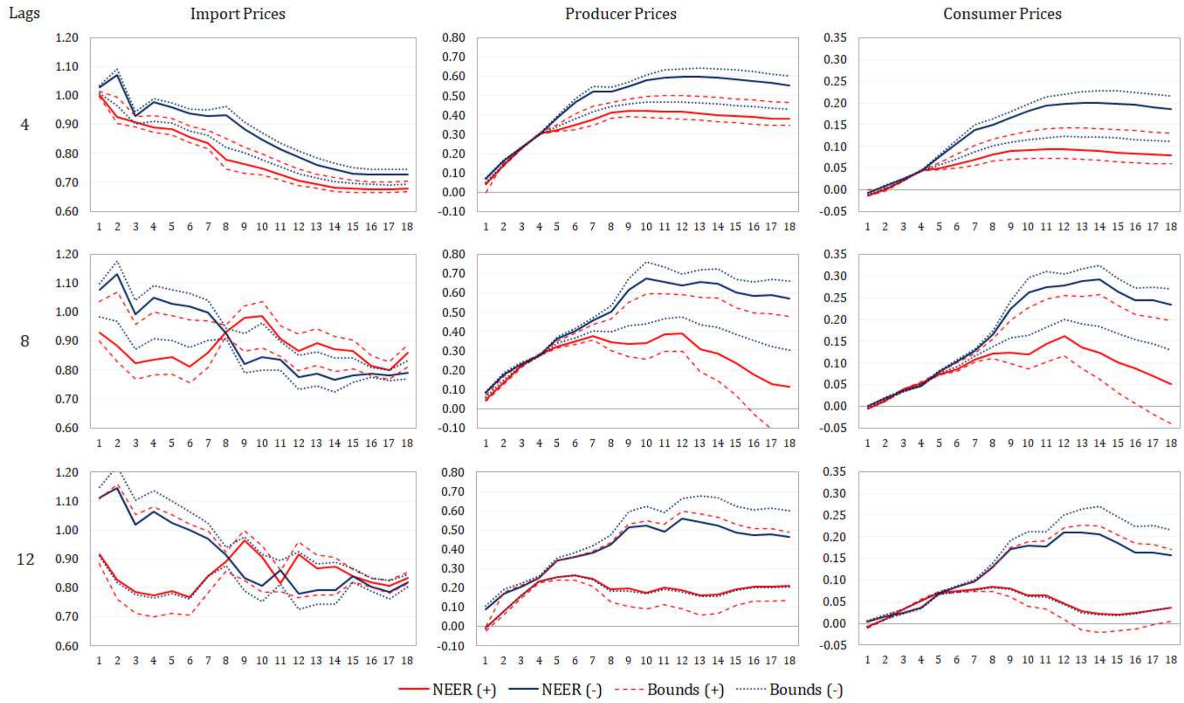
7-variable VAR models

Ordering e) p^w, e, y^m, r, p^m, p^p, p^c;

Ordering f) p^w, y^m, r, p^m, p^p, p^c, e.

Results suggest that ERPT is lower for import prices and somewhat higher for producer and consumer prices in the price chain model. Remarkably, CPI median responses after an appreciation shock are higher than in previous models, although similar after a depreciation shock. This could be due to the different effects that exchange rate and upstream (p^m and p^p) prices have on final inflation. Thus, an appreciation shock transmits more to final CPI inflation when we account for the whole price chain. Observe that the empirical distribution of the IRFs yield wider bands based on the 5th and 95th percentiles. Finally, we do not report results for ordering f as it delivers unstable VAR models with nonsensical wide percentiles for the IRFs bootstrap distribution.

Figure 5 – Response of domestic price indices to asymmetric shock in the nominal effective exchange rate
Equation 1.19 (VAR model); Ordering e) p^w, e, y^m, r, p^m, p^p, p^c.



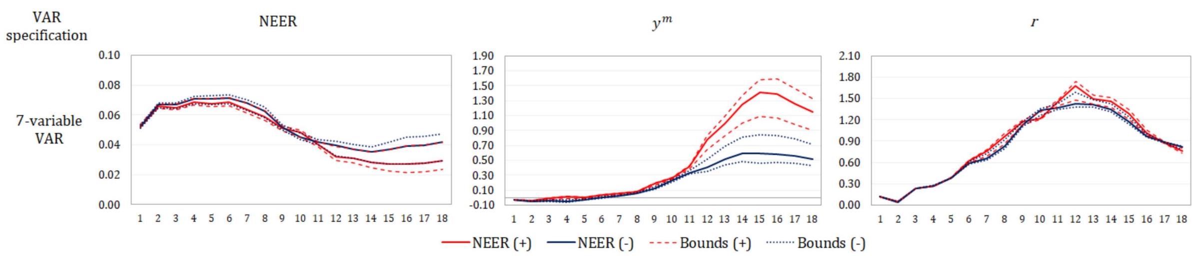
i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

There are possible real effects of exchange rate adjustments. Notably, the change in relative prices of foreign goods, services and assets due to a currency shock might embed substitution and wealth effects, which depend on the price-elasticity of the tradable sector. A wealth effect is present when the appreciation leads to positive growth in demand and, thus, output. Indeed, in the matrix of slope lagged terms, there are highly significant positive coefficients for the exchange rate at $t-2$ in the y^m equation, which could capture potential wealth effects. As such, our system model seems able to capture an increase in output after an appreciation and a decrease after a depreciation. Notice that we achieve this result without imposing specifically asymmetric coefficients in the y^m equation. The asymmetric response of output (and whichever asymmetric response that might appear in the whole system) derives from the asymmetry imposed in the price equations (i.e., the p^m, p^p, and p^c). Without this restriction, the system is a textbook VAR where all the responses are symmetric[32]. To better illustrate what we so far assumed by hypothesis to

[32] The responses are symmetric in the analytic solution to capture IRFs. If we still use the (KILIAN;

be a wealth effect, figure 3 shows the CIRFs for the three other endogenous variables of the VAR model, namely, the nominal effective exchange rate (NEER) and the output gap (y^m). Again, we normalise the responses in terms of elasticity (i.e., the ratio between the actual CIRF and the exchange rate CIRF). The exchange rate and the interest rate respond roughly symmetrically to an exchange rate shock, whereas the output has an asymmetric reaction. For the latter, an appreciation shock increases output to a higher level than a depreciation shock decreases, regardless of what causes these shocks. If this effect yields large demand and thus large prices, it helps us explain why in figure 2 the pass-through after an appreciation is alleviated after the 12^{th} month.

Figure 6 – Responses of nominal effective exchange rate, output and interest rate to an asymmetric shock in the nominal effective exchange rate.
Equation 1.19 (VAR model); Ordering e) p^w, e, y^m, r, p^m, p^p, p^c;



i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

1.6.2 Testing for Asymmetry

Significant asymmetry in the slope coefficients shows that specifying the model with asymmetric decomposition renders a better fit than the symmetric case. Asymmetric dynamic responses are likely to represent better asymmetric effects stemming from structural sources, which could be further confirmed in a fully theoretical structural representation. Even if there is asymmetry in the slope parameters of the equations, it does not necessarily imply large effects on the implied impulse response function.[33] The other way around is also possible, as argued in section 1.4.4.

One relevant topic is assessing if the asymmetric model is a good representation of the DGP. Put in other words, one needs to assess the significance of the asymmetric

VIGFUSSON, 2011a) algorithm (or other bootstrap methods) in a symmetric VAR, some asymmetry can pop up, but they will likely be too small to have any significance.

[33] See appendix B for a deeper discussion.

patterns, which justifies the choice for an asymmetric modelling.[34] Pimentel, Luporini and Modenesi (2016), for example, do not provide such analysis. There are two ways of assessing it. The most traditional one is through confidence intervals. However, there is no obvious way to compute confidence intervals for nonlinear IRFs, as several have been proposed.[35] The procedure in Forero, Vega *et al.* (2016), which relies on Koop, Pesaran and Potter (1996) and Potter (2000), builds confidence intervals by changing the algorithm 1 to embed uncertainty regarding the parameter values. A further loop simulates a bootstrap sample and re-estimates the parameters. After this, one computes percentiles covering the $1-\alpha$ confidence region.[36]

A second option, perhaps more appropriate to asymmetric VAR models, is an asymptotic valid chi-squared Wald test offered by Kilian and Vigfusson (2011a), with the null of symmetric impulse responses. So far in the literature, it is not clear what are the main differences, pros and cons between building confidence intervals for nonlinear impulse responses and testing the joint null hypothesis of symmetric responses to positive and negative exchange rate shocks. Nonetheless, their results could lie close to each other as both use bootstrap techniques.

For the baseline single-equation regression models, we have already shown the slope-based testing of the nulls H_0^5 and H_0^6. Their results are in Table 2. Then, for the baseline VAR models, we report the test of the joint null hypothesis of symmetric responses:

$$I_p(h,\delta) = -I_p(h,\delta) \tag{1.21}$$

where $I_p(h,\delta)$ is the impulse response of the price index p at horizon h after a shock of size δ.

Results show that assessing asymmetries by slope coefficients in single-equation models can lead to different conclusions to assessing them by dynamic long-run responses in a system model. For import prices, for example, Table 2 reports symmetric coefficients, whereas Table 6 shows asymmetry for the 12-lag dynamic models. Only the VAR model has significant asymmetries at the 95% level for producer and consumer prices. Moreover, as seen in appendix I, for other VAR models (e.g., with 4 lags), asymmetry for CPI

[34] Observe that the bounds drawn at each plot only represent the 5th and 95th percentiles over the empirical distribution of all LxM impulse response functions.

[35] Sims and Zha (1999), Lütkepohl (2000), Inoue and Kilian (2013), Winker, Helmut and Staszewska-Bystrova (2014)

[36] The idea is traditional on the literature of bootstrapping confidence bands and follows Efron (1982), Efron (1992), Efron (1992), Efron and Tibshirani (1994), and Runkle (1987).

models do not hold. Thus, significant asymmetry in the slope coefficients does not imply asymmetry in all the dynamic nonlinear responses, which was overlooked in Pimentel, Luporini and Modenesi (2016).

Table 4 – *p-values* for the dynamic response-based asymmetry test

	AR(1)								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
H	1	2	4	1	2	4	1	2	4
1	0.61	0.55	0.58	0.13	0.12	0.12	0.36	0.32	0.32
2	**0.05**	**0.00**	**0.00**	0.32	0.29	0.28	0.63	0.61	0.61
3	*0.06*	**0.01**	**0.01**	0.27	0.20	0.20	0.78	0.75	0.75
4	*0.08*	**0.01**	**0.01**	0.19	0.09	0.09	0.82	0.81	0.81
5	0.14	**0.02**	**0.02**	0.16	*0.06*	*0.05*	0.54	0.53	0.53
6	0.17	**0.03**	**0.03**	0.21	*0.09*	*0.08*	0.36	0.36	0.38
7	0.17	**0.03**	**0.03**	0.21	*0.07*	*0.07*	0.43	0.43	0.45
8	*0.06*	**0.01**	**0.00**	0.25	*0.07*	*0.06*	0.26	0.20	0.22
9	*0.06*	**0.00**	**0.00**	0.33	0.10	*0.09*	0.28	0.19	0.21
10	*0.08*	**0.01**	**0.01**	0.41	0.15	0.13	0.36	0.26	0.28
11	*0.09*	**0.00**	**0.00**	0.46	0.17	0.15	0.44	0.33	0.35
12	*0.09*	**0.00**	**0.00**	0.40	*0.08*	*0.07*	0.33	0.15	0.15
13	0.12	**0.01**	**0.00**	0.47	0.10	*0.09*	0.41	0.18	0.18
14	0.16	**0.01**	**0.01**	0.55	0.14	0.12	0.48	0.23	0.23
15	0.20	**0.01**	**0.01**	0.60	0.17	0.15	0.54	0.28	0.29
16	0.25	**0.02**	**0.02**	0.67	0.22	0.20	0.61	0.34	0.35
17	0.30	**0.02**	**0.02**	0.72	0.26	0.24	0.68	0.41	0.42
18	0.36	**0.04**	**0.03**	0.77	0.32	0.29	0.73	0.47	0.48
	5-variable VAR (ordering c)								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
H	1	2	4	1	2	4	1	2	4
1	0.71	0.21	0.23	**0.03**	**0.02**	**0.04**	*0.06*	0.28	0.33
2	0.21	**0.00**	**0.00**	**0.01**	*0.06*	0.11	0.17	0.54	0.61
3	0.28	**0.00**	**0.00**	**0.00**	**0.01**	**0.03**	*0.08*	0.35	0.44
4	0.32	**0.00**	**0.01**	**0.00**	**0.01**	**0.03**	0.12	0.51	0.61
5	0.39	**0.01**	**0.01**	**0.00**	**0.01**	**0.02**	**0.03**	0.17	0.26
6	0.50	**0.01**	**0.02**	**0.00**	**0.01**	**0.04**	**0.05**	0.26	0.36
7	0.55	**0.00**	**0.01**	**0.00**	**0.00**	**0.00**	**0.05**	0.16	0.25
8	0.66	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	**0.05**	*0.07*	0.12
9	0.75	**0.00**	**0.01**	**0.00**	**0.00**	**0.00**	**0.04**	**0.01**	**0.01**
10	0.82	**0.00**	**0.01**	**0.00**	**0.00**	**0.01**	*0.06*	**0.01**	**0.02**
11	0.85	**0.00**	**0.00**	**0.01**	**0.00**	**0.00**	*0.06*	**0.01**	**0.03**
12	0.83	**0.00**	**0.00**	**0.01**	**0.00**	**0.00**	**0.03**	**0.00**	**0.01**
13	0.88	**0.00**	**0.00**	**0.01**	**0.00**	**0.00**	**0.04**	**0.00**	**0.01**
14	0.91	**0.00**	**0.01**	**0.01**	**0.00**	**0.00**	*0.06*	**0.01**	**0.01**
15	0.94	**0.00**	**0.01**	**0.02**	**0.00**	**0.00**	*0.09*	**0.01**	**0.02**
16	0.96	**0.01**	**0.01**	**0.03**	**0.00**	**0.00**	0.10	**0.01**	**0.03**
17	0.97	**0.01**	**0.02**	**0.04**	**0.00**	**0.01**	0.14	**0.02**	**0.04**
18	0.98	**0.02**	**0.03**	**0.04**	**0.00**	**0.01**	0.16	**0.03**	*0.06*

The significance level regards the Wald test with null $H_0 : I_p(h,\delta) = -I_p(h,\delta)$.
Based on 10,000,000 simulations (see algorithm 2, displayed in appendix B).
Bold values indicate p-value < 0.05.
Italic values indicate p-value < 0.10.
Both models have $p = 12$.

1.6.3 Cointegration

A portion of the single-equation literature relies on the first differences of the data to form stationary series. The typical ERPT equation by Campa and Goldberg (2005) with first-differenced data can be estimated with several lags on the first-differentiated variables to allow a gradual adjustment of the exchange rate as we did. However, if cointegration between long-run level variables is present, it could lead to bias in the ERPT estimates, especially in the long-run adjustment. There are some ways to embed long-run relationships in the ERPT estimates. In the single-equation approach, works like Aron, Macdonald and Muellbauer (2014) and Brun-Aguerre, Fuertes and Phylaktis (2012) include lagged and somewhat parsimonious *ECM* terms. Other studies like Delatte and López-Villavicencio (2012), Jammazi, Lahiani and Nguyen (2015), and Lourenço and Vasconcelos (2018) apply the bound-test approach to cointegration via ARDL models. System cointegration is the choice by Karoro, Aziakpono and Cattaneo (2009) and Aron *et al.* (2014). Regardless of the strategy, pre-testing for cointegration is necessary. Notice, however, that the actual empirical differences in ERPT estimates after controlling for cointegration are not discussed in the literature, and this is a vague subject. Generally, if the true relationship is in levels and these levels covariates are cointegrated, first differencing may fail to capture long-run information. However, if short-run ERPT is of interest (which is often the case for policymakers), estimating in differences is likely to be more robust than in levels, especially if structural breaks shifting the mean of the process is present (ARON; MACDONALD; MUELLBAUER, 2014).

The main disadvantage of working with cointegration is the well-known issue with the power of unit root tests.[37] Nonetheless, the popularity of the autoregressive distributed lag (ARDL) approach comes from its applicability irrespective of whether the underlying regressors are purely I(0), purely I(1) or mutually cointegrated (NARAYAN; NARAYAN, 2005) - which makes the interpretation of results less sensitive to the power of unit root tests.

One question that may arise is if it is possible that differencing each component of the system individually (like in 1.11) distorts interesting features of the relationship between the original variables (LÜTKEPOHL, 2013). The exclusion of long-run adjust-

[37] The references are manifold. Cochrane (1991), for example, states that tests for unit roots have arbitrarily low power in finite samples because of the random walk component that every unit root series have. Also, a well-known rule discussed in Campbell and Perron (1991) is that a nonrejection of the unit root hypothesis may be due to the misspecification of the deterministic components included as regressors.

ments when it is present in the DGP could lead to biased estimates.[38] To account for possible long-run effects, we modify the structure in equation 1.11 in order to reach a nonlinear autoreressive distributed lag (NARDL) model:

$$\Delta p_t = \beta_0 + \gamma ECM_{t-1} + \sum_{i=1}^{p} \beta_{1,i} \Delta p_{t-i} + \sum_{i=0}^{p} \left[\beta_{2,i}^{+} \Delta e_{t-i}^{+} + \beta_{2,i}^{-} \Delta e_{t-i}^{-} + \beta_{3,i} x_{t-i} \right] + u_{2t},$$

$$\text{where, } ECM_{t-1} = \alpha_1 \, e_{t-1}^{+} + \alpha_2 \, e_{t-1}^{-} + \alpha_3 x_{t-1} - p_{t-1} \tag{1.22}$$

The NARDL approach to cointegration is a simple and flexible nonlinear dynamic framework capable of simultaneously modelling asymmetries both in the underlying long-run relationship and in the short-run dynamic adjustment. Following the linear case by Pesaran and Shin (1998) and Pesaran, Shin and Smith (2001), the work by Shin, Yu and Greenwood-Nimmo (2014) proposes a pragmatic bounds-testing procedure for the existence of a stable long-run relationship which is valid irrespective of whether the underlying regressors are $I(0)$, $I(1)$ or mutually cointegrated. However, one caveat regarding NARDL models is often omitted in empirical applications. For the single-equation analysis to be valid, the set of covariates in the error correction model must be exogenous. Accordingly, the (N)ARDL approach is feasible when there is *one* cointegrating vector within the system. Pesaran, Shin and Smith (2001) shows that for the correct identification of Δp_t in the system in 1.22, one should not have a bi-directional relationship between p and the right-hand side variables (e^{+}, e^{-}, x). This result is the same as stating that the system must hold at most one conditional level cointegrating relationship between p and (e^{+}, e^{-}, x) (MCNOWN; SAM; GOH, 2018). On the contrary, multiple cointegrating vectors calls upon the traditional VECM techniques by Engle and Granger (1987), Johansen, Juselius *et al.* (1990), Phillips and Ouliaris (1990), and Johansen *et al.* (1995). Hence, to test the adequacy of the NARDL equation, we further formulate a new hypothesis. Consider first the following VAR-system written in the form of a vector ECM, where we now assume a matrix notation:

$$\Delta \boldsymbol{z}_t = \boldsymbol{a}_0 + \boldsymbol{a}_1 t + \boldsymbol{\Pi} \boldsymbol{z}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{z}_{t-i} + \boldsymbol{\epsilon}_t, \tag{1.23}$$

[38] Kilian and Lütkepohl (2017) argues that not using all available information about the model variables' cointegrating structure reduces the estimator's accuracy in small samples. There is a new wide field of the time-series modelling that focus on the long-run behaviour of nonstationary variables that might hold asymmetries in their adjustments (PARK; PHILLIPS, 2001; SCHORDERET *et al.*, 2003; GRANGER; YOON, 2002; SAIKKONEN; CHOI, 2004; ESCRIBANO; SIPOLS; APARICIO, 2006; BAE; JONG, 2007).

where $\boldsymbol{z}_t$ is a vector of variables in the form $[y_t, x_t^1, ..., x_t^k]$, a_0 and a_1 are parameters, t is a deterministic trend and $\boldsymbol{\Pi}$ is a matrix of parameters of the form:

$$\boldsymbol{\Pi} = \begin{pmatrix} \pi_{yy} & \boldsymbol{\pi}_{yx} \\ \boldsymbol{\pi}_{xy} & \boldsymbol{\Pi}_{xx} \end{pmatrix} \tag{1.24}$$

Pesaran, Shin and Smith (2001) shows that the correct identification of y_t in the system in 1.23 requires the absence of a bi-directional relationship between x_t and y_t, which implies that the k-vector $\boldsymbol{\pi}_{xy}$ must be equal to zero. This result is the same as stating that the system must hold at most one conditional level cointegrating relationship between y_t and $\boldsymbol{x}_t$. We can transform 1.23 into its conditional ECM form for the vector $\boldsymbol{x}_t$ conditional on $\boldsymbol{x}_{t-1}, y_{t-1}, \Delta y_{t-i}$, and $\Delta \boldsymbol{x}_{t-i}$:

$$\Delta \boldsymbol{x}_t \ = \ \boldsymbol{a}_{x0} \ + \ \boldsymbol{a}_{x1} t \ + \ \boldsymbol{\Pi}_{\boldsymbol{xx}} \boldsymbol{x}_{t-1} \ + \pi_{xy} y_{t-1} \ + \ \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_{xi} \Delta \boldsymbol{z}_{t-i} \ + \ \boldsymbol{\epsilon}_t \tag{1.25}$$

From 1.25, we test the exclusion of y_{t-1}. In case of non-rejection of the null $H_0^{xy} : \boldsymbol{\pi}_{xy} = 0$, we fulfil Pesaran's assumption 3 and the UECM for $\boldsymbol{x}_t$ is:

$$\Delta \boldsymbol{x}_t \ = \ \boldsymbol{a}_{x0} \ + \ \boldsymbol{a}_{x1} t \ + \ \boldsymbol{\Pi}_{\boldsymbol{xx}} \boldsymbol{x}_{t-1} \ + \ \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_{xi} \Delta \boldsymbol{z}_{t-i} \ + \ \boldsymbol{\epsilon}_t, \tag{1.26}$$

which is equivalent to 1.22.

There are different ways of testing and then including such long-run components in the ERPT equation. We first estimate the model 1.22 using the same unrestricted set of covariates as before. Pesaran's bound test over the system indicates either multiple cointegration relationships or no cointegration at all. Thus, we do not carry on with this analysis.[39] The only exception was a very parsimonious ECM term without the domestic variables, with only the exchange rate and international producer prices (w^x). The ECM is:

$$ECM_{t-1} \ = \ \alpha_1 \, e_{t-1}^{+} \ + \ \alpha_2 \, e_{t-1}^{-} \ + \ \alpha_3 w_{t-1}^{x} - p_{t-1} \tag{1.27}$$

It is also possible to impose theoretical restrictions on how one builds the long-run structure in 1.22 and, particularly, which variables of the vector x_t to include in it. Consider that a log-linear equation of the form can represent the function 1.7:

[39] Results and are in appendix F.

$$p_t^m = \beta E_t + \alpha_1 w_t^x + \alpha_2 p_t^{comm} + \alpha_3 w_t^m + \alpha_4 y_t^x + \alpha_5 y_t^m + \varepsilon \tag{1.28}$$

Long-run price homogeneity in the form of lack of money illusion, for example, states that for a *fixed* exchange rate, doubling foreign prices eventually doubles import prices (i.e., $\alpha_1 + \alpha_2 + \alpha_3 = 1$). In the long run, import prices cannot react more than proportionally after changes to foreign prices. The second restriction states that doubling the exchange rate at given foreign prices, for instance, is equivalent to doubling foreign prices at a given exchange rate, i.e., $\beta = -(\alpha_1 + \alpha_2)$. Imposing this last restriction yields the following ECM:

$$ECM_{t-1} = \alpha_1(w_{t-1}^x - E_{t-1}) + \alpha_2(p_{t-1}^{comm} - E_{t-1}), \tag{1.29}$$

at which $\alpha_3 = 0$. Imposing it requires that the long-run elasticities sum to unity, ensuring that, when the exchange rate does not change, an equal increase in each of these price variables results in the same proportionate increase in import prices. Observe that one can estimate 1.29 by changing w^x and p^{comm} from Dollars to domestic currency. We can also impose asymmetry in the long run adjustment of 1.29:

$$\begin{aligned} ECM_{t-1} = \alpha_1^{(+)}(w_{t-1}^x - E_{t-1})^{(+)} + \alpha_1^{(-)}(w_{t-1}^x - E_{t-1})^{(-)} + \\ + \alpha_2^{(+)}(p_{t-1}^{comm} - E_{t-1})^{(+)} + \alpha_2^{(-)}(p_{t-1}^{comm} - E_{t-1})^{(-)} \end{aligned} \tag{1.30}$$

where now the price homogeneity restriction becomes $\beta^{(+)} = -(\alpha_1^{(+)} + \alpha_2^{(+)})$ and $\beta^{(-)} = -(\alpha_1^{(-)} + \alpha_2^{(-)})$.

Unlike the original set of covariates, models with ECM structures 1.29 and 1.30 are valid NARDL models. To see why, we show the main cointegration tests in table 5. Observe that F_{yx} is significant, while $F_{xy(i)}, i = [1, 2, 3]$ are not, with one exception. Moreover, the Johansen trace test also yields one cointegration vector, favouring NARDL models.

We estimate these three NARDL models by OLS and compute impulse response functions of shocks drawn from an AR(1) for the log difference of the NEER. Results are in table 6 and figure 7. They all show cumulative short-run asymmetry for import prices. This means the pass-through after the first one to four months is larger for depreciations. The IRFs confirm this in all three cases. Results diverge when it comes to long-run adjustment. For the third model, the long-run slope elasticity of depreciation is larger. In contrast, the impulse response after 18 months converges to a scenario where the

Table 5 – NARDL models (1) - Cointegration

	Import Prices		
Model:	1.27	1.29	1.30
F_{yx}	$6.220^{\dagger}$	$7.944^{\ddagger}$	$9.971^{\ddagger}$
Symmetry			
F_{yx} short-run symmetry	$5.652^{\dagger}$	$7.692^{\ddagger}$	$7.772^{\ddagger}$
F_{yx} long-run symmetry	$7.660^{\ddagger}$	-	$10.866^{\ddagger}$
F_{yx} short- and long-run symmetry	$7.335^{\dagger}$	$7.692^{\ddagger}$	$6.905^{\dagger}$
Cointegration			
$F_{xy(1)}$	3.746	3.014	1.509
$F_{xy(2)}$	2.229	5.268	2.197
$F_{xy(3)}$	$7.635^{\ddagger}$	-	-
t_{bounds}	$-4.661^{\dagger}$	$-4.380^{\dagger}$	$-5.481^{\ddagger}$
Johansen test			
Trace	1	1	1
Auxiliary tests			
F_1	7.129	11.010	11.318
(p-value)	(0.000)	(0.000)	(0.000)
Error Correction	-0.159	-0.160	-0.172
t-statistic	-5.022	-4.905	-7.124
BDM critical value	-4.190	-3.980	-4.380

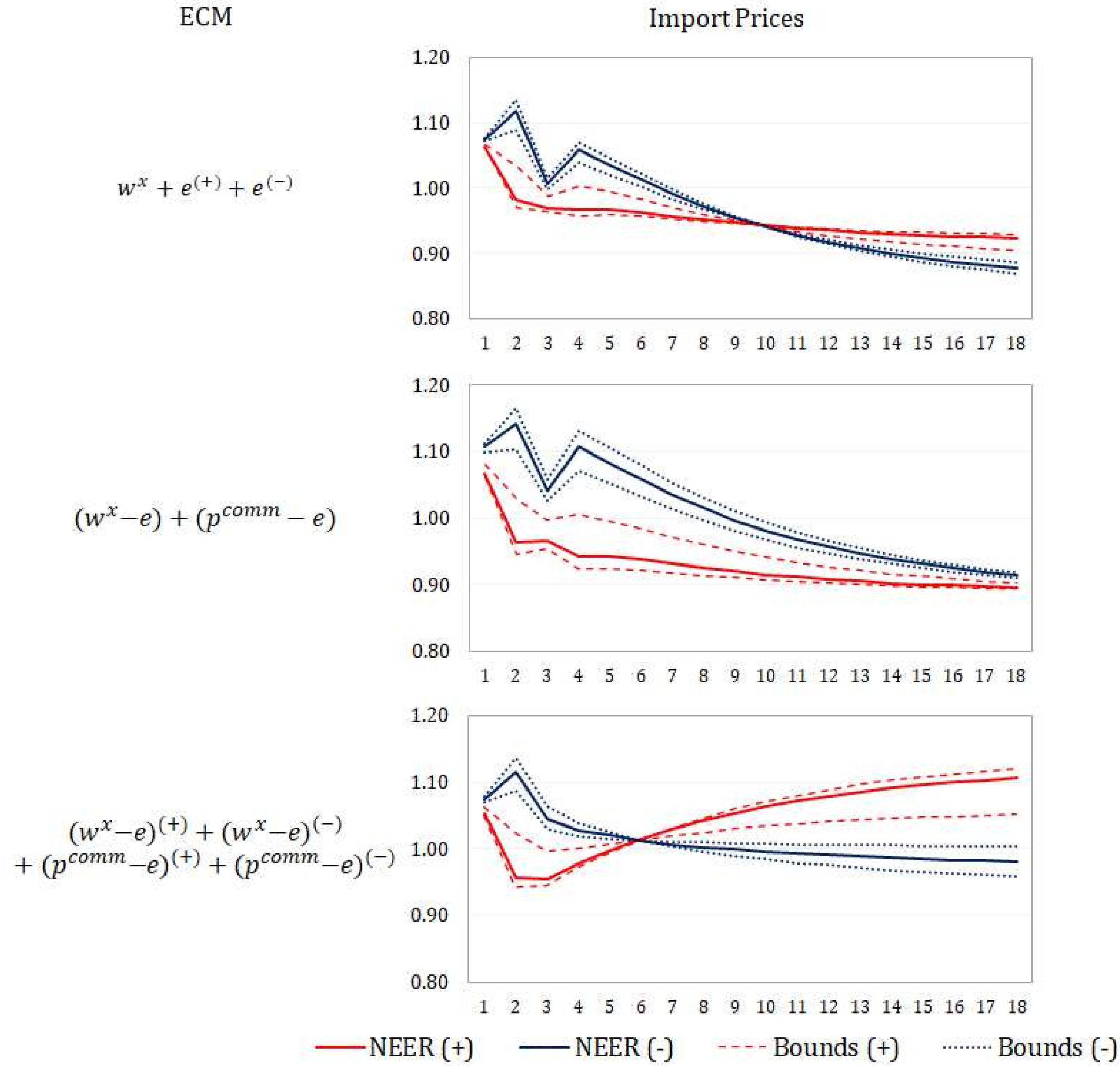
effect of an appreciation shock is larger. For the other two models, long-run ERPT tends to be symmetric in the long run, especially in the second, where symmetry is imposed in the ECM. Compared to the baseline models shown in the previous pages, NARDL models show similar short-run adjustments, but their convergence can differ from models with only first-differenced variables. Nonetheless, the degree of long-run ERPT does not change much across specifications. The lower ones come from VAR models with long adjustments (12 lags), with ERPT around 0.8 after 18 months. Lower-lag VAR models, single-equation with first-differenced data, and NARDL models all have ERPT between 0.9 and 1.1.

Table 6 – NARDL models (2) - ERPT coefficients

	Long-run relationship					Short-run relationship					
	Speed of adjustment	Long-run ERPT			Wald test	Impact ERPT		Wald test	Cumulative ERPT		Wald test
ECM	γ	β	α_1^+	α_1^-	$H_0: \alpha_1^+ = \alpha_1^-$	β^+	β^-	$H_0: \beta^+ = \beta^-$	$\sum \beta^+$	$\sum \beta^-$	$H_0: \sum \beta^+ = \sum \beta^-$
B.1	-0.159	-	-0.939	-0.826	0.813	-0.954	-0.969	0.008	-0.954	-1.201	**5.425**
B.2	-0.041	-0.886	-	-	-	-0.817	-0.865	0.252	-0.791	-1.143	**4.743**
B.3	-0.172	-	-0.918	-1.196	0.001	-0.690	-0.772	1.018	-0.690	-1.034	**15.707**

Dependent variable: import prices (p^m). The short-run structure has up to 3 lags in all models. Negative coefficients for depreciations (signal -) indicate increase in the dependent variable. Negative coefficients for appreciations (signal +) indicate decrease in the dependent variable. Diagnostics and stability tests are in appendix F.

Figure 7 – Response of domestic price indices to asymmetric shock in the nominal effective exchange rate
Equation 1.22 (NARDL model); Maximum lags = 6.



i) The bounds are the 5th and 95th percentiles of all the LxM response functions conditional on Ω^i; ii) NEER(+) and NEER(-) account for nominal effective exchange rate appreciation and depreciation, respectively; iii) IRFs computed based on algorithm 1.

1.6.4 Size asymmetry

Within the portion of the literature that studies asymmetric effects, only a few works assess size effects in the ERPT. As shown in section 2.1, these effects represent nonlinear responses often attributed to the menu costs hypothesis.

There is no clear way to define a size asymmetry hypothesis. The first is to disentangle large variations from small ones, regardless their direction. The second way interacts size with directional asymmetry, i.e., it first assesses if there are differences between large and small changes and then compares if large appreciations are more prominent than large depreciations (or vice-versa). This would indicate a sign asymmetry that is only triggered on higher quantiles of the distribution of Δe_t. Theoretically, one can assume it is a mix between one of the channels for sign asymmetry and a menu costs hypothesis.

The usual approach for assessing size asymmetry effects consists of building dummies for large exchange rate changes. Consider $I[.]$ the Heaviside function. Large appreciations and depreciations are, respectively:

$$\Delta e_t^{large(+)} = \Delta e_t \; I \; [\Delta e_t > \Delta e_{c_1}] \tag{1.31}$$

$$\Delta e_t^{large(-)} = \Delta e_t \; I \; [\Delta e_t < \Delta e_{c_2}], \tag{1.32}$$

where $c1$ and $c2$ are the thresholds based on symmetric quantiles such that $c_2 = 1 - c_1$. As such, Δe_{c_1} and Δe_{c_2} are the data values where the empirical cumulative distribution function crosses c_1 and c_2, respectively. When defining the small exchange rate changes, we choose two methods. The first defines as small all "interquantile" exchange rate variations, i.e., between c_1 and c_2:

$$\Delta e_t^{small} = \Delta e_t \; I \; [c2 \leq \Delta e_t \leq c_1] \tag{1.33}$$

The second decomposes 1.33 into small and large appreciations and depreciations:

$$\Delta e_t^{small(-)} = \Delta e_t \; I \; [c2 \leq \Delta e_t < 0] \tag{1.34}$$

$$\Delta e_t^{small(+)} = \Delta e_t \; I \; [0 < \Delta e_t \leq c_1] \tag{1.35}$$

As the empirical distribution of Δe_t shows negative skewness (i.e., depreciations are more frequent and in average higher than appreciations)[40], the values Δe_{c_1} and Δe_{c_2} differ. A final simpler approach consists of assuming the threshold value that triggers menu-costs adjustments to be the same regardless of their sign:

$$\Delta e_t^{large(+)} = \Delta e_t \; I \; [\Delta e_t > k] \tag{1.36}$$

$$\Delta e_t^{large(-)} = \Delta e_t \; I \; [\Delta e_t < -k], \tag{1.37}$$

The choice of k should sort large exchange rate values appropriately. As we do not have a consensual value that defines what is large or small, a plausible value captures the main episodes of abrupt exchange rate changes and not the casual ones. Aron *et al.* (2014), for example, define $k = 0.03$. We set k as 0.05 and 0.1 because 0.3 would yield very similar dummies to the quantile case where $c_1 = 0.9$ and $c_2 = 0.1$. This way, the large NEER changes capture some known episodes in the Brazilian economy: the 6-month devaluation spiral in the political campaign of 2002, the slump in emerging currencies in late 2008 after the sub-prime crisis, the 9% appreciation in the political campaign of 2018, the 2-month appreciation in 2005, caused by a combination of trade surplus and high interest rates, among other less notable episodes of high exchange rate changes.

Results[41] reported in table 3 show that there is no evidence of size asymmetry on the ERPT for import prices. The coefficients are generally larger for casual exchange rate changes, but this difference is not relevant at the 5% level. In the case of producer prices, results vary on the quantile choice. There is evidence of size asymmetry when choosing

[40] See appendix G.

[41] We analyse slope coefficients for two reasons. First, in the computation of the nonlinear impulse response function of asymmetric equations, the degree of asymmetry depends on the size of the shock, as demonstrated by Kilian and Vigfusson (2011a). Therefore, size asymmetry would emerge even in a system that does not impose it in structural terms. The second is that many large exchange rate changes are one-off events. When they occur, they do not seem to influence subsequent months. Even an offsetting movement in the opposite direction could happen, which is hardly captured by the dynamics of an AR(1) (for example, in January 2003 the NEER appreciated by 3.7% and in February 2003, it depreciated by 4.7%). The only periods with a high exchange rate change followed by changes of similar magnitude were in 2002 and 2008. Another way of seeing this is noticing that the AR(1) residuals are larger when the exchange rate change is also large. As such, our counterfactual paths for the NEER would not represent realistic movements throughout the h following months. One way of correcting this would be imposing the counterfactual paths of residuals (step M of the algorithm) conditional on episodes of high NEER change. Nonetheless, we still expose in appendix H the nonlinear IRFs by the size of the shock, confirming that asymmetries grows with the size of the shock, which does not imply economic meaning by itself though.

the pairs $c_1 = 0.7; c_2 = 0.3$ and $c_1 = 0.8; c_2 = 0.2$, although it vanishes when choosing $c_1 = 0.9; c_2 = 0.1$. This indicates that this type of asymmetry might be conditional on a certain band and that they are not assured to happen every time the exchange rate surpasses a given threshold. When assessing the sample values where the dummy was activated in the first two cases and not activated in the last, we conclude that the intervals $[-4.07\%; -2.67\%]$ for depreciations and $[2.42\%; 3.60\%]$ for appreciations seem relevant for size asymmetry effects in the pass-through effect for producer prices. For consumer prices, when choosing equal thresholds as 0.05, there is evidence of asymmetry. This result indicates that inflation at the consumer level is more responsive to both increases and decreases in the nominal effective exchange rate that surpasses the ballpark of 0.05. However, notice again that there might be a band for this interaction between large exchange rate changes and inflation. When setting $k = 0.1$, the asymmetry switches size, and the small currency changes are now more important - although this happens only in model 2.

Table 7 – Size Asymmetry

	Model 1			*Model 2*			
Import Prices	$\Delta e^{(+)}_{large}$	$\Delta e^{(-)}_{large}$	Δe_{small}	$\Delta e^{(+)}_{large}$	$\Delta e^{(-)}_{large}$	$\Delta e^{(+)}_{small}$	$\Delta e^{(-)}_{small}$
Quantile threshold							
$c_1 = 0.7$; $c_2 = 0.3$	-1.0473	-0.9610	-1.2901	-1.0611	-0.9388	-1.4521	-1.1425
$c_1 = 0.8$; $c_2 = 0.2$	-1.1167	-0.9059	-1.2396	-1.0676	-0.9236	-1.0283	-1.4515
$c_1 = 0.9$; $c_2 = 0.1$	-0.9690	-0.8922	-1.0300	-0.9557	-0.8984	-0.9799	-1.0724
Equal threshold							
0.05	-0.9312	-0.8874	-1.0547	-0.9310	-0.8875	-1.0536	-1.0558
0.1	-0.8694	-0.9271	-1.0266	-0.7162	-0.9226	-1.0940	-0.9295
Producer Prices							
Quantile threshold							
$c_1 = 0.7$; $c_2 = 0.3$	-0.4519†	-0.3825†	0.1906	-0.4149†	-0.4003†	0.4943	-0.1132
$c_1 = 0.8$; $c_2 = 0.2$	-0.4596†	-0.3957†	-0.0356	-0.4494†	-0.3988†	0.0073	-0.0815
$c_1 = 0.9$; $c_2 = 0.1$	-0.5158	-0.3964	-0.2786	-0.5114	-0.4032	-0.2064	-0.2107
Equal threshold							
0.05	-0.5381	-0.4107	-0.3563	-0.5672	-0.3904	-0.2573	-0.2696
0.1	0.0947	-0.1625	-0.4742†	0.2124	-0.1290	-0.3714†	-0.5190†
Consumer Prices							
Quantile threshold							
$c_1 = 0.7$; $c_2 = 0.3$	-0.0715	-0.1118	0.0157	-0.0553	-0.1203	0.1488	-0.1173
$c_1 = 0.8$; $c_2 = 0.2$	-0.0604	-0.1128	-0.0622	-0.0709	-0.1284	-0.0394	-0.1019
$c_1 = 0.9$; $c_2 = 0.1$	-0.0626	-0.1173	-0.0392	-0.0655	$-0.1125^{\dagger}$	-0.0577	-0.0104
Equal threshold							
0.05	-0.1523†	-0.1282†	-0.0141	-0.1652†	-0.1395	0.0290	-0.0439
0.1	-0.2989	-0.1559	-0.1025	0.0540	-0.0515	0.0243	-0.1680†

† indicates size asymmetry with a 5% significance level. If it is placed on a *large* superscript variable, it means that its coefficient is larger than the *small* variable and vice-versa. Those rows without it have no evidence of size asymmetry.
Model 1 treats all *small* exchange rate changes (i.e., in between c_1 and c_2) in one variable; whereas *Model 2* breaks them into appreciations ($\Delta e^{(+)}_{small}$) and depreciations ($\Delta e^{(-)}_{small}$).

1.6.5 Decomposition of the CPI

One further interest in the ERPT literature is whether the pass-through coefficients vary across different categories of goods and services. The general knowledge states that the nontradable category has lower pass-through than the tradable. This comes mainly from the fact that nontradables do not compete internationally, so there are no forces such as the *law of one price* and the PPP equalising them across countries. Nominal exchange rates react to accommodate shocks provoking changes in the real exchange rate (RER). If ERPT is similar for tradables and nontradables, the real exchange rate will not adjust to shocks, and wealth effects dominate substitution effects. This could be a costly adjustment to the country since adjustment will likely occur primarily through "expenditure reducing" rather than through "expenditure switching", which could lead to high unemployment costs. For example, after a NER depreciation, if the RER remains stable and if the pass-through is positive, the price index will be more expensive, and reallocation mechanisms within the consumption basket (*expenditure switching*) will be absent. On its place, an *expenditure reducing* mechanism rises (EDWARDS; CABEZAS, 2021). A similar relation exists between the import prices and the tradables at the domestic (PPI or CPI) level. Suppose the pass through of tradables is too high and equal to pass through of imports. In that case, the substitution channel from expensive imported goods to cheaper (and often lower quality) domestic goods blurs. Asymmetry plays an interesting role in this analysis. If its degree varies between these sectors, the substitution effects and expenditure reallocation tend to differ when the exchange rate appreciates or depreciates.

This section performs estimation of some price decomposition of the Brazilian CPI. There are two more usual ways of breaking apart the aggregate index. The first is the more typical one and relates to the "tradability" of the CPI components (tradables *versus* nontradables). Besides being a binary classification, the tradability is more likely to be continuous, as each good has a certain level of tradability. As such, the higher the cost of transportation and the shorter the shelf life, the less tradable a good is. Therefore, not only are services inside the nontradable classification but all kinds of prepared food are perishable to some extent. Economists have built different indices for nontradables and services, where the former embodies the latter.

The second disaggregation regards the mechanism of price fluctuation. In Brazil, it is common to analyse the behaviour of CPI considering the pattern of prices that are freely determined within the markets and those that are considered administered by regulatory agencies and political power in general. Those goods with prices allowed to fluctuate by supply and demand forces freely are classified as "free". In contrast, whichever goods with any alternative price structure are classified as "administered". These goods often face price controls of two types: long-run contracts and control by the public administration. The contracts are often indexed to past inflation. Thus they do not change rapidly to supply and demand forces. The public prices encompass public services[42] and taxes. However, even within the public administration, the mechanism

[42] In a nutshell, public services in Brazil comprehend: household utilities (water, sewage, electricity and natural gas - whether piped or in a cylinder), healthcare (not only the service itself but pharmaceutical

beyond the price adjustments can differ - across goods and throughout time. As we stress in the second essay of this thesis, fuel prices have had different phases of price adjustment - from strict control to daily fluctuation based on international markets. Regardless of the phase, little to no forces link those prices to domestic supply and demand forces. Therefore, observe that pass-through can be high even in goods considered to be administered.[43]

To make the analysis readily readable, we display in figures 8 and 9 the impulse response functions for the baseline models only, for 4, 8, and 12 lags. We confirm that the pass-through to tradables is indeed more significant. The first columns of these figures show an average price change of 0.26 after a depreciation shock and 0.09 after an appreciation shock. The following two columns show that the pass-through to nontradables - which embeds services and mainly perishable food - is larger than to services only. The direction of asymmetry varies according to the lag choice, which did not happen in the aggregate indices discussed previously. In most cases, the ERPT of depreciation to both nontradables and services have a negative sign in the first semester, indicating that the prices of these classes of goods lower after a depreciation shock. Regardless, asymmetry is hardly confirmed after considering the error bands. The average ERPT for nontradables and services after 12 months is 6% and 2%, respectively, and the responses are roughly symmetric.

goods), transportation (public transportation fares, four types of fuel, toll roads), and communication (telephone lines and mail services).

[43] Fuel prices have always been regarded as administered in our sample period, even in phases of international wholesale price fluctuation. See the estimates for phase 03 of price readjustment for gasoline in the third chapter (notice, however, that the stages of the price distribution are different though: the third chapter assesses the wholesale price charged at the refinery, whereas the administered price in the CPI decomposition is at the consumer level, set at the local stations).

Figure 8 – Response of the CPI decomposition by tradability to asymmetric shock in the nominal effective exchange rate
Equation 1.15, $x = [p^w, w^m]$; $p = 4, 8$ and 12.

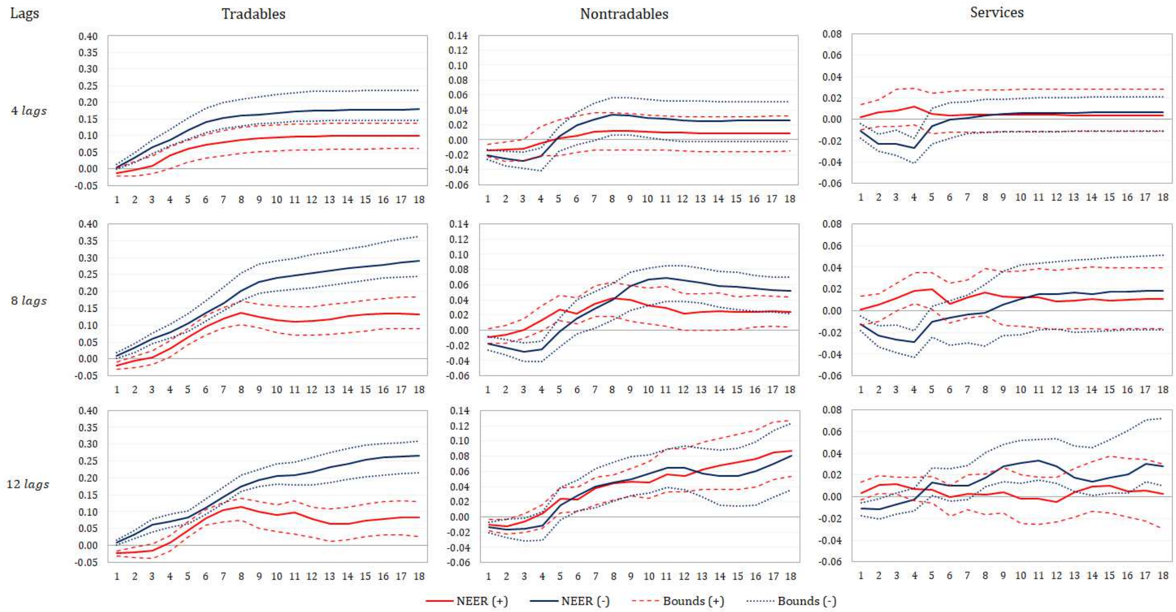
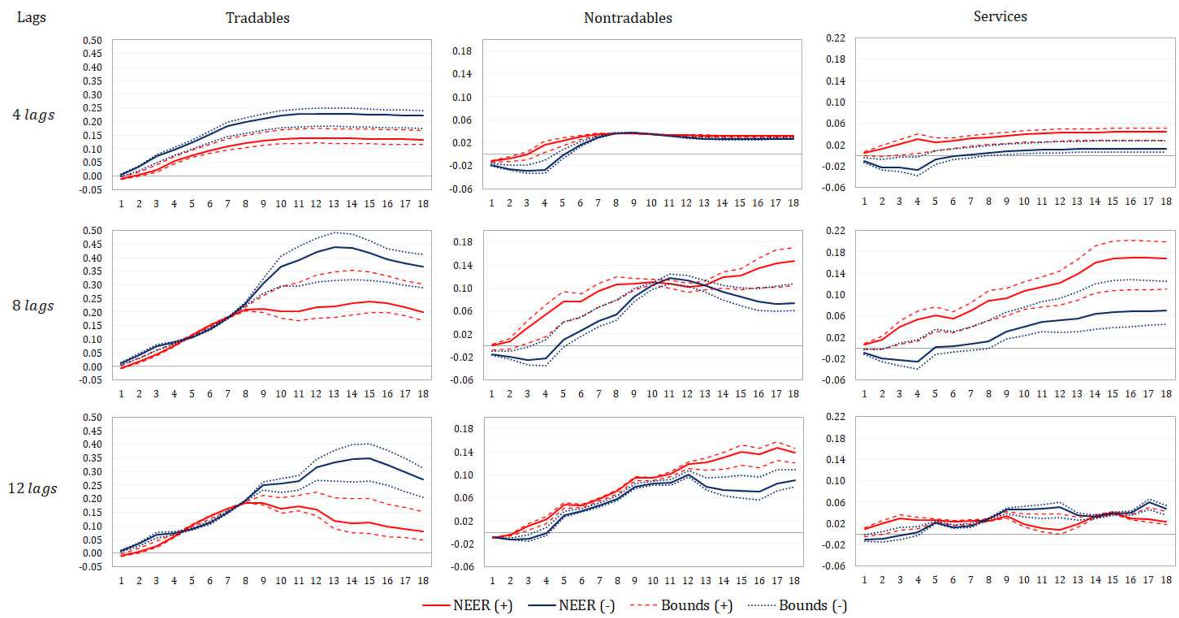


Figure 9 – Response of the CPI decomposition by tradability to asymmetric shock in the nominal effective exchange rate.
5-variable VAR model; $p = 4, 8$ and 12.



The second analysis, concerning the response of free determined prices versus those that are "administered", is in figures 10 and 11. The free prices embed both tradables and nontradables prices that fluctuate according to supply and demand mechanisms. Therefore, the result in terms of asymmetry is likely a weighted average of the responses of tradables and nontradables. As the weight of the latter is larger[44], we expect that a substantial part of the asymmetry seen in the tradables (first column of figures 8 and 9) to fade. The pass-through is symmetric within the first eight months in both single-equation and VAR models and around 9.9% for appreciations and 12.3% for depreciations, but with overlapping confidence intervals. The asymmetric pattern becomes visible after 12 months. However, the degree of pass-through is indeed minor than the tradables. Notice that the 8- and 12-lagged VAR models have higher pass through, indicating that the adjustment might have long lags.

After a year, the administered prices have higher pass-through after a devaluation, although lower after an appreciation, that is:

$$\begin{aligned} ERPT^{(-)}_{adm,h=12} > ERPT^{(-)}_{free,h=12}, \\ ERPT^{(+)}_{adm,h=12} < ERPT^{(+)}_{free,h=12}. \end{aligned} \tag{1.38}$$

Moreover, the effect of an appreciation shock is equal to zero in all but one model (the 4-lag VAR model in figure 11). This leads to a high degree of asymmetry for administered prices (model $C.5$ has, for example, $ERPT^{(-)}_{18} = 31\%$ and $ERPT^{(+)}_{18} \simeq 0\%$). Therefore, whichever the price structure chosen as optimal in these sectors is, they seem to import the depreciation to the final consumer inflation to a larger extent than the tradables do. Hence, an exchange rate shock in both directions could change relative prices between those that are free and those that are administered. As shown by Freitas and Bugarin (2007), if the policymaker wants to partially neutralise the relative price changes caused by the exchange rate shock, then its policy should allow for some inflationary pressure in the free sector. If monetary policy pursues stable relative prices, the administered sector's price response could play a role in adjusting all prices in the free sector. This finding increases the caution and watchfulness of monetary policy, as the degree of asymmetry - which causes domestic inflation to spark more after a depreciation - stems mainly from administered prices, whose price structures are designed mainly by policymakers in public agencies.

44 Nontradables represent 59% of the free prices, whereas tradables represent 41%.

Figure 10 – Response of the CPI decomposition by mechanism of fluctuation to asymmetric shock in the nominal effective exchange rate. Equation 1.15, $x = [p^w, w^m]$; $p = 4, 8$ and 12.

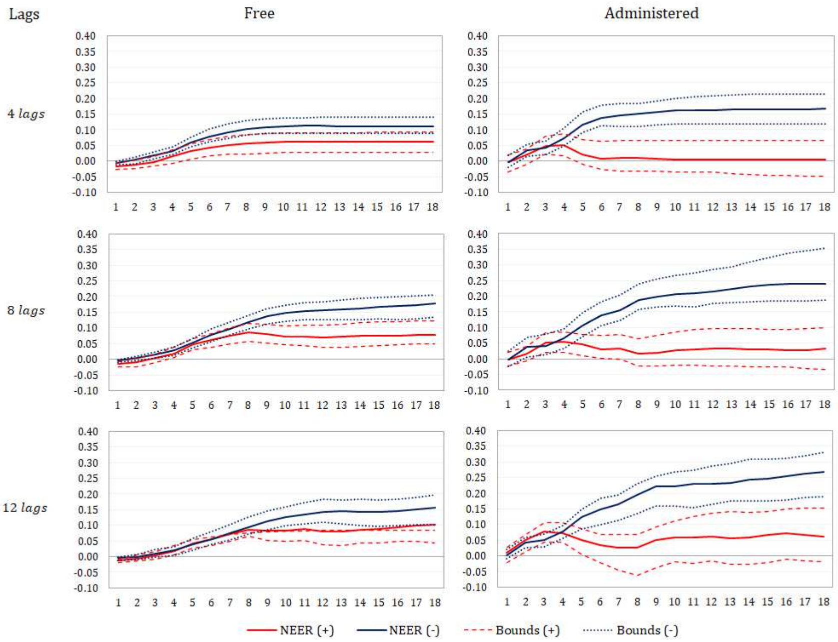
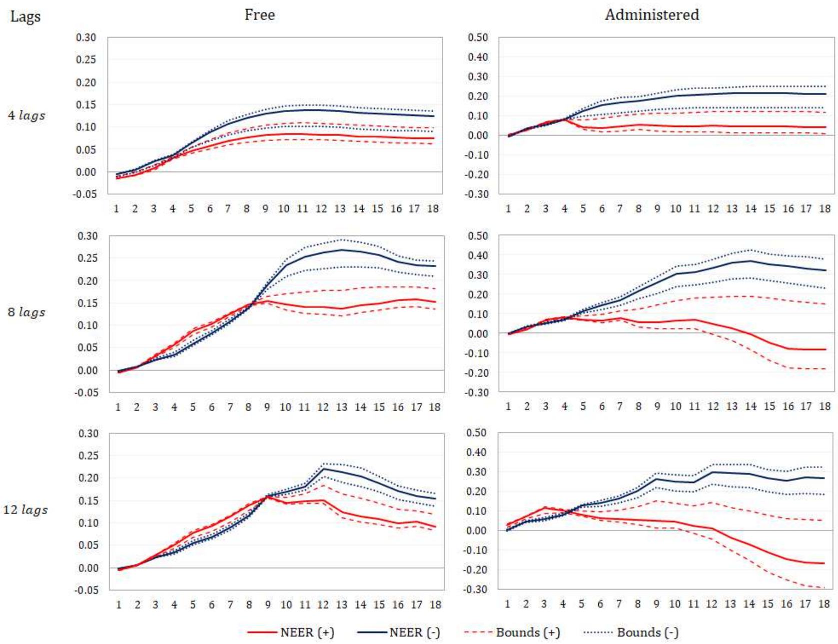


Figure 11 – Response of the CPI decomposition by mechanism of fluctuation to asymmetric shock in the nominal effective exchange rate. 5-variable VAR model; $p = 4, 8$ and 12.



1.7 Conclusions

We investigate asymmetric exchange-rate pass-through in Brazil between 1999 and 2018 of three price indexes at the imports, producer, and consumer levels. We argue that asymmetry is relevant because it yields positive inflation as the net outcome of depreciation and appreciation episodes of the same magnitude, everything else constant, which is of interest for monetary policy. We believe our work brings some novelties. First, in methodological terms, we bring to the ERPT phenomena the debate between James Hamilton and Lutz Kilian on how simple slope-based regression analysis is or is not outperformed by structural or semi-structural dynamic models. To do so, we resort to the ERPT literature to estimate standard regression models with variables in first differences and compare them to dynamic responses drawn from more “systemic” settings. In this exercise, we show that, despite similar in some cases, the degree of ERPT and its asymmetry can vary in modelling choices. Particularly, asymmetry is more evident in VAR models with longer lag structures. Even in more straightforward settings, where system models are not the researcher’s interest, similar responses can have different conclusions in terms of symmetry tests.

Secondly, it is the first to study and compare asymmetric effects throughout the price distribution chain. In this regard, we show that asymmetry is more robust to the modelling choice for consumer prices. This is solid evidence that the CPI inflation in Brazil has an asymmetric response to an exchange rate shock: it tends to go up more after a depreciation than it tends to shrink after an appreciation. For import prices and PPI, conclusions depend more on the model.

Thirdly, we decompose the CPI index to track at which level of disaggregation prices tend to suffer from higher ERPT. The “tradability” of the CPI decomposition indicates that the positive asymmetry in the aggregate CPI is caused mainly by the tradables, compared to nontradables and services. Moreover, the positive asymmetry in the aggregate CPI seems to stem mainly from the administered prices, which embeds long-run contracts and public goods and services prices. As such, we might conclude that the optimal price structure in these sectors seems to import the depreciation to the final consumer inflation to a larger extent than the tradables do. Therefore, there might be mechanisms that help offset asymmetric effects in the tradables sector that are not fully functional within the price structure of administered goods. For monetary policy purposes, these goods seem to be interesting sources for identifying possible channels of international shocks that are transmitted via the exchange rate. Whenever the economy faces a depreciation shock, administered prices grow larger than tradables, conditional on our sample period.

Lastly, we have also seen evidence of size asymmetry for producer and consumer price indices. However, the effect does not increase linearly for higher exchange rate changes. It seems that the size asymmetry is conditional on certain quantiles of the distribution of the exchange rate changes, i.e., it is conditional on *certain* episodes of high exchange rate variability, not *all* above a certain threshold. Computing conditional impulse responses could identify, within the sample period, which episodes sparked higher inflation and which did not. Studying policy variables and the context in which each episode occurred could help diagnose what influences size asymmetry in the pass-through. We let this exercise for future research;

There is a lot yet to be investigated. In methodological terms, if the dynamic asymmetry is indeed present, as in the VAR models, it calls for a better understanding of its sources at the theoretical level. As such, fully structural general equilibrium models could, for example, identify the microeconomic sources of asymmetry that are so far assumed in very simplistic theoretical models. Still, in the methodological field, the power of the asymmetry test and the methods for building confidence intervals also seem to be interesting improvements.

2 ASYMMETRIC PRICE TRANSMISSION OF INTERNATIONAL SHOCKS IN THE BRAZILIAN FUEL MARKET

2.1 Introduction

It is not a rare assertion that final consumers perceive price increases more often than decreases, and the Brazilian automotive fuel market has not been any different in recent years. In the process that made automotive gasoline step up from 2.75 BRL per litre in 2012 to 6.00 BRL in late 2021, consumers often alleged that the pass-through from upstream price increases was larger than from decreases.[45] The transmittal of a price shock from international prices to the final retail prices depends on the response of many intermediate players. In this study, we seek to understand price transmission at the very first stage of the price distribution, i.e., from international prices - petroleum, fuel, and currency - to the wholesale prices charged at the refineries in Brazil.

We focus our attention on the major Brazilian oil company Petrobras and its current price policy adopted in 2016, which states that the wholesale prices of gasoline and diesel would vary vis-à-vis the international fundamentals - mainly the exchange rate, the import prices of oil, and the import parity price of foreign fuel. Accordingly, the fuel prices are a direct function of these international prices, facilitating the specification of a fairly simple regression model. The company and policymakers aimed for such a policy to deliver more transparency and a competitive design for the domestic fuel market. However, it is still unclear how prices precisely vary to international fundamentals. Particularly, a company with such an outstanding marketing share may indulge in raising prices more easily and frequently than lowering them. The demand for fuel is acknowledged as inelastic, and a decrease in prices could be offset by a low-opportunity-cost decrease in the profit margin.

The relevance of price asymmetry in the fuel market is manifold. First, automotive fuel is widely used in different economic sectors as input and affects the transportation margins of arguably all manufactured goods within the economy. Noncompetitive increases in fuel prices affect relative prices and increase costs across industries.[46] Fuel price dynamics are relevant at both the micro and macroeconomic levels. For the former, asymmetrical changes in fuel prices might lead to asymmetrical changes in relative prices. Persistent imbalances can arise as consequences, resulting in energy-use inefficiencies compared to a competitive benchmark. From the macroeconomic perspective, there are

[45] Appendix K shows several media headlines with the increasing perception of price asymmetry.
[46] See Harun *et al.* (2018) for an applied example using input-output models.

concerns regarding the effects of oil shocks on inflation - once automotive and jet fuels are important transmission channels of such oil shocks. For example, Nordhaus (2007) noted that oil shocks boost inflation via fuel prices, which account for about half of the overall effect, and indirect effects on firms' costs, which increase airfares and shipping costs.[47] If fuel prices change asymmetrically, they are more likely to remain higher than the equilibrium level after a positive shock than to remain lower after a negative shock. This conjecture, combined with inelastic demand, a lack of substitutability between alternative fuel types, and transportation means might spark inflation.[48] Also, standard monetary policy may not have its full desired effects on fuel prices in a market responsive only to international exogenous variables, as in Brazil.

Second, such asymmetric behaviour could be an essential regulatory issue, as authorities are concerned if whether large oil companies exercise their market power to charge distributors higher prices than necessary. From the perspective of normative standard economic theory, asymmetric price transmissions lead to consumer welfare losses which should be avoided (FASOULA; SCHWEIKERT, 2018).

Lastly, the company's price policy is the root of debates in both academia and policy arenas. The company has mixed stock ownership. The state owns a bloc of its share, while private investors own the remaining shares. This means that the company has to fulfil goals that benefit both public and private shareholders. However, their objectives often vary. While private investors seek profitability based on some traditional concepts of the shareholder theory (FRIEDMAN, 1962), the public agents seek more diverse social goals, recalling the stakeholder theory (CHOI; WANG, 2009). Remarkable trade-offs characterise this dual aspect of the price policy. On one side, under a standard profit-maximisation goal, *ad-hoc* political interference in the prices is inefficient. On the other side, the growth prospects of oil production in the Pre-salt layer that launched Brazil as a preeminent player raised the attention of particular groups. They are mainly interested in the welfare gains the whole society could attain. One of their main arguments is that fuel prices should not be so sensitive to international volatility in an oil-extracting and fuel-producing country. Economists and policymakers from these groups often

[47] A more recent literature (HOOKER, 2002; GREGORIO *et al.*, 2007; BACHMEIER; CHA, 2011) shows that the pass-through of oil shocks to overall inflation has been decreasing for several possible reasons, although most of this literature is concentrated on developed countries, mainly the United States.

[48] There are, of course, other factors that determine whether such shocks cause economy-wide increases in prices. A crucial channel is when labour becomes more expensive in all sectors of the economy as workers adjust their inflation expectations in the wake of the shock (GREGORIO *et al.*, 2007). The perception of whether the shock was temporary or persistent is also important for the long-run effects on overall inflation.

propose price-adjustment mechanisms that could cushion the local economy from foreign shocks. Exploring these proposals is out of the scope of this paper, but evidence of price asymmetry under a so-called competitive policy adds new information to the debate.

We propose a simple adjustment function to test whether the pass-through from international to wholesale prices is asymmetric. To reach that aim, we use daily data from 2016 to 2019. The database contains automotive fuel prices at the (Petrobras) refineries level, the nominal exchange rate (NER), international crude oil prices, and gasoline and diesel prices. Our empirical model suggests whether the company still uses its dominant power to keep prices above the equilibrium level, even after claiming that the market is now competitive (i.e., domestic prices of refined fuel follow import prices). More precisely, it can be observed if the estimates for the increases and decreases of international prices are statistically different from each other - and the effect of an increase is larger.

With this in mind, our research is relevant in both academic and policy spheres because i) price asymmetry is a recurring topic of discussion as consumers often have the impression that positive changes in oil prices are easier passed through to retail fuel prices than negative ones; ii) the rule that drives readjustments and the company's pricing policy is unknown in terms of how frequent and to which extent the changes in international price are embedded in fuel prices; iii) the company has a history of political and regulatory interventions on its price policy, and it is still uncertain how appropriately the new pricing policy dealt with such an issue; iv) any sign of asymmetric pass-through might indicate the extent of market power held by Petrobras and finally v) to our knowledge, this topic has not been assessed at the wholesale stage in Brazil yet. Beyond these direct relevance, studying fuel price dynamics originating from non-renewable sources also has implications on the economic incentives to promote alternative energy sources (ATIL; LAHIANI; NGUYEN, 2014), especially if current market traits point to consumer welfare losses.

The results indicate that during at least one year, the company exerted positive price asymmetry in the gasoline market. This indicates that a counterfactual path for gasoline prices with symmetric competitive transmittal would render lower prices. Moreover, such asymmetry is evident in both short- and long-run structures, which indicates that the daily policy was asymmetric and that the long-run reaction to exogenous shocks was also asymmetric. One possible mechanism we conjecture from these results is that the company had more leeway to exert price asymmetry in the gasoline market as the diesel market had more interventions and agreements with public players. The takeaway is that the company performed a noncompetitive price practice that diverged from its goals at the time. This has implications in terms of transparent governance and adds relevant information to the optimal pricing policy debate.

Besides this introduction, in the next section, the essay presents a short overview of the main findings in the empirical literature; section three describes the Brazilian fuel market, its institutional background, and the main overturns in the pricing policy. Section four describes the methodological approach, section five depicts how we gathered and treated the data, and section six discusses the results. Finally, we conclude the study in section seven, highlighting the main contributions and addressing some limitations and future investigations.

2.2 Literature Overview

The literature on asymmetric price transmission (APT) is broad within the areas of agricultural and energy economics. The survey by Meyer and Cramon-Taubadel (2004) is a natural guideline. Firstly, one should note that there are different sorts of symmetry on price transmission. It could be whether vertical or horizontal/spatial: the former regards different degrees of transmission throughout the marketing chain – from wholesalers to retailers, while the latter comprehends different transmissions among different firms/regions, but in the same level of distribution. It could also have different magnitude and speed. The asymmetry of magnitude is defined as the difference in intensity of response to an increase or a decrease in the upstream prices and the asymmetry of speed is the difference between response times of new readjustments.

To this day, surprisingly there were not much research on the economics of fuel price transmission in the Brazilian market. Serigati (2014) argued that the price policy at the time (i.e, before 2014) brought an overwhelming competitive pressure to bear on the ethanol sector. As the ethanol is a substitute for gasoline, the sector's competitiveness and profitability were being hindered by the controlled price policy for gasoline and diesel. After simulations of hypothetical paths for the domestic fuel prices, he concluded that an ideal price policy to foster the ethanol sector would be the one that associates the readjustments to the international price of oil and exchange rate, which would be the actual policy only by 2016, as we mentioned.

Silva's (2003) major concern was to analyse how the price strategy was being carried out since the market liberalisation in 2002. Among her conclusions was the fact that the prices of gasoline, diesel, and LPG (liquid petroleum gas) were not in line with international prices, especially in the periods of elections (2002) and the Iraq war (2003), when was natural to expect higher international volatility. Therefore, the discretionary price policy was said to be a non-neutral mechanism used to cushion international shocks, as it reinforced the company's massive market power. A further conclusion regards new possibilities for price policy at the time, including the French mechanism based on a "trigger" for fuel taxes whenever the international prices rise above a certain threshold. The point argued was that a shock-absorber mechanism is necessary, although it could be properly formalised to keep a certain level of competitiveness inwards and avoid political interests.

Rodrigues, Losekann and Filho (2018) is an important correlated work. They also studied the asymmetric behaviour in Brazilian fuel market, but instead of price transmission, the authors focused on price response, which is how fuel demand respond to price variations. They sought to estimate demand functions for automotive fuel (gasoline, ethanol and compressed natural gas - CNG)[49] to show that the inclusion of asymmetric price transmission (APR) enhances the ability to predict how fuel demand will respond to a certain policy that affects fuel prices.

When it comes specifically to the asymmetry of price transmissions, Uchôa (2008), estimating Threshold Autoregressive (TAR) models, confirmed the hypothesis of positive asymmetry in the transmission of international shocks at the retail level: when gasoline prices are above its long-run equilibrium path, it tends to remain there longer than when it is below. Silva *et al.* (2014) sought to analyse the pass-through from distributors to the retail sector, which is the downstream part of the distribution chain. With disaggregated data for municipalities, they showed that the asymmetry does not happen nationally, but only in a portion of the cities (30% of the sample), as the conclusions were in terms of symmetry in the transmittal.

In the international literature, several works studied the asymmetric responses of fuel prices, as we will only discuss few of them and only acknowledge others. Borenstein, Cameron and Gilbert (1997) tested and confirmed that retail gasoline prices respond more quickly to increases than to decreases in crude oil prices in three points of the distribution chain. The evidence was that the adjustment of spot gasoline markets to changes in crude oil prices appears to be responsible for a fraction of the asymmetry. The asymmetry in the adjustment of retail gasoline to terminal price changes also contributes with a significant fraction. Among the possible sources of asymmetry are production/inventory adjustment

[49] Diesel is not a typical small automotive fuel in Brazil, as the fleet is composed majority by flex vehicles that run with either gasoline and ethanol. Diesel is instead more commonly used in trucks.

lags and market power of some sellers. Bachmeier and Griffin (2003) challenged the results in Borenstein, Cameron and Gilbert (1997) by estimating an error-correction model with daily spot gasoline and crude-oil price data over the period 1985-1998. Adopting a more standard estimation approach and using daily instead of weekly data was sufficient to eliminate most of the evidence of asymmetry.

In terms of empirical strategy, Atil, Lahiani and Nguyen (2014) hold some similarities to what we propose. The authors gathered data on crude oil, gasoline, and natural gas from 1997 to 2012 to show asymmetric price reaction in the short run for the gasoline and in the long run for the natural gas. Moreover, negative oil shocks tended to have greater effects than positive ones. They attributed this result to downward price expectation spirals that affect fuel prices during periods of collapsing economic conditions.

Kpodar and Abdallah (2017) gathered retail prices from 162 countries from 2000 to 2014 to conclude that declines in crude oil prices lead to smaller effects on retail gasoline prices than increases in crude oil prices, pointing to a positive asymmetry in the fuel price pass-through. Some of the recent works are also Asane-Otoo and Schneider (2015) for Germany; Fasoula and Schweikert (2018) for Austria; Balaguer and Ripollés (2012) for Spain; Bettendorf, Geest and Varkevisser (2003) for the Netherlands; Wlazlowski (2001) and Bermingham and O'Brien (2011) for the UK; Meyler (2009) and Salles (2014) for selected European countries and Radchenko and Tsurumi (2006), Deltas (2008) and Honarvar (2009) for the US.

Diversely from the recent domestic literature, we focus on the production stage, i.e., the wholesale prices. This gives a clearer time series data - that depends mainly on input costs and the market design - free from distributors margins, transportation costs, and taxes. The usage of the recently available Petrobras' daily data is also a novelty, as we will be able to check the behavior of a specific prominent company, instead of the average behavior of several companies, captured by the usual aggregated data from market fuel provided by ANP. Finally, the empirical strategy is a recent improvement on cointegration literature and provides an intuitive tool for assessing asymmetry. In this matter, we followed Atil, Lahiani and Nguyen (2014) to investigate the transmittal of international prices to fuel prices in a NARDL-based framework. We discuss the methods in section 4.

2.3 Characteristics of the Brazilian fuel market

The Brazilian fuel market has two preeminent players. The ANP (National Agency of Petroleum, Natural Gas and Biofuels) is the federal government agency linked to the Ministry of Mines and Energy responsible for the regulation of the oil sector. Its operations encompass several functions: oil and natural gas exploration and production; refining, processing, transportation and storage of these elements and their derivatives; distribution and trade, monitoring and inspecting the market, among many others. The Petrobras (Petróleo Brasileiro S.A.) is a mixed company operating on the oil, natural gas, and energy industries. It was the 8th largest energy company in the world in 2016, with a production of 2.55 million barrels of oil (BOE) per day and an enterprise value of $132 billion.[50] In general, the government can influence the fuel market by the interplay of both institutions. The National Agency via economic regulation and the Petrobras by holding common stocks and choosing its president.

In this work, we focus on the production stage, i.e., the price of wholesale automotive fuel charged at the refineries. To comprehend the formation of wholesale prices in Brazil, we need to characterise the institutional background, in which the market structures, technical features of the refining process and the pricing policy help explain how prices are formed.

The first characterisation regards the market structure, which is marked by two main features: in practical terms, it approximates a monopoly design and it is hugely influenced by the state. Created in 1953, the Petrobras used to have a monopoly over the Brazilian oil market. In 1997, however, a congress law[51] permitted other companies to operate in the same activities performed by Petrobras. Thereafter, several companies began the extraction and production of petroleum, such as Shell, Chevron, Statoil, Repsol, among others. Nevertheless, in the refining stage, the domestic company remained with a considerable share, holding 81% of the oil refineries[52]. Moreover, the three refineries not owned by Petrobras have low capacity and produce less than 4% of the total gasoline and diesel produced domestically.[53] Albeit the company had lost a part of its share in the domestic market of fuel in recent years, it still has considerable market power, supplying

50 See <https://bit.ly/2LRdzl5>

51 See law number 9,478, from 6th August 1997.

52 For this statistics, we counted only the refineries that produce automotive fuel, as a few of them produce only some refined products (for example, one of the units refines shale oil and produces mainly naphtha). Thus, between 2016 and 2019, Petrobras held 13 out of 16 fuel-producing refineries (81%).

53 Source: ANP. See also the appendix L.

87% and 79% of the domestic market of gasoline and diesel, respectively. These facts show how Petrobras still has quite a considerable share in the energy market, despite there being more than 20 years of openness to foreign competition.

Secondly, a technical feature of the refining process that helps understand how prices are set and why they depend on multiple international mechanisms. Despite being a major crude oil producer, Brazil still needs to supply its refineries with imported oil. The main reason for this is that the oil extracted in the Brazilian basins is too heavy, which makes the refining inefficient, given the technical specifications of the country's refineries. As such, lighter imported oil is required to blend domestic oil as to achieve an efficient chemical composition that suits the refineries. More recently though, Petrobras alleged that the Pre-salt oil is lighter and is highly used on its refineries, which supposedly lowers the dependence for imported oil. Regardless, the domestic market depends on imported oil as an input to the refining process.

The final aspect is the pricing policy. Before 2016, the government used to intervene in the price strategy, providing subsidies in order to cushion the local economy against shocks in the international markets and also to hamper inflation, as the derivatives are inputs entrenched in the whole structure of the economy. This practice matched a stylised fact addressed in Sterner (1989), which points that oil-producing countries with state monopolies tend to have lower domestic fuel prices. In 2016, the fuel prices strategy became more flexible as the wholesale prices would be the result of international fundamentals, such as fuel, oil, and the exchange rate. This was the so-called import parity price (IPP) policy. Accordingly, the parity price - the international prices of fuel traded in global markets, plus transportation costs – is the price importers had to pay if they were to supply the domestic market with foreign fuel. Hence, the policy allowed international prices to compete with the ones charged by Petrobras, which in theory would make foreign competition feasible. The following description was taken from Petrobras official website:

> *"Our price policy for gasoline and diesel sold to distributors is based on the import parity price, formed by these products' international prices plus the costs that importers would have, such as transportation and port fees, for example. Parity is necessary because the Brazilian fuel market is open to free competition, and distributors may choose to import the products. In addition, the average price includes a margin that covers risks (such as exchange rate and price volatility)." [. . .]*

> *"Using international market prices as a benchmark, we analyse our share of the domestic market and periodically decide whether the prices practised at the refineries will be maintained, reduced or increased."*

The main takeaway from the institutional background is that wholesale prices depend on the price of oil, which is the main input, but it also depends on the international parity price, which is a policy goal without much connection to input costs - it is rather an attempt to change the market design and foster competitiveness. In short, figure 12 tries to draw a simplified scheme of price transmission for fuel. The scheme draws these two main mechanisms for the price composition - one that entails the costs of inputs and other that depicts the international parity pursuit by the domestic players. We regard three main branches of price shocks within the international environment.

As explained earlier in this section, the refined oil is a blend of domestic and imported sources. This leads us to the first transmission channel, given by the the international price of oil[54], which is the main input for this market. Everything remaining constant, an increase in the international price of oil sparks the import prices for the next contracts. For this mechanism to hold in a partial equilibrium setting at which we study only prices and not quantities, there cannot be much substitutability between foreign and domestic oil, i.e., the firm does not change its blend composition when relative prices change. Moreover, there are more entrenched structural effects of increases in the price of crude oil. In a general equilibrium context, the transmission of a shock to the price of crude oil affects domestic's price chain, as the production of other refined products and several other manufactures are linked to the price of oil. A price increase of crude oil can also spark the demand for alternative and perhaps cleaner energy sources. We omit these more complex economy-wide pass-through effects.

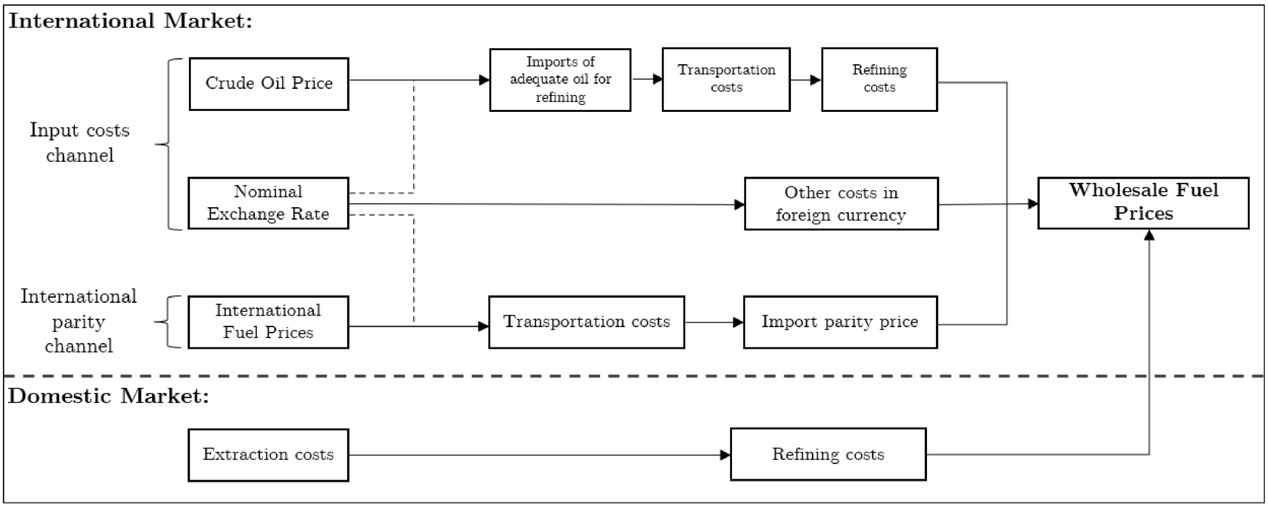
Next, there is the currency channel. Aside from affecting the other branches (dashed line), the nominal exchange rate represents itself a broad chain of costs, as diverse extraction and refining costs can be denominated in foreign currencies. Notice, however, that this mechanism can be, first, sluggish, as the imports of related equipment and machinery occur from time to time and immediate swings in the exchange rate may not affect the cost structure in the short run, and second, partially offset by the interplay in future markets via hedging. Everything else constant, a depreciation of the BRL

[54] The price of oil appropriate for refining is given by both WTI - originated from U.S. oil fields, primarily in Texas, Louisiana, and North Dakota - and the Brent Crude - from oil fields in the North Sea.

increases the prices of oil and fuel in domestic currency (dashed line) and increase the costs of imported inputs via pass-through effects.

Finally, international fuel prices are the main benchmark for the IPP policy. It allows minor players to import fuel from other countries and, after adding transportation costs, sell them at competitive prices, which was not possible when monopolistic practices kept prices below the equilibrium level.[55] From what the company claimed, the wholesale prices would, to some extent, keep track of the international fuel prices. However, the two first channels might interpose themselves in the process of daily readjustments. Next, we try to shed some light on the asymmetry behind these mechanisms. This would indicate possible noncompetitive practises when ANP and Petrobras pursue a more competitive market.

Figure 12 – Mechanisms of price transmission for wholesale fuel in Brazil



When the IPP policy began by October 2016, the readjustments on prices were characterised - with exceptions - by a new price per month, generally defined in the middle of the month. This new price should embed the variations on international prices and other refinery-level costs since the last readjustment. This phase lasted for almost 6 months and had 9 price adjustments for the gasoline and 10 for diesel. In what follows, we will call it phase 1.

In July 2017, the company adjusted its strategy in order to have daily adjustments (phase 2). The official explanation was that the previous frequency was incapable of

55 This scheme has also a locational justification - in a big country like Brazil, refineries are not distributed evenly across regions and, for some enterprises, it could be more efficient to collect imported fuel from the ports rather domestic fuel from the refineries. Another justification would be...

tracking international volatility. Thus, prices now could be changed whenever it's necessary, but the adjustments are limited to a band of +- 7%. On the company's official website, one can find a description of the price policy, in line with what we just mentioned. Firstly, in the section "10 answers to your questions about gasoline price":

> *"The fuels derived from the petroleum are commodities and has its prices attached to the international markets, with quotations varying daily. [...] In an environment of open economy and price freedom, we face the competition of fuel importers, whose prices are also attached to the international market. Thus, the price variations at the refinery are important for us to effectively compete in the domestic market."*

This new scheme lingered for almost a year and, in practice, it was not totally daily, as in 10% and 7% of the working days in this period saw the gasoline and diesel prices respectively remaining constant. Nevertheless, the standard deviations increased in a threefold proportion compared to phase 1, as table 1 shows.

Table 8 – Descriptive statistics by phases of the price policy

(a) - Gasoline

		Price level		**Price adjustments**					
	Working days	Average	SD.	N	Max.	Min.	Increases	Decreases	Zero
Phase 1	178	1.493	0.060	9	8.10%	-5.90%	3	6	169
Phase 2	249	1.597	0.201	223	5.10%	-3.93%	121	102	26
Phase 3	233	1.848	0.243	85	5.61%	-7.16%	47	38	148
Total	*660*	*1.657*	*0.242*	*317*	*8.10%*	*-7.16%*	*171*	*146*	*343*

(b) - Diesel

		Price level		**Price adjustments**					
	Working days	Average	SD.	N	Max.	Min.	Increases	Decreases	Zero
Phase 1	178	1.670	0.074	10	9.50%	-10.40%	4	6	168
Phase 2	227	1.802	0.182	212	4.40%	-10.00%	119	93	15
Phase 3*	147	2.089	0.173	8	13.03%	-15.28%	4	4	139
Phase 4	108	2.128	0.129	16	4.84%	-6.00%	12	4	92
Total	*660*	*1.884*	*0.237*	*246*	*13.03%*	*-15.28%*	*139*	*107*	*414*

Source: author's own calculations with data from Petrobras and ANP.

In May 2018, the Brazilian economy was hit by a massive strike carried out by truck drivers in the whole country. Among their claims were the increasing costs with diesel when compared to the prices charged for the freight service, which was held constant.[56] Indeed, the diesel price was consecutively readjusted upward as a result of the increasing costs of crude oil, the devalued exchange rate and the several taxes charged throughout its distribution chain. This situation was squeezing their margins up to the point that some drivers had no positive profit after deducting all their costs. Moreover, adjusting to a new price every day was a costly process. After negotiations between government agencies and drivers union, the price policy changed again. This time though, gasoline and diesel would follow different patterns.

From late May on, gasoline prices remained in phase 2 (daily readjustments), but diesel price was constrained by a legal deal established with drivers, which we call phase 3*. The agreement, called "subsidy program in diesel commercialisation" initially guaranteed that the retail price would fall 0.46 BRL per liter and would remain at this level for sixty days and it would be readjusted once a month therefrom. The wholesale prices fell in a different amount though: during the ten-day period of negotiations, the wholesale prices successively fell 0.036 BRL (-1.54%), 0.23 BRL (-10.0%), and 0.07 BRL (-3.33%). The subsidy kept valid the remainder of 2018. For gasoline, phase 2 lingered a month further than diesel, until early July. Despite not facing an opposing and organised opposition as the truck drivers exerted on diesel prices, the gasoline prices seemed to be swayed by the context. Following the inverse of what was done between phases 01 and 02, this time the company slowed the pace of readjustments, although this change was not officially released in the media. Gasoline price's phase 3 had an increase in the frequency of short periods without new prices, with new adjustments, on average, every 2.75 working days, which is approximately two prices per week.

Finally, with the subsidy program coming to an end by December 31st, 2018, the diesel price entered into phase 4, which is analogous to gasoline's phase 3. The rule for this phase was marked by the company's claim that it would hold the diesel price fixed for up to 7 days, regardless of how volatile the markets are. Indeed, the data show consistency in such a claim, as a new price was chosen every 6.75 days on average throughout this period.

Figures 13 and 14 below shows the patterns of fuel prices as described above. What differs them was the fact that the diesel was the main ingredient of a social deadlock caused by a strike, which made the policies different from that moment on. Thus, phase 3* was the main divergence of fuel prices from the international prices on the period October/2016 - July/2019. Apart from this phase, automotive fuel prices in

[56] The freight prices were not set in terms of the operational costs. As a main outcome after the negotiations, the congress approved minimum freight prices.

Brazil followed the fundamentals provided by the international markets, with different frequencies in the readjustments, as we shall see in section 5.

Figure 13 – Wholesale gasoline prices and phases of the price policy

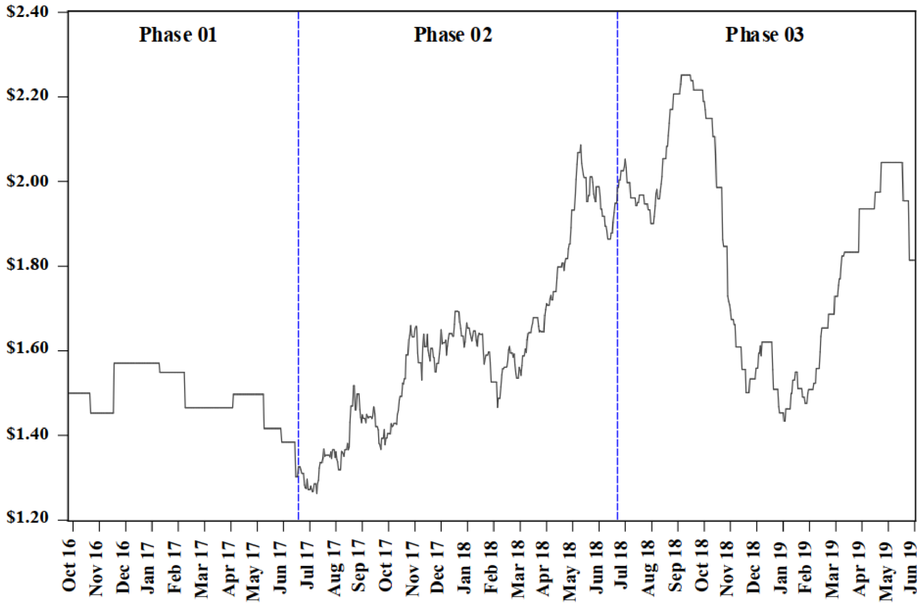
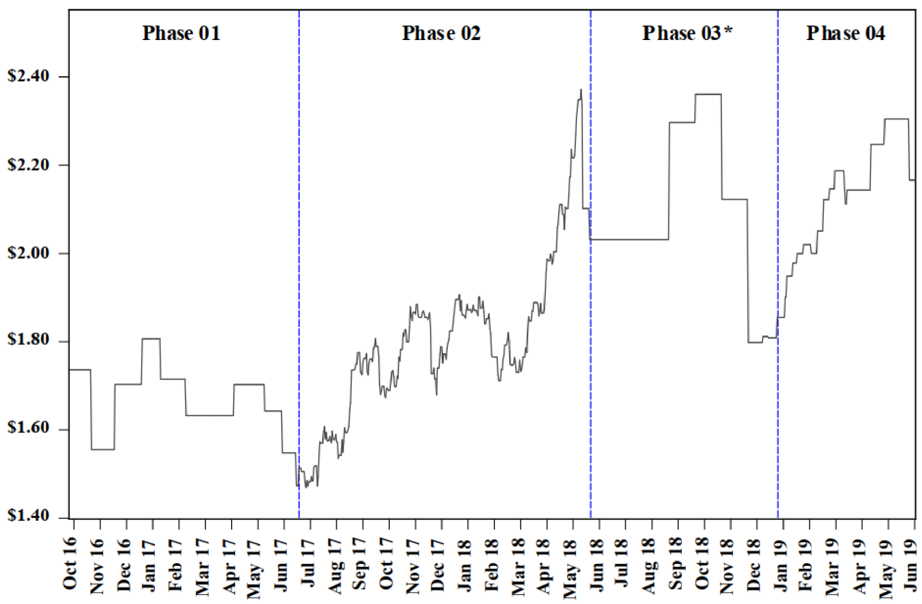


Figure 14 – Wholesale diesel prices and phases of the price policy



2.4 Methodology

The literature on the topic uses a wide range of empirical tools to address symmetry responses on fuel markets. Frey and Manera (2007) thoroughly identified them, stressing that each econometric model is specialised to capture a subset of asymmetries and each asymmetry is properly assessed with a subset of econometric models. They can be divided into five broad classes: autoregressive distributed lag (ARDL) model, partial adjustment models (PAM), error correction model (ECM), the whole class of regime-switching models, vector autoregressive (VAR) and error correction models (ECM). A recent improvement on cointegration techniques that were not available at the time of Frey's and Manera's survey is the NARDL model by Shin, Yu and Greenwood-Nimmo (2014), which i) allows cointegration tests without the typical concerns on unit roots testing, ii) allows the investigation of at least three types of asymmetries and iii) has desirable small-sample properties and its lag structure helps to deal with residual serial correlation and endogenous regressors (NARAYAN, 2005). Firstly, we show the derivation of an error correction model applicable to the transmission of international shocks to fuel prices.

2.4.1 Cumulative adjustment functions

To test for the possibility that fuel prices varies diversely towards increases and decreases on its international determinants, we work with a function that captures the rate at which the wholesale fuel prices adjust to international parity price, crude oil, and exchange rate changes, similar to Borenstein, Cameron and Gilbert (1997). To do this, let us first assume a simple linear long-run relationship:

$$P \;=\; \phi_0 \;+\; \phi_1 P^* \;+\; \phi_2 C \;+\; \epsilon \tag{2.1}$$

where P^* is the fuel price at the refinery, in BRL per liters, P^* is the international price that drives the readjustments according to the IPP policy and expressed in domestic currency. C is a set of input costs that influence the price adjustments and will encompass both the price of crude oil and the nominal exchange rate, and ϵ is a normal and i.i.d. error term. An important feature of this function is that it assumes a time-invariant adjustment process during the sample period (i.e., the adjustment coefficients are constant regardless the absolute magnitude of the change on oil prices and exchange rates, the period of the month and year and so on). Defining $\Delta P_t = P_t - P_{t-1}; \Delta P_t^* = P_t^* - P_{t-1}^*$

and $\Delta C_t = C_t - C_{t-1}$ and considering ΔP_t^{τ} as the price adjustment in t due to changes in P^* in τ yields:

$$\begin{aligned} \Delta P_t^{\tau} &= \alpha_0 \Delta P_t^* + \beta_0 \Delta C_t \\ \Delta P_{t+1}^{\tau} &= \alpha_1 \Delta P_t^* + \beta_1 \Delta C_t \\ &\vdots \\ \Delta P_{t+n}^{\tau} &= \alpha_n \Delta P_t^* + \beta_n \Delta C_t \end{aligned} \tag{2.2}$$

The adjustment dynamics depicted above shows that it takes n periods to the shocks on independent variables in t to be fully passed-through to the fuel prices. The total change in refinery fuel price in any period t will thus depend on the changes on prices in the previous n periods:

$$\Delta P_t \ = \ \Delta P_t^{\tau} \ + \ \Delta P_t^{\tau-1} \ + \ ... \ + \Delta P_t^{\tau-n} \ = \ \sum_{i=0}^{n} \alpha_i \Delta P_{t-i}^* \ + \ \beta_i \Delta C_{t-i}, \tag{2.3}$$

where, for simplification, we assume that both international prices and the cost variable have the same n lags of adjustment. Equation 2.3, however, implies symmetric adjustment. In order to capture different effects from increases and decreases on international prices, we can instead assume:

$$\begin{aligned} \Delta P_t^{\tau} &= \alpha_0^+ \Delta P_t^* + \beta_0 \Delta C_t \\ \Delta P_{t+1}^{\tau} &= \alpha_1^+ \Delta P_t^* + \beta_1 \Delta C_t \\ &\vdots \\ \Delta P_{t+n}^{\tau} &= \alpha_n^+ \Delta P_t^* + \beta_n \Delta C_t, \end{aligned} \tag{2.4}$$

if $\Delta P_t^* > 0$, and:

$$\begin{aligned} \Delta P_t^{\tau} &= \alpha_0^- \Delta P_t^* + \beta_0 \Delta C_t \\ \Delta P_{t+1}^{\tau} &= \alpha_1^- \Delta P_t^* + \beta_1 \Delta C_t \\ &\vdots \\ \Delta P_{t+n}^{\tau} &= \alpha_n^- \Delta P_t^* + \beta_n \Delta C_t, \end{aligned} \tag{2.5}$$

if $\Delta P_t^* < 0$. Defining:

$$\begin{aligned} \Delta P_t^{*+} &= max \left\{ \Delta P_t^*, 0 \right\}; \\ \Delta P_t^{*-} &= min \left\{ \Delta P_t^*, 0 \right\}, \end{aligned} \tag{2.6}$$

and assuming different lags of adjustment for international prices and costs, a straightforward empirical model for assessing asymmetric adjustments in the gasoline price would then be:

$$\Delta P_t \;=\; \sum_{i=0}^{n}(\alpha_i^{+}\Delta P_{t-i}^{*+} \;+\; \alpha_i^{-}\Delta P_{t-i}^{*-}) \;+\; \sum_{s=0}^{m}(\beta_s \Delta C_{t-s}) \tag{2.7}$$

The key feature of an asymmetric lag response structure is its intertemporal independence. Allowing for different effects of increases and decreases implies that an increase in an explanatory variable followed by a decrease of the same amount does not imply a reversal on gasoline prices, as it could continue to rise in $t+1$ (this happens when $\alpha_0^{+} > \alpha_1^{-}$).

The following step is to change the lag adjustment model in 2.7 into a partial adjustment model. Doing so implies that there is a long-run relationship towards which the changes on fuel prices tend to revert at a low speed. The most common partial adjustment structure has an error correction term, which is the lagged residual from the long-run relationship described in 2.1. The error correction model is:

$$\begin{aligned}\Delta P_t \;=\; & \sum_{i=0}^{n}(\alpha_i^{+}\Delta P_{t-i}^{*+} \;+\; \alpha_i^{-}\Delta P_{t-i}^{*-}) \;+\; \sum_{s=0}^{m}(\beta_s \Delta C_{t-s}) \;+ \\ & +\theta_1(P_{t-1} \;-\; \phi_0 \;-\; \phi_1 P_{t-1}^{*} \;-\; \phi_2 C_{t-1}) \;+\; \epsilon_t\end{aligned} \tag{2.8}$$

For the asymmetric adjustment to a long-run equilibrium to be valid, the variables should be cointegrated (FASOULA; SCHWEIKERT, 2018). We address it in the next section.

2.4.2 NARDL model

A simple way to reach an empirical identification from the adjustment function in equation 2.8 is via the cointegration approach of a NARDL model. Assuming a symmetric relationship as implied in 2.1, one can right the following conditional error correction model (CECM):

$$\begin{aligned}&\Delta p_t = \alpha_0 \;+\; \alpha_1 t \;+\; \sum_{i=1}^{p}\gamma_{1i}\Delta p_{t-i} \;+\; \sum_{i=0}^{q_2}\gamma_{2i}\Delta p_{t-i}^{*} \;+\; \sum_{i=0}^{q_3}\gamma_{3i}\Delta c_{t-i} \;-\; \theta EC_{t-1} \;+\; u_t, \\ &EC_{t-1} \;=\; \frac{\beta_1}{\theta}p_{t-1} - \left(\frac{\beta_2}{\theta}p_{t-1}^{*} \;+\; \frac{\beta_3}{\theta}c_{t-1}\right).\end{aligned} \tag{2.9}$$

where lower case variables are now the natural logarithms of the same variables of the last section and EC is an error correction term. It measures the cointegrating relationship between p_t, p_t^* and c_t. Finally, θ is the speed of adjustment at which any short-run disequilibrium converges towards the long-run relationship. The unrestricted model is:

$$\begin{aligned}\Delta p_t = \alpha_0 \ + \ \alpha_1 t \ + \ \beta_1 p_{t-1} \ + \ \beta_2 p_{t-1}^* \ + \ \beta_3 c_{t-1} \ + \\ + \sum_{i=1}^{p} \gamma_{1i} \Delta p_{t-i} \ + \ \sum_{i=0}^{q_2} \gamma_{2i} \Delta p_{t-i}^* \ + \ \sum_{i=0}^{q_3} \gamma_{3i} \Delta c_{t-i} \ + \ u_t,\end{aligned} \tag{2.10}$$

which is the CEC form of a VAR model. In traditional time series econometrics, the cointegration of the system depicted above typically all variables in the VAR to be I(1) (ENGLE; GRANGER, 1987; PHILLIPS; OULIARIS, 1990). Therefore, the standard procedure requires a series of unit root tests in each of the variables, which is subject to misclassification, as the tests might suffer from different weaknesses. In order to avoid such problems, the cointegration approach developed by Pesaran, Shin and Smith (2001) is robust to whether the variables are $I(0)$ or $I(1)$. The null hypothesis is of no cointegration: $\beta_1 = \beta_2 = \beta_3 = 0$. The bound test has two critical significance values: the lower one is based on the assumption that all variables are $I(0)$ and the upper bound critical value is based on the assumption that they are all $I(1)$. Therefore, if the F_{PSS} test lies above the upper bound at a given significance level, one can assert the presence of a long-run relationship in the system. If it is beyond the lower bound, the null is not rejected and there is no cointegration. Finally, if the values lie in between the bounds, the test is inconclusive. In this case, most empirical works (for example, Brun-Aguerre, Fuertes and Greenwood-Nimmo (2013)) check the significance of the EC term given by θ, comparing its t-statistics to the critical values of Banerjee, Dolado and Mestre (1998).

Shin, Yu and Greenwood-Nimmo (2014) extends the standard model depicted in equation 2.10 to encompass asymmetric adjustments at both short and long run. The asymmetric decomposition of the first-difference variables is as in equation 2.6. We obtain the asymmetric decomposition in levels by the cumulative sum of 2.6:

$$p_t^{*(+)} \ = \ \sum_{s-1}^{t} \Delta p_s^{*(+)} \ = \ \sum_{s=1}^{t} max(\Delta p_s^*, 0), \tag{2.11}$$

$$p_t^{*(-)} \ = \ \sum_{s-1}^{t} \Delta p_s^{*(-)} \ = \ \sum_{s=1}^{t} min(\Delta p_s^*, 0). \tag{2.12}$$

Equations 2.11 and 2.12 are non-stationary time series representing the positive and negative decomposition of international fuel prices, respectively. The final specification of a NARDL model is then:

$$\Delta p_t = \alpha_0 + \alpha_1 t + \beta_1 p_{t-1} + \beta_2^{(+)} p_{t-1}^{*(+)} + \beta_2^{(-)} p_{t-1}^{*(-)} + \beta_3 c_{t-1} + \\ + \sum_{i=1}^{p} \gamma_{1i} \Delta p_{t-i} + \sum_{i=0}^{q_2} \gamma_{2i}^{(+)} \Delta p_{t-i}^{*(+)} + \sum_{i=0}^{q_3} \gamma_{2i}^{(-)} \Delta p_{t-i}^{*(-)} + \sum_{i=0}^{q_4} \gamma_{3i} \Delta c_{t-i} + u_t. \tag{2.13}$$

According to Nieh and Wang (2005), owing to its advantages of solving the typical problem of integration that has been a substantial focus for the time series literature and for dealing with the small-sample issue, the ARDL bound test has been largely applied in recent years.[57] Moreover, according to (HONARVAR, 2009), many of the studies that employ ECMs in the APT literature presume that asymmetry is present only in the adjustments process to the equilibrium, but not in the cointegration relationship, i.e., there is no long-run asymmetry imposed. The NARDL has the advantage of assessing also the long-run symmetry, given by $\beta_2^{(+}$ and $\beta_2^{(-)}$ in equation 2.13.

Finally, we build a few specifications regarding the nature of c_t. As seen in figure 12, there are different mechanisms affecting wholesale prices. Noticeably, the third branch given by the IPP policy encompass the other two, which allow us to specify a fairly simple model with only international prices in domestic currency. Therefore, we start with a parsimonious choice of covariates, wherein wholesale price adjustments depend fully on the international parity prices (i.e., $c_t = \phi$). In the following three specifications, we add the first two branches of figure 12 - the transmittal of crude oil and exchange rates - separately, but we keep them in the symmetric linear notation. For the latter, we decide to include the price of oil in domestic currency, as it is the actual cost of importing it. Doing so allow us to assess if there are any remaining pass-through effects of import prices of oil and exchange rate conditional on the international parity price and thus enhance the fitness of the models. This leads to four specifications:

Model A: $(\mathrm{p}_t^{*(+)},\ p_t^{*(-)})$

Model B: $(\mathrm{p}_t^{*(+)},\ p_t^{*(-)},\ e_t)$

Model C: $(\mathrm{p}_t^{*(+)},\ p_t^{*(-)},\ p_t^{oil})$

Model D: $(\mathrm{p}_t^{*(+)},\ p_t^{*(-)},\ p_t^{oil},\ e_t)$

[57] See e.g. Verheyen (2012), Bahmani-Oskooee and Aftab (2017), Nusair (2017), Bahmani-Oskooee and Gelan (2018), Lourenço and Vasconcelos (2018), Lourenço and Vasconcelos (2019).

We test the null of symmetry by a standard Wald test: in the long run (between coefficients β_2), and in the short run, which could be the contemporaneous effect (between coefficients γ_2, i.e., when $i = 0$) and the cumulative effect (between the short-run summation: $\sum_i \gamma_{2i}$). When lag selection is larger than one, they have different interpretation. For example, in periods of high-frequency changes in prices (phase 2), we expect that the full adjustment takes place within one day so that there is not much contribution of older lags - the prices of international fuel on Monday are fully passed-through to domestic wholesale prices on Tuesday and it is unlikely that the prices on the previous Friday still affects the new adjustments conditional on Monday. On the other hand, in phases of weekly adjustments (phase 3 for gasoline and 4 for diesel), we expect a more sluggish adjustments and allow the selection of up to 6 lags. We will come back to lag-selection details in section 3.6.

2.5 Data Description

In early 2018, Petrobras gave open access to daily prices of fuel charged on its refineries, i.e., free of taxes, transportation costs, mark-ups of the final supplier, and other costs in the distribution chain that might enter final retail prices. These prices are our measure of wholesale prices, as it represents the upstream stage that distributors buy from whenever they opt for domestic fuel. They correspond to 96.2% and 98.8% of the total gasoline and diesel produced in the country's refineries in the period, respectively.[58]

Besides fuel prices, we gathered the spot nominal exchange rate (NER) between the Brazilian Real (BRL) and US Dollars (USD), expressed as $\frac{BRL}{USD}$ and the West Texas Intermediate crude oil in Dollars. As a proxy for the international prices of automotive fuel, we selected the prices of regular gasoline and ultra-low-sulfur no. 2 diesel as reported by the U.S. Energy Information Administration (EIA). These prices are collected in two different spots - the Golf Coast and the New York Harbor. As they showed to be quite similar and in order to work with a single time series, we averaged them. We plot the time series in figure 15 and they show[59] - especially in the period of daily adjustments (phase 2) - considerably similar trends. Therefore, there are strong visual evidence of pass-through from international prices to wholesale domestic price. The similar pattern between the domestic currency and the barrel prices is due to the fact that the BRL is

[58] Source: ANP.

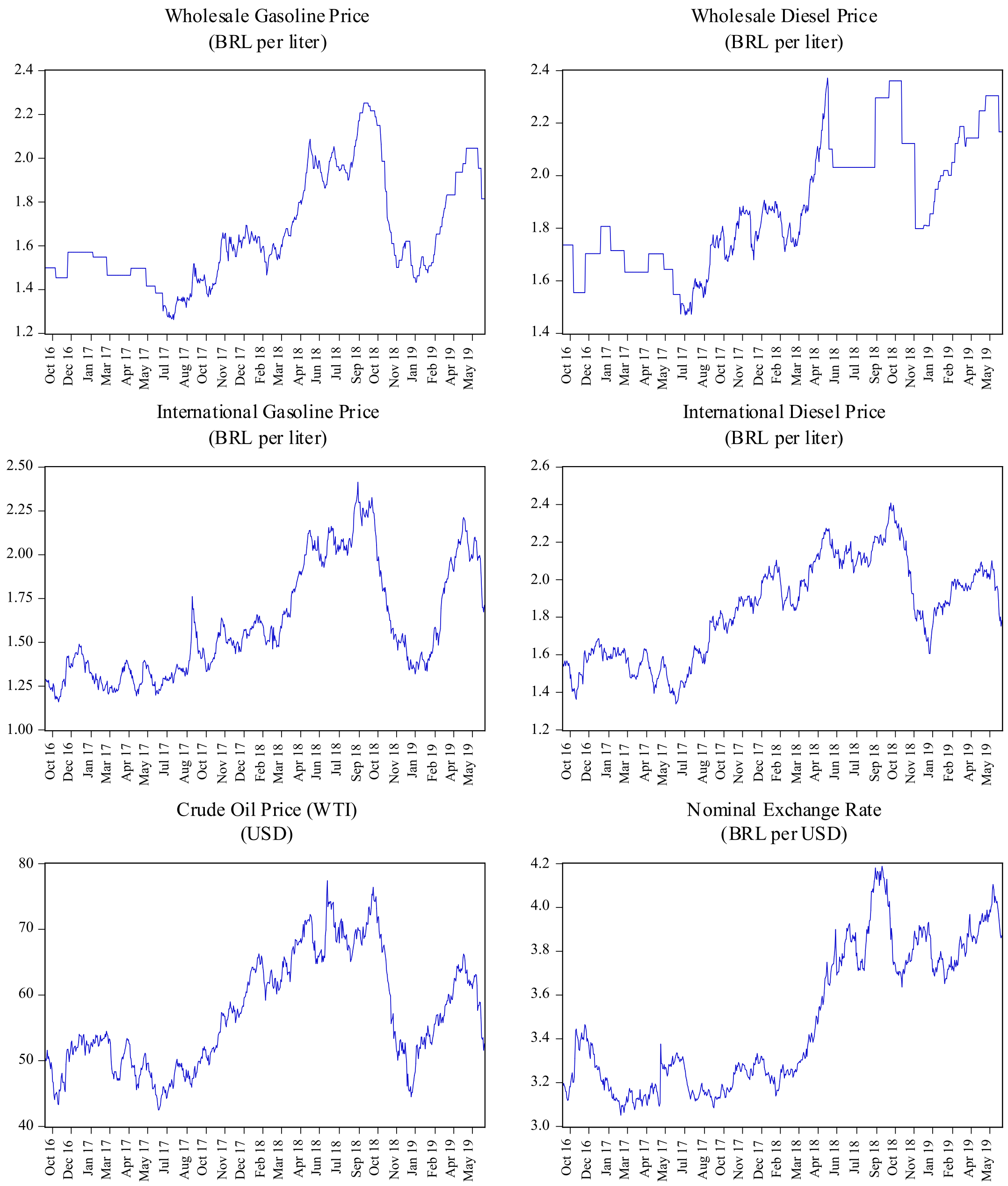
[59] The original data on international price of fuel is given by USD per gallon. In figure 15 we show them in BRL per liter (1 US gallon $\approx$ 3.785 liter).

considered a commodity currency.[60]

In the sample period (October 2016 to June 2019), the exchange rate depreciated from around 3.15 BRL/USD up to a peak of 4.18 by July 2018. The barrel price showed a similar increasing trend from mid-2017 to late-2018. Since November 2018, both decreased and stabilised in the first months of 2019, but at a new level when compared to 2016. Particularly, the exchange rate shows a typical sign of a structural break. The period that runs from October 2018 until the end of the sample is notorious for showing some signs of dissociation between the oil price and the NER. That is, the strong downturn seen in oil prices in the last three months of that year was not accompanied by a similar trend in the exchange rate. Lastly, the international prices of fuel had similar trends and accompanied the movements in the crude oil. However, after the peak in October 2018, the prices of gasoline plummeted next to the sample minimum and a few months later peaked again to over 2.20 BRL. The diesel price remained more unwavering, somewhere slightly above the sample average (1.71 BRL). In terms of correlation, the gasoline price remained slightly closer to the oil prices when compared to diesel (0.88 *vs* 0.85).

[60] A commodity currency co-moves with the world prices of primary commodity products, due to the domestic countries' specialisation on the export of raw materials. As such, when commodity prices increase, exporters' revenue also tend to increase. The domestic currency appreciates if they realise their gains in domestic currency. Moreover, a part of the literature investigates the feedback and spill-over effects between them, as the close relationship between currencies and commodities is a known stylised fact in international finance, although its causes demand further investigations. See Albert, Zelati and Araújo (2014) and Beckmann, Czudaj and Arora (2017).

Figure 15 – Time Series



2.6 Results

The daily sample on domestic fuel prices has a particular feature: it has data from Tuesdays to Saturdays. This means that the prices on Mondays are the same over the past weekend and the changes in markets on Monday are taken into account in the new price only by Tuesdays. Keeping the sample with a 6-day range would generate a lot of missing data on Saturdays, as the financial markets are closed and we would not have data for exchange rates and barrel prices. Therefore, we dragged the gasoline price series back one day to match the weekdays. Thus, the period Tue-Sat becomes Mon-Fri, as if the new price released on Tuesday morning was actually released on Monday night. This makes sense as the idea behind daily readjustments is that they depend on the figures seen in the financial market when it closes the day before. In terms of interpretation, the contemporaneous effect in the error correction form (in t) captures the one-day-lag effect. Analysing the short-run structure allows the understanding of how quickly the prices are adjusted.

Firstly, a few comments on how to interpret the short (γ_2^+ and γ_2^-) and long (β_2^+ and β_2^-) run coefficients. In the case of high-frequency price readjustments, the short run coefficients are central. They describe the asymmetric effects of positive or negative input price changes on wholesale prices. These are estimates of the theoretical price adjustments parameters sought by the company during the IPP policy. In contrast, the long run coefficients evaluate reaction times, length of fluctuations and the speed of adjustment towards a long-run equilibrium between international and wholesale prices (FASOULA; SCHWEIKERT, 2018). They show how fast and to what extent wholesale prices return to a long-run stable relationship with the international prices. Of course that, for a better understanding of the difference between them, we need a clear definition of what a shock would be in this market. International parity prices in domestic currency have two main sources of shocks. The first stems from nominal exchange rate and could be the outcome of diverse global and domestic shocks. The second comes from energy-market supply shocks that affect oil prices. For example, the highest residuals (above two standard deviations) in a regression of Δp_t^* against $\Delta Oil_t, \Delta Oil_{t-1}$ indicate that positive shocks are related to supply disruptions. One is the result of the Harvey hurricane that shut down refineries in August and September 2017, and the other two are related to the shift from producing winter-grade fuel to summer-grade fuel, which also shuts down refineries for cleaning and plant adequacy. Negative shocks, however, are harder to evaluate and may encompass both supply and demand sources.[61] As we are not focused on identifying exogenous shocks but rather high-frequency adjustments, we simply assume that shocks to the international parity price stem from multiple sources, but the positive ones are likely supply shocks.

The daily time series offers enough information to estimate the models in different time frames, without the need to resource to the small-sample properties of the Pesaran's bounds test. Thus, based on figures 13 and 14, we computed the NARDL models in three periods for the gasoline and four for the diesel. For each period, the equation assumes different specifications regarding the choices of covariates, as already mentioned. Tables 9 and 10 below depicts the results for the diesel and gasoline, respectively. As one can see, we show three sets of parameters based on equation 2.13. This representation of the NARDL model shows how short-run adjustments converge to a possible long-run equilibrium. Thus, the long-run coefficients can be taken into account if the model cointegrates, i.e., if the value for F_{PSS} is above a certain bound as calculated by Pesaran, Shin and Smith (2001), which means that the processes are governed by a common factor. In this regard, the information on price practises provided before leads us to assume that the company's behaviour is dictated by what happens in both the short- and long-run structure. The former are the daily reaction, while the latter are the itensity and speed of reaction after shocks.

For gasoline prices, the estimates for phase 1 showed mild evidence of cointegration, which means one should assess carefully the long-run terms. This result is expected as prices in that phase used to remain at the same level for approximately 20 days, that is, they were not converging to an equilibrium generated by a long-run relationship between international and wholesale prices. Moreover, as expected, the short-run coefficients do not tell much on how daily changes in international prices influenced the new wholesale prices chosen every moment. This points to the hypothesis that phase 1 for both diesel and gasoline holds little correlation with international prices.

Nevertheless, phase 2 shows a different pattern, as the four models had distinctive evidence of cointegration. This was when the company started to pursue daily readjustments. Interestingly, these models present pieces of evidence in favour of long-run asymmetry in this period: the average point estimate for increases in international prices were 0.87, compared to only 0.51 to the decreases and, what is more important, they

[61] For example, Boroumand, Porcher and Urom (2021) highlights three main negative oil shocks between 2005 and 2020: 2008 (associated with the global financial crisis), 2014 (associated with the sharp drop in prices due to the boom of shale oil), and 2020 (associated with the pandemic economic downturn). As such, these shocks have supply, demand, and financial sources.

showed to be statistically different. What this unveils is that potentially almost all the effect of a 1% increase in international gasoline prevails in the long run, whilst the transmission is incomplete after a 1% decrease. The short-run structure of models B and D shows a similar picture in the contemporaneous effect, i.e., the effect seen one day after the change in international prices. The cumulative short-run effect, however, has different patterns. In models B and D, the lag structure shows that the short run effects die out within one day (that is why we place a slash in the cumulative coefficients). However, model C tells us that after two days, the company refreshed its wholesale prices at around 25% after both increases and decreases of international gasoline (i.e., symmetrically).

Then, in phase 3, the pace of adjustments slackened (one new price every three days) and the long-run asymmetry vanished. The short-run structure reveals, however, some signs of asymmetric price policy. The cumulative effect had a negative sign for decreases in the international price, which means that they are being passed-through in the opposite direction[62]. That is, in this period a decrease in the international price of gasoline resulted in an increase in wholesale prices, in what is a sign of departure from what the company usually claims. In this period, at some point wholesale prices could have gone up even after a decrease in international price. In other to control for the political context, we have searched in official media releases for possible exogenous interference in this period, but have found none.

Diesel's phase 1 model is not much different from gasoline's. As expected, there is no clear short-run asymmetry, although the specifications show slightly stronger evidence of cointegration. Phase 2 shows consistent symmetry in both long and short run, with the exception of model C. A relevant feature was that, in phase 2, the price policy for diesel was more sensitive to the international prices than the gasoline - the cumulative effect in the short run represent a pass-through of 65% in average. Indeed, diesel price was adjusted more often in that phase (93% of the days, against 89%) and had a larger average change (0.21% *vs* 0.17%), which indicates that the pass-through from international volatility was larger compared to gasoline.

Following, phase 3* was characterised by steady prices for longer periods as a result of the government's subsidy program. The short-run coefficients show some evidence of positive asymmetry, firstly with positive values for $p^{*(+)}$ in models A and B and, secondly, with negatives values for $p^{*(-)}$ in models C and D. As prices slumped

[62] Recall how one should interpret signs in the NARDL model: the negative threshold decomposition must be read with opposite signs, because its marginal change is always negative. So a negative coefficient combined to a negative marginal change yields a positive final effect over the dependent variable.

after the strike in May 2018 (see figure 14), it is natural to see the positive asymmetry in the period, due to the fact that prices were set below its equilibrium level. Therefore, decreases in the international prices needed not to be accounted for in new prices, but increases did.

Finally, phase 4 followed a scheme similar to gasoline's phase 3. However, the price policy again shows some differences. This time, the models have evidence of negative asymmetry, i.e., pass-though from decreases in international prices were larger. This possible shows that the price police for diesel still suffered pressures from particular organised groups in the society, as the company failed to keep a policy of symmetric readjustments.

Overall, the estimates regarding the whole sample show that the positive asymmetry form phase 3 in the market for gasoline prevails in some cases even accounting for a larger period (see models B and D). Notwithstanding, the price policy for diesel is symmetric throughout the total sample. The positive asymmetry of phase 3*, as a result of the price level being below the equilibrium, is potentially offset by a negative asymmetry in the ensuing phase.

Although the roles played by the exchange rate and price of oil are secondary in the asymmetric price transmission, their inclusion and exclusion in the models change the estimates. In the second phase of gasoline, for example, the gasoline price holds both short- and long-run relationship with oil prices. For the former, 16% of a change in the WTI oil price was embedded into the wholesale price of gasoline[63].

We can summarise our findings as follows:

i) Short run coefficients have little to no effect in price readjustments in phases that they were sluggish;

ii) There is evidence of positive asymmetry in Gasoline's phases 02 and 03 (July 2017 to June 2019). Decreases in international prices did not seem to have a grip on Petrobras' wholesale prices from July 3rd 2018 to June 6th 2019;

iii) In the run-up to the strike in May 2018, the degree of pass-through from international prices were almost complete in the diesel market (60%-70% in just two days, although in a symmetric fashion);

iv) There is evidence of negative asymmetry in diesel's phase 4, which also shown up in the estimates for the whole sample. This might be a consequence of the pressure

[63] For simplicity, we did not show the coefficients of both p^{oil} and e. The complete output are available from the author.

exerted by certain classes of consumers to hold diesel price at lower levels. It is difficult to disentangle these effects within the model though.

Table 9 – Results (1) - Asymmetry on the price transmission of *diesel*

Specification			Cointegration		Long run			Short run contemporaneous			Short Run cumulative		
Phase	**Model**	**Lags**	F_{PSS}	**ECM**	$p^{*(+)}$	$p^{*(-)}$	**Wald**	$\Delta\mathbf{p}^{*(+)}$	$\Delta\mathbf{p}^{*(-)}$	**Wald**	$\sum\Delta\mathbf{p}^{*(+)}$	$\sum\Delta\mathbf{p}^{*(-)}$	**Wald**
1	A	(1,6,6)	4.632 ±	-0.088***	0.978***	0.956***	0.249	-0.133	0.090	-	-0.244	-0.277	-
1	B	(1,6,6,0)	3.538	-0.094***	0.472	0.460	-	-0.171*	0.050	1.483	-0.269	-0.309	-
1	C	(1,6,6,2)	4.547 †	-0.120***	0.925***	0.914***	0.075	-0.119	0.096	-	-0.194	-0.463	-
1	D	(1,6,6,0,2)	3.617 ±	-0.120***	1.015	1.003	-	-0.108	0.106	-	-0.184	-0.462	-
2	A	(1,2,2)	3.748	-0.106***	0.930***	1.030***	6.916***	0.021	0.138*	0.969	0.673***	0.772***	0.318
2	B	(1,2,2,2)	3.636	-0.132***	1.108***	1.140***	0.582	0.118	0.222**	0.805	0.604***	0.708***	0.365
2	C	(1,2,2,2)	2.752	-0.102***	0.907***	1.001***	6.666***	0.027	0.156**	1.176	0.664***	0.753***	0.258
2	D	(1,2,2,2,0)	3.725 ±	-0.147***	1.102***	1.109***	0.031	0.098	0.204**	0.862	0.549***	0.667***	0.483
3*	A	(1,0,2)	7.451 ‡	-0.158***	1.007***	0.933***	3.791*	0.154***	0.231*	0.252	0.154***	-0.045	0.840
3*	B	(1,0,2,0)	5.561 †	-0.160***	1.068***	1.003***	1.176	0.162**	0.246*	0.275	0.162**	-0.028	0.767
3*	C	(2,1,6,1)	8.517 ‡	-0.270***	1.091***	1.014***	10.260**	-0.075	0.290*	1.641	-0.075	-1.216**	4.690**
3*	D	(2,1,6,0,1)	6.846 ‡	-0.270***	0.989***	0.896***	6.488**	-0.098	0.237	1.337	-0.098	-1.341**	4.982**
4	A	(1,2,5)	6.622 ‡	-0.227***	0.851***	0.846***	0.004	-0.075	0.200*	1.961	0.141	-0.413	-
4	B	(1,5,0,1)	6.917 ‡	-0.301***	1.097***	0.932***	23.279***	-0.177	0.000	-	-0.681**	0.000	4.038**
4	C	(1,1,5,2)	7.380 ‡	-0.263***	0.867***	0.843***	0.092	-0.035	0.161	-	-0.035	-0.619	2.361
4	D	(1,5,0,1,0)	5.517 ‡	-0.301***	1.116***	0.943***	19.460***	-0.169	0.000	-	-0.675*	0.000	3.925*
Total	A	(1,4,2)	6.080 †	-0.053***	0.556***	0.549***	0.140	-0.068	0.212***	6.890***	0.445***	0.312***	0.583
Total	B	(1,4,2,0)	4.553 †	-0.053***	0.549*	0.543*	0.101	-0.068	0.212***	6.831***	0.445***	0.312***	0.577
Total	C	(1,4,2,0)	4.570 †	-0.053***	0.572***	0.564***	0.198	-0.066	0.211***	6.648**	0.449***	0.310***	0.625
Total	D	(1,4,2,0,0)	3.651 ±	-0.053***	0.559**	0.550**	0.155	-0.066	0.210***	6.566**	0.449***	0.309***	0.623

Phase 1: October 14th 2016 - July 3rd 2017; **Phase 2:** July 4th 2017 - May 21st 2018; **Phase 3*:** June 1st 2018 - December 31st 2018; **Phase 4:** January 1st 2019 - June 6th 2019

Table 10 – Results (2) - Asymmetry on the price transmission of *gasoline*

Specification			Cointegration		Long run			Short run contemporaneous			Short run cumulative		
Phase	**Model**	**Lags**	F_{PSS}	ECM	$p^{*(+)}$	$p^{*(-)}$	**Wald**	$\Delta p^{*(+)}$	$\Delta p^{*(-)}$	**Wald**	$\sum \Delta \mathbf{p}^{*(+)}$	$\sum \Delta \mathbf{p}^{*(-)}$	**Wald**
1	A	(1,6,0)	2.213	-0.052**	0.486*	0.596**	5.141**	-0.084	0.031*	3.599*	-0.025	0.031*	0.105
1	B	(1,6,0,0)	1.989	-0.067***	0.131	0.208	-	-0.108*	0.013	3.846*	-0.066	0.013	-
1	C	(1,4,5,6)	2.136	-0.070***	0.408**	0.505**	9.251***	-0.075	0.063	-	0.263*	-0.354**	5.663**
1	D	(1,5,0,0,6)	1.479	-0.070***	0.049	0.126	-	-0.107*	0.008	3.125*	0.074	0.008	-
2	A	(1,2,2)	6.770 ‡	-0.124***	1.001***	0.623***	7.177**	0.087	0.101	-	0.233***	0.343***	0.584
2	B	(1,0,0,2)	11.356 ‡	-0.198***	0.773***	0.421***	17.370***	0.153***	0.083***	12.908***	-	-	-
2	C	(1,2,2,0)	5.733 ‡	-0.134***	0.935***	0.500***	9.664***	0.087	0.084	-	0.230***	0.322***	0.402
2	D	(1,0,0,2,0)	9.561 ‡	-0.204***	0.756***	0.376***	18.416***	0.160***	0.077***	12.844***	-	-	-
3	A	(6,4,5)	7.974 ‡	-0.111***	0.763***	0.763***	0.001	-0.103	0.058	-	0.304**	-0.197	3.699*
3	B	(6,4,5,0)	7.879 ‡	-0.137***	0.415***	0.389***	3.644*	-0.133*	0.000	1.280	0.278*	-0.336*	5.538**
3	C	(6,4,5,2)	6.987 ‡	-0.118***	0.738***	0.741***	0.066	-0.065	0.068	-	0.400**	-0.170	4.845**
3	D	(6,4,5,0,2)	7.330 ‡	-0.147***	0.378***	0.354***	3.334*	-0.100	0.005	-	0.367**	-0.316*	7.009***
Total	A	(1,4,2)	8.172 ‡	-0.040***	1.135***	0.679***	2.596	-0.013	0.075*	1.685	0.406***	0.239***	2.570
Total	B	(1,4,2,0)	8.906 ‡	-0.061***	0.664***	0.174	6.841***	-0.018	0.034	-	0.399***	0.178***	4.459**
Total	C	(1,3,2,0)	8.524 ‡	-0.054***	1.053***	0.536***	5.905**	-0.016	0.070*	1.634	0.324***	0.231***	0.926
Total	D	(1,4,2,0,6)	8.421 ‡	-0.004***	0.673***	0.151	11.744***	-0.008	0.034	-	0.374***	0.167***	3.863*

Phase 1: October 14th 2016 - July 3rd 2017; **Phase 2:** July 4th 2017 - July 2nd 2018; **Phase 3:** July 3rd 2018 - June 6th 2019.
Notes:
Wald tests are not computed if both variables are non-significant at 10% level; when the short-run lag structure is 0 for a given variable, its contemporaneous effect is the same as the cumulative, so the coefficients are repeated; the order of variables in the lag structure is the dependent followed by the order as reported in the end of section 4;
10% significance (*)
5% significance (**)
1% significance (***)
Cointegration at 10% level (±)
Cointegration at 5% level (†)
Cointegration at 1% level (‡)

2.7 Robustness

2.7.1 Mean reversion

One possible explanation for positive asymmetry in gasoline phase 2 is an imbalance between domestic and international prices at the beginning of this phase. As such, if domestic prices were too low compared to international prices, there must have been some positive asymmetry for the catching up to occur, i.e., positive asymmetry would be the result of a mean reversion mechanism. To test for it, we would have to know how smooth the company would perform such a transition. As there is not much clue on this, we assume that with daily adjustments it could be performed quite quickly, within weeks. Thus, we re-estimated the baseline models splitting the beginning of phase 2, first by trimming half a month, then increasing the split window by two weeks until we leave the first two months (8 weeks) out of the sample.

Table 11 – Testing for mean reversion

N of weeks trimmed	Model A		Model B		Model C		Model D	
	$p^{*(+)}$	$p^{*(-)}$	$p^{*(+)}$	$p^{*(-)}$	$p^{*(+)}$	$p^{*(-)}$	$p^{*(+)}$	$p^{*(-)}$
2 weeks	$0.996^{\ddagger}$	0.639	$0.891^{\ddagger}$	0.471	$0.757^{\ddagger}$	0.426	$0.702^{\ddagger}$	0.319
4 weeks	$0.997^{\dagger}$	0.636	$0.863^{\ddagger}$	0.455	$0.714^{\ddagger}$	0.430	$0.641^{\ddagger}$	0.309
6 weeks	$1.049^{\dagger}$	0.601	$0.891^{\dagger}$	0.441	$0.725^{\dagger}$	0.439	$0.596^{\dagger}$	0.306
8 weeks	1.078	0.646	$0.882^{\ddagger}$	0.517	$0.776^{\dagger}$	0.150	$0.630^{\dagger}$	0.258

We trimmed 2 to 8 weeks at the beginning of the gasoline phase 2 in order to account for a possible adjustment mechanism that would render positive asymmetry.
Notes:
i) we omitted the chi-squared statistics from the Wald test and signalled the significance of the relevant asymmetries. Accordingly:
‡ 1%;
† 5%.
ii) we report only the long-run coefficients because in all the 16 models the lag selection for $p^{*(+)}$ and $p^{*(-)}$ were 0.

Results indicate persistence of the patterns first shown in table 2. As we rule out positive asymmetry stemming from a catching up mechanism, we support that there must be an adjustment function that leans towards positive asymmetry in the gasoline market for that particular period.

2.7.2 Endogeneity

For the estimations of the single-equation NARDL model to be valid, the variables included in the ECM must constitute no more than one long-run relationship. Hence, one should not have a bi-directional causation between domestic wholesale prices and the right-hand size variables $(p^{*(+)}, p^{*(-)}, p^{oil}, e)$ (PESARAN; SHIN; SMITH, 2001; MCNOWN; SAM; GOH, 2018). Consider the following ECM:

$$\Delta \boldsymbol{x}_t \;=\; \boldsymbol{a}_{x0} \;+\; \boldsymbol{a}_{x1} t \;+\; \boldsymbol{\Pi}_{\boldsymbol{xx}} \boldsymbol{x}_{t-1} \;+ \pi_{xy} y_{t-1} \;+\; \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_{xi} \Delta \boldsymbol{z}_{t-i} \;+\; \boldsymbol{\epsilon}_t, \tag{2.14}$$

where $\boldsymbol{x}_t = (p^{*(+)}, p^{*(-)}, p^{oil}, e)$ and $y_t = pt$. From 2.14, we test the exclusion of y_{t-1}. In case of non-rejection of the null $H_0^{xy} : \boldsymbol{\pi}_{xy} = 0$, we fulfil Pesaran's assumption 3. Table 12 below has five modified F statistics from bound tests for the whole sample. The first (F_{yx}) is our baseline value, already seen in tables 9 and 10. The following four statistics are related to equations where $p^{*(+)}, p^{*(-)}, p^{oil}$, and e are the dependent variable, respectively. The results indicate that the NARDL model is suitable to our data. We stress the relevance of such tests as they are not often seen in the applied literature. High significance level in any $F_{xy}(i), i = [1, 4]$ would change our interpretation of coefficients.

Table 12 – Testing for multiple cointegration vectors

	Gasoline				**Diesel**			
F_{PSS}	A	B	C	D	A	B	C	D
F_{yx}	8.172‡	8.906‡	8.524‡	8.421‡	6.080†	4.553†	4.570†	3.651±
$F_{xy(1)}$	3.831	4.273	2.936	3.741	1.737	1.570	2.692	1.907
$F_{xy(2)}$	2.116	3.286	1.644	2.641	0.649	3.198	1.003	2.894
$F_{xy(3)}$	-	3.155	-	2.651	-	2.774		2.053
$F_{xy(4)}$	-	-	1.786	1.499	-	-	1.865	1.489

F_{yx} is the baseline bound-test statistic.
$F_{xy}(i)$, $i = [1, 4]$ are statistics derived from equations where $p^{*(+)}, p^{*(-)}, p^{oil}$, and e are the dependent variable, respectively.
Cointegration at 10% level (±)
Cointegration at 5% level (†)
Cointegration at 1% level (‡)

2.7.3 Deterministic terms

According to Pesaran, Shin and Smith (2001), the NARDL allows five specifications of deterministic terms within the Conditional Error Correction (CEC) form. In this matter, theory and practice can both help out in the proper choice of model specification. Thus, whenever the price policy was characterised by staggered price dynamics, we decided to leave the deterministic trend out of the model (this was the case in phase 1 for gasoline and 1 and 3* for diesel). This seems to be a sensible choice as the deterministic trend shows to be non-significant and does not alter the baseline results. On the other hand, when prices showed clear signs of a slope, we tested the inclusion of a time trend (phases 2, 3, and total for gasoline and 4 for diesel). This way, we assumed that, in periods where the fuel prices are climbing up or slumping down, the price dynamics is not fully explained by the variables within the model. If there is an unobserved covariate influencing this upward/downward dynamics, the deterministic trend component can roughly account for it. A positive coefficient for a time trend, for example, could indicate persistence stemming from domestic inflation. The crucial point is that the persistence of cointegration after including a deterministic term shows that the co-movement between international and domestic fuel prices is not spurious.

The five specifications of a NARDL model provided in Pesaran, Shin and Smith (2001) regards how the deterministic terms affect the cointegrating relationship. Accordingly, they could be restricted to a linear combination of the elements in the cointegrating vector or simply enter the CEC form without affecting the long-run dynamics (in this case, they are unrestricted). Results depicted in tables 9 and 10 have an unrestricted time trend (PSS's case V). To test the robustness of the cointegration test, we estimate the restricted model (PSS's case IV).

Table 13 – Restricted and unrestricted deterministic terms

Phase	Model	F_{PSS}	
		Restricted	*Unrestricted*
2	A	$6.770^{\ddagger}$	$9.025^{\ddagger}$
2	B	$11.356^{\ddagger}$	$14.183^{\ddagger}$
2	C	$5.733^{\ddagger}$	$7.165^{\ddagger}$
2	D	$9.561^{\ddagger}$	$11.463^{\ddagger}$
Total	A	$8.172^{\ddagger}$	$6.145^{\ddagger}$
Total	B	$8.906^{\ddagger}$	$7.137^{\ddagger}$
Total	C	$8.524^{\ddagger}$	$6.833^{\ddagger}$
Total	D	$8.421^{\ddagger}$	$7.028^{\ddagger}$

The restricted model has a deterministic time trend in the long-run structure (equation 7), whereas the unrestricted model does not (see cases *iv* and *v* in Pesaran, Shin and Smith (2001).

2.7.4 Asymmetry at the international level

The discussion so far points to positive asymmetry in the price adjustments of gasoline, specially between 2017 and 2018. To assess if this price policy leaning towards positive asymmetry is peculiar within fuel markets, we can perform a few tests of asymmetry at the international level. If positive asymmetry is asserted elsewhere in a range of plausible regression models, one can argue that it is a typical price behaviour at the global level. This would whether imply that asymmetry is plausible in competitive markets or that global markets consistently have some degree of market power at different price stages. It could, for example, happen because prices are sticky downward because previous prices are a focal point for oligopolistic sellers. A second assumption regards the costs of adjusting fuel inventories. If the marginal costs of decreasing inventories is high (for example, sellers have limited resources to keep inventories when demand increases due to logistic bottlenecks), then it is likely that price decreases are less desirable. Lastly, consumers could have search costs when retail prices increase. This happens when they believe a price increase in a given station is common to all other alternative stations, rather than a change in relative prices between stations. Hence, the consumer will not search for a lower price in a competitor and supply curves are extremely inelastic in the short run to price increases (BORENSTEIN; CAMERON; GILBERT, 1997).

To make a comparable exercise, we look at the fuel market in the United States. Borenstein, Cameron and Gilbert (1997) provide a thorough guideline on how prices are formed in the American market. For our interest in particular, the most remarkable difference is that in the USA there is a closer competitive design for the market of wholesale automotive fuel. As such, as standard microeconomic theory states (reference), a more competitive market is more likely to have symmetric price adjustments.

According to how the American market is designed, we have two ways of estimating price transmission. The most similar to the Brazilian case would be assessing the transmission from spot prices (settled in the NY Harbor and GC markets) to wholesale prices. However, they are roughly the same, that is, wholesalers' margins hardly change spot prices. The estimations with these two series would render nearly complete and symmetric pass-through coefficients. To avoid such a naive regression, we estimate the transmission from crude oil to spot prices. Differently, this regression captures the main cost structure as it embeds the major input to the production of fuel.

2.7.5 Import Parity Price with transportation costs

The ideal measure of international parity price must also embeds transportation and storage costs incurred to deliver foreign fuel in the domestic ports. Until now, we simply multiplied the international fuel prices to the nominal exchange rate and assumed these costs to be somewhat constant within our 3-year span or - in models C and D - assumed that the international price of oil accounts for these costs. Thus, an exogenous cost shock that is independent to the oil prices would affect the actual parity price although it would be unnoticeable in our measure of international prices. An alternative solution regards the IPP data calculated by the *S&P Platts* for a group of Brazilian ports, which is also the reference price in Petrobras' bulletins. Unfortunately, they are available[64] only at the weekly frequency and from November 2018 to June 2019.

[64] The historic time series is private data and the author has not managed to access it yet. However, a glimpse of how this data looks like is available in the Petrobras' website.

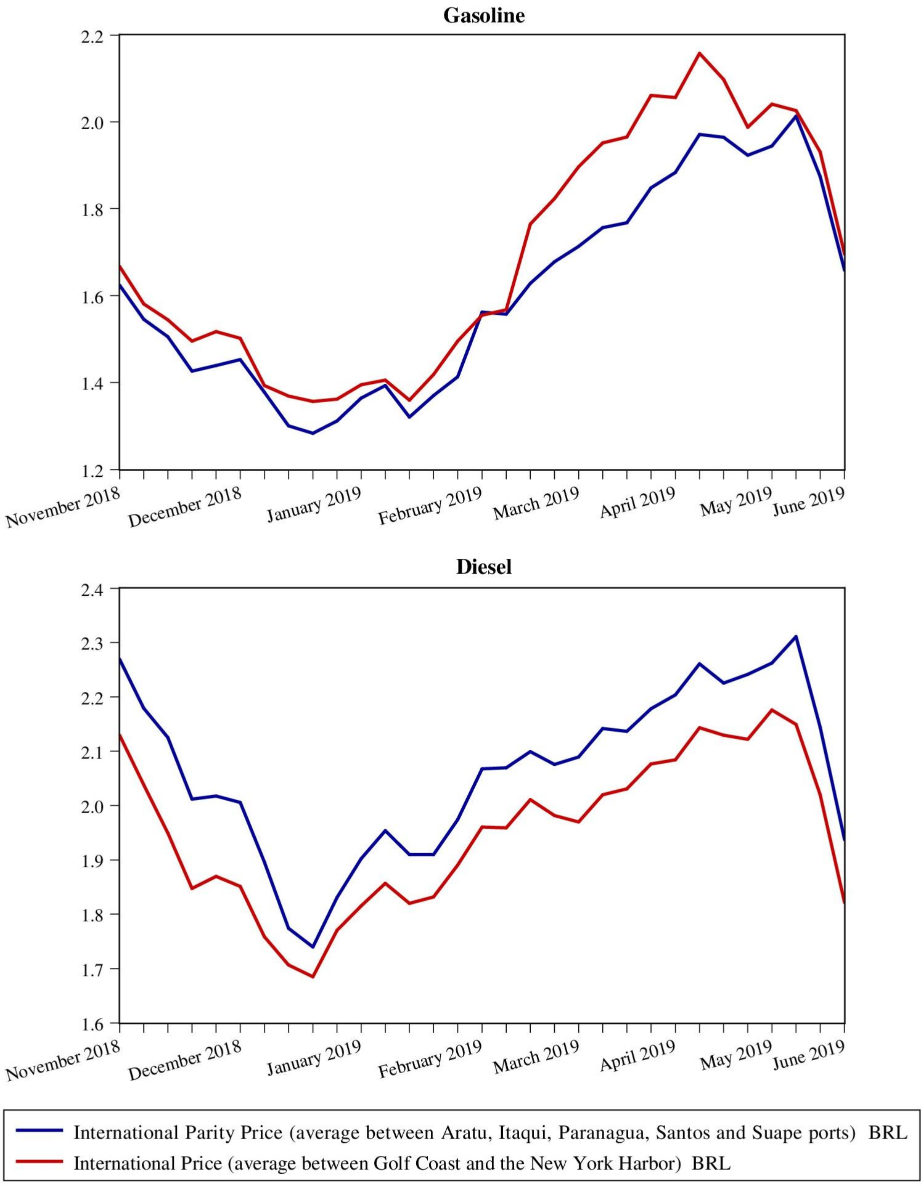
Table 14 – Asymmetry in foreign markets

Panel A: restricted sample (2016-2019)									
Gasoline	N	Lag selection	F_{PSS}	oil^+	oil^-	Δoil^+	Δoil^-	$\sum \Delta oil^+$	$\sum \Delta oil^-$
Daily	661	(3, 1, 3)	3.587	0.737	0.740	0.666	0.674	0.666	0.648
Weekly	218	(2, 2, 1)	6.656‡	0.742	0.746	0.672	0.800	0.447	**0.800**
Monthly	50	(2, 3, 3)	6.027†	0.692	0.669	0.849	0.962	0.121	1.199
Diesel									
Daily	661	(1, 1, 2)	2.866	**0.746**	0.713	0.662	0.647	0.662	0.716
Weekly	218	(2, 2, 2)	4.092	**0.885**	0.842	0.852	0.705	0.701	0.589
Monthly	50	(2, 3, 1)	3.166	0.869	0.810	0.959	0.771	0.500	0.771
Panel B: unrestricted sample (2009-2020)									
Gasoline									
Daily	2,805	(4, 2, 1)	9.712‡	0.707	0.711	0.337	0.361	0.254	**0.361**
Weekly	583	(3, 4, 1)	6.644‡	0.905	0.897	0.767	0.642	0.760	0.642
Monthly	134	(3, 1, 4)	4.064†	0.923	0.907	0.786	0.862	0.786	0.675
Diesel									
Daily	2,805	(4, 4, 4)	6.298†	*0.969*	0.962	0.254	0.251	0.329	0.324
Weekly	583	(4, 4, 3)	7.831‡	**1.147**	1.115	*0.791*	0.667	*0.769*	0.537
Monthly	134	(1, 1, 3)	5.187†	**1.093**	1.050	0.731	0.841	0.731	0.640

Panel A shows the results of estimations after restricting the sample to make the estimations comparable to the results for the Brazilian market (2016-2019);
Panel B shows the results of estimations allowing for a larger sample running between two major breaks in the oil market (the financial crisis in late 2008 and the pandemic surge in March 2020). Hence, the sample runs from January 2009 to February 2020;
For simplicity, we omit the significance level from each coefficient and highlight in **bold** those that reject the null of symmetry with at least 5% of significance an in *italic* those that reject with 10%.
Cointegration at 10% level (±)
Cointegration at 5% level (†)
Cointegration at 1% level (‡)

The diesel prices follow what we expected: they have similar trends and vary in the same proportion, and the shift between them can be attributed to transportation and storage costs. However, the gasoline data show that our measure of international prices could be overestimated to some extent and even the difference between the slopes seen uneven between March and April 2019. The steeper slope could be a source of measurement error, demanding some caution when interpreting the short-run coefficients for gasoline phase 3, where some positive asymmetry was found.

Figure 16 – International Parity Price



2.8 Conclusion

This study intends to investigate asymmetry in the transmittal of international fundamentals to the wholesale price of two important fuels in Brazilian behemoth oil company Petrobras from 2016 to 2019. What motivated the work was the company's claim that the price policy would strictly follow international fuel prices. In this setting, the domestic market would become more competitive, which means that foreign companies could import refined fuel and sell at a competitive level, regardless of the still outstanding market share held by the major domestic company.

The subject of our study differs from typical asymmetric price transmission (APT) analysis in agricultural economics. Differently from these typical works, the upstream and downstream prices in our case do not refer to the *same* good - but to close substitutes. The domestic wholesale price charged at the refineries is the price of domestic fuel obtained from refining a *mix of domestic and international* oil. Put differently, they are not composed of imported international fuel but are produced domestically. Therefore, the price transmission mechanism is not defined in direct terms, as global fuel prices are only *reference prices* for the domestic market. As such, it is plausible to assume - although we lack a theoretical model to support it - that in this design, domestic prices could divert from upstream prices easier because the latter is not explicitly part of the cost structure of the former. As such, there is no close economic link supporting that the downstream prices should follow upstream prices closely. However, they are and should be correlated because they are determined by common factors - such as the international price of oil and refining costs, which could be similar even in different countries.

We employed a simple asymmetric adjustment function that can be straightforwardly identified by a nonlinear autoregressive distributed lag model (NARDL) to achieve the goal. With daily wholesale and international fuel prices, nominal exchange rates and crude oil, we estimated adjustment functions for three different phases of the price policy for gasoline and four for diesel. In short, the estimates point to Petrobras' gasoline price policy being symmetric from October 2016 to mid-2018. From there, we found consistent evidence that the transmittal was asymmetric. Considering the whole sample, international shocks in gasoline prices were transmitted to wholesale prices within the next two to six working days in a proportion of 35-40% for positive shocks and 18-24% for negative ones. Moreover, in the long run, the prices seem to converge to an unbalanced relationship, which can be interpreted as prices not returning to the same level after the occurrence of positive and negative shocks in the same proportion.

On the other hand, the price of diesel showed a looser adherence to international prices in the long run. However, in the run-up to the strike in May 2018, international prices were almost fully passed-through to wholesale prices (60%-70% in just two days). In the subsequent period - from May to December 2018 - the price policy was swayed by a governmental program, which settled deals with consumers, represented mainly by truck drivers. In this phase, as the price was way below its equilibrium level, positive shocks on international diesel were easily passed through. The results showed that, even after this period, diesel prices seemed to hold negative asymmetry, meaning that decreases in international prices are passed-through to a larger extent when compared to increases. Overall, 30-45% of international shocks were transmitted to wholesale prices in a symmetric fashion.

This evidence shows that the company probably has more leeway to exert positive asymmetric transmission in the gasoline market, which, despite its importance, is less organised in terms of classes of workers and unions. We also consider the hypothesis that the company might have used the gasoline price to offset the losses (if there were any) from the subsidy program in diesel commercialisation.

There is still a lot to be studied to answer if there is an asymmetry in the final retail prices consumers pay at fuel stations. As we go down the distribution chain, more complexities appear, and more players are involved, so international shocks tend to be more offset depending on the market structures. In Brazil, retail prices vary considerably on a municipality basis, which might be the subject of regional and geographical studies. Last but not least, one interesting topic of discussion in both society and academia is the presence of fuel cartels, i.e., informal organisations that uses their power as oligopolistics to control fuel prices (SILVEIRA *et al.*, 2019) tacitly. As an intersection with what we studied here, one can perceive spatial clusters of positive price asymmetry as a sign of large market power and, along with other evidence, collusive behaviour. We leave such investigations to future studies.

Bibliography

ALBERT, F. M. de C.; ZELATI, P. E. C.; ARAÚJO, D. L. A. de. Is the brazilian real a commodity currency? large sample empirical evidence. *Independent Journal of Management & Production*, Instituto Federal de Educação, Ciência e Tecnologia de São Paulo (IFSP), v. 5, n. 1, p. 1–24, 2014.

AMITI, M.; ITSKHOKI, O.; KONINGS, J. International shocks, variable markups, and domestic prices. *The Review of Economic Studies*, Oxford University Press, v. 86, n. 6, p. 2356–2402, 2019.

AMITI, M.; ITSKHOKI, O.; KONINGS, J. *Dominant currencies: How firms choose currency invoicing and why it matters*. [S.l.], 2020.

ARON, J.; FARRELL, G.; MUELLBAUER, J.; SINCLAIR, P. Exchange rate pass-through to import prices, and monetary policy in south africa. *Journal of Development Studies*, Taylor & Francis, v. 50, n. 1, p. 144–164, 2014.

ARON, J.; MACDONALD, R.; MUELLBAUER, J. Exchange rate pass-through in developing and emerging markets: A survey of conceptual, methodological and policy issues, and selected empirical findings. *Journal of Development Studies*, Taylor & Francis, v. 50, n. 1, p. 101–143, 2014.

ASANE-OTOO, E.; SCHNEIDER, J. Retail fuel price adjustment in germany: A threshold cointegration approach. *Energy Policy*, Elsevier, v. 78, p. 1–10, 2015.

ATIL, A.; LAHIANI, A.; NGUYEN, D. K. Asymmetric and nonlinear pass-through of crude oil prices to gasoline and natural gas prices. *Energy Policy*, Elsevier, v. 65, p. 567–573, 2014.

BACCHETTA, P.; WINCOOP, E. V. Why do consumer prices react less than import prices to exchange rates? *Journal of the european economic association*, Oxford University Press, v. 1, n. 2-3, p. 662–670, 2003.

BACCHETTA, P.; WINCOOP, E. V. A theory of the currency denomination of international trade. *Journal of international Economics*, Elsevier, v. 67, n. 2, p. 295–319, 2005.

BACHMEIER, L. J.; CHA, I. Why don't oil shocks cause inflation? evidence from disaggregate inflation data. *Journal of Money, Credit and Banking*, Wiley Online Library, v. 43, n. 6, p. 1165–1183, 2011.

BACHMEIER, L. J.; GRIFFIN, J. M. New evidence on asymmetric gasoline price responses. *Review of Economics and Statistics*, MIT Press, v. 85, n. 3, p. 772–776, 2003.

BAE, Y.; JONG, R. M. D. Money demand function estimation by nonlinear cointegration. *Journal of Applied Econometrics*, Wiley Online Library, v. 22, n. 4, p. 767–793, 2007.

BAHMANI-OSKOOEE, M.; AFTAB, M. On the asymmetric effects of exchange rate volatility on trade flows: New evidence from us-malaysia trade at the industry level. *Economic Modelling*, Elsevier, v. 63, p. 86–103, 2017.

BAHMANI-OSKOOEE, M.; GELAN, A. Exchange-rate volatility and international trade performance: Evidence from 12 african countries. *Economic Analysis and Policy*, Elsevier, v. 58, p. 14–21, 2018.

BALAGUER, J.; RIPOLLÉS, J. Testing for price response asymmetries in the spanish fuel market. new evidence from daily data. *Energy Economics*, Elsevier, v. 34, n. 6, p. 2066–2071, 2012.

BALL, L.; MANKIW, N. G. Asymmetric price adjustment and economic fluctuations. *The Economic Journal*, Oxford University Press Oxford, UK, v. 104, n. 423, p. 247–261, 1994.

BANERJEE, A.; DOLADO, J.; MESTRE, R. Error-correction mechanism tests for cointegration in a single-equation framework. *Journal of time series analysis*, Wiley Online Library, v. 19, n. 3, p. 267–283, 1998.

BARDE, S. *et al.* A generalised variable elasticity of substitution model of new economic geography. *Documents de Travail de l'OFCE*, Observatoire Francais des Conjonctures Economiques (OFCE), v. 2033, 2008.

BASHER, S. A.; HAUG, A. A.; SADORSKY, P. Oil prices, exchange rates and emerging stock markets. *Energy Economics*, Elsevier, v. 34, n. 1, p. 227–240, 2012.

BECKMANN, J.; CZUDAJ, R. Is there a homogeneous causality pattern between oil prices and currencies of oil importers and exporters? *Energy Economics*, Elsevier, v. 40, p. 665–678, 2013.

BECKMANN, J.; CZUDAJ, R.; ARORA, V. The relationship between oil prices and exchange rates: theory and evidence. *US Energy Information Administration working paper series*, p. 1–62, 2017.

BELAISCH, M. A. *Exchange rate pass-through in Brazil*. [S.l.]: International Monetary Fund, 2003.

BERMINGHAM, C.; O'BRIEN, D. Testing for asymmetric pricing behaviour in irish and uk petrol and diesel markets. *The Energy Journal*, JSTOR, p. 1–26, 2011.

BERNANKE, B.; GERTLER, M.; WATSON, M. W. Oil shocks and aggregate macroeconomic behavior: The role of monetary policy: A reply. *Journal of Money, Credit, and Banking*, The Ohio State University Press, v. 36, n. 2, p. 287–291, 2004.

BERNANKE, B. S. Irreversibility, uncertainty, and cyclical investment. *The quarterly journal of economics*, MIT Press, v. 98, n. 1, p. 85–106, 1983.

BETTENDORF, L.; GEEST, S. A. Van der; VARKEVISSER, M. Price asymmetry in the dutch retail gasoline market. *Energy Economics*, Elsevier, v. 25, n. 6, p. 669–689, 2003.

BORENSTEIN, S.; CAMERON, A. C.; GILBERT, R. Do gasoline prices respond asymmetrically to crude oil price changes? *The Quarterly journal of economics*, MIT Press, v. 112, n. 1, p. 305–339, 1997.

BORENSZTEIN, E.; QUEIJO, V. H. V. Exchange rate pass-through in south america: An overview. IDB Working Paper Series, 2016.

BOROUMAND, R. H.; PORCHER, T.; UROM, C. Negative oil price shocks transmission: The comparative effects of the gfc, shale oil boom, and covid-19 downturn on french gasoline prices. *Research in International Business and Finance*, Elsevier, v. 58, p. 101455, 2021.

BRUN-AGUERRE, R.; FUERTES, A.-M.; GREENWOOD-NIMMO, M. Heads i win, tails you lose: New evidence of long-run asymmetry in exchange rate pass-through by exporting firms. *Available at SSRN 2365275*, 2013.

BRUN-AGUERRE, R.; FUERTES, A.-M.; GREENWOOD-NIMMO, M. Heads i win, tails you lose: Asymmetry in exchange rate pass-through into import prices. *Journal of the Royal Statistical Society (A)-Forthcoming*, 2016.

BRUN-AGUERRE, R.; FUERTES, A.-M.; PHYLAKTIS, K. Exchange rate pass-through into import prices revisited: what drives it? *Journal of international Money and Finance*, Elsevier, v. 31, n. 4, p. 818–844, 2012.

BURSTEIN, A.; EICHENBAUM, M.; REBELO, S. *Why are Rates of Inflation so low after large devaluations?* [S.l.], 2002.

BURSTEIN, A.; EICHENBAUM, M.; REBELO, S. Large devaluations and the real exchange rate. *Journal of political Economy*, The University of Chicago Press, v. 113, n. 4, p. 742–784, 2005.

BURSTEIN, A. T.; NEVES, J. C.; REBELO, S. Distribution costs and real exchange rate dynamics during exchange-rate-based stabilizations. *Journal of monetary Economics*, Elsevier, v. 50, n. 6, p. 1189–1214, 2003.

BUSSIERE, M. Exchange rate pass-through to trade prices: The role of nonlinearities and asymmetries. *Oxford Bulletin of Economics and Statistics*, Wiley Online Library, v. 75, n. 5, p. 731–758, 2013.

CAÍRES, M. L. Análise do pass-through da taxa de câmbio para preços de bens importados no brasil. 2013.

CAMPA, J. M.; GOLDBERG, L. S. Exchange rate pass-through into import prices. *Review of Economics and Statistics*, MIT Press, v. 87, n. 4, p. 679–690, 2005.

CAMPBELL, J. Y.; PERRON, P. Pitfalls and opportunities: what macroeconomists should know about unit roots. *NBER macroeconomics annual*, MIT press, v. 6, p. 141–201, 1991.

CARNEIRO, D. D.; MONTEIRO, A. M. d.; WU, T. Y. H. *Mecanismos não-lineares de repasse cambial para o IPCA*. [S.l.], 2002.

CASELLI, F. G.; ROITMAN, A. Nonlinear exchange-rate pass-through in emerging markets. *International Finance*, Wiley Online Library, 2016.

CA'ZORZI, M.; HAHN, E.; SÁNCHEZ, M. Exchange rate pass-through in emerging markets. ECB Working paper, 2007.

CHARI, V.; KEHOE, P. J.; MCGRATTAN, E. qcan sticky prices generate volatile and persistent real exchange rates?, rreview of economic studies 69. *63*, p. 633, 2002.

CHEN, Y.; YANG, S. Time-varying effect of international iron ore price on china's inflation: A complete price chain with tvp-svar-sv model. *Resources Policy*, Elsevier, v. 73, p. 102200, 2021.

CHOI, J.; WANG, H. Stakeholder relations and the persistence of corporate financial performance. *Strategic management journal*, Wiley Online Library, v. 30, n. 8, p. 895–907, 2009.

COCHRANE, J. H. A critique of the application of unit root tests. *Journal of Economic Dynamics and Control*, Elsevier, v. 15, n. 2, p. 275–284, 1991.

CORREA, A. d. S.; MINELLA, A. Nonlinear mechanisms of the exchange rate pass-through: A phillips curve model with threshold for brazil. *Revista Brasileira de Economia*, SciELO Brasil, v. 64, n. 3, p. 231–243, 2010.

CORSETTI, G.; DEDOLA, L.; LEDUC, S. High exchange-rate volatility and low pass-through. *Journal of Monetary Economics*, Elsevier, v. 55, n. 6, p. 1113–1128, 2008.

DELATTE, A.-L.; LÓPEZ-VILLAVICENCIO, A. Asymmetric exchange rate pass-through: Evidence from major countries. *Journal of Macroeconomics*, Elsevier, v. 34, n. 3, p. 833–844, 2012.

DELGADO, F. A. Hysteresis, menu costs, and pricing with random exchange rates. *Journal of Monetary Economics*, Elsevier, v. 28, n. 3, p. 461–484, 1991.

DELTAS, G. Retail gasoline price dynamics and local market power. *The Journal of Industrial Economics*, Wiley Online Library, v. 56, n. 3, p. 613–628, 2008.

DEVEREUX, M. B.; ENGEL, C. *Fixed vs. floating exchange rates: How price setting affects the optimal choice of exchange-rate regime.* [S.l.]: National Bureau of Economic Research Cambridge, Mass., USA, 1998.

DEVEREUX, M. B.; ENGEL, C. Exchange rate pass-through, exchange rate volatility, and exchange rate disconnect. *Journal of Monetary economics*, Elsevier, v. 49, n. 5, p. 913–940, 2002.

DEVEREUX, M. B.; ENGEL, C. Monetary policy in the open economy revisited: Price setting and exchange-rate flexibility. *The review of economic studies*, Wiley-Blackwell, v. 70, n. 4, p. 765–783, 2003.

DORNBUSCH, R. *Exchange rates and prices.* [S.l.], 1985.

EDWARDS, S. *The relationship between exchange rates and inflation targeting revisited.* [S.l.], 2006.

EDWARDS, S.; CABEZAS, L. *Exchange rate pass-through, monetary policy, and real exchange rates: Iceland and the 2008 crisis.* [S.l.], 2021.

EFRON, B. *The jackknife, the bootstrap and other resampling plans.* [S.l.]: SIAM, 1982.

EFRON, B. Bootstrap methods: another look at the jackknife. In: *Breakthroughs in statistics.* [S.l.]: Springer, 1992. p. 569–593.

EFRON, B.; TIBSHIRANI, R. J. *An introduction to the bootstrap.* [S.l.]: CRC press, 1994.

ENDERS, W.; GRANGER, C. W. J. Unit-root tests and asymmetric adjustment with an example using the term structure of interest rates. *Journal of Business & Economic Statistics*, Taylor & Francis Group, v. 16, n. 3, p. 304–311, 1998.

ENGEL, C. Expenditure switching and exchange-rate policy. *NBER macroeconomics annual*, MIT Press, v. 17, p. 231–272, 2002.

ENGLE, R. F.; GRANGER, C. W. Co-integration and error correction: representation, estimation, and testing. *Econometrica: journal of the Econometric Society*, JSTOR, p. 251–276, 1987.

ESCRIBANO, A.; SIPOLS, A. E.; APARICIO, F. Nonlinear cointegration and nonlinear error correction: Record counting cointegration tests. *Communications in Statistics-Simulation and Computation*, Taylor & Francis, v. 35, n. 4, p. 939–956, 2006.

FASOULA, E.; SCHWEIKERT, K. *Price regulations and price adjustment dynamics: Evidence from the Austrian retail fuel market.* [S.l.], 2018.

FERRUCCI, G.; JIMÉNEZ-RODRÍGUEZ, R.; ONORANTE, L. Food price pass-through in the euro area-the role of asymmetries and non-linearities. ECB working paper, 2010.

FORBES, K.; HJORTSOE, I.; NENOVA, T. Shocks versus structure: explaining differences in exchange rate pass-through across countries and time. Bank of England Working Paper, 2017.

FORBES, K.; HJORTSOE, I.; NENOVA, T. The shocks matter: improving our estimates of exchange rate pass-through. *Journal of international economics*, Elsevier, v. 114, p. 255–275, 2018.

FORERO, F. J. P.; VEGA, M. *et al. Asymmetric Exchange Rate Pass-through: Evidence from Nonlinear SVARs.* [S.l.], 2016.

FOSTER, H.; BALDWIN, R. Marketing bottlenecks and the relationship between exchange rates and prices. *Department of Economics, Massachusetts Institute of Technology, Cambridge. Processed*, 1986.

FREITAS, P. S. de; BUGARIN, M. N. S. A study on administered prices and optimal monetary policy: the brazilian case. *Premio de Banca Central Rodrigo Gómez/Central Banking Award" Rodrigo Gómez"*, Centro de Estudios Monetarios Latinoamericanos, CEMLA, 2007.

FREY, G.; MANERA, M. Econometric models of asymmetric price transmission. *Journal of Economic surveys*, Wiley Online Library, v. 21, n. 2, p. 349–415, 2007.

FRIEDMAN, M. The case for flexible exchange rates. *Essays in positive economics*, Chicago, v. 157, n. 203, p. 33, 1953.

FRIEDMAN, M. Capitalism and freedom. *Ethics*, v. 74, n. 1, 1962.

FROOT, K. A.; KLEMPERER, P. D. *Exchange rate pass-through when market share matters.* [S.l.]: National Bureau of Economic Research Cambridge, Mass., USA, 1988.

GELOS, G.; USTYUGOVA, Y. Inflation responses to commodity price shocks–how and why do countries differ? *Journal of International Money and Finance*, Elsevier, v. 72, p. 28–47, 2017.

GIOVANNINI, A. Exchange rates and traded goods prices. *Journal of international Economics*, Elsevier, v. 24, n. 1-2, p. 45–68, 1988.

GRANGER, C. W.; YOON, G. Hidden cointegration. *U of California, Economics Working Paper*, n. 2002-02, 2002.

GREGORIO, J. D.; LANDERRETCHE, O.; NEILSON, C.; BRODA, C.; RIGOBON, R. Another pass-through bites the dust? oil prices and inflation [with comments]. *Economia*, JSTOR, v. 7, n. 2, p. 155–208, 2007.

GUST, C.; LEDUC, S.; VIGFUSSON, R. Trade integration, competition, and the decline in exchange-rate pass-through. *Journal of Monetary Economics*, Elsevier, v. 57, n. 3, p. 309–324, 2010.

HA, J.; STOCKER, M.; YILMAZKUDAY, H. *Inflation and exchange rate pass-through.* [S.l.]: The World Bank, 2019.

HAHN, E. Pass-through of external shocks to euro area inflation. ECB Working Paper, 2003.

HAMILTON, J. D. This is what happened to the oil price-macroeconomy relationship. *Journal of monetary economics*, Elsevier, v. 38, n. 2, p. 215–220, 1996.

HAMILTON, J. D. What is an oil shock? *Journal of econometrics*, Elsevier, v. 113, n. 2, p. 363–398, 2003.

HAMILTON, J. D. *Causes and Consequences of the Oil Shock of 2007-08.* [S.l.], 2009.

HAMILTON, J. D. Nonlinearities and the macroeconomic effects of oil prices. *Macroeconomic dynamics*, Cambridge University Press, v. 15, n. S3, p. 364–378, 2011.

HAMILTON, J. D. Import prices and inflation. *28th issue (March 2011) of the International Journal of Central Banking*, 2018.

HARUN, M.; MAT, S. H. C.; FADZIM, W. R.; KHAN, S. J. M.; NOOR, M. S. Z. The effects of fuel subsidy removal on input costs of productions: leontief input-output price model. *International Journal of Supply Chain Management*, v. 7, n. 5, p. 529–534, 2018.

HELBLING, D. L.; MERCER-BLACKMAN, V.; DAO, C. T.-N.; ERBIL, N.; OAKES, E. Is inflation back? commodity prices and inflation. *World Economic Outlook*, International Monetary Fund Washington, DC, 2008.

HONARVAR, A. Asymmetry in retail gasoline and crude oil price movements in the united states: an application of hidden cointegration technique. *Energy Economics*, Elsevier, v. 31, n. 3, p. 395–402, 2009.

HOOKER, M. A. Are oil shocks inflationary? asymmetric and nonlinear specifications versus changes in regime. *Journal of money, credit and banking*, JSTOR, p. 540–561, 2002.

INOUE, A.; KILIAN, L. Inference on impulse response functions in structural var models. *Journal of Econometrics*, Elsevier, v. 177, n. 1, p. 1–13, 2013.

INTERNATIONAL MONETARY FUND. Europe: Coping with high commodity prices. *Regional Economic Outlook: Europe*, p. 19–36, October 2008.

ITO, T.; SATO, K. Exchange rate pass-through and domestic inflation: A comparison between east asia and latin american countries. *Research institute of Economy, Trade and Industry, RIETI Discussion Papers*, v. 7040, 2007.

ITO, T.; SATO, K. Exchange rate changes and inflation in post-crisis asian economies: Vector autoregression analysis of the exchange rate pass-through. *Journal of Money, Credit and Banking*, Wiley Online Library, v. 40, n. 7, p. 1407–1438, 2008.

JAMMAZI, R.; LAHIANI, A.; NGUYEN, D. K. A wavelet-based nonlinear ardl model for assessing the exchange rate pass-through to crude oil prices. *Journal of International Financial Markets, Institutions and Money*, Elsevier, v. 34, p. 173–187, 2015.

JOHANSEN, S.; JUSELIUS, K. Maximum likelihood estimation and inference on cointegration—with appucations to the demand for money. *Oxford Bulletin of Economics and statistics*, v. 52, n. 2, p. 169–210, 1990.

JOHANSEN, S.; JUSELIUS, K. *et al. Some structural hypotheses in a multivariate cointegration analysis of the purchasing power parity and the uncovered interest parity for UK*. [S.l.], 1990.

JOHANSEN, S. *et al. Likelihood-based inference in cointegrated vector autoregressive models*. [S.l.]: Oxford University Press on Demand, 1995.

JORDÀ, Ò. Estimation and inference of impulse responses by local projections. *American economic review*, v. 95, n. 1, p. 161–182, 2005.

KARORO, T. D.; AZIAKPONO, M. J.; CATTANEO, N. Exchange rate pass-through to import prices in south africa: is there asymmetry? 1. *South African Journal of Economics*, Wiley Online Library, v. 77, n. 3, p. 380–398, 2009.

KILIAN, L. The economic effects of energy price shocks. *Journal of Economic Literature*, v. 46, n. 4, p. 871–909, 2008.

KILIAN, L. Not all oil price shocks are alike: Disentangling demand and supply shocks in the crude oil market. *American Economic Review*, v. 99, n. 3, p. 1053–69, 2009.

KILIAN, L.; KIM, Y. J. How reliable are local projection estimators of impulse responses? *Review of Economics and Statistics*, The MIT Press, v. 93, n. 4, p. 1460–1466, 2011.

KILIAN, L.; LÜTKEPOHL, H. *Structural vector autoregressive analysis*. [S.l.]: Cambridge University Press, 2017.

KILIAN, L.; TAYLOR, M. P. Why is it so difficult to beat the random walk forecast of exchange rates? *Journal of International Economics*, Elsevier, v. 60, n. 1, p. 85–107, 2003.

KILIAN, L.; VIGFUSSON, R. J. Are the responses of the us economy asymmetric in energy price increases and decreases? *Quantitative Economics*, Wiley Online Library, v. 2, n. 3, p. 419–453, 2011.

KILIAN, L.; VIGFUSSON, R. J. Nonlinearities in the oil price-output relationship. *FRB International Finance Discussion Paper*, n. 1013, 2011.

KILIAN, L.; VIGFUSSON, R. J. The role of oil price shocks in causing us recessions. *Journal of Money, Credit and Banking*, Wiley Online Library, v. 49, n. 8, p. 1747–1776, 2017.

KIM, S.; ROUBINI, N. Exchange rate anomalies in the industrial countries: A solution with a structural var approach. *Journal of Monetary economics*, Elsevier, v. 45, n. 3, p. 561–586, 2000.

KLEMPERER, P. Competition when consumers have switching costs: An overview with applications to industrial organization, macroeconomics, and international trade. *The review of economic studies*, Wiley-Blackwell, v. 62, n. 4, p. 515–539, 1995.

KNETTER, M. M. Is export price adjustment asymmetric?: evaluating the market share and marketing bottlenecks hypotheses. *Journal of International Money and Finance*, Elsevier, v. 13, n. 1, p. 55–70, 1994.

KOOP, G.; PESARAN, M. H.; POTTER, S. M. Impulse response analysis in nonlinear multivariate models. *Journal of econometrics*, Elsevier, v. 74, n. 1, p. 119–147, 1996.

KPODAR, K.; ABDALLAH, C. Dynamic fuel price pass-through: Evidence from a new global retail fuel price database. *Energy Economics*, Elsevier, v. 66, p. 303–312, 2017.

KRUGMAN, P. *Pricing to Market when the Exchange Rate Changes*. [S.l.], 1986.

KRUGMAN, P. R. *Exchange-rate instability*. [S.l.], 1989.

LOPEZ-VILLAVICENCIO, A.; MIGNON, V. *et al. Exchange rate pass-through in emerging countries: Do the inflation environment, monetary policy regime and institutional quality matter?* [S.l.]: CEPII, Centre d'etudes prospectives et d'informations internationales, 2016.

LOURENÇO, L. d. S.; VASCONCELOS, C. R. F. Nonlinear exchange rate pass-through in latin america. *Economics Bulletin*, v. 38, n. 3, p. 1566–1582, 2018.

LOURENÇO, L. d. S.; VASCONCELOS, C. R. F. Impacts of exchange rate non-linearity on brazilian foreign trade. *International Economics and Economic Policy*, Springer, p. 1–21, 2019.

LÜTKEPOHL, H. Bootstrapping impulse responses in var analyses. In: SPRINGER. *COMPSTAT*. [S.l.], 2000. p. 109–119.

LÜTKEPOHL, H. *Introduction to multiple time series analysis*. [S.l.]: Springer Science & Business Media, 2013.

MACDONALD, R. *Exchange rate economics: theories and evidence*. [S.l.]: Routledge, 2007.

MACIEL, L. F. P. *Pass-through cambial: uma estimação para o caso brasileiro*. Tese (Doutorado), 2006.

MARSTON, R. C. Pricing to market in japanese manufacturing. *Journal of International Economics*, Elsevier, v. 29, n. 3-4, p. 217–236, 1990.

MCCALLUM, B.; NELSON, E. Monetary policy for an open economy: an alternative framework with optimizing agents and sticky prices. *Oxford Review of Economic Policy*, Oxford University Press, v. 16, n. 4, p. 74–91, 2000.

MCCARTHY, J. Pass-through of exchange rates and import prices to domestic inflation in some industrialized economies. *Eastern Economic Journal*, JSTOR, v. 33, n. 4, p. 511–537, 2007.

MCNOWN, R.; SAM, C. Y.; GOH, S. K. Bootstrapping the autoregressive distributed lag test for cointegration. *Applied Economics*, Taylor & Francis, v. 50, n. 13, p. 1509–1521, 2018.

MEYER, J.; CRAMON-TAUBADEL, S. V. Asymmetric price transmission: a survey. *Journal of agricultural economics*, Wiley Online Library, v. 55, n. 3, p. 581–611, 2004.

MEYLER, A. The pass through of oil prices into euro area consumer liquid fuel prices in an environment of high and volatile oil prices. *Energy Economics*, Elsevier, v. 31, n. 6, p. 867–881, 2009.

MIRDALA, R. *et al.* Exchange rate pass-through to domestic prices under different exchange rate regimes. *Journal of Applied Economic Sciences (JAES)*, Reprograph, v. 8, n. 26, p. 466–489, 2013.

MUSSA, M. L. *The Theory of Exchange Rate Determination.* [S.l.], 1984.

NARAYAN, P. K. The saving and investment nexus for china: evidence from cointegration tests. *Applied economics*, Taylor & Francis, v. 37, n. 17, p. 1979–1990, 2005.

NARAYAN, P. K.; NARAYAN, S. Estimating income and price elasticities of imports for fiji in a cointegration framework. *Economic Modelling*, Elsevier, v. 22, n. 3, p. 423–438, 2005.

NIEH, C.-C.; WANG, Y.-S. Ardl approach to the exchange rate overshooting in taiwan. *Review of Quantitative Finance and Accounting*, Springer, v. 25, n. 1, p. 55–71, 2005.

NORDHAUS, W. D. Who's afraid of a big bad oil shock? *Brookings Papers on Economic Activity*, JSTOR, v. 2007, n. 2, p. 219–238, 2007.

NUSAIR, S. A. The j-curve phenomenon in european transition economies: A nonlinear ardl approach. *International Review of Applied Economics*, Taylor & Francis, v. 31, n. 1, p. 1–27, 2017.

OBSTFELD, M. *International macroeconomics: beyond the Mundell-Fleming model.* [S.l.]: National Bureau of Economic Research Cambridge, Mass., USA, 2001.

OBSTFELD, M.; ROGOFF, K. Exchange rate dynamics redux. *Journal of political economy*, The University of Chicago Press, v. 103, n. 3, p. 624–660, 1995.

OBSTFELD, M.; ROGOFF, K. New directions for stochastic open economy models. *Journal of international economics*, Elsevier, v. 50, n. 1, p. 117–153, 2000.

OBSTFELD, M.; ROGOFF, K. S. *Risk and exchange rates.* [S.l.]: National bureau of economic research Cambridge, Mass., USA, 1998.

PARK, J. Y.; PHILLIPS, P. C. Nonlinear regressions with integrated time series. *Econometrica*, Wiley Online Library, v. 69, n. 1, p. 117–161, 2001.

PELTZMAN, S. Prices rise faster than they fall. *Journal of political economy*, The University of Chicago Press, v. 108, n. 3, p. 466–502, 2000.

PESARAN, M. H.; SHIN, Y. An autoregressive distributed-lag modelling approach to cointegration analysis. *Econometric Society Monographs*, Cambridge University Press, v. 31, p. 371–413, 1998.

PESARAN, M. H.; SHIN, Y.; SMITH, R. J. Bounds testing approaches to the analysis of level relationships. *Journal of applied econometrics*, Wiley Online Library, v. 16, n. 3, p. 289–326, 2001.

PHILLIPS, P. C.; OULIARIS, S. Asymptotic properties of residual based tests for cointegration. *Econometrica: Journal of the Econometric Society*, JSTOR, p. 165–193, 1990.

PIMENTEL, D. M.; LUPORINI, V.; MODENESI, A. d. M. Assimetrias no repasse cambial para a inflação: Uma análise empírica para o brasil (1999 a 2013). *Estudos Econômicos (São Paulo)*, SciELO Brasil, v. 46, p. 343–372, 2016.

POLLARD, P. S.; COUGHLIN, C. C. Size matters: asymmetric exchange rate pass-through at the industry level. *University of Nottingham Research Paper*, n. 2004/13, 2004.

POTTER, S. M. Nonlinear impulse response functions. *Journal of Economic Dynamics and Control*, Elsevier, v. 24, n. 10, p. 1425–1446, 2000.

RADCHENKO, S.; TSURUMI, H. Limited information bayesian analysis of a simultaneous equation with an autocorrelated error term and its application to the us gasoline market. *Journal of econometrics*, Elsevier, v. 133, n. 1, p. 31–49, 2006.

RODRIGUES, N.; LOSEKANN, L.; FILHO, G. S. Demand of automotive fuels in brazil: Underlying energy demand trend and asymmetric price response. *Energy Economics*, Elsevier, v. 74, p. 644–655, 2018.

ROGOFF, K. The purchasing power parity puzzle. *Journal of Economic literature*, JSTOR, v. 34, n. 2, p. 647–668, 1996.

ROWLAND, P. Exchange rate pass-through to domestic prices: the case of colombia. *Borradores de Economía; No. 254*, Banco de la República, 2003.

RUNKLE, D. E. Vector autoregressions and reality. *Journal of Business & Economic Statistics*, Taylor & Francis, v. 5, n. 4, p. 437–442, 1987.

SAIKKONEN, P.; CHOI, I. Cointegrating smooth transition regressions. *Econometric theory*, JSTOR, p. 301–340, 2004.

SALLES, A. A. D. Some evidence on the asymmetry between gasoline and crude oil prices in selected countries. *International Journal of Energy Economics and Policy*, v. 4, n. 4, p. 670–678, 2014.

SCHORDERET, Y. *et al. Asymmetric cointegration.* [S.l.]: Université de Genève/Faculté des sciences économiques et sociales Genève, 2003.

SERCU, P.; UPPAL, R.; HULLE, C. V. The exchange rate in the presence of transaction costs: implications for tests of purchasing power parity. *The Journal of Finance*, Wiley Online Library, v. 50, n. 4, p. 1309–1319, 1995.

SERIGATI, F. C. Petrobras: a vilã poderia ser a solução. *AgroANALYSIS*, v. 34, n. 02, p. 25–28, 2014.

SHAMBAUGH, J. A new look at pass-through. *Journal of International Money and Finance*, Elsevier, v. 27, n. 4, p. 560–591, 2008.

SHIN, Y.; YU, B.; GREENWOOD-NIMMO, M. Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ardl framework. In: *Festschrift in Honor of Peter Schmidt.* Nova York: Springer, 2014. p. 281–314.

SILVA, A. S. D.; VASCONCELOS, C. R. F.; VASCONCELOS, S. P.; MATTOS, R. S. de. Symmetric transmission of prices in the retail gasoline market in brazil. *Energy Economics*, Elsevier, v. 43, p. 11–21, 2014.

SILVA, C. M. d. S. Estratégia de preços da petrobras no mercado de combustíveis brasileiro pós-liberalização e instrumentos de amortecimento de variações internacionais. *Trabalho de dissertação de Mestrado. Rio de Janeiro: UFRJ*, 2003.

SILVEIRA, D.; VASCONCELOS, S.; BOGOSSIAN, P.; NETO, J. H. V. Cartel screening in the brazilian fuel retail market. *Available at SSRN 3421185*, 2019.

SIMS, C. A.; ZHA, T. Error bands for impulse responses. *Econometrica*, Wiley Online Library, v. 67, n. 5, p. 1113–1155, 1999.

STERNER, T. Oil products in latin america: the politics of energy pricing. *The Energy Journal*, International Association for Energy Economics, v. 10, n. 2, 1989.

TURNER, P. Power properties of the cusum and cusumsq tests for parameter instability. *Applied Economics Letters*, Taylor & Francis, v. 17, n. 11, p. 1049–1053, 2010.

UCHÔA, C. F. A. Testando a assimetria nos preços da gasolina brasileira. *Revista Brasileira de Economia*, SciELO Brasil, v. 62, n. 1, p. 103–117, 2008.

VERHEYEN, F. Bilateral exports from euro zone countries to the us—does exchange rate variability play a role? *International Review of Economics & Finance*, Elsevier, v. 24, p. 97–108, 2012.

WARE, R.; WINTER, R. Forward markets, currency options and the hedging of foreign exchange risk. *Journal of International Economics*, Elsevier, v. 25, n. 3-4, p. 291–302, 1988.

WEBBER, A. G. Newton's gravity law and import prices in the asia pacific. *Japan and the World Economy*, Elsevier, v. 12, n. 1, p. 71–87, 2000.

WINKER, P.; HELMUT, L.; STASZEWSKA-BYSTROVA, A. Confidence bands for impulse responses: Bonferroni versus wald. Kiel und Hamburg: ZBW-Deutsche Zentralbibliothek für . . . , 2014.

WLAZLOWSKI, S. Petrol and crude oil prices: asymmetric price transmission. Aston University, 2001.

Appendix A - Summary of the literature for Brazil

Work	Belaisch (2003)	Caíres (2013)	Carneiro, Monteiro and Wu (2002)
Sample	1999-2002	1996-2013	1994-2011
Methodology	VAR	ADL	OLS and NLLSQ
Asymmetries	No	Yes	No
Nonlinearities	No	Yes	Yes
Long Run	No		
Price decomposition	Consumer & Wholesale; Free & Administered; Tradables & Nontradables.	30 groups of imported goods	Consumer level (aggregate and decomposed)
ERPT definition	IRF-based exchange rate elasticity	Slope coefficients	Slope coefficients
ERPT coefficients	12-month ERPT: 17% (CPI); 120%(Wholesale); 15% (Tradables); 12% (Nontradables)	Varying coefficients and structural breaks were identified. ERPT to import prices varies from 0.5 to 0.8	6.3% - 11%
General Conclusions	ERPT to CPI is limited but rapid (full effect in two quarters). ERPT to wholesale prices was complete in that period.	There are structural breaks in the ERPT coefficients that could be associated with changes in the currency regime.	Nonlinear models improve the estimation of linear models.

Continuation

Work	Maciel (2006)	Pimentel, Luporini and Modenesi (2016)	Correa and Minella (2010)
Sample	2000-2005	1999-2013	1998-2005
Methodology	GLS	SVAR	SETAR (with endogenous variables)
Asymmetries	No	Yes	Yes
Nonlinearities Long Run	No	No	Yes
Price decomposition	Consumer level (different levels of disaggregation)	Consumer level	Consumer level
ERPT definition	Slope coefficients	IRF-based exchange rate elasticity	Slope coefficients
ERPT coefficients	After 10 months: 40% for tradables; 6% for nontradables; 14.6% for administered prices	Symmetric model: 5.36% - 7.15%. Asymmetric model: 2.22% - 2.46% for appreciations and 5.01% - 7.72 for depreciations.	0% and 11% for low and high output gap; 2% and 11% for appreciations and depreciations; 80% and 7% for low and high exchange rate volatility.
General Conclusions	Goods can be reclassified from a binary definition (tradables vs nontradables) to a continuous classification, based on the ERPT coefficient.	Strong positive asymmetry fir all specifications.	Presence of nonlinear mechanisms in the ERPT. It is higher when the economy is booming, when the exchange rate depreciates above a certain threshold, and when exchange rate volatility is lower.

ADL: Augmented Distributed Lag; NLLSQ: Nonlinear Least Squares; GLS: Generalised Least Squares; SETAR: Self-exciting Threshold Autoregressive; SVAR: Structural VAR. This table does not exhaust the Brazilian literature. We recommend also: a, b, c

Appendix B - Comments on the asymmetric VAR model

In the text, we extended the 2-variable VAR model by Kilian and Vigfusson (2011a) (KV) to 4-, 5- and 7-variable models in order to build a partially structural relation that allows domestic variables to feed back the exchange rate equation after asymmetric shocks.

The exposition in the text and the discussion in Kilian and Vigfusson (2011a) show that the IRFs are nonlinear functions of the i) structural parameters; ii) the uncertainty over the paths of innovations; iii) and the history of the process. To see these issues, consider the following 2-variable asymmetric VAR model in which x_t is split into positive and negative changes in the y_t equation:

$$\begin{aligned} x_t &= b_{10} + \sum_{i=1}^{p} b_{11,i} x_{t-i} + \sum_{i=1}^{p} b_{12,i} y_{t-i} + \varepsilon_{1,t} \\ y_t &= b_{20} + \sum_{i=0}^{p} g^{+}_{21,i} x^{+}_{t-i} + \sum_{i=0}^{p} g^{-}_{21,i} x^{-}_{t-i} + \sum_{i=1}^{p} b_{22,i} y_{t-i} + \varepsilon_{2,t} \end{aligned} \tag{B.1}$$

- *Structural parameters*

Observe that the degree of asymmetry depends on how the model is specified on its *entirety*, not only on the asymmetric slope coefficients (g^{+}_{21} and g^{-}_{21}). In the two-variable case, ignoring all the uncertainty around the stochastic realisation of the innovations and setting the constant terms (b_{10} and b_{20}) equal to zero for simplicity, the degree of asymmetry after $h = 1$ is given simply by $g^{+}_{21,0} - g^{-}_{21,0}$. However, when $h = 2$, the degree of asymmetry depends on *all* slope coefficients and is given by:

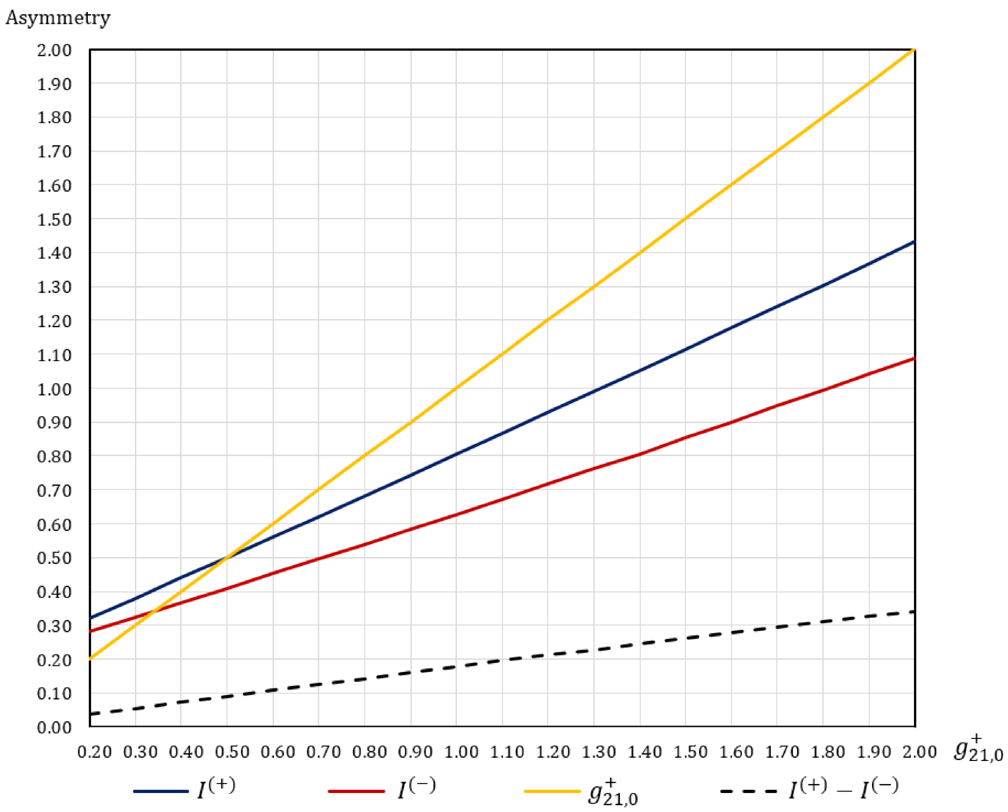
$$b_{11}(g^{+}_{21,0} - g^{-}_{21,0}) + b_{12}({g^{+}_{21,0}}^{2} - {g^{-}_{21,0}}^{2}) + b_{21,1} - g_{21,1} \tag{B.2}$$

As all the structural coefficients matter when assessing asymmetry over the predictive horizon, the choice between either a shock in a AR(1) and a shock in a semi-structural VAR setting has implication on the degree of asymmetry. Notice that imposing an AR(1) is equivalent to restricting $b_{12,i}$, $i = [1, ..., p]$ to zero.

In fact, this is the central issue driving the difference between slope parameters of a single equation and impulse response function of system models. To show how different interpretation can be when analysing slope parameters and dynamic responses, even in a non-economic artificial setting, we simulate the model B.1. Figure B.1 depicts the inferred asymmetry when analysing slope parameters and dynamic impulse responses.

Observe that the dynamic degree of asymmetry grows less than proportionally when the asymmetric slope coefficient grows. The yellow line is a 45º-degree line that associates the values of $g_{21,0}$ to the implied asymmetry, which in this case is simply $g_{21,0}$. The other lines show the degree of asymmetry implied by dynamic structural responses. The result for this is obvious: structural shocks trigger all the coefficients within the system, whereas in simple regression models, asymmetry is slope-based and, thus, linear. On other words, the patterns depicted in the blue, red and black-dashed lines can change if we change *any* structural parameter.

Figure B.1 - Dynamic asymmetries by size of the slope asymmetry
1,000 simulated time series of 1,000 observations each



- *Innovations' paths*

Much of the discussion on the importance of accounting for the uncertainty in the innovations and the history of the variables (Ω^i) have already been exposed in Kilian and Vigfusson (2011a). Accordingly, they propose an algorithm that shows how to obtain the correct conditional and unconditional impulse responses, respectively:

$$
\begin{aligned}
&I_y\left(h,\delta,\Omega^i\right) \\
&I_y\left(h,\delta\right) = \int I_y(h,\delta,\Omega^i)\, d\,\Omega^i
\end{aligned} \tag{B.7}
$$

The unconditional response is also what we work upon in our text. If the interest is policy or forecasting, the conditional is also useful, because is sets the response for given initial conditions.

Notice that in B.2, we show a closed form for the asymmetry by holding $\varepsilon_{1,t} = 0$ and $\Omega^i = 0$. When considering the uncertainty in the innovations, finding the algebraic solution for the asymmetry becomes more complex. Consider first that the mean of the process x_t conditional on the history being zero, is also zero, such that:

$$
\begin{aligned}
&E(x_t|\Omega^i = 0) = 0, \text{ which implies,} \\
&E(x_t^+|\Omega^i = 0) = E(x_t^-|\Omega^i = 0) = 0
\end{aligned} \tag{B.3}
$$

Consider a shock of size δ in $\varepsilon_{1,t}$. When we add the uncertainty in ε_{1t}, there are three possible paths for x_t^+ in $h = 1$: if the innovation is lower than $-\delta$ before the shock, the expected effect on x_t^+ is zero (for example, if $\varepsilon_{1t} = -1.5$ and $\delta = 1$, the positive shock had no effect on x_t^+). If ε_{1t} is between $-\delta$ and 0, x_t^+ goes from 0 to $\varepsilon_{1t} + \delta$ and, finally, if $\varepsilon_{1t} > 0$, the effect on x_t^+ is δ:

$$
E(x_t^+|\Omega^i = 0, \varepsilon_{1,t}, \delta) - E(x_t^+|\Omega^i = 0, \varepsilon_{1,t}) = \begin{cases} \delta, & \text{if } \varepsilon_{1,t} \geq 0 \\ \varepsilon_{1,t} + \delta, & \text{if} -\delta < \varepsilon_{1,t} < 0 \\ 0, & \text{if } \varepsilon_{1,t} < -\delta \end{cases} \tag{B.4}
$$

Each of these three realisation yields different degree of asymmetry. When h grows larger, the asymmetry becomes a nonlinear function of both the parameters and the expected values of *all* $x_t^+, t = 1, ..., h$. The severity of ignoring the uncertainty on the innovations lowers if the distribution of $\varepsilon_{1,t}$ is well-behaved and has zero mean. In that case, $E(x_t^+|\Omega^i = 0, \varepsilon_{1,t} = 0)$ converges to δ. If the shape of the distribution is not known though, a bootstrap method is necessary to build the actual paths for x_t^+ (and, consequently, x_t^-).

One last caveat remains and it potentially brings more complexity to the analysis. Similar to Kilian and Vigfusson (2011a)'s exercise in their appendix, we show how asymmetry depends on the expected value of x_t, the variable that enters the second equation with its asymmetric decomposition. To do so, we build a loop over the intercept of x_t's equation, b_{10}, so it varies inside the range $[-5.0, 5.0]$. Figure H.3 shows that for large absolute expected values of x_t, the system is symmetric. Asymmetry becomes somewhat relevant when $E[x_t]$ lies between -2.0 and 2.0. Interestingly, there is *negative* asymmetry when $E[x_t]$ is around ± 1.0. Take $n_y^{(+)}$ and $n_y^{(-)}$ as the total amount of positive and negative asymmetries respectively over all simulated series. Figure B.3 depicts in its vertical axis the proportion of positive asymmetries (i.e., $n_y^{(+)}/n$, where n is the total number of simulations).

Luckily, in economic models with first-differenced data, the average value of the covariates is usually nearer than zero than the ones needed to yield counter-intuitive asymmetry.

Figure B.2 - Asymmetric shocks by different values of $E[x_t]$
1,000 simulated time series of 1,000 observations each

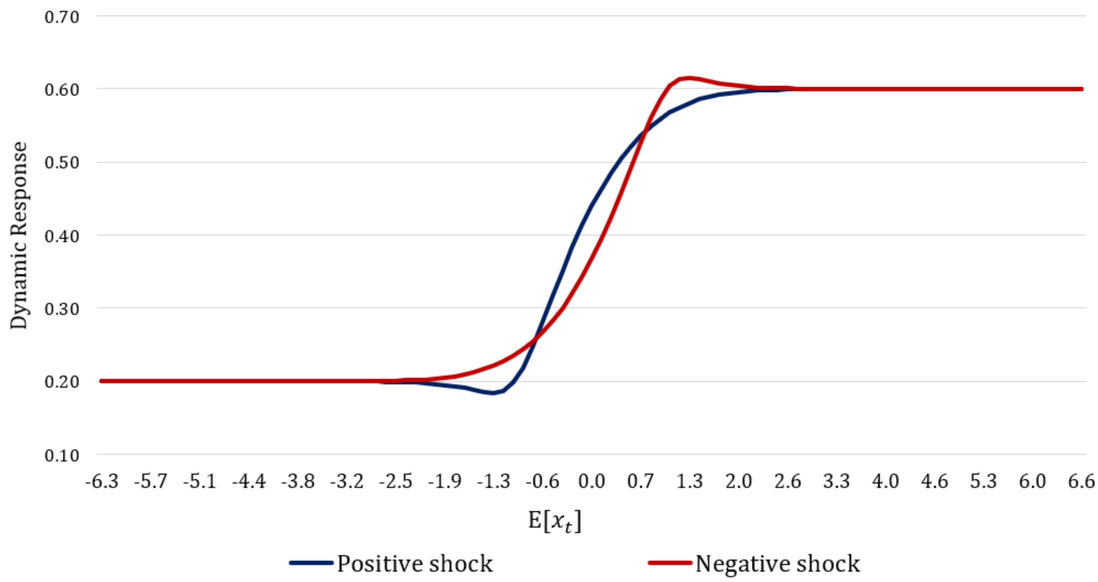
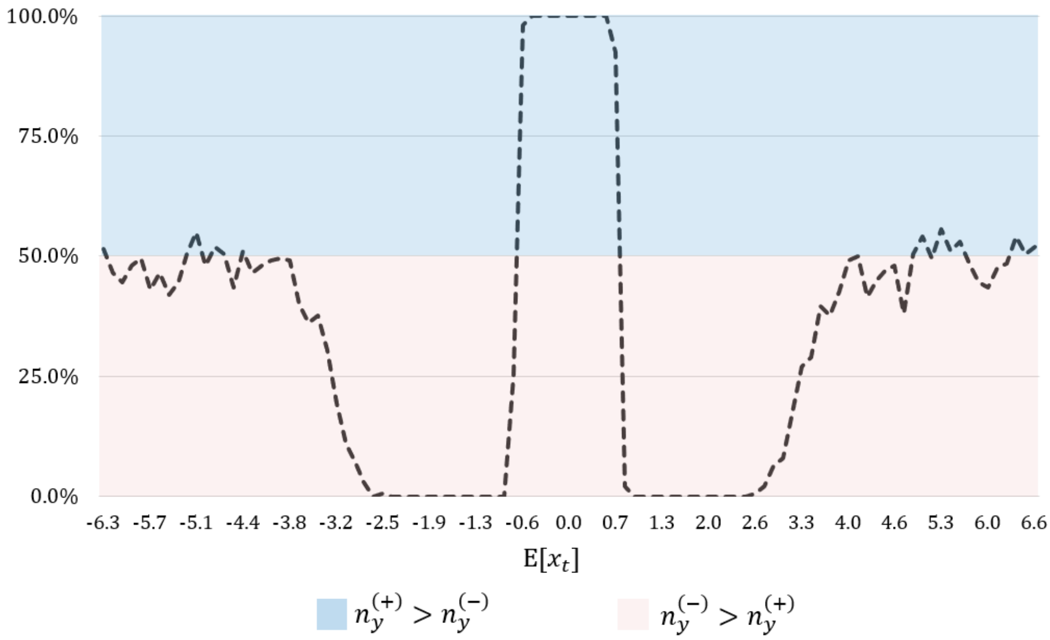


Figure B.3 - Proportion of positive and negative asymmetries by values of $E[x_t]$

1,000 simulated time series of 1,000 observations each



B.1. Computing asymmetries

Most studies in the ERPT literature that assess asymmetry use slope-based Wald tests to determine if asymmetry plays a role in the regression. In our case, for example, this would imply testing either $H_0 : \beta_{2i}^{(+)} = \beta_{2i}^{(-)}$, $i = [0, ..., p]$ or $H_0 : \sum_{i=0}^{p} \beta_{2i}^{(+)} = \sum_{i=0}^{p} \beta_{2i}^{(-)}$ in 1.4.3. However, as Kilian and Vigfusson (2011a) states:

> *"While slope-based tests are useful in assessing the symmetry of the slope parameters of single-equation regression models, they are not informative about the degree of symmetry of the impulse response function obtained from a fully specified dynamic structural model."*

This suggests that a proper approach is testing for dynamic asymmetry directly on the impulse response functions.

The main issue in systems like these is answering if the asymmetries in the reduced-form parameters lead to significant asymmetries in the impulse response function. The plots we show in the text only show the distribution of the LxM responses conditional on the original data and the parameters estimated by OLS. To test for the significance of the asymmetric patterns, bootstrap simulation over algorithm 1 is used:

Algorithm 2

---

1. Estimate either models depicted in the text by OLS using the whole sample period and store the vector of residuals $\hat{\varepsilon}_t$;

2. Set the numbers:
H: *the horizon for IRFs;* **L:** *number of draws for different initial forecasting points;* **M:** *number of draws for different shock realisations;* **N:** *number of draws for the parameter uncertainty.*

3. For each $n = 1, ..., N$:

A) Take a block of p consecutive values of p_t, $e_t^{(+)}$, $e_t^{(-)}$ and x_t to serve as initial conditions. This is not a history Ω^i;
B) Randomise a sequence of residuals from $\hat{\varepsilon}_t$ and call it $\hat{\varepsilon}_t^*$;
C) From the initial conditions and $\hat{\varepsilon}_t^*$, simulate the bootstrapped sample p_t^*, $e_t^{(+)*}$, $e_t^{(-)*}$, and x_t^*.
D) From the bootstrap sample obtained in C) and the original parameters estimated in 1., re-estimate the model;
E) For each $l = 1, ..., L$:

a) Take a block of p consecutive values of p_t, $e_t^{(+)}$, $e_t^{(-)*}$, and x_t. This defines a history Ω^i.
b) For each $m = 1, ..., M$:

i) Simulate three realisations for ε_{t+h}, for $h = 0, ..., H$, by drawing with replacement from the empirical distribution of the residuals. The realisations are identical, except for that $\varepsilon_{j,t}$ (i.e., $h = 0$) equal to δ and $-\delta$ in two of them, where j is the variable we want to shock. The other one is the baseline.
ii) Use the three bootstrapped realisations from i), the estimated coefficients from D) and the history Ω^i to simulate the paths $\left\{\hat{PT}_t\right\}_T^{T+h}$, $\left\{\hat{PT}_t^+\right\}_T^{T+h}$, and $\left\{\hat{PT}_t^-\right\}_T^{T+h}$.
iii) Calculate:
. $\left\{\hat{PT}_t^+\right\}_T^{T+h} - \left\{\hat{PT}_t\right\}_T^{T+h}$, and call it IRF_m^+;
. $\left\{\hat{PT}_t^-\right\}_T^{T+h} - \left\{\hat{PT}_t\right\}_T^{T+h}$, and call it IRF_m^-;

c) Average IRF_m^+ and IRF_m^- to get IRF_l^+ and IRF_l^-.

F) Average IRF_l^+ and IRF_l^- to get IRF_n^+ and IRF_n^-.

4. Compute a joint covariance matrix and perform the Wald test with the null $H_0 : IRF_n^+ = IRF_n^-$ for each step h following the χ^2_{h+1} distribution.

This routine is similar to the one used in Forero, Vega *et al.* (2016) to build confidence intervals. Compared to algorithm 1, it adds a N-step loop to account for the parameter uncertainty. Following the usual methods on bootstrap literature, we chose $N = 1,000$.[65]

However, notice that the bootstrap simulation Wald test of the joint null hypothesis of symmetric responses to positive and negative shocks used in Kilian and Vigfusson (2011a) *depends on the magnitude of δ, so that the evidence against symmetry depends on the magnitude of the shock considered.* It turns out that the Wald test between the asymmetric IRFs in their work tends to underestimates asymmetry, even when it is imposed structurally. We first show that for a range of structural asymmetric slope coefficients, asymmetry only arises in the KV's test for large asymmetric coefficients. As such, it could reject asymmetry of small degree even when it is imposed. Consider again the same 2-equation system:

$$\begin{aligned} x_t &= b_{10} + \sum_{i=1}^{p} b_{11,i} x_{t-i} + \sum_{i=1}^{p} b_{12,i} y_{t-i} + \varepsilon_{1,t} \\ y_t &= b_{20} + \sum_{i=0}^{p} b_{21,i} x_{t-i} + \sum_{i=0}^{p} g^{+}_{21,i} x^{+}_{t-i} + \sum_{i=1}^{p} b_{22,i} y_{t-i} + \varepsilon_{2,t}, \end{aligned} \tag{2.15}$$

where now we assume: $g^{+}_{21,i} = \phi\, b_{21,i}$

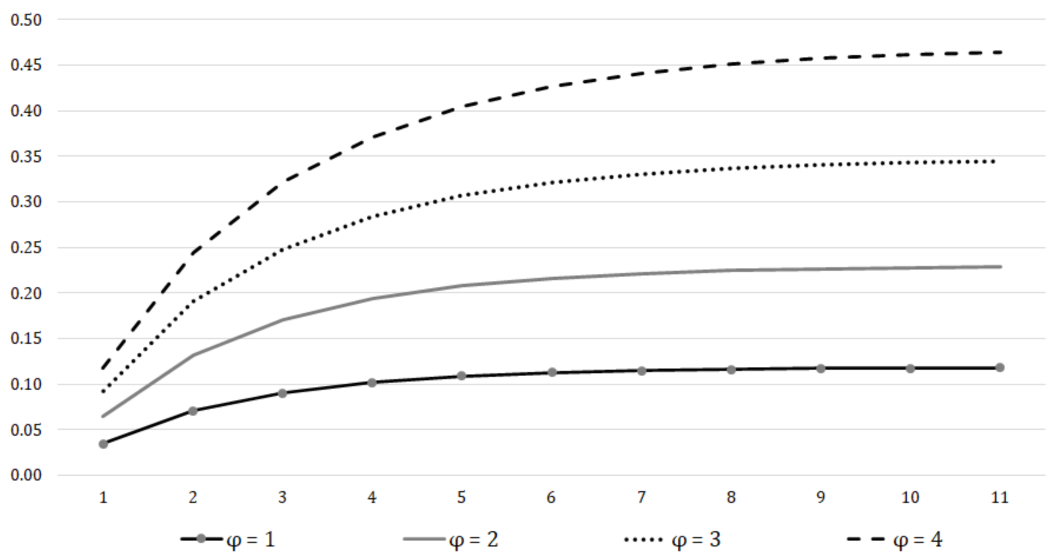
where $\phi > 0$ is the degree of asymmetry. To show the implication of a small ϕ in the tests proposed by KV, we simulate the system 2.15 for $N = 1,000$ and let ϕ assume integer values from 1 to 4. To do it, we arbitrarily chose the following parameters:

$$\begin{aligned} &b_{10} = b_{20} = 0; \quad b_{11,1} = 0.2; \quad b_{12,1} = 0.1; \\ &b_{21,0} = 0.2, \quad b_{21,1} = 0.1, \quad b_{22,1} = 0.5; \\ &g^{+}_{21,0} = \phi\, b_{21,0}; \quad g^{+}_{21,0} = \phi\, b_{21,0}. \end{aligned}$$

We set $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$ as random draws from $N(0,1)$. Figure H.2 below shows the structural asymptotic unconditional impulse response function after a 1 s.d. shock in the simulated system. We generated it by 1,000 simulations of time series with 1,000 observations each. For one realisation, the impulse response is the average between all responses from 1 to $N - h$. The degree of asymmetry is $I(h, \delta, \Omega^i) - I(h, -\delta, \Omega^i)$ and indicate positive asymmetry in all cases, even when the shock is said to be *small*.

[65] For traditional percentile techniques, see Efron (1992). For more on nonlinear impulse responses, see Koop, Pesaran and Potter (1996) and Potter (2000).

Figure B.4 - Degree of structural asymmetry
$I(h, \delta, \Omega^i) - I(h, -\delta, \Omega^i)$;
1,000 simulated time series of 1,000 observations each



To assess the chi-squared asymmetry test proposed by KV, we estimate the parameters by OLS and test nonlinear impulse responses with algorithm 2. Results are in table B.1 below and show that a combination of a small degree of asymmetry to small-sized shocks produce results that over-reject the null of asymmetric dynamic responses. Notice that for small lags, even the model with a fourfold asymmetry shows evidence of symmetric response. This result conflicts with the ones shown in table B.2, regarding the slope coefficients. If KV's argumentation follows that slope-based analysis do not tell much about the dynamics and tend to overstate asymmetry, their test based on impulse responses could overstate symmetry in models with subtle asymmetric parameters, especially after shocks of typical size (1 s.d.).

Table B.2 - Test of symmetric slope coefficient

$$H_0 : g_{21,0} = ... = g_{21,p} = 0$$

	F	**Marginal Significance Level**
$\phi = 1$	9.11	0.0001
$\phi = 2$	8.43	0.0002
$\phi = 3$	31.99	0.0000
$\phi = 4$	67.70	0.0000

This test is the same as the one reported in KV's Table 2.

Table B.1 - Testing the symmetry of the dynamic response

$\phi = [1, 2, 3, 4]$; $H = 10$ and shock sizes from 1 to 4 standard deviations

$\boldsymbol{\phi = 1}$	1 s.d.	2 s.d.	3 s.d.	4 s.d.	$\boldsymbol{\phi = 3}$	1 s.d.	2 s.d.	3 s.d.	4 s.d.
1	0.104	**0.001**	**0.001**	**0.001**	1	0.185	**0.000**	**0.000**	**0.000**
2	0.217	**0.001**	**0.001**	**0.001**	2	0.377	**0.000**	**0.000**	**0.000**
3	0.125	**0.004**	**0.002**	**0.002**	3	0.082	**0.000**	**0.000**	**0.000**
4	0.071	**0.008**	**0.005**	**0.005**	4	**0.013**	**0.000**	**0.000**	**0.000**
5	0.115	**0.017**	**0.011**	**0.009**	5	**0.026**	**0.000**	**0.000**	**0.000**
6	0.165	**0.026**	**0.018**	**0.017**	6	**0.032**	**0.000**	**0.000**	**0.000**
7	0.201	**0.044**	**0.032**	**0.030**	7	0.055	**0.000**	**0.000**	**0.000**
8	0.236	0.058	**0.050**	**0.046**	8	0.080	**0.000**	**0.000**	**0.000**
9	0.307	0.084	0.071	0.067	9	0.114	**0.000**	**0.000**	**0.000**
10	0.384	0.093	0.103	0.100	10	0.162	**0.000**	**0.000**	**0.000**
$\boldsymbol{\phi = 2}$	1 s.d.	2 s.d.	3 s.d.	4 s.d.	$\boldsymbol{\phi = 4}$	1 s.d.	2 s.d.	3 s.d.	4 s.d.
1	0.237	**0.045**	**0.046**	**0.047**	1	0.056	**0.000**	**0.000**	**0.000**
2	0.472	**0.001**	**0.000**	**0.000**	2	0.161	**0.000**	**0.000**	**0.000**
3	0.590	**0.002**	**0.001**	**0.001**	3	0.194	**0.000**	**0.000**	**0.000**
4	0.479	**0.005**	**0.002**	**0.002**	4	**0.014**	**0.000**	**0.000**	**0.000**
5	0.611	**0.010**	**0.004**	**0.004**	5	**0.028**	**0.000**	**0.000**	**0.000**
6	0.730	**0.018**	**0.008**	**0.007**	6	**0.040**	**0.000**	**0.000**	**0.000**
7	0.794	**0.031**	**0.013**	**0.011**	7	0.066	**0.000**	**0.000**	**0.000**
8	0.835	**0.047**	**0.020**	**0.017**	8	0.061	**0.000**	**0.000**	**0.000**
9	0.773	**0.025**	**0.015**	**0.016**	9	0.057	**0.000**	**0.000**	**0.000**
10	0.823	**0.037**	**0.023**	**0.027**	10	0.085	**0.000**	**0.000**	**0.000**

10,000,000 iterations ($LxNxM$).

As asymmetric exchange rate pass-though tends to be very subtle in the slope parameters, we raise attention to the possibility of over-rejection of dynamic asymmetry with KV's test. Notwithstanding, compared to the oil shock literature, it seems that the exchange rate pass-through phenomena has a better asymmetric representation.

Appendix C - Correlation Matrix

	p^m	p^p	p^c	w^m	w^x		p^{comm}	neer	y^x		y^m		r
	IPI	**PPI**	**CPI**	**W**	**FPPI**	**OIL**	**CEPI**	**NEER**	**KILIAN**	**WGDP**	**PROD**	**GAP**	**R**
IPI	1.000												
CPI	0.897	1.000											
PPI	0.901	0.986	1.000										
W	0.471	0.731	0.714	1.000									
FPPI	0.773	0.951	0.946	0.834	1.000								
OIL	0.390	0.689	0.625	0.824	0.809	1.000							
CEPI	0.347	0.654	0.582	0.795	0.775	0.992	1.000						
NEER	-0.623	-0.268	-0.266	0.301	-0.013	0.343	0.358	1.000					
KILIAN	-0.366	-0.222	-0.315	-0.130	-0.162	0.181	0.252	0.212	1.000				
WGDP	-0.117	-0.012	0.031	0.025	0.078	0.159	0.184	0.178	0.269	1.000			
PROD	0.340	0.647	0.569	0.842	0.749	0.938	0.942	0.342	0.201	0.120	1.000		
GAP	0.203	0.100	0.064	-0.001	0.135	0.105	0.106	-0.131	-0.132	-0.263	0.117	1.000	
R	-0.526	-0.715	-0.728	-0.792	-0.796	-0.721	-0.695	-0.177	0.100	-0.069	-0.665	0.165	1.000

Variables' codes are in table 1.

Appendix D - Trade Weights

Trade partner	**1999**	**2000**	**2001**	**2002**	**2003**	**2004**	**2005**	**2006**	**2007**	**2008**	**2009**	**2010**	**2011**	**2012**	**2013**	**2014**	**2015**	**2016**	**2017**	**2018**
Algeria	0.000	0.041	0.000	0.000	0.037	0.048	0.061	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Argentina	0.172	0.186	0.168	0.155	0.152	0.138	0.133	0.142	0.140	0.128	0.137	0.125	0.118	0.117	0.109	0.099	0.096	0.103	0.099	0.102
Canada	0.030	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Chile	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.051	0.047	0.038	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.033	0.036	0.000
China	0.000	0.000	0.036	0.051	0.070	0.092	0.114	0.141	0.170	0.194	0.193	0.222	0.229	0.244	0.247	0.261	0.288	0.264	0.286	0.319
France	0.058	0.051	0.056	0.058	0.058	0.057	0.058	0.050	0.048	0.045	0.044	0.042	0.038	0.042	0.043	0.040	0.042	0.042	0.039	0.036
Germany	0.138	0.120	0.131	0.145	0.137	0.126	0.131	0.115	0.117	0.116	0.120	0.109	0.106	0.101	0.101	0.097	0.097	0.103	0.097	0.097
India	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.037	0.042	0.000	0.042	0.046	0.000	0.000	0.000	0.000
Italy	0.076	0.059	0.059	0.058	0.057	0.051	0.049	0.046	0.045	0.045	0.045	0.042	0.043	0.044	0.044	0.044	0.044	0.042	0.041	0.041
Japan	0.076	0.080	0.083	0.077	0.082	0.071	0.073	0.068	0.062	0.066	0.065	0.061	0.055	0.055	0.047	0.041	0.046	0.040	0.039	0.040
Mexico	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.034	0.000	0.000	0.043	0.000	0.000	0.041	0.040	0.044	0.045
Nigeria	0.000	0.000	0.037	0.036	0.050	0.087	0.056	0.069	0.071	0.065	0.058	0.051	0.059	0.057	0.064	0.066	0.043	0.000	0.000	0.000
Rep. of Korea	0.031	0.039	0.043	0.035	0.000	0.043	0.050	0.055	0.046	0.052	0.059	0.073	0.070	0.065	0.063	0.060	0.051	0.062	0.055	0.049
Spain	0.034	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
United Kingdom	0.036	0.034	0.034	0.044	0.039	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
USA	0.349	0.354	0.353	0.342	0.318	0.286	0.275	0.263	0.255	0.250	0.246	0.237	0.239	0.232	0.240	0.247	0.251	0.272	0.263	0.270
Venezuela	0.000	0.036	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

Yearly trade weights used in the construction of the series FPPI (foreign PPI) and CEPI (coomodity export price index).
$FPPI_t = \sum_{i=1}^{c} w_{iy} PPI_{it}$, where w_{iy} identifies the country with the i-th larger weight on Brazilian imports, at year y.

Appendix E - Robustness and alternative models

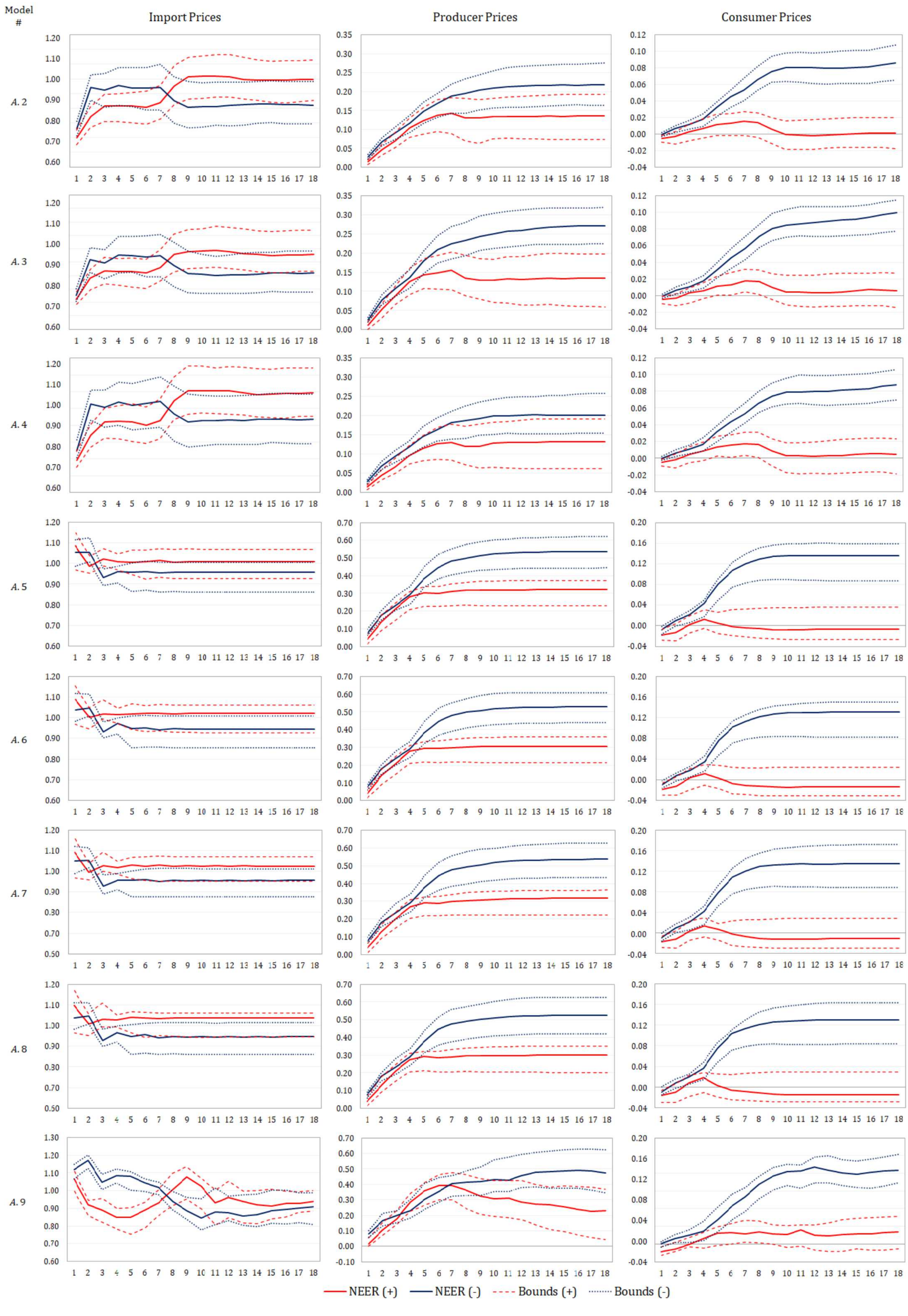
Baseline

Model	*Covariates*	*Empirical strategy*	Lags
A.1	$p^w; y^m$	SE in differences	8
	ECM		
B.1	$w^x + e^{(+)} + e^{(-)}$	SE with ECM term	AIC, $p_{max} = 8$
B.2	$(w^x - e) + (p^w - e)$	SE with ECM term	AIC, $p_{max} = 8$
B.3	$(w^x - e)^{(+)} + (w^x - e)^{(-)}$ $+(p^w - e)^{(+)} + (p^w - e)^{(-)}$	SE with ECM term	AIC, $p_{max} = 8$
	Ordering		
C.1	a) p^w, e, y^m, p	4-variable VAR	4, 8, and 12
C.2	b) p^w, y^m, p, e	4-variable VAR	4, 8, and 12
C.3	e) p^w, e, y^m, r, p^m, p^p, p^c	7-variable VAR	4, 8, and 12

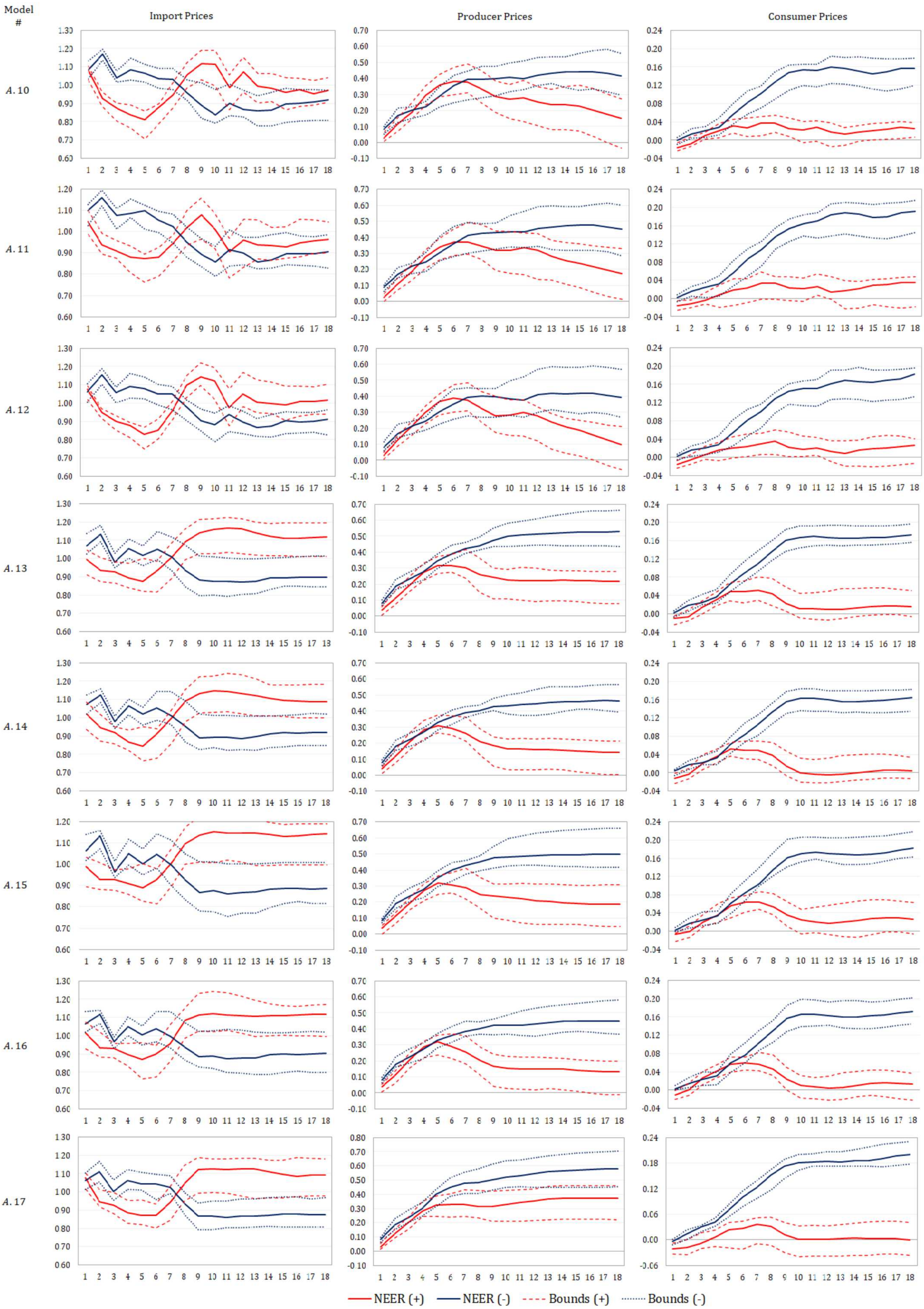
Robustness

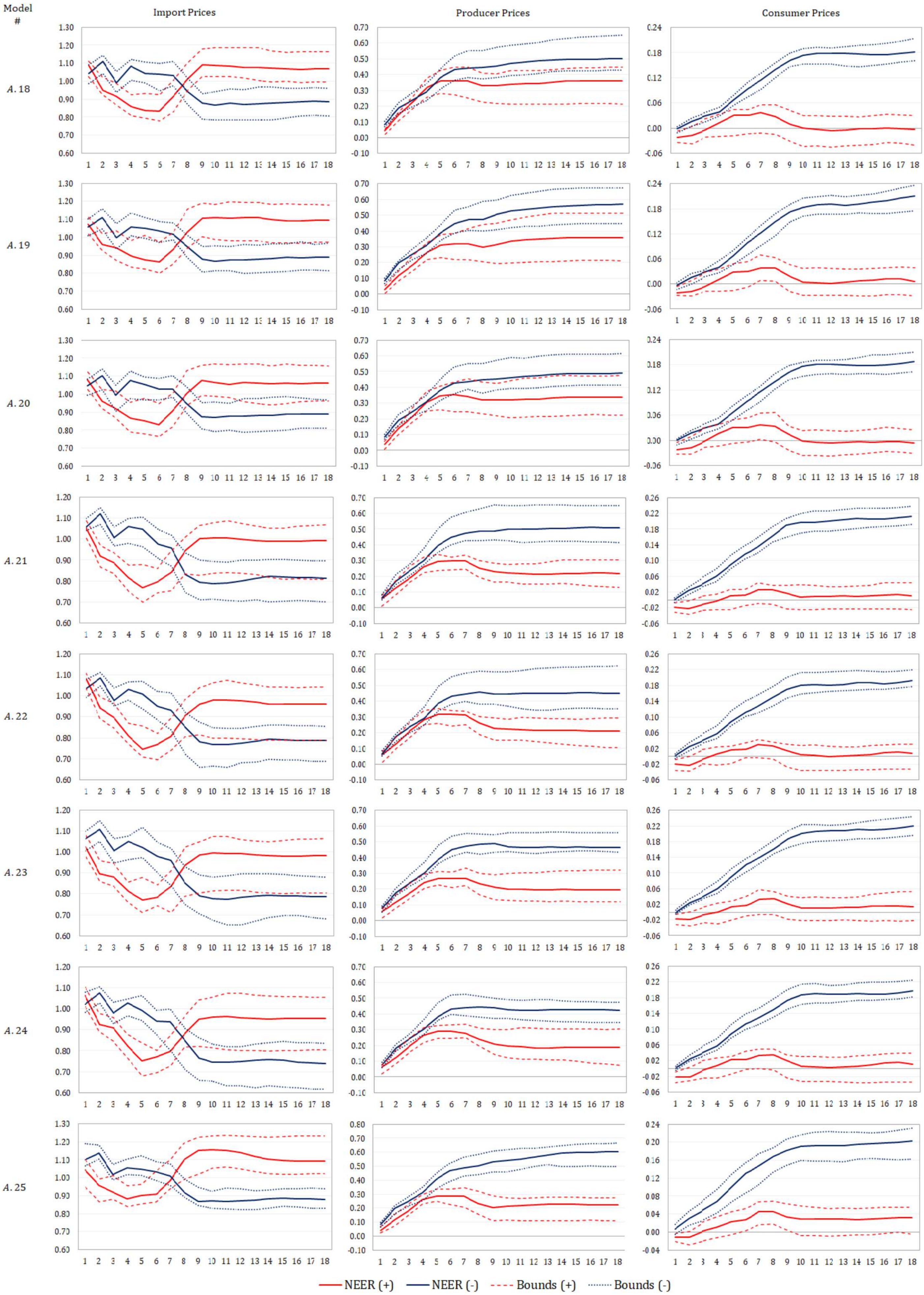
Model	*What changed*	*Model*	*What changed*
A.2	$p^w; y^m; w^m$	B.4	B.1 with fixed lags (8)
A.3	$p^w; y^m; y^x$	B.5	B.2 with fixed lags (8)
A.4	$p^w; w^m; y^m; y^x$	B.6	B.3 with fixed lags (8)
A.5	A.1 with 4 lags	B.7	B.1 with fixed lags (12)
A.6	A.2 with 4 lags	B.8	B.2 with fixed lags (12)
A.7	A.3 with 4 lags	B.9	B.3 with fixed lags (12
A.8	A.4 with 4 lags	B.10	B.1 with FPPI instead of OIL
A.9	A.1 with 12 lags	B.11	B.2 with FPPI instead of OIL
A.10	A.2 with 12 lags	B.12	B.3 with FPPI instead of OIL
A.11	A.3 with 12 lags	C.4	Ordering c)p^w, e, y^m, r, p
A.12	A.4 with 12 lags	C.5	Ordering d)p^w, y^m, r, p, e
A.13	A.1 with the addition of r	C.6	Ordering f)p^w, y^m, r, p^m, p^p, p^c, e
A.14	A.2 with the addition of r	C.7	C.2 with ordering 2
A.15	A.3 with the addition of r	C.8	C.3 with ordering 2
A.16	A.4 with the addition of r	C.9	C.2 with ordering 3
A.17	A.1 with CIPI instead of OIL	C.10	C.3 with ordering 3
A.18	A.2 with CIPI instead of OIL	C.11	C.2 with 4 lags
A.19	A.3 with CIPI instead of OIL	C.12	C.3 with 4 lags
A.20	A.4 with CIPI instead of OIL	C.13	C.2 with 12 lags
A.21	A.1 with FPPI instead of OIL	C.14	C.3 with 12 lags
A.22	A.2 with FPPI instead of OIL	C.15	C.2 with CIPI instead of OIL
A.23	A.3 with FPPI instead of OIL	C.16	C.3 with CIPI instead of OIL
A.24	A.4 with FPPI instead of OIL	C.17	C.2 with FPPI instead of OIL
A.25	A.1 with PROD instead of GAP	C.18	C.3 with FPPI instead of OIL
A.26	A.2 with PROD instead of GAP	C.19	C.2 with PROD instead of GAP
A.27	A.3 with PROD instead of GAP	C.20	C.3 with PROD instead of GAP
A.28	A.4 with PROD instead of GAP		
A.29	A.3 with WGDP instead of KILIAN		
A.30	A.4 with WGDP instead of KILIAN		

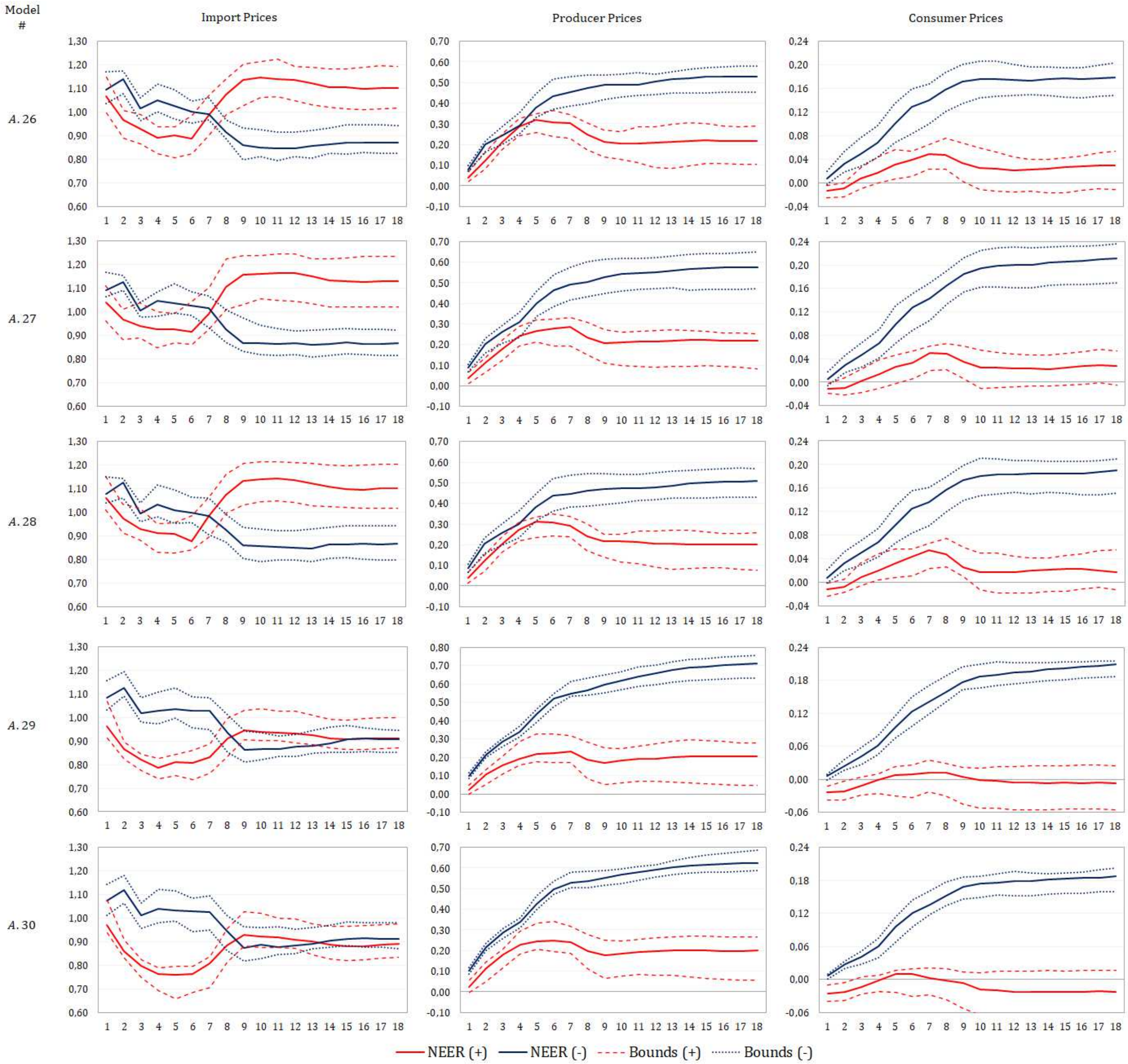
Notes:
i) The baseline models are in the text;
ii) The robustness models are below;
iii) We report only the IRFs. All the coefficients and further diagnostics are available from the authors.
iv) For simplicity of notation, we define *model A:* single equation; *model B:* single equation with ECM term; *model C:* VAR model.

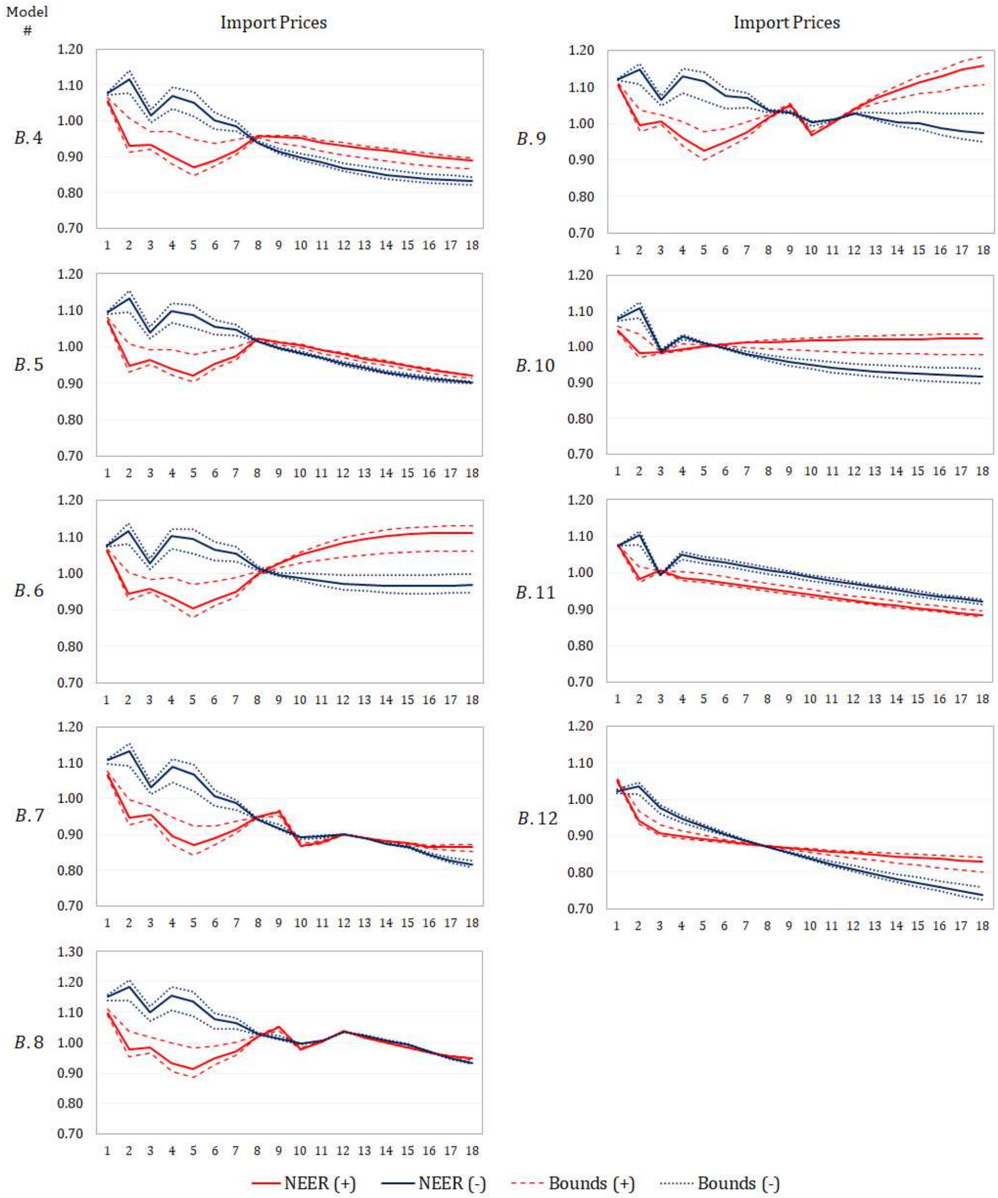


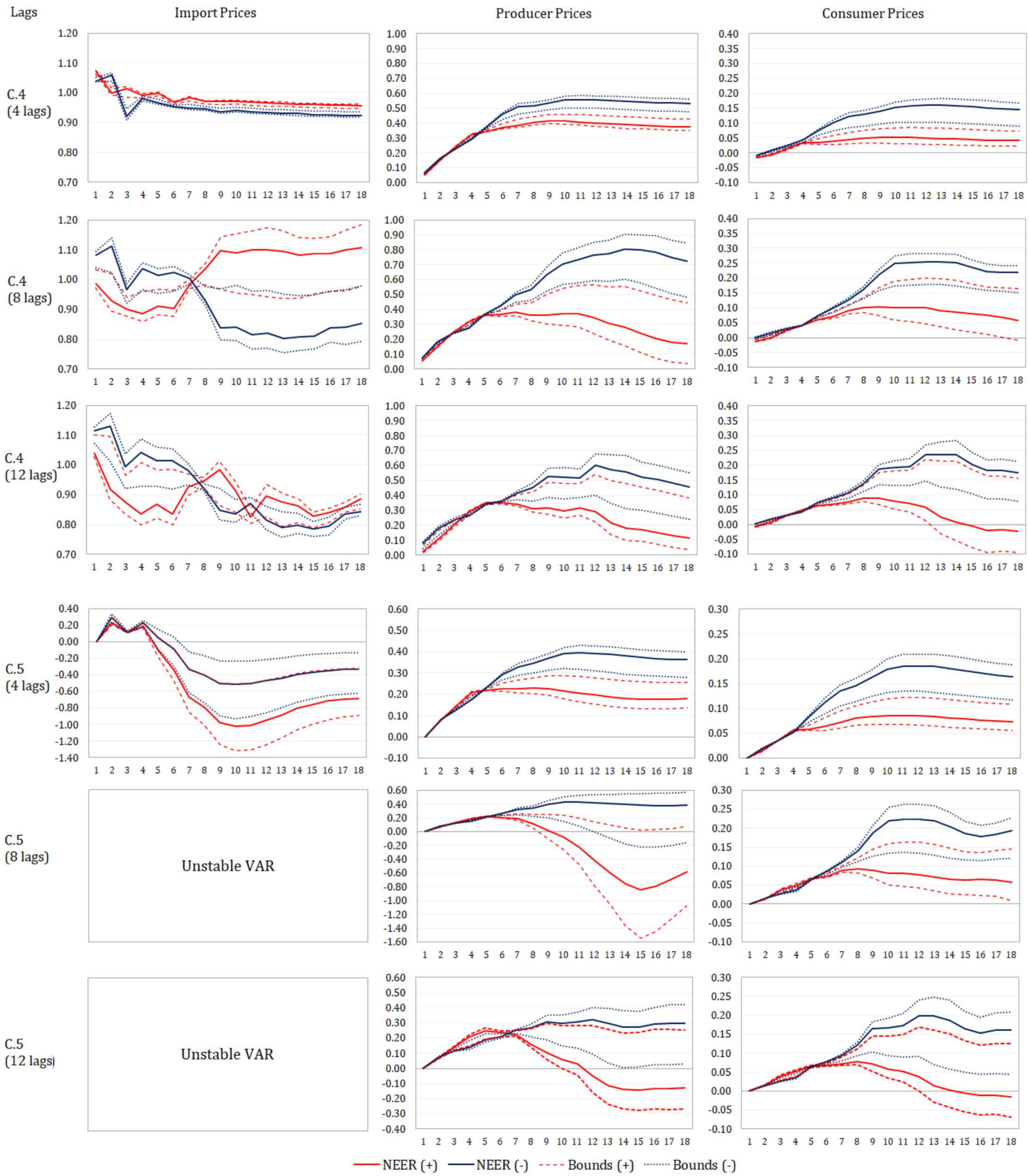
Model #
Import Prices
Producer Prices
Consumer Prices
A.2
A.3
A.4
A.5
A.6
A.7
A.8
A.9
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)

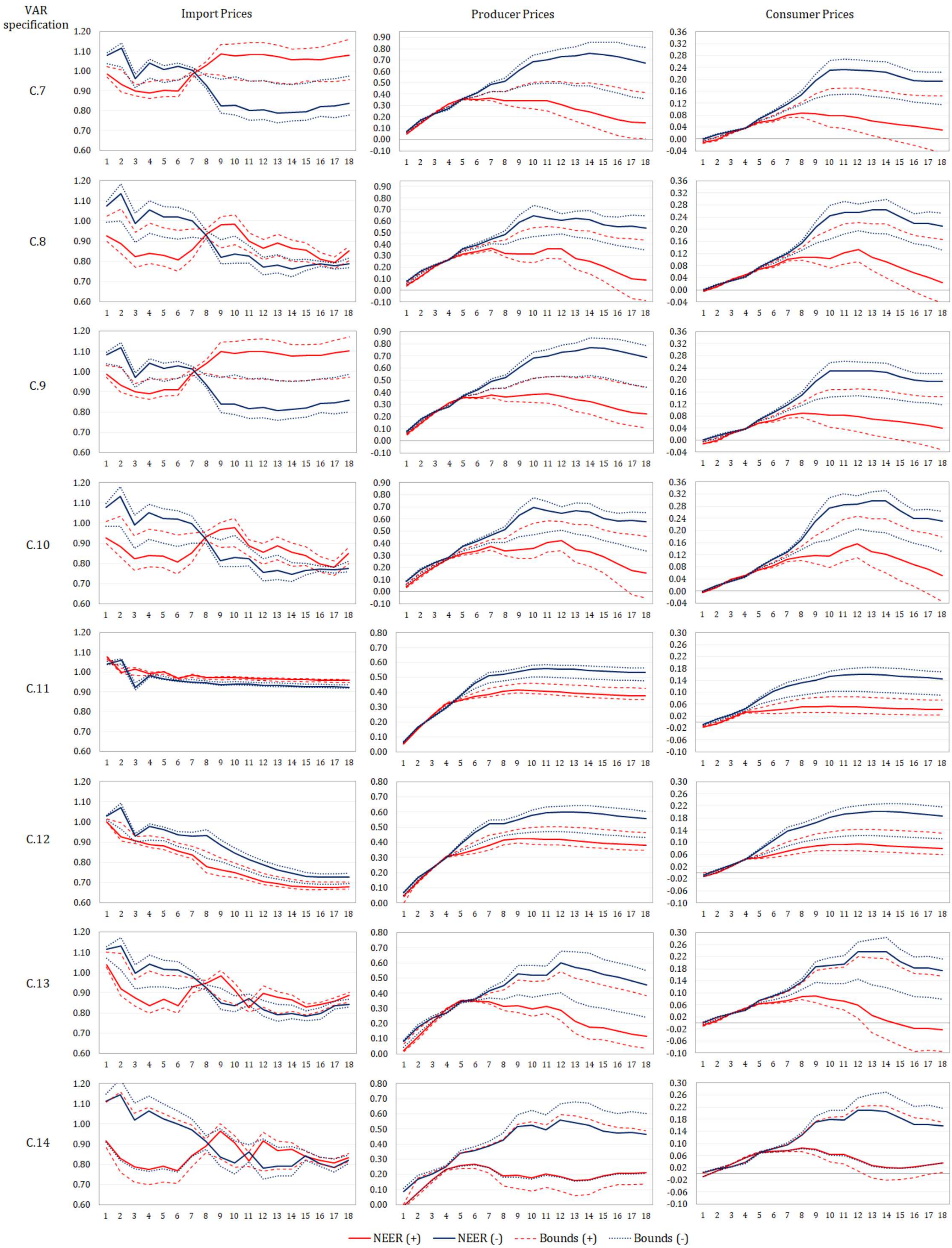

Model #
Import Prices
Producer Prices
Consumer Prices
A.10
A.11
A.12
A.13
A.14
A.15
A.16
A.17
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)

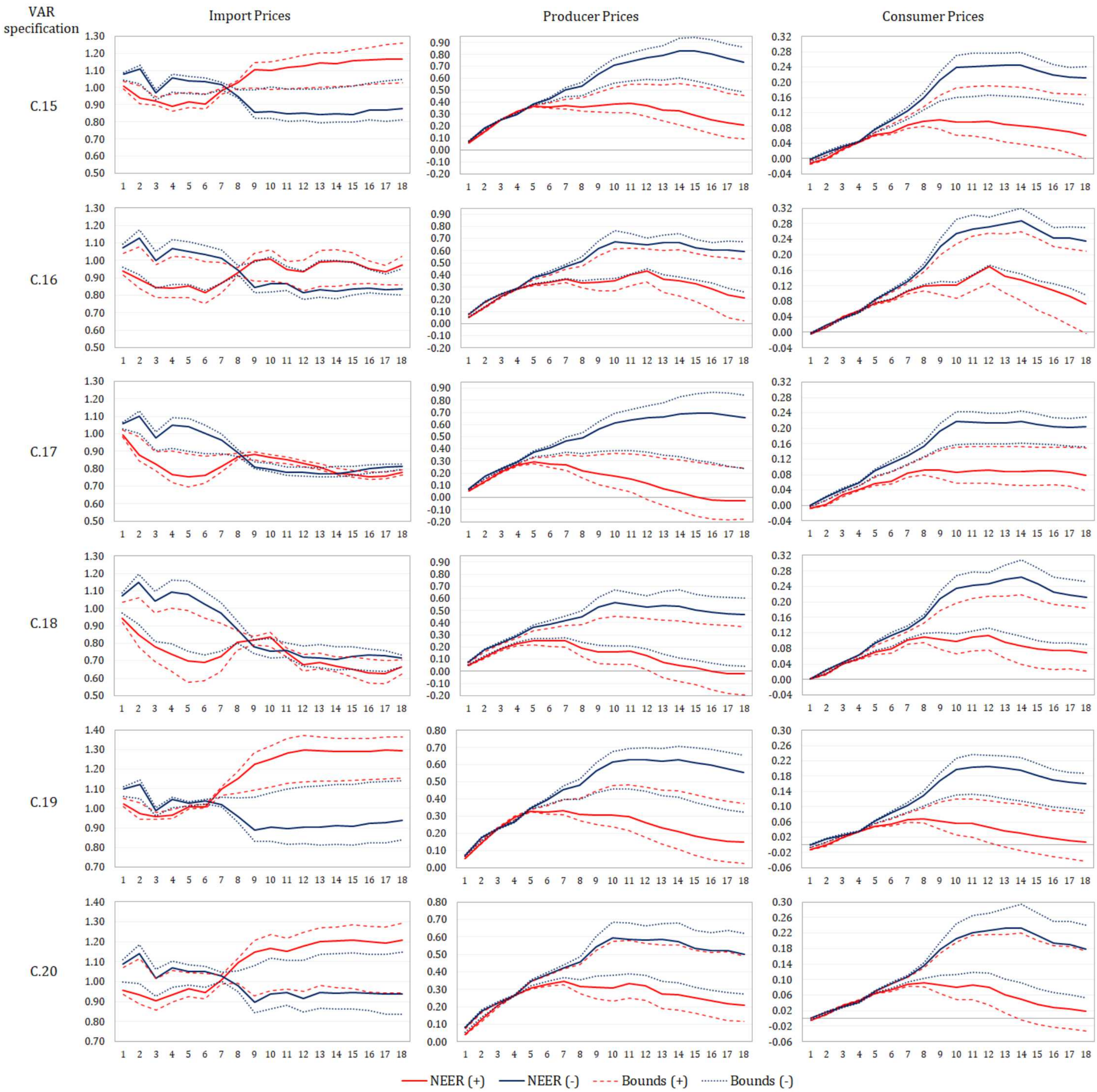

Model
#
Import Prices
Producer Prices
Consumer Prices
A.18
A.19
A.20
A.21
A.22
A.23
A.24
A.25
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)


Model #
Import Prices
Producer Prices
Consumer Prices
A. 26
A. 27
A. 28
A. 29
A. 30
NEER (+) NEER (-) Bounds (+) Bounds (-)


Model
#
Import Prices
Import Prices
B.4
B.5
B.6
B.7
B.8
B.9
B.10
B.11
B.12
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)


Lags
Import Prices
Producer Prices
Consumer Prices
C.4
(4 lags)
C.4
(8 lags)
C.4
(12 lags)
C.5
(4 lags)
C.5
(8 lags)
Unstable VAR
C.5
(12 lags)
Unstable VAR
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)


VAR specification
Import Prices
Producer Prices
Consumer Prices
C.7
C.8
C.9
C.10
C.11
C.12
C.13
C.14
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)


VAR specification
Import Prices
Producer Prices
Consumer Prices
C.15
C.16
C.17
C.18
C.19
C.20
NEER (+)
NEER (-)
Bounds (+)
Bounds (-)

Appendix F - NARDL models

F.1 - Following the baseline specification

	Imports				Producer				Consumer			
Model:	A.1	A.2	A.3	A.4	A.1	A.2	A.3	A.4	A.1	A.2	A.3	A.4
Cointegration												
PSS(2001) tests												
F_{yx}	7.400‡	6.562‡	6.518‡	6.318‡	2.855	2.520	2.659	2.404	3.056	2.657	6.032‡	5.733‡
$F_{xy(1)}$	3.593$^{\mp}$	3.759$^{\mp}$	2.988	2.924	4.164$^{\mp}$	3.802†	4.462†	4.106†	3.504	3.505$^{\mp}$	4.070†	3.414$^{\mp}$
$F_{xy(2)}$	1.981	3.651$^{\mp}$	1.796	3.235	3.732$^{\mp}$	3.239	6.841	2.792	2.110	1.777	2.101	1.804
$F_{xy(3)}$	8.020‡	7.323‡	7.140‡	9.620‡	1.646‡	1.374	3.915	2.049	1.867	1.491	2.092	1.788
$F_{xy(4)}$	12.852‡	11.138‡	5.628‡	4.980‡	10.719‡	8.951‡	3.271‡	3.322‡	8.350‡	6.540‡	6.336‡	8.687‡
$F_{xy(5)}$	-	8.010‡	3.860$^{\mp}$	6.495‡	-	8.644‡	2.395†	5.926‡	-	5.456‡	4.017†	5.465‡
$F_{xy(6)}$	-	-	-	3.267	-	-	-	7.402$^{\mp}$	-	-	-	3.494$^{\mp}$
t_{bounds}	-	-	-	-	-3.037	-3.114	-3.129	3.322	-	-	-	-
Johansen test												
Nº of coint. vectors (Trace)	2	2	2	2	2	2	1	2	1	1	1	1
Auxiliary tests												
F_1	8.679	7.542	7.436	7.124	2.756	2.375	2.539	2.261	3.927	3.195	7.109	6.599
(p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.029)	(0.040)	(0.029)	(0.039)	(0.004)	(0.008)	(0.000)	(0.000)
F_2	5.221	3.347	3.205	3.238	1.362	1.109	1.502	1.294	1.621	1.469	2.250	2.940
(p-value)	(0.023)	(0.037)	(0.042)	(0.023)	(0.244)	(0.332)	(0.225)	(0.277)	(0.106)	(0.232)	(0.1079)	(0.034)
Error Correction	-0.165	-0.172	-0.157	-0.190	-0.036	-0.037	-0.038	-0.031	-0.051	-0.050	-0.060	-0.058
t-statistic	-6.740	-6.872	-6.849	-7.227	-3.814	-3.935	-4.042	-3.838	-4.332	-4.373	-6.587	-6.882
BDM critical value	-4.380	-4.550	-4.380	-4.750	-4.380	-4.550	-4.380	-4.750	-4.380	-4.550	-4.380	-4.750

NARDL models using the baseline covariates should be carefully applied. For the single-equation approach to be valid, the regressors should be weakly exogenous (MCNOWN; SAM; GOH, 2018). This means that F_{yx} should be significant and $F_{xy(i)}$, $i = [1, ..., k]$ should not. A similar way of testing this is assessing the number of cointegration vectors with traditional techniques (JOHANSEN; JUSELIUS, 1990). Results show that, while holding significant long-run terms, the import price models have multiple vectors of cointegration, which call upon the use of VEC models. The models for producer and consumer prices generally do not reject the null of no cointegration.

Levels of significance: ‡ indicates 1%; † indicates 5%; ∓ indicates 10%;

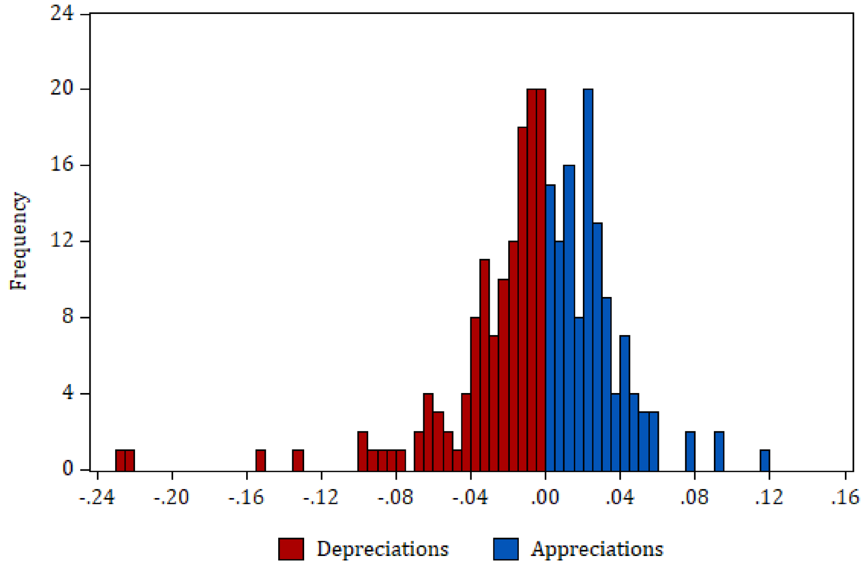
F.2 - Valid NARDL models

	Imports		
Model:	B.1	B.2	B.3
Cointegration			
$F_{PSS}(2001)$ *test*			
F_{yx}	6.220†	7.944‡	9.971‡
$F_{xy(1)}$	3.746	3.014	1.509
$F_{xy(2)}$	2.229	5.268	2.197
$F_{xy(3)}$	7.635‡	-	-
t_{bounds}	-4.661†	-4.380†	-5.481‡
Johansen test			
Trace	1	1	1
Auxiliary tests			
F test (independent)	7.129	11.010	11.318
(p-value)	(0.000)	(0.000)	(0.000)
Error Correction	-0.159	-0.160	-0.172
t-statistic	-5.022	-4.905	-7.124
BDM critical value	-4.190	-3.980	-4.380
Diagnostics			
Serial Correlation			
Breusch-Godfrey LM Test	0.718	0.836	3.644
(p-value)	(0.542)	(0.504)	(0.028)
Heteroskedasticity			
BPG	1.276	1.414	1.809
(p-value)	0.223	0.120	0.039
Stability			
CUSUM	Stable	Stable	Stable
CUSUMSQ	Stable	Stable	Stable

The NARDL models with price homogeneity restrictions are valid cointegrating models because there is only one cointegrating vector confirmed by both F_{PSS} and Johansen tests in the first panel. Moreover, the auxiliary tests show that the significance of the tests just mentioned does not come from the deterministic trend, i.e., the F-statistic is high for the set of covariates in the EC term. The Error Correction term is also negative and highly significant, as the rule of thumb states.
Notes:
i) BDM stands for Banerjee, Dolado and Mestre (1998);
ii) The null for the LM test is is that there is no serial correlation.
iii) BPG stands for Breusch-Pagan-Godfrey. The null is
iv) CUSUM and CUSUMSQ stands for cumulative sum and cumulative sum of squares, respectively. See Turner (2010) for more details.

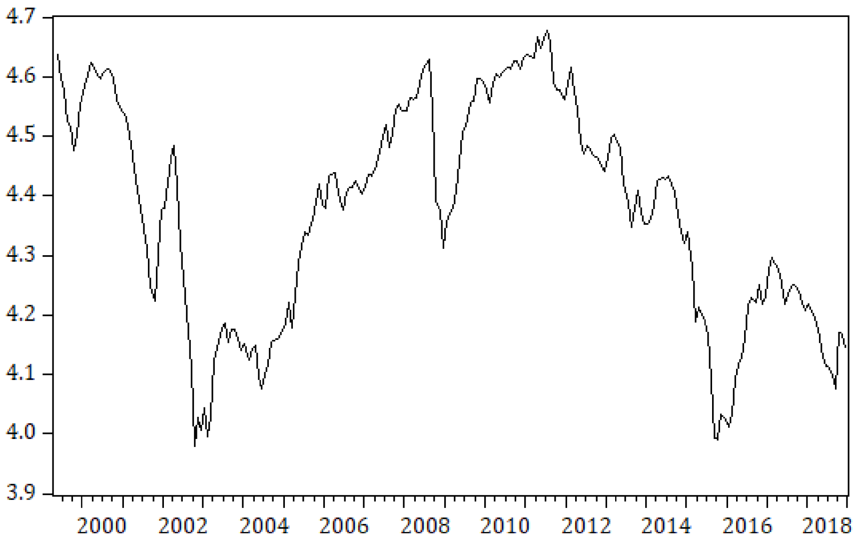
Appendix G - Nominal Effective Exchange Rate - Plots

Figure G.1 - Histogram of Δe_t



In our sample, the change of the nominal exchange rate has negative median and a left tail, i.e., depreciations are more frequent and in average larger than appreciations. Throughout the sample, there are more prominent depreciation episodes, especially in global crisis that hit emerging economies harder and in political crisis with outcomes on the country's risk.

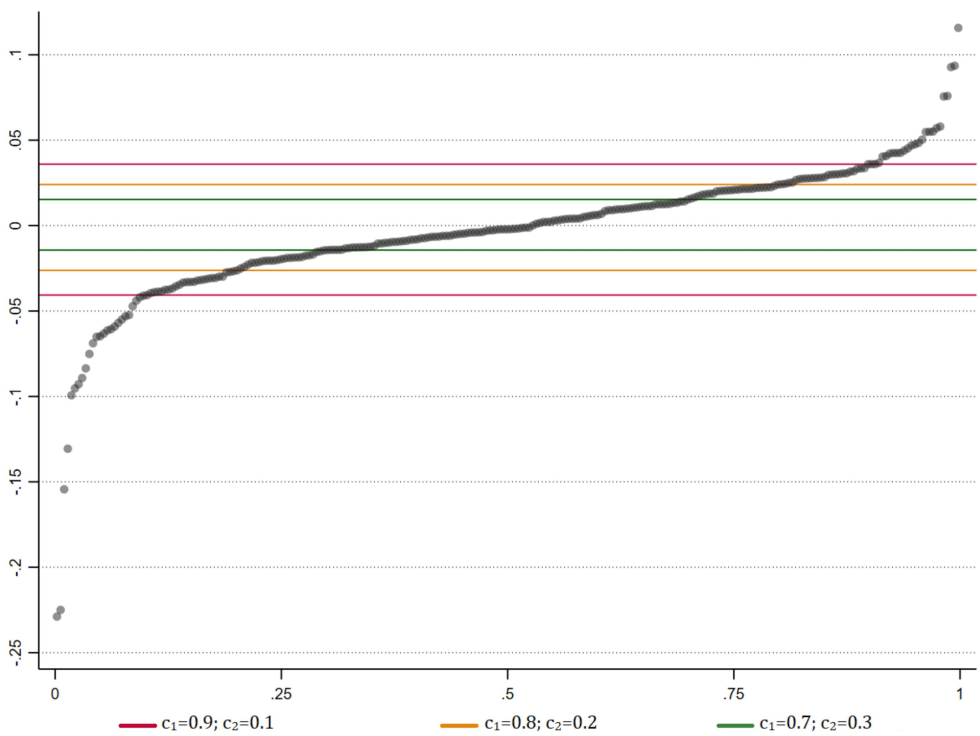
Figure G.2 - Log of the Nominal Effective Exchange Rate (NEER)



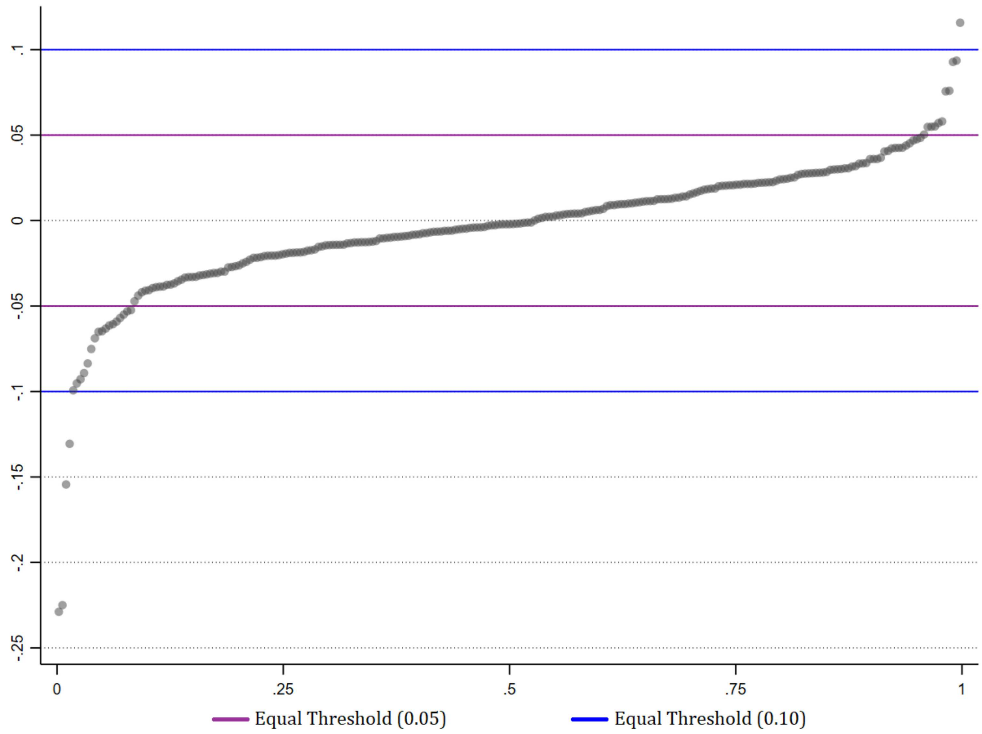
An increase is an appreciation.
Source: BIS

Figure G.3 - Quantiles of $\Delta neer_t$

Threshold Quantile



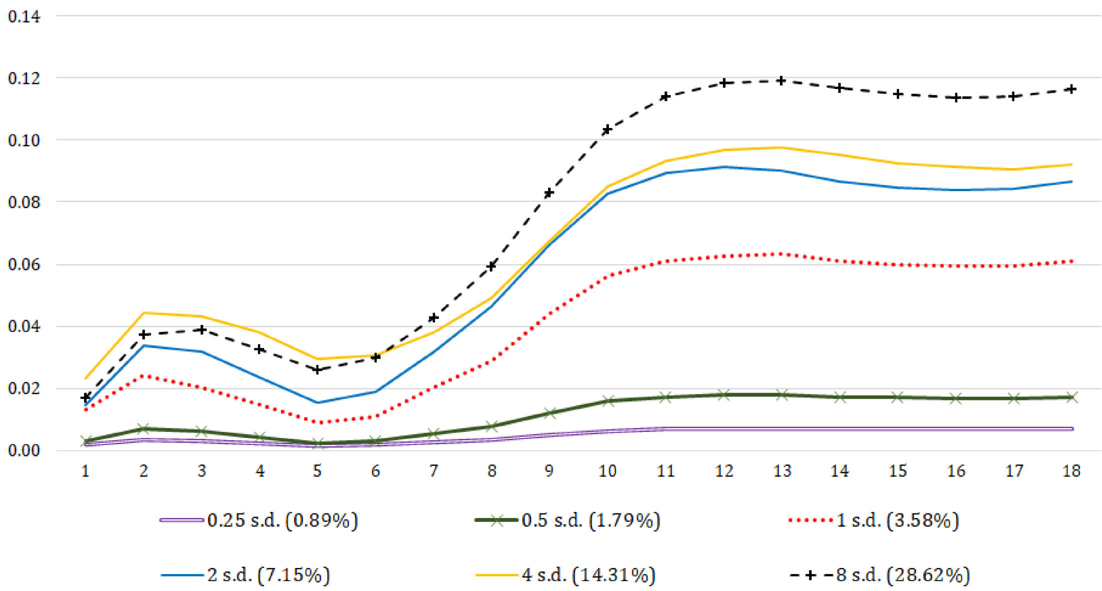
Equal Quantiles



The quantile distribution of the nominal effective exchange rate and the thresholds used in section 1.6.4.

Appendix H - Degree of asymmetry by size of the shock

Model A.1 with the NEER changes decomposed by size: $\Delta e_t^{large(+)}$, $\Delta e_t^{large(-)}$, and Δe_t^{small}
$c_1 = 0.9$; $c_2 = 0.1$;
Dependent variable: CPI



One standard deviation change in the error term of the AR(1) is equivalent to a 3.58% change in the NEER. The figure shows that the larger the size of the shock, the higher the degree of asymmetry, given by $ERPT_t^{(-)} - ERPT_t^{(+)}$. When the shock is too low (e.g. < 1 s.d.), it is unable to activate the size dummy, which has threshold values of 3.6% and 4%. As the shock increases, this dummy is activated. However, as discussed in the text, the exchange rage path after a large shock might not be well represented by an AR(1), because these shocks tend to be an one-off event, not followed by high changes in the following months. Moreover, the degree of asymmetry growing with the size of the shock is an statistic feature of the nonlinear impulse responses as discussed in Kilian and Vigfusson (2011a). Even without imposing structural size asymmetry in economic terms, the model could produce size asymmetry. Notice that this is true even after dividing the response by the size of the shock (such that the underlying response is in terms of "one standard deviation") and dividing it by the exchange rate response (so that the interpretation is an elasticity, see equation 1.16.).

Appendix I - Testing for asymmetry: further models

	AR(1); p=4								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
H	1	2	4	1	2	4	1	2	4
1	0.89	0.69	0.61	0.52	0.58	0.60	0.57	0.51	0.51
2	0.36	0.34	0.35	0.80	0.86	0.87	0.64	0.64	0.64
3	0.15	*0.10*	*0.10*	0.82	0.78	0.77	0.76	0.71	0.70
4	0.24	0.16	0.16	0.92	0.89	0.89	0.81	0.77	0.76
5	0.36	0.25	0.25	0.15	*0.07*	*0.07*	*0.08*	**0.01**	**0.01**
6	0.47	0.36	0.35	0.19	0.11	0.11	0.11	**0.02**	**0.02**
7	0.59	0.46	0.46	0.27	0.17	0.17	0.17	**0.04**	**0.04**
8	0.69	0.57	0.57	0.36	0.23	0.23	0.24	*0.06*	*0.06*
9	0.78	0.67	0.67	0.46	0.31	0.31	0.32	*0.10*	*0.09*
10	0.84	0.75	0.75	0.55	0.40	0.40	0.41	0.14	0.13
11	0.89	0.81	0.81	0.64	0.48	0.49	0.50	0.18	0.18
12	0.92	0.86	0.86	0.70	0.55	0.55	0.57	0.24	0.23
13	0.95	0.90	0.90	0.77	0.62	0.62	0.64	0.29	0.28
14	0.97	0.93	0.93	0.83	0.70	0.70	0.68	0.34	0.33
15	0.98	0.96	0.95	0.87	0.76	0.76	0.74	0.41	0.40
16	0.99	0.97	0.97	0.90	0.80	0.80	0.80	0.47	0.46
17	0.99	0.98	0.98	0.93	0.85	0.85	0.84	0.54	0.53
18	1.00	0.99	0.99	0.95	0.88	0.88	0.88	0.60	0.59

	AR(1); p=8								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	1	2	4	1	2	4	1	2	4
H	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
1	0.40	0.41	0.46	0.39	0.39	0.39	0.37	0.32	0.31
2	0.19	*0.07*	*0.07*	0.69	0.69	0.69	0.50	0.47	0.47
3	0.15	*0.07*	*0.07*	0.51	0.47	0.46	0.62	0.55	0.54
4	0.20	*0.07*	*0.08*	0.63	0.59	0.58	0.65	0.59	0.58
5	0.30	0.12	0.13	0.20	0.10	*0.10*	0.50	0.42	0.42
6	0.40	0.19	0.20	0.22	0.12	0.12	0.34	0.23	0.24
7	0.42	0.23	0.24	0.31	0.18	0.18	0.45	0.33	0.33
8	*0.09*	**0.02**	**0.03**	0.35	0.20	0.19	0.24	*0.06*	*0.05*
9	*0.07*	**0.01**	**0.02**	0.45	0.27	0.27	0.24	**0.05**	**0.05**
10	*0.10*	**0.02**	**0.03**	0.54	0.36	0.35	0.32	*0.07*	*0.07*
11	0.14	**0.04**	**0.04**	0.63	0.44	0.43	0.38	0.10	*0.10*
12	0.19	*0.06*	*0.07*	0.71	0.53	0.51	0.46	0.14	0.14
13	0.24	*0.08*	*0.09*	0.78	0.60	0.59	0.54	0.18	0.18
14	0.30	0.11	0.13	0.83	0.68	0.67	0.62	0.23	0.23
15	0.36	0.14	0.17	0.88	0.74	0.73	0.69	0.29	0.28
16	0.43	0.18	0.21	0.91	0.80	0.79	0.75	0.36	0.35
17	0.50	0.23	0.26	0.94	0.85	0.84	0.80	0.42	0.41
18	0.57	0.29	0.32	0.96	0.88	0.88	0.84	0.47	0.46

	4-variable VAR (ordering a); p=4								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
H	1	2	4	1	2	4	1	2	4
1	0.92	0.64	0.58	0.38	0.55	0.57	0.34	0.42	0.43
2	0.30	0.35	0.35	0.67	0.83	0.84	0.53	0.55	0.55
3	*0.08*	*0.07*	*0.08*	0.61	0.71	0.71	0.55	0.62	0.63
4	0.15	0.14	0.14	0.76	0.84	0.84	0.66	0.71	0.73
5	0.24	0.21	0.23	*0.10*	*0.07*	*0.06*	*0.06*	**0.01**	**0.01**
6	0.34	0.31	0.33	0.14	0.10	0.10	0.10	**0.03**	**0.03**
7	0.45	0.42	0.43	0.20	0.16	0.16	0.16	**0.04**	**0.05**
8	0.54	0.52	0.54	0.26	0.21	0.21	0.22	*0.07*	*0.07*
9	0.61	0.60	0.61	0.33	0.28	0.28	0.27	0.11	0.11
10	0.70	0.69	0.70	0.37	0.36	0.36	0.34	0.15	0.16
11	0.76	0.76	0.78	0.43	0.44	0.45	0.33	0.19	0.21
12	0.82	0.83	0.84	0.50	0.53	0.53	0.41	0.25	0.27
13	0.86	0.87	0.89	0.40	0.60	0.61	0.28	0.30	0.33
14	0.90	0.91	0.92	0.47	0.67	0.69	0.33	0.36	0.39
15	0.93	0.94	0.95	0.55	0.74	0.75	0.33	0.43	0.47
16	0.95	0.96	0.97	0.62	0.78	0.80	0.38	0.50	0.54
17	0.97	0.97	0.98	0.68	0.81	0.83	0.42	0.55	0.61
18	0.98	0.98	0.99	0.74	0.85	0.87	0.49	0.61	0.67

	4-variable VAR (ordering a); p=8								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	1	2	4	1	2	4	1	2	4
H	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
1	0.51	0.38	0.43	0.38	0.30	0.31	*0.06*	0.17	0.20
2	0.12	*0.06*	*0.07*	0.59	0.58	0.59	*0.07*	0.28	0.31
3	0.11	*0.06*	*0.06*	0.37	0.36	0.38	0.12	0.39	0.43
4	0.15	*0.08*	*0.09*	0.52	0.48	0.50	0.13	0.40	0.45
5	0.24	0.14	0.16	0.18	*0.08*	*0.09*	0.10	0.28	0.34
6	0.32	0.20	0.23	0.26	0.12	0.13	**0.02**	*0.10*	0.13
7	0.40	0.27	0.30	0.36	0.18	0.19	**0.04**	0.15	0.19
8	0.13	**0.04**	*0.05*	0.41	0.19	0.21	**0.01**	**0.02**	**0.03**
9	0.16	**0.04**	**0.05**	0.51	0.26	0.28	**0.01**	**0.02**	**0.03**
10	0.21	*0.06*	*0.07*	0.58	0.32	0.34	**0.02**	**0.04**	**0.05**
11	0.27	*0.09*	0.11	0.66	0.40	0.42	**0.03**	*0.06*	*0.07*
12	0.30	0.12	0.14	0.73	0.49	0.51	**0.04**	*0.08*	0.11
13	0.38	0.15	0.18	0.79	0.56	0.58	*0.06*	0.11	0.15
14	0.45	0.20	0.23	0.69	0.63	0.66	*0.06*	0.14	0.18
15	0.52	0.24	0.28	0.76	0.70	0.73	*0.07*	0.17	0.21
16	0.58	0.30	0.34	0.79	0.76	0.79	*0.09*	0.16	0.19
17	0.61	0.36	0.40	0.81	0.81	0.83	*0.10*	0.20	0.24
18	0.68	0.43	0.47	0.85	0.86	0.87	0.13	0.25	0.29

	4-variable VAR (ordering a); p=12								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
H	1	2	4	1	2	4	1	2	4
1	0.30	0.30	0.44	0.19	*0.09*	0.10	*0.09*	0.18	0.20
2	**0.00**	**0.00**	**0.00**	0.26	0.22	0.26	0.24	0.40	0.45
3	**0.00**	**0.00**	**0.01**	0.22	0.12	0.14	0.36	0.55	0.60
4	**0.00**	**0.01**	**0.01**	0.16	**0.04**	*0.06*	0.40	0.63	0.68
5	**0.01**	**0.01**	**0.02**	0.12	**0.01**	**0.02**	0.29	0.37	0.42
6	**0.01**	**0.02**	**0.04**	0.19	**0.02**	**0.04**	0.12	0.11	0.16
7	**0.02**	**0.03**	**0.05**	0.22	**0.03**	**0.04**	0.18	0.16	0.22
8	**0.02**	**0.01**	**0.01**	0.27	**0.02**	**0.02**	0.13	**0.02**	**0.03**
9	**0.02**	**0.01**	**0.02**	0.36	**0.03**	**0.04**	*0.09*	**0.01**	**0.02**
10	**0.04**	**0.01**	**0.03**	0.45	**0.04**	*0.06*	0.12	**0.02**	**0.03**
11	*0.05*	**0.01**	**0.03**	0.41	**0.04**	*0.07*	0.14	**0.03**	**0.05**
12	*0.05*	**0.02**	**0.03**	0.42	**0.01**	**0.01**	0.13	**0.01**	**0.01**
13	*0.08*	**0.02**	**0.04**	0.44	**0.01**	**0.01**	0.17	**0.01**	**0.02**
14	*0.10*	**0.02**	**0.04**	0.52	**0.01**	**0.02**	0.17	**0.02**	**0.03**
15	*0.10*	**0.04**	*0.06*	0.60	**0.02**	**0.03**	0.22	**0.03**	**0.05**
16	0.13	**0.05**	*0.09*	0.63	**0.03**	**0.04**	0.26	**0.04**	*0.06*
17	0.12	*0.07*	0.11	0.67	**0.04**	*0.06*	0.28	**0.05**	*0.08*
18	0.15	*0.09*	0.14	0.66	**0.05**	*0.08*	0.34	*0.06*	0.10

	5-variable VAR (ordering a); p=4								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	1	2	4	1	2	4	1	2	4
H	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
1	0.92	0.69	0.67	0.64	0.65	0.66	0.43	0.49	0.50
2	0.28	0.33	0.35	0.80	0.88	0.89	0.64	0.67	0.67
3	*0.08*	*0.06*	*0.07*	0.41	0.66	0.69	0.61	0.67	0.68
4	0.13	0.10	0.12	0.57	0.81	0.83	0.77	0.82	0.83
5	0.22	0.17	0.19	**0.02**	*0.05*	*0.07*	0.11	**0.03**	**0.03**
6	0.31	0.26	0.28	**0.03**	*0.09*	0.12	0.18	**0.05**	**0.05**
7	0.42	0.36	0.39	**0.05**	0.14	0.18	0.26	*0.08*	*0.08*
8	0.53	0.46	0.49	*0.08*	0.20	0.25	0.34	0.12	0.12
9	0.63	0.56	0.59	0.11	0.27	0.33	0.38	0.16	0.16
10	0.71	0.65	0.68	0.14	0.35	0.41	0.46	0.22	0.21
11	0.75	0.74	0.76	0.16	0.43	0.50	0.48	0.28	0.28
12	0.79	0.80	0.83	0.19	0.52	0.59	0.55	0.36	0.35
13	0.85	0.86	0.88	0.16	0.60	0.67	0.38	0.41	0.43
14	0.89	0.90	0.92	0.19	0.67	0.74	0.44	0.49	0.48
15	0.92	0.93	0.94	0.24	0.74	0.80	0.42	0.55	0.56
16	0.95	0.96	0.96	0.30	0.78	0.84	0.49	0.60	0.63
17	0.97	0.97	0.98	0.36	0.83	0.88	0.56	0.65	0.68
18	0.98	0.98	0.99	0.41	0.86	0.91	0.63	0.71	0.74

	5-variable VAR (ordering a); p=8								
	Import Prices			**Producer Prices**			**Consumer Prices**		
	Shock size (s.d.)			Shock size (s.d.)			Shock size (s.d.)		
H	1	2	4	1	2	4	1	2	4
1	0.45	0.12	0.15	0.81	0.50	0.51	*0.07*	0.22	0.25
2	*0.05*	**0.02**	**0.03**	0.76	0.79	0.79	0.20	0.46	0.50
3	**0.01**	**0.01**	**0.01**	*0.06*	0.23	0.29	*0.07*	0.20	0.25
4	**0.01**	**0.01**	**0.02**	*0.08*	0.32	0.39	0.13	0.32	0.39
5	**0.02**	**0.02**	**0.03**	**0.01**	**0.02**	**0.04**	0.13	0.21	0.26
6	**0.03**	**0.04**	*0.06*	**0.02**	**0.04**	*0.05*	0.12	0.18	0.23
7	**0.03**	**0.03**	**0.04**	**0.03**	**0.03**	**0.05**	0.18	0.24	0.29
8	**0.01**	**0.01**	**0.02**	**0.03**	**0.02**	**0.04**	0.10	**0.02**	**0.03**
9	**0.01**	**0.01**	**0.01**	**0.05**	**0.03**	*0.05*	0.11	**0.01**	**0.01**
10	**0.02**	**0.01**	**0.02**	*0.06*	**0.05**	*0.08*	0.15	**0.01**	**0.01**
11	**0.02**	**0.01**	**0.03**	*0.09*	*0.07*	0.12	0.20	**0.01**	**0.02**
12	**0.03**	**0.02**	**0.04**	0.12	0.11	0.16	0.25	**0.02**	**0.03**
13	**0.05**	**0.03**	*0.05*	0.15	0.14	0.21	0.30	**0.03**	**0.05**
14	*0.06*	**0.04**	*0.07*	0.11	0.18	0.27	0.30	**0.05**	*0.07*
15	*0.08*	**0.05**	*0.09*	0.15	0.24	0.33	0.33	*0.06*	*0.09*
16	0.11	*0.07*	0.11	0.19	0.29	0.40	0.34	*0.09*	0.12
17	0.14	*0.09*	0.15	0.23	0.36	0.47	0.40	0.12	0.15
18	0.18	0.12	0.19	0.29	0.42	0.54	0.47	0.15	0.20

Appendix J - CPI decomposition and cointegration

Model:	Tradables	Administered
Cointegration		
$F_{PSS}(2001)$		
F_{yx}	7.6120‡	7.4315†
$F_{xy(1)}$	2.3139	5.4679
$F_{xy(2)}$	1.7685	-
$F_{xy(3)}$	4.3069	-
t_{bounds}	-4.5020†	-3.8177†
Johansen test		
Trace	1	1
Auxiliary Tests		
F test (independent)	10.2145	-
(p-value)	0.0000	-
Error Correction	-0.0493	-0.0386
t-statistics	-5.55	-3.86
BDM critical value	-4.15	-3.71
ERPT slope coefficients		
Short-run ERPT (+)	-0.0271	0.0000
Short-run ERPT (-)	-0.3136‡	-0.0825‡
Long-run ERPT (+)	-0.0903	0.0000
Long-run ERPT (-)	-0.3136‡	-0.5075‡
Diagnostics		
Serial Correlation		
Breusch-Godfrey LM Test	0.0903	1.3785
(p-value)	0.9137	0.2423
Heteroskedasticity		
BPG	4.1642	3.8638
(p-value)	0.0001	0.0000
Stability		
CUSUM	Stable	Stable
CUSUMQ	Unstable	Unstable

Notes:
i) The models with nontradables, services and free prices do not show evidence of cointegration in multiple specifications;
ii) The NARDL models are:

$$\Delta p_t^{trade} = c + \beta t + \alpha_0\, p_{t-1}^{trade} + \alpha_1\, e_{t-1}^{(+)} + \alpha_2\, e_{t-1}^{(-)} + \alpha_3\, p_{t-1}^{w} + \sum_{i=0}^{1} \beta_{0i}\, \Delta e_{t-i}^{(+)} + \sum_{i=0}^{2} \beta_{1i}\, \Delta p_{t-i}^{trade} + \varepsilon_t$$

$$\Delta p_t^{admin} = c + \beta t + \alpha_0\, p_{t-1}^{admin} + \alpha_1\, e_{t-1}^{(-)} + \sum_{i=0}^{1} \beta_{0i}\, \Delta e_{t-i}^{(+)} + \sum_{i=0}^{5} \beta_{1i}\, \Delta e_{t-i}^{(-)} + \sum_{i=0}^{2} \beta_{2i}\, \Delta p_{t-i}^{admin} + \varepsilon_t$$

iii) Levels of significance:
‡ indicates 1%; † indicates 5%;

Appendix K - Media releases and increasing perception of price asymmetry

Headline	**Date**	**Stage**	**News media company**
"Why can the price of gasoline fall at the refinery, but rise at the gas station?"	January 27th 2017	Retail	UOL
"The drop in fuel prices does not reach the stations."	November 28th 2018	Retail	G1
"If oil and Dollar are stable, why did Petrobras increase gasoline?"	September 27th 2019	Wholesale	UOL
"Gasoline drops 40% but the final price doesn't reach the pumps."	March 27th 2020	Retail	FDR
"Reduction in the price of gasoline should not reach the pumps."	March 19th 2021	Retail	Valor Investe
"Why the Dollar falls but gasoline and diesel continue to rise?"	June 16th 2021	Wholesale	UOL
"Reduction in refineries does not reach the pumps and the price of gasoline exceeds BRL 6."	June 27th 2021	Retail	O Progresso
"Why did the price of gasoline rise even with stable exchange rate?"	July 26th, 2021	Wholesale	AUTO Esporte

Translations are our own.

Appendix L - Domestic production by refinery and ownership (2016-2019)

volume in thousands of m^3

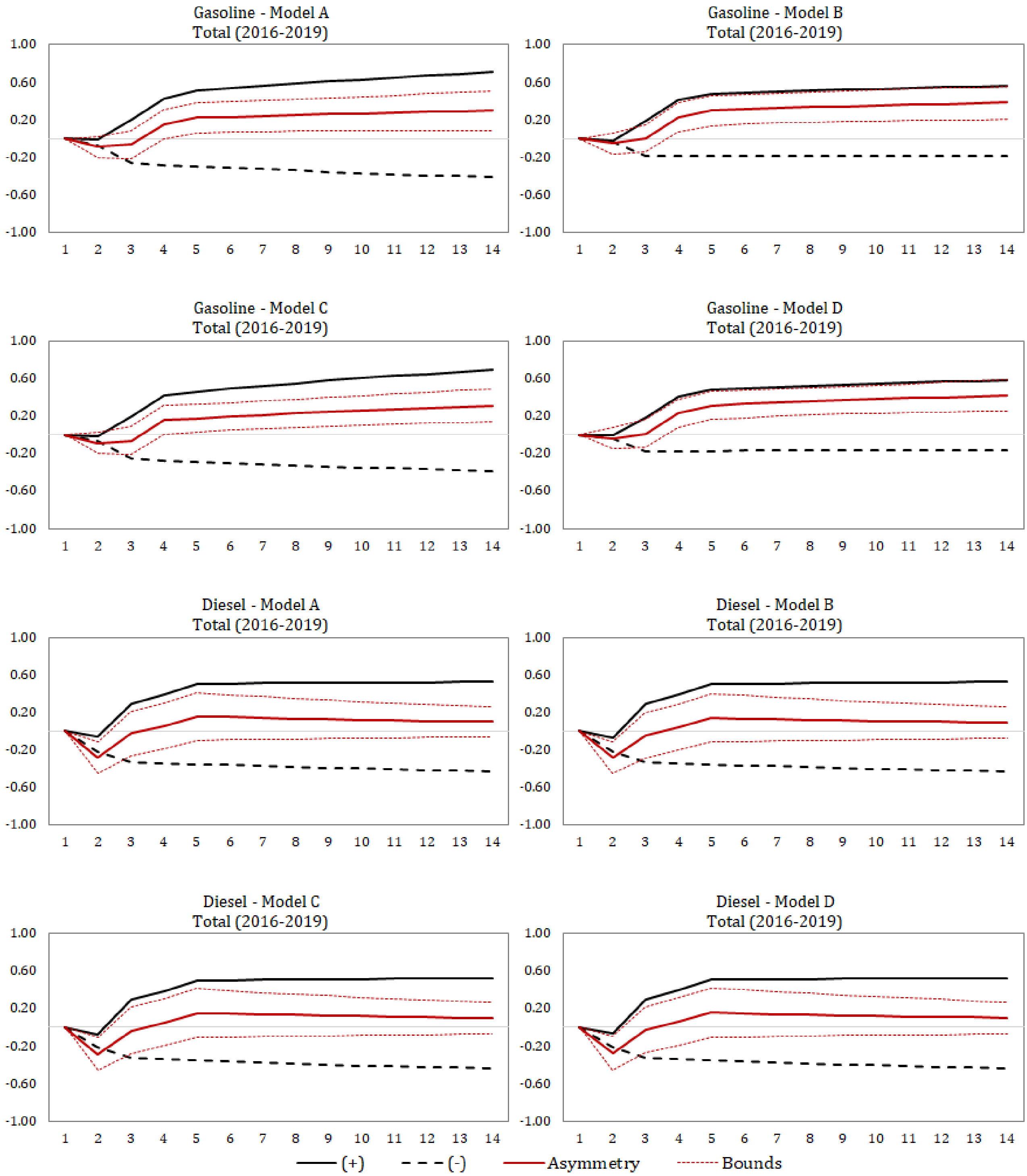
Refinery	*Gasoline* Volume	*Gasoline* Share (%)	*Diesel* Volume	*Diesel* Share (%)
RNEST	-	-	12,796	7.57%
RPCC	1,249	1.18%	1,174	0.69%
Lubnor	-	-	176	0.10%
Recap	4,515	4.26%	5,725	3.39%
Reduc	7,145	6.75%	11,823	7.00%
Refap	8,593	8.11%	16,204	9.59%
Regap	8,635	8.15%	15,381	9.10%
Reman	2,176	2.05%	2,301	1.36%
Repar	11,445	10.81%	18,156	10.75%
RPBC	8,908	8.41%	16,253	9.62%
Replan	20,288	19.15%	35,160	20.81%
Revap	12,576	11.87%	14,468	8.56%
RLAM	11,265	10.64%	17,545	10.38%
Manguinhos	2,114	2.00%		
Riograndense	1,640	1.55%	1,754	1.04%
Dax Oil	59	0.06%	36	0.02%
Petrobras' Total	96,795	91.38%	167,161	98.94%
Non-Petrobras' Total	3,812	3.60%	1,790	1.06%
Other non-refinery sources	5,314	5.02%	-	-
Total	105,921	100.00%	168,951	100.00%

Black indicates Petrobras' refineries and **red** indicates refineries owned by other companies.

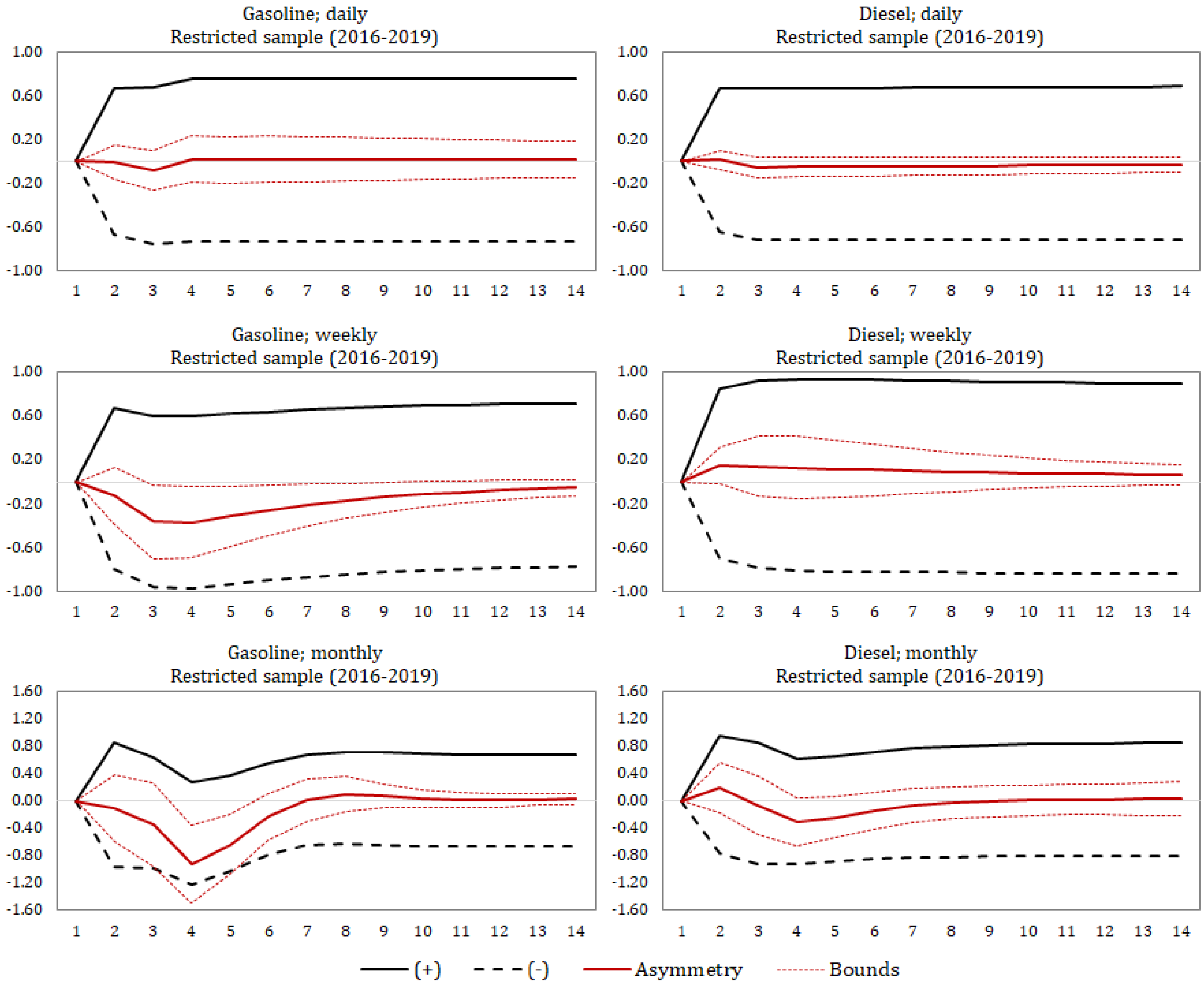
Non-refinery sources encompass petrochemical stations and blending stations.

Appendix M - Dynamic multipliers

Appendix M.1 - Brazilian fuel market



Appendix M.2 - North-American fuel market



Continuation