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Essays on Monetary Economics

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Resumo

Esta tese é composta por três capítulos. No primeiro capítulo, revisitamos o programa de estabilização Plano Real de 1994, que colocou um fim no período de hiperinflação no Brasil. Esse episódio histórico teve como característica o lançamento de uma moeda virtual, a unidade real de valor (URV), que teoricamente teria quebrado a inércia inflacionária e ao mesmo tempo sincronizado os reajustes de preços na economia. Em nossa análise, utilizamos microdados do índice de preços ao consumidor da FIPE para documentar estatísticas de preços ao longo do processo de estabilização. Os nossos resultados reforçam que o Plano Real alterou as características de precificação no país. Há evidências de que os reajustes de preços se tornaram menos frequentes, menores e mais sincronizados, enquanto a distribuição de reajustes de preços se tornou mais simétrica e concentrada em torno de zero, porém somente após a introdução da nova unidade monetária, o real. Assim, não encontramos evidências decisivas de mudança no processo de formação de preços durante o período da URV.

No segundo capítulo, complementamos o estudo do programa de estabilização Plano Real de 1994 à luz de um modelo de custos de menu com regras de preços dependentes do estado. Dadas as estatísticas de preços calculadas com microdados do índice de preços ao consumidor da FIPE e as evidências de mudanças no processo de formação de preços somente após a introdução da nova unidade monetária, calibra-se um modelo de custos de menu com o objetivo de ilustrar a dinâmica de precificação ao longo do processo de estabilização. Dessa maneira, a nossa análise traz uma base teórica para caracterizar esse período de transição econômica do Plano Real, que culminou no fim da hiperinflação no Brasil. Em especial, é surpreendente a ausência de alterações significativas durante o período da URV apesar de incentivos teóricos para uma conversão de preços em URV em detrimento da moeda oficial.

No terceiro capítulo, avaliamos o impacto de choques de política monetária de alta frequência nos preços de ativos financeiros no Brasil. Primeiro, construímos uma ampla base de dados com informações sobre anúncios de política monetária, que inclui não somente as decisões do COPOM e as respectivas atas das reuniões, mas principalmente as participações de autoridades do Banco Central do Brasil (BCB) em diferentes eventos. Há evidências de que as decisões do COPOM e os discursos de autoridades exercem uma influência considerável nos ativos financeiros em comparação com outras fontes de choques monetários. Segundo, utilizamos um modelo de vetor autorregressivo estrutural (SVAR) com identificação por heterocedasticidade, sob a hipótese de que períodos com eventos monetários são relativamente mais voláteis do que períodos sem tais eventos, para obtermos séries de choques de política monetária. Com isso, estimamos funções impulso-resposta por meio de projeções locais (LPs). Os nossos resultados mostram que os choques monetários têm impacto significativo e persistente nos preços de ativos analisados. Além disso, a transmissão desses choques é fortemente condicionada pela postura da política monetária, com efeitos mais duradouros durante os ciclos de alta de juros. Por fim, os choques de aperto monetário que identificamos também afetam indicadores macroeconômicos, reduzindo a inflação em horizontes médios, ao mesmo tempo em que impõem custos relativamente pequenos sobre o produto.

Palavras-chave: Plano Real, hiperinflação, formação de preços, microdados, custos de menu, choques de política monetária, vetor autorregressivo estrutural, projeções locais.

Abstract

This thesis is comprised of three chapters. The first chapter revisits the 1994 Real Plan stabilization program that successfully ended hyperinflation in Brazil. This historical episode features the launch of a virtual currency, the unit of real value (URV), which theoretically broke up inflationary inertia while synchronizing price adjustments in the economy. We use product-level data from a non-publicly disclosed Brazilian consumer price index (CPI) to document price-setting statistics throughout the stabilization process. Our results reinforce that the Real Plan altered the price-setting characteristics in the country. There is evidence that prices changed less frequently, by a lower amount, and in a more synchronized way, while the distribution of price adjustments became more symmetric and concentrated around zero, but only after introducing a new monetary unit, the Brazilian real. Therefore, we did not find decisive evidence of a price dynamics change during the URV phase of the plan.

In the second chapter, we extend the study of the 1994 Real Plan stabilization program by examining it through the lens of a menu cost model with state-dependent pricing rules. Given the price statistics calculated from the FIPE CPI microdata and the evidence of price formation changes only after introducing the new currency unit, we calibrate a menu cost model aimed at illustrating the pricing dynamics throughout the stabilization process. In this way, our analysis provides a theoretical foundation to characterize this transitional economic period of the Real Plan, ultimately leading to the end of hyperinflation in Brazil. Notably, it is surprising to observe the absence of significant changes during the URV period, despite theoretical incentives for converting prices into URV to the detriment of the official currency.

In the third chapter, we evaluate the impact of high-frequency monetary policy shocks on financial asset prices in Brazil. First, we construct an extensive database containing information on monetary policy announcements, which includes not only the decisions of the Monetary Policy Committee (MPC) and the corresponding meeting minutes but, most importantly, the participation of Central Bank of Brazil (BCB) officials in various events. Evidence suggests that MPC statements and speeches by BCB authorities exert a considerable influence on financial assets compared to other sources of monetary shocks. Second, we use a structural vector autoregressive (SVAR) model identified through heteroskedasticity, under the assumption that periods with monetary events are relatively more volatile than otherwise, to obtain series of monetary policy shocks. Then, we estimate impulse-response functions through local projections (LPs). Our results indicate that monetary shocks exert a significant and persistent impact on the analyzed asset prices. Moreover, the transmission of these shocks is strongly conditioned by the policy stance, with more sustained effects during tightening phases. Finally, our identified tightening shocks also affect macroeconomic indicators, reducing inflation over medium horizons while imposing comparatively small output costs.

Keywords: Real plan, hyperinflation, price setting, microdata, menu-cost, monetary policy shocks, structural vector autoregressive, local projections.

Summary

Chapter 1

Micro Evidence on the Real Plan Stabilization Program in Brazil: Data and Price-Setting Statistics[1]

1.1. Introduction

Hyperinflation promotes a rapid erosion of purchasing power and rising uncertainty over economic conditions. It causes people to make unsound economic choices since it distorts the informational content embedded in prices. In other words, relative prices become a misleading guide to consumer choices so it reduces welfare (e.g., Cagan, 1956; Sargent, 1982; Fischer et al., 2002; Reinhart & Rogoff, 2009). Consequently, people lose faith in the currency and start seeking alternative methods to store value until a government implements a successful stabilization plan. This paper revisits the Brazilian hyperinflationary experience and provides micro evidence of how the country overcame hyperinflation in the context of the Real Plan stabilization program in 1994.

Among fiscal and monetary disarrays, Brazil suffered from extremely high inflation issues in the late 1980s and early 1990s, but none of the stabilization attempts was able to bring inflation back to low levels in a definitive way. Brazilian inflation had a structural component given by fiscal deficits and an inertial one associated with an unsynchronized mechanism of indexation of contracts.[2] Table provides information

[1] This is joint work with Julia Araujo, Marco Bonomo, and Mauro Rodrigues.

[2] For a detailed description of indexation in the Brazilian economy, see Simonsen (1995). Besides, Gustavo Franco and Edmar Bacha, who participated directly in the Real Plan, share their impressions of the Brazilian economy and the stabilization plan in Franco (1995) and Bacha (2012), respectively. More recently, Ayres et al (2019) provided a historical and economic analysis of the Brazilian fiscal and monetary interactions from 1960 to 2016.

regarding the governments, economic stabilization plans, monetary units, and their conversion rate to the previous currencies during the 1980s and 1990s. Specifically, under the presidency of José Sarney (1985-1989), Brazil implemented four plans: Cruzado I (Feb/1986), Cruzado II (Nov/1986), Bresser (Jun/1987), and Verão (Jan/1989). The next president, Fernando Collor de Mello (1990-1992) implemented two programs: Collor I (Mar/1990) and Collor II (Jan/1991). In common, all of these attempts tried to control hyperinflation by freezing prices and wages, with the caveat that Collor I also confiscated household savings. Subsequently, amid a corruption scandal at the end of 1992, Collor resigned and his vice president, Itamar Franco, remained in office until 1994. Under his administration, Brazil executed the Real Plan led by his Minister of Finance and succeeding President Fernando Henrique Cardoso.

Table 1 - Currencies and Economic Plans in BrazilTable

Currency	Symbol	Period	Conversion Rate	President	Plan
Cruzado	Cz$	28/Feb/1986 - 15/Jan/1989	-	José Sarney	Cruzado I (Feb/1986), Cruzado II (Nov/1986), and Bresser (Jun/1987)
Cruzado Novo	NCz$	16/Jan/1989 - 15/Mar/1990	NCz$ 1 = Cz$ 1000		Verão I (Jan/1989)
Cruzeiro	Cr$	16/Mar/1990 - 31/Jul/1993	Cr$ 1 = NCz$ 1	Fernando Collor	Collor I (Mar/1990) and Collor II (Jan/1991)
Cruzeiro Real	CR$	01/Aug/1993 - 30/Jun/1994	CR$ 1 = Cr$ 1000	Itamar Franco	-
Real	R$	01/Jul/1994 -	R$ 1 = CR$ 2750		Real (Jul/1994)

The Real Plan has origins in the Larida Plan, named after an economic stabilization proposal outlined in Arida and Lara-Resende (1985). In this seminal work, the authors propose the creation of an indexed currency, the Readjusted Obligations of the National Treasury (ORTN) pro-rata, to synchronize all price adjustments. Their idea later reappeared in the Real Plan with the unit of real value (URV) – the program's crucial feature – an alternative unit of account inspired by the ORTN.

The Real Plan had three distinguishable phases. The preliminary one started with the introduction of a new currency, the cruzeiro real (CR$), on August 1, 1993, and was characterized by fiscal adjustment with austerity policies under the Immediate Action Program (PAI) in July 1993 and the Emergency Social Fund (FSE) in February 1994. It marked a commitment to address the structural component of inflation by reducing fiscal deficits.

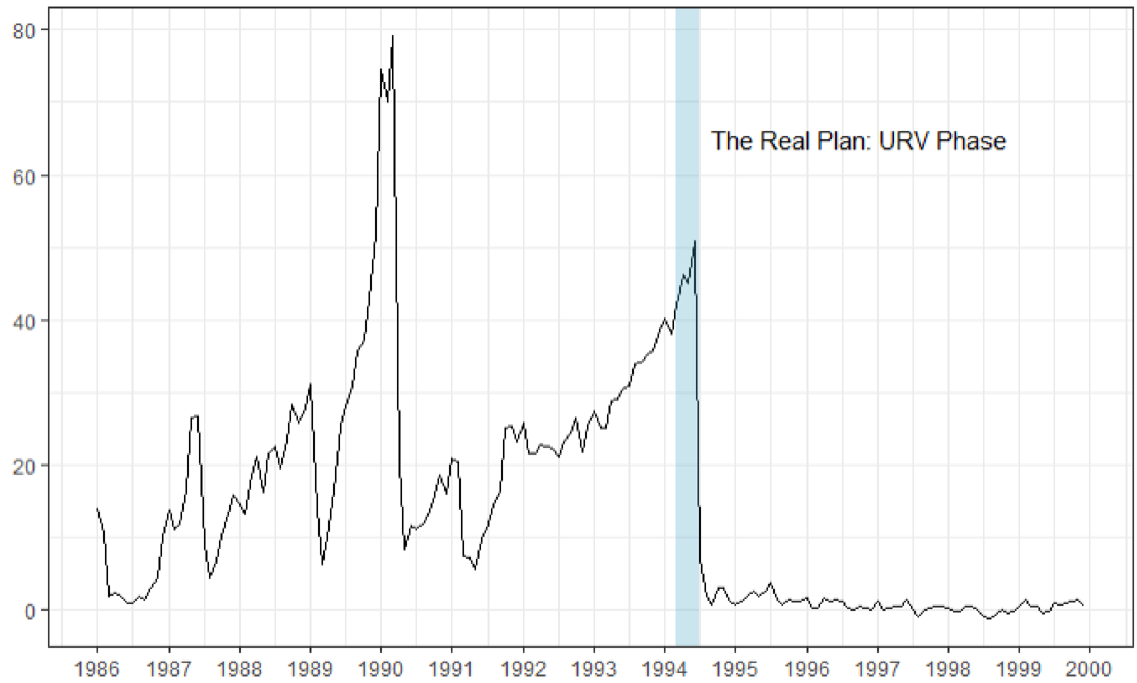
The following phase – the plan itself – lasted 4 months (from March 1 until June 30, 1994) in which a temporary unit of account with daily correction, the URV, was launched and coexisted with the CR$. Although both the CR$ and the URV served as units of account, prices in URV were designed to be more stable. Specifically, while the CR$ absorbed the effects of hyperinflation, the URV was pegged to the U.S. dollar at parity and daily updated (to the CR$) according to inflation indexes. As a result, the URV represents an estimate of the CR$ loss of purchasing power. For example, given a CR$ inflation of 10%, the value of the CR$/URV would increase by 10%. This would lead to a CR$ rise in both prices and salaries, but theoretically, they would remain relatively stable in URV, which could voluntarily stimulate its usage as the main unit of account. The rationale is that, as people notice money holds its value, they would start thinking of prices in URV (as opposed to CR$) and stop expecting prices to always go up. Therefore, the URV daily indexation would synchronize price adjustments (i.e., promote the alignment of relative prices) in the economy, thus breaking up the inertial component of inflation.

At the time of the URV implementation, prices could be quoted in URV but had to be paid in CR$ whereas the conversion was mandatory for salaries. Overall, free negotiations determined the terms of the conversion with the exemption of salaries, rent, school fees, and public utilities. The existing contracts had an option to convert until the launch of a new official currency (which would be pre-announced 30 days in advance), whereas the new ones should be established in URV. However, at first, the quotation of goods and services prices for prompt payment had to be exclusively in CR$ until the end of April (to avoid reducing menu costs of daily price adjustments), then it was allowed URV quotation so that people could get used to stable prices. In parallel, as a rule, salaries were initially readjusted based on their average value over the past 4 months, then daily adjusted and monthly paid in CR$, thus preserving its purchasing power.

In the plan's final phase, the URV became the Brazilian real (R$), not only serving as a unit of account but also as a medium of exchange, substituting the CR$ at a fixed CR$/R$ conversion rate of CR$ 2,750 on July 1, 1994 (see Table 1). Henceforth, the R$ became the new official currency.

Figure 1 illustrates the monthly inflation, measured by the Consumer Price Index of the Economic Research Institute Foundation (FIPE CPI), from 1985 until the end of 1999. In the Appendix, Figure A1 exhibits the yearly FIPE CPI behavior throughout that period. In both figures, we highlighted the URV phase of the Real Plan, which is the period that the stabilization took place. As Figure 1 suggests, Brazilian inflation was continuously deteriorating even during the URV phase (also stimulated by fears that the government would reintroduce freezing price measures before the monetary change). Thereafter, the Real Plan accomplished enormous disinflation and attained sustainable price stability. For instance, the annual inflation rate peaked at 5,167.34% in June of 1994 while the monthly rate dropped from 50.8% to 6.9% in the following month. Therefore, the plan succeeded with a pre-announced approach, without household savings confiscation or price freezes of previous stabilization attempts. Almeida and Bonomo (2002) argue that the Real Plan is an example of how fast disinflation is possible without high costs, even in a context of low credibility, conditional on a mechanism of price alignment – the URV – to eliminate distribution asymmetry.

Figure 1 - FIPE CPI: Monthly Inflation (%)



Notes: The highlighted area represents the Real Plan's unit of real value (URV) phase.

In this paper, we examine price adjustment during the implementation of the Real Plan stabilization program in Brazil, which marked a transition period between hyperinflation and relatively low inflation, using the non-publicly disclosed FIPE CPI microdata of prices at the store level. This dataset of prices is the only one that is has been available for the years of hyperinflation in Brazil.

Our main interest is verifying whether URV-converted prices behaved according to the predictions of URV policies. Since the stabilization plan succeeded, one could expect that the URV highly influenced price setting in the sense that URV-converted price adjustments should be increasingly more synchronized besides being relatively closer to zero and more symmetric. As a result, we document the price-setting dynamics at the time of the Real Plan by looking at micro price adjustment in CR$ and its relationship with the URV to access the mechanisms through which the plan succeeded. Accordingly, we compute and discuss the price-setting statistics of inflation, frequency, absolute size, distribution of price changes, and brand-level intra-market relative price variability (RPV). Specifically, we compare price settings in Brazil across different periods: pre-URV (from August 1993 until February 1994, with prices in CR$), URV (from March until June 1994, with prices in both CR$ and URV), and post-URV (from July until December 1994, with prices in R$).

Our paper relates to the literature on price setting using microdata and how pricing varies under different inflationary scenarios. Klenow and Malin (2010) and Nakamura and Steinsson (2013) provide a comprehensive review of studies that shed new light on micro price adjustment features and how price adjustment's sluggishness translates into the effects of demand and monetary shocks on output in the short-run.[3] In this context, several papers explore the low and stable inflation environment of developed regions, such as Bils and Klenow (2004), Klenow and Kryvtsov (2008), and Nakamura and Steinsson (2008) for the U.S., and Dhyne et al (2006) and Álvarez et al (2006) for Europe. However, high inflation or even a hyperinflationary environment is a common feature in developing countries and some developed ones, which offers invaluable opportunities to access the dynamics of price adjustments under different inflation scenarios. For example, Lach and Tsiddon (1992) for Israel, Konieczny and Skrypacz (2005) for Poland, Gagnon (2009) for Mexico, Wulfsberg (2016) for Norway, and Nakamura et al. (2018) for the U.S., and Alvarez et al (2018) for Argentina.[4] Contrastingly, to our knowledge, this is the first study to provide an in-depth analysis of micro price adjustment focused on the mechanics of a successful stabilization program that fought hyperinflation. Specifically, instead of comparing the characteristics before and after the disinflation, we explore explicitly the effects of the stabilization policies on prices.

[3] On one side, a better assessment of pricing characteristics contributes to enhancing behavioral assumptions on structural models of price adjustment. On the other hand, it helps central banks to make informed decisions in the conduct of the monetary policy since it provides information on inflation persistence, that is, the speed with which inflation moves toward the target after a shock in the economy.

[4] Nakamura et al. (2018) and Wulfsberg (2016) address the small inflation variability for the U.S. and Norway, respectively.

Some studies focus on Brazil like Gouvea (2007), Lopes (2008), and Barros et al. (2009) that analyze price setting after the economic stabilization, and Angelis (2012), Araujo and Rodrigues (2018), and Araujo (2019) who show that the Real Plan completely changed price setting characteristics.[5] Nonetheless, they do not explore the role of the URV in the stabilization process. In contrast, we add to the literature by covering the transition from a hyperinflationary to a relatively low inflation period on a narrower window of analysis (1993-1995) in which we explore the relationship of the microdata on prices at the store level and the URV information.

Our results reinforce that the Real Plan altered the price-setting characteristics in Brazil. In spite of that, interestingly, we do not find decisive evidence of price dynamic change during the URV phase of the plan, which indicates the stabilization mainly occurred with the monetary reform that introduced the R$. As a result, prices changed less frequently, by a lower amount, and in a more synchronized way, while the distribution of price adjustments became more symmetric and concentrated around zero only after introducing a new monetary unit.

There are a few explanations for these findings. First, although the FIPE microdata structure is weekly-based, it is a monthly dataset since the quotation of a specific unit of analysis is usually once a month so it possesses a relatively low frequency. Second, the price quotation in URV effectively started in May while the price surveys reflected inflation with a 2-4 weeks delay. Third, a positive association between rising inflation and greater RPV especially before the introduction of the R$ in July in a context of greater uncertainty. Fourth, the exchange rate policy contributed to the resulting outcome since real exchange rate indexes remained relatively stable during the first half of 1994 while experienced a real appreciation afterwards that facilitated imports, aligning with the strategy to exert downward pressure on domestic prices and combat inflation. These features contribute to finding muted supportive evidence of stabilization effects during the 4-month URV period.

In conclusion, based on micro price adjustments, our results do not support the claim that the URV had a decisive role in altering the pricing dynamics and attaining sustainable price stability. In other words, there is low evidence of price dynamic changes during the URV period. Notwithstanding, these outcomes suggest by no means that the URV was insignificant since the mechanism enabled a milder transition to a new currency

[5] Additionally, Correa et al. (2016) present a qualitative price survey conducted by the Central Bank of Brazil (BCB) with local firms.

while promoting an immediate shift of the distribution of prices towards zero, besides guaranteeing real wage purchasing power during the stabilization period.

The rest of the paper is organized as follows: data, price-setting statistics, and concluding remarks.

1.2. Data

We focus our analysis on the period that starts in January 1993 and goes until December 1995. As a result, our window of analysis encompasses the pre-stabilization period, the URV phase of the Real Plan (from March until June 1994), and the post-introduction of the R$ as the new monetary unit (from July 1994 onwards). To contrast these periods and access the URV effect on the Brazilian price-setting dynamics, we use the non-publicly disclosed FIPE CPI microdata and the URV official information. In this section, we discuss the main data features.

The FIPE CPI computes price change in the municipality of Sao Paulo, which is the most populous and the largest GDP per capita city in Brazil.[6] The index measures inflation according to a 4-week accumulated period (that is, the price average of four consecutive weeks is compared with the price average of the previous ones) for those who are in the salary bracket between 1 and 10 minimum wages. For that purpose, the FIPE conducts four index updates every month (hereafter, each month is assumed to possess four weeks) and the last one gives its CPI monthly variation.[7] Table A1 in the Appendix summarizes the information on the number of days in each week and month during the three non-leap years in our analysis.

Price surveys compose the FIPE microdata and possess information on price quotes of consumer goods and services at the store level. This is such a unique dataset because it is the only one available for the years of hyperinflation in Brazil.[8] It does not impose product substitution, nor price imputation, and it does not have a sales indicator on its products so the data contains posted prices (as opposed to regular prices, which

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[6] Brazilian Institute of Geography and Statistics (IBGE) estimates for 2020.
[7] The FIPE divides each month into 4 periods of 7 to 8 days.
[8] Lopes (2008), Angelis (2012), Araujo and Rodrigues (2018), and Araujo (2019) also use the FIPE microdata in their price-setting studies.

exclude sales).[9] In the dataset, the most disaggregated unit is a specific brand of product, whose price was collected in a given week at a particular store (henceforth, an item). Even though this is weekly-based data, a specific item is usually quoted only once a month.[10] Consequently, in practice, we have quasi-monthly data for item price trajectories. Table 2 provides an example of the data structure using the information of a 600 ml bottle of Antarctica (brand) beer (product) sold at an undisclosed but specified supermarket (a type of outlet) during the 4th week of August 1993.[11]

Table 2 - The Unit of Analysis: Item

	Product	**Brand**	**Type of Outlet**	**Outlet**	**Week**
Example	Beer	600 ml Antarctica Bottle	Supermarket	-	4ª Week of August
FIPE Code	412	412000	1	615	32

Notes: The table exemplifies the most disaggregated unit of analysis – an item. Each price quote of an item has a specific code that carries information on its product, brand, outlet, and week of the survey.

Over the 1993 to 1995 window, the FIPE changed the index methodology once at the end of 1993 to incorporate changes in consumption behavior. For the sake of consistency, we use the weights based on the Consumer Expenditure Survey (POF) of 1990-1991 to aggregate products while excluding those dropped from the index because of that change. In the Appendix, Table A2 provides an overview of the weighting scheme across the official FIPE groups of products – food, housing, transportation, personal expenses, apparel, healthcare, and education.

Table 3 brings alternative product classifications (alongside its weights) commonly used by the Brazilian Central Bank (BCB) that aggregates products into different categories such as: free market (food at home, industrial goods, and services) and regulated prices; and tradables, non-tradables, and regulated prices. Importantly, we do not exclude regulated prices, controlled by the government, such as electricity, gas and sewer utilities, gasoline, public transportation fares, and pharmaceutical drugs, to

[9] For a detailed explanation of price adjustment features that interfere with price rigidity statistics, see Klenow and Malin (2011) and Nakamura and Steinsson (2013).
[10] On average, only 10% of the items have prices quoted more than once per month.
[11] Each item is identified by a specific code that carries information on its product, brand, outlet, and week that price was surveyed.

investigate whether they have behaved differently than the non-regulated ones concerning the URV conversion.

Table 3 - Alternative Product Classifications and Weights

Product Classifications	**POF 90-91 1994-1999**
General Index	100.0
Free Market Prices	78.0
Food at Home	29.6
Industrial Goods	25.7
Services	22.7
Tradables	35.8
Non-Tradables	42.2
Regulated Prices	22.0

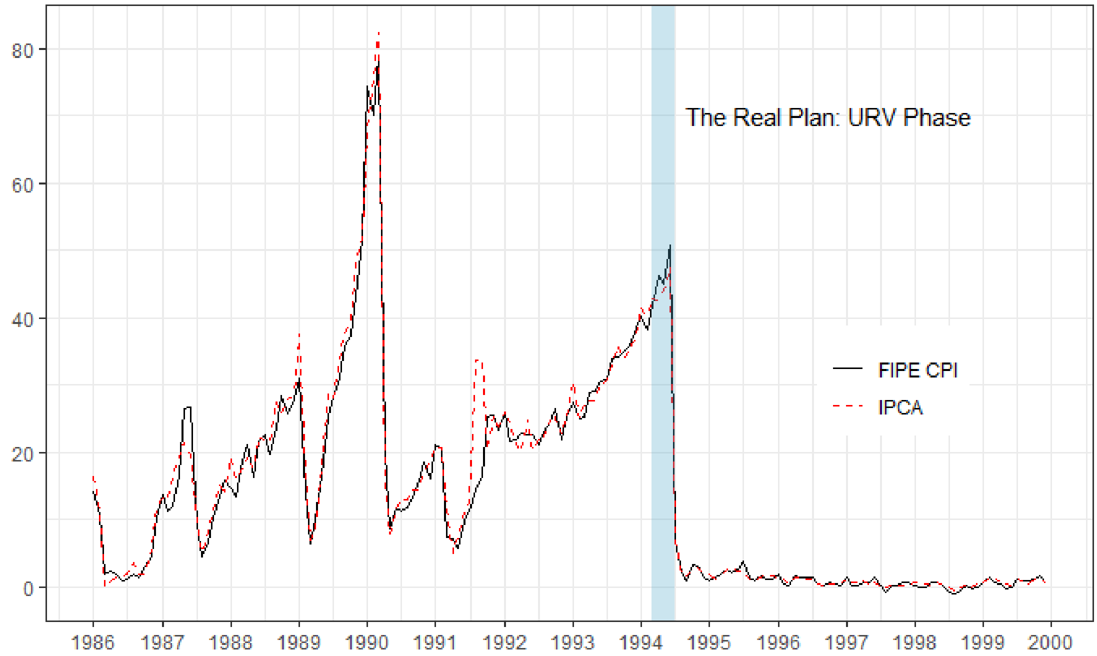
Notes: The weighting scheme is based on the Consumer Expenditure Survey (POF) of 1990-1991.

Before discussing the URV, we clarify that the Broad National Consumer Price Index (IPCA) of the Brazilian Institute of Geography and Statistics (IBGE) is the official inflation index used in the inflation-targeting framework conducted by the BCB, but it does not have microdata available for the early 1990s.[12] Nevertheless, despite not having a national coverage as the IPCA, the FIPE CPI possesses a similar structure so that both indexes capture the same inflation movements over time as shown in Figure 2.

The BCB set different methodologies for the pre-URV (from January 1993 to February 1994) and the URV phase of the Real Plan (from March until June 1994) through Act 1066 in February 1994. By the time the guidelines were released, the government published all of the past daily Cr$/URV and CR$/URV exchange rates retroactively to explain and induce its usage as a unit of value to update contracts while describing how it would be computed afterward.

[12] The CPI of the Brazilian Institute for Statistics of the Getulio Vargas Foundation (IBRE-FGV) is another popular Brazilian inflation indicator. A crucial difference between the IBRE-FGV CPI and the FIPE CPI is that the former possesses information only after the economic stabilization in Brazil (from March 1996 onwards) whereas the latter has price quotes for the years of hyperinflation. Gouvea (2007) and Barros et al. (2009) use the IBRE-FGV microdata in their studies.

Figure 2 - Monthly Inflation (%) in Brazil: FIPE CPI and IPCA



Notes: It compares the Broad National Consumer Price Index (IPCA) monthly inflation and the FIPE CPI. Overall, both indexes capture the same inflation movements over time.

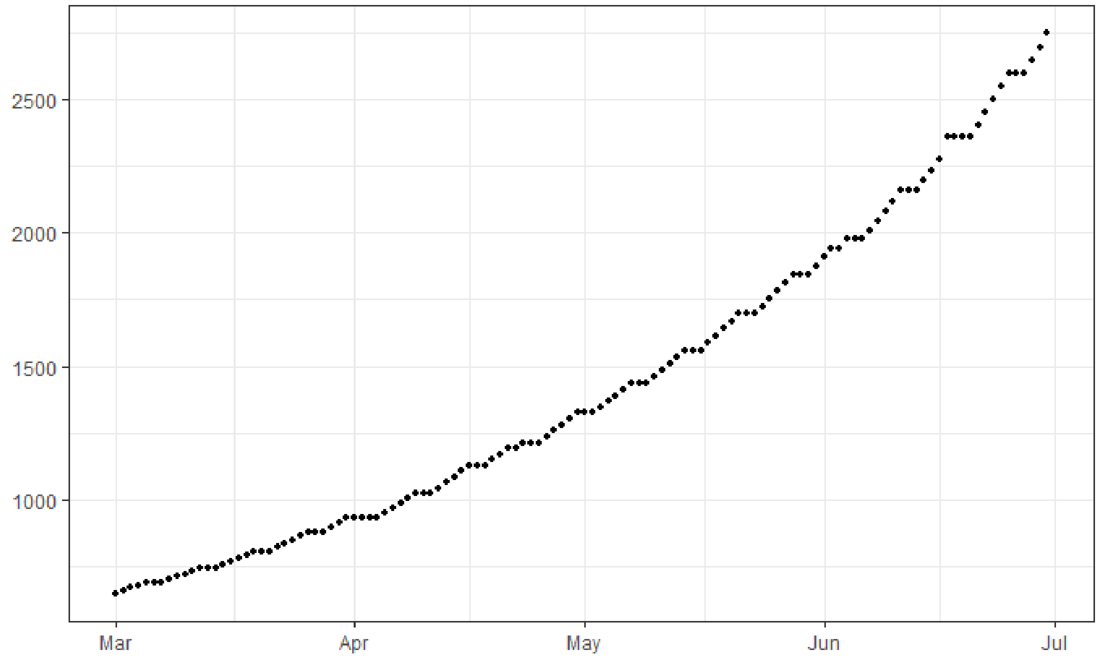
For the pre-URV period, the CR$/URV monthly rate was defined (ex-post) as the arithmetic mean of three existing monthly inflation indexes: the FIPE CPI (measured at the 3rd week), the IBGE IPCA-E, and the FGV IGP-M.[13] Hence, the CR$/URV reflected a measurement of the CR$ loss of purchasing power. Additionally, the CR$/URV daily values were computed using a monthly factor $mf = \sqrt[T]{1 + \frac{(CR\$/URV\ monthly\ rate)}{100}}$, where T is the number of working days in a month and *CR$/URV monthly rate* is the resulting average inflation of those underlying components. As a result, the CR$/URV daily quote was the previous one multiplied by its correspondent monthly factor so that $(\text{CR\$/URV})_t = (\text{CR\$/URV})_{t-1} * mf$ for $t = 1, \dots, T$.

For the URV period, the BCB would announce the CR$/URV quote for the following working day based on price surveys of those three inflation indicators which

[13] The IBGE Special Extended National Consumer Price Index (IPCA-E) is currently the cumulative quarterly result of the IPCA-15. The IPCA-15 differs from the IPCA exclusively in terms of geographic coverage (it covers fewer urban areas) and data collection period (from day 16 of the previous month to day 15 of the month of reference instead of using the full month of reference). In practice, the IPCA-15 is a preliminary reading of the IPCA. In contrast with the other indexes, the FGV General Price Index – Market (IGP-M) is not a CPI, but it contains a CPI component, and it was created to update financial values amid the deterioration of Brazilian inflation in 1989

reflected 2-4 weeks delayed inflation. Besides, as a rule, the CR$/URV monthly rate should be within the interval given by those indexes' highest and lowest monthly inflation. The unit value of URV was initially set at CR$ 647.5 on March 1, then daily updated leading up to CR$ 2,750 on June 30 (see Figure 3). Hereafter, the government not only extinguished the CR$ but also transformed the URV into the R$, which was still pegged to the U.S. dollar at parity (i.e., R$ 1 = URV 1 = CR$ 2,750).[14] In the Appendix, Figure A2 reveals that the URV followed the nominal foreign exchange rate (CR$/U.S. Dollar) at that time.

Figure 3 - CR$/URV Daily Quotes during the URV Period



Notes: It illustrates the CR$/URV daily quotes, announced by the BCB, during the URV phase of the Real Plan (from March to June 1994).

Table 4 compares the CR$/URV monthly inflation with its underlying components' behavior. For the pre-URV period, the URV inflation was precisely the arithmetic mean of its components as expected. For the subsequent period, even though the CR$/URV inflation did not have to be the mean of these indexes, it was closely related to it. Therefore, the URV monthly variation closely followed the Brazilian inflation.

[14] In the transition from the URV to the real, the government established that inflation in July of 1994 would be computed through the comparison of prices in R$, collected at that month, with prices in URV of the previous month (after being converted from CR$).

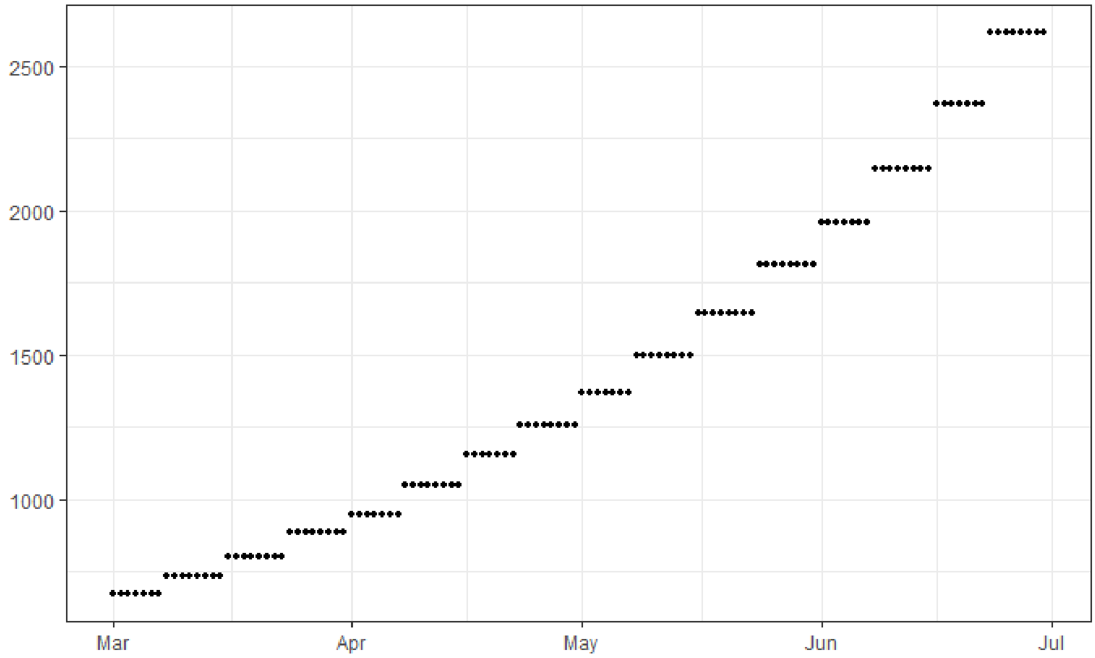
Table 4 - Monthly Inflation (%): CR$/URV and its Components

Period	CR$/URV	Avg. Infl.	FIPE-3	IPCA-E	IGP-M
Pre-URV					
Jan-93	27.71	27.71	27.83	29.47	25.83
Feb-93	26.73	26.73	25.06	26.72	28.41
Mar-93	25.74	25.74	25.02	25.96	26.25
Apr-93	27.98	27.98	27.78	27.34	28.83
May-93	29.19	29.19	29.26	28.61	29.70
Jun-93	29.80	29.80	30.28	27.61	31.50
Jul-93	30.71	30.71	30.21	30.67	31.25
Aug-93	32.39	32.39	33.40	31.99	31.79
Sep-93	34.55	34.55	34.00	34.38	35.28
Oct-93	34.90	34.90	34.48	35.17	35.04
Nov-93	35.19	35.19	35.51	33.90	36.15
Dec-93	37.59	37.59	37.76	36.69	38.32
Jan-94	39.73	39.73	40.94	39.17	39.07
Feb-94	39.17	39.17	37.04	39.70	40.78
URV					
Mar-94	43.26	43.55	41.31	43.63	45.71
Apr-94	42.60	42.53	45.43	41.25	40.91
May-94	44.00	43.82	44.66	44.21	42.58
Jun-94	46.60	46.28	48.97	44.65	45.21

Notes: The table provides the monthly inflation (%) of the CR$/URV and its underlying components, that is, the FIPE-3 CPI (measured in the 3rd week), the IBGE IPCA-E, and the FGV IGP-M for both the pre-URV and the URV periods. Because the Cr$ was the Brazilian official currency until July 1993, Cr$/URV quotes were converted to CR$/URV at $CR\$1 = Cr\$1000$.

Given our interest in the mechanics of the Real Plan stabilization program that ended hyperinflation in Brazil, we combine the FIPE microdata, that is, weekly price quotes expressed in the official monetary unit (Cr$, CR$, or R$) with daily URV information. For this purpose, the moment that price is surveyed is crucial. A problem that arises is that FIPE does not have the exact date of price surveys on their files. However, the organization said they usually collect almost all the prices in the first three working days of each week and there is no price survey on non-working days. We incorporate these pieces of evidence when converting all prices to URV (or R$) through a weekly weighted average of the CR$/URV quote by assigning double chances that prices were collected in each of the first three working days of a week. Hence, we have a point estimate for the CR$/URV quote that prevailed at the time that each item was surveyed. Figure 4 shows these weekly estimates for the URV period.

Figure 4 - Weekly Weighted Avg. of the CR$/URV during the URV Period



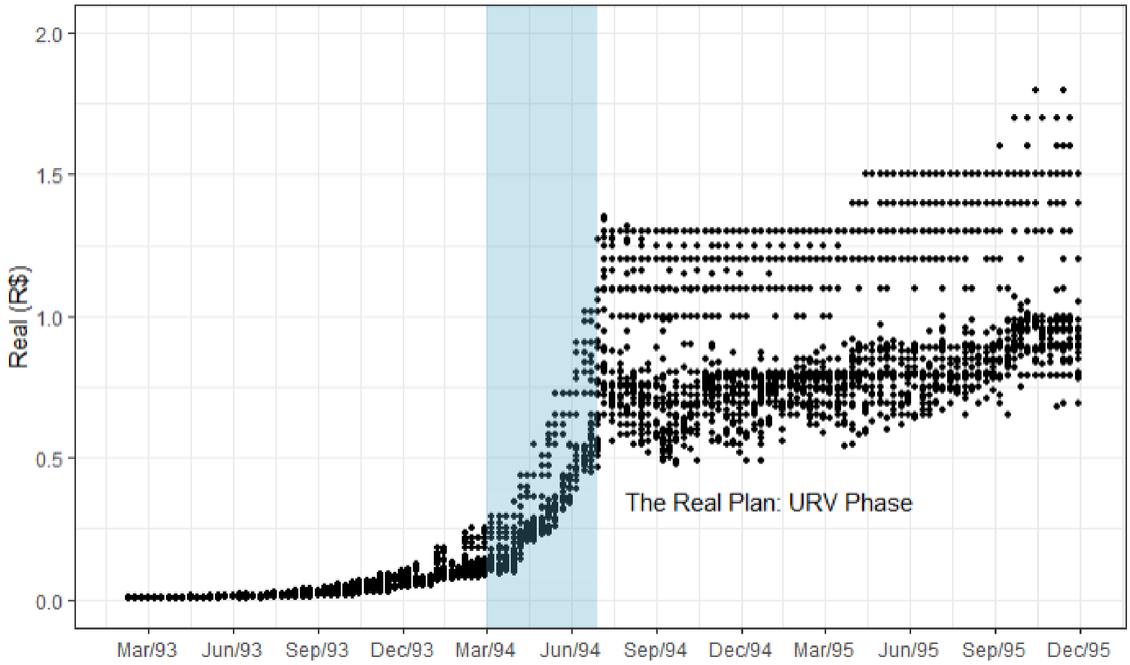
Notes: It illustrates the weekly weighted average of the CR$/URV during the URV phase of the Real Plan (from March to June 1994). Based on FIPE evidence, it assigns double chances that prices were collected in each of the first three working days of a week.

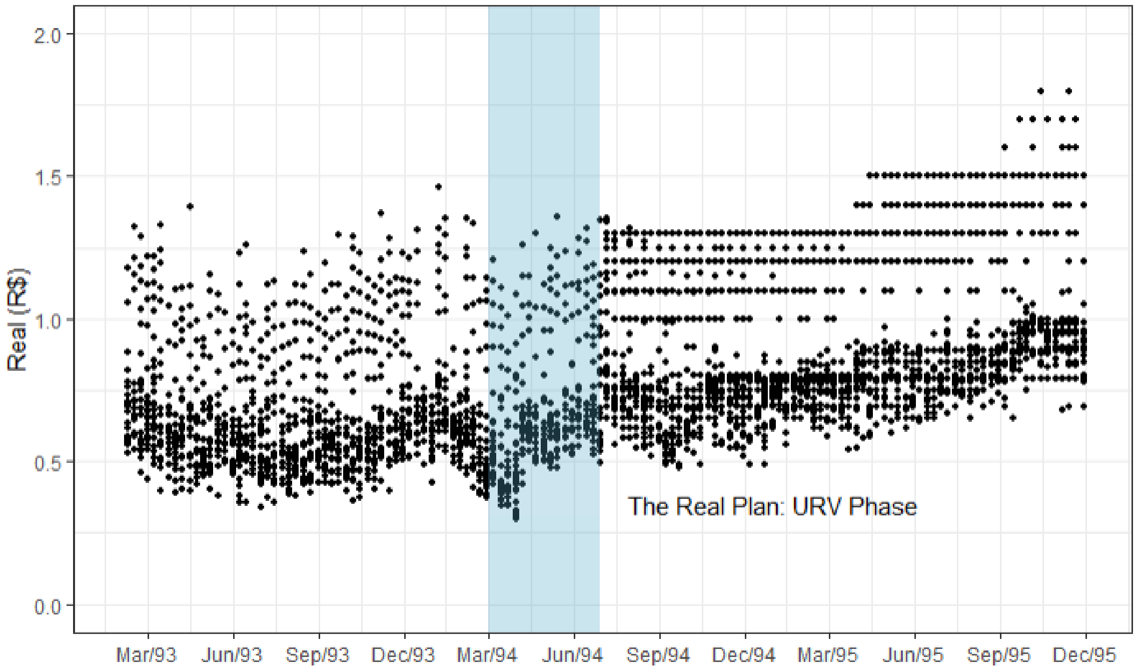
Contrastingly, all the previous studies on price setting during the hyperinflationary period in Brazil only use the official fixed conversion rates to have all prices in R$, that is, R$ 1 = CR$ 2,750 = Cr$ 2,750.000 (see Table 1), without regard to the URV information.[15] To highlight and differentiate the CR$/URV conversion from the fixed one, in Figure 5, we illustrate the price trajectories of the brand exemplified in Table 2 whereby each point represents a price quoted in a specific outlet in a given week. While the top graph uses prices in R$ under the fixed conversion procedure, the bottom one shows prices in R$ based on the weekly weighted average of the CR$/URV. The former shows a dramatic rise in CR$ prices in the first half of 1994, but the latter indicates a less pronounced URV price change which ultimately signals that URV prices would be much

[15] See Angelis (2012), Araujo and Rodrigues (2018), and Araujo (2019).

more stable than CR$ prices despite evidence of some URV price inflation during the stabilization phase.

Figure 5 - Price Trajectory under the Fixed and the CR$/URV Conversion





Notes: The graphs illustrate all the price trajectories of the brand exemplified in Table 2 whereby each point represents a price quoted in a specific outlet in a given week. The top graph uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000). The bottom graph uses the weekly weighted average of the CR$/URV to convert prices to R$.

In the Appendix, Figure A3 confirms it through the corresponding item monthly inflation and suggests how the distribution of inflation became more symmetric and concentrated toward zero with the URV conversion. Consequently, the increased likelihood of using the URV as an alternative unit of account could have led policymakers to introduce the R$ as the new monetary unit substituting both the CR$ and the URV after only 4 months.

Before concluding this section, a question that emerges is how imprecise is the estimative for the CR$/URV. This is important since there are two sources of measurement error on it – uncertainty about the price survey exact date and accuracy of the point estimate.

Let $P_{i,t}^{CR\$}$ and $\pi_{i,t}^{CR\$}$ be the CR$ price and inflation of item *i* at time *t*, V_t^{URV} and π_t^{URV} the weekly weighted average of the CR$/URV and its resulting inflation at *t*, and $P_{i,t}^{URV}$ and $\pi_{i,t}^{URV}$ the converted URV price and inflation of *i* at *t*. Then, we can relate $\pi_{i,t}^{URV}$ with $\pi_{i,t}^{CR\$}$ in the following way:

$$\pi_{i,t}^{URV} = \frac{P_{i,t}^{URV}}{P_{i,t-4}^{URV}} - 1 = \frac{P_{i,t}^{CR\$}/V_t^{URV}}{P_{i,t-4}^{CR\$}/V_{t-4}^{URV}} - 1 = \frac{P_{i,t}^{CR\$}/P_{i,t-4}^{CR\$}}{V_t^{URV}/V_{t-4}^{URV}} - 1 = \frac{1+\pi_{i,t}^{CR\$}}{1+\pi_t^{URV}} - 1 = \frac{\pi_{i,t}^{CR\$}-\pi_t^{URV}}{1+\pi_t^{URV}} \quad (1)$$

To evaluate the CR$/URV estimate, first, we compute how different the weekly weighted average of the CR$/URV ($V_t^{URV}$) is from the actual CR$/URV announced quotes (on working days) during the URV period. We find that it is on average 0.64% lower than the actual CR$/URV, which is reasonable since the CR$/URV values are usually lower during the first days of a week.[16] Additionally, we compare the monthly inflation of the weekly weighted average of the CR$/URV ($\pi_t^{URV}$) with those based on the actual CR$/URV values. The average difference is -0.03 percentage points (p.p.), thus very close to zero. Second, given $\pi_{i,t}^{CR\$}$, we compute how much $\pi_{i,t}^{URV}$ that uses π_t^{URV} is different from those that use the actual CR$/URV inflation. We find an average difference of 0.02 p.p. (or 0.83 p.p. in absolute value) during the URV period. Third, we assume that CR$ prices follow the actual CR$/URV variation through timely monthly updates during the URV period. Then we calculate $\pi_{i,t}^{URV}$ for these changing values of $\pi_{i,t}^{CR\$}$ given the

[16] The variance and the standard deviation of the differences between the weekly weighted average of the CR$/URV and the actual CR$/URV values are 0.07% and 2.65%, respectively.

corresponding π_t^{URV}. On average, we find that $\pi_{i,t}^{URV}$ varies by 0.02 p.p. (or 0.78 p.p. in absolute value), which is very similar to the previous estimative. Taken together, all of these simulations point out the weekly weighted average of the CR$/URV is a satisfactory point estimate given its relatively small deviation from actual values.

Furthermore, in Equation 1, note that if $\pi_{i,t}^{CR\$} = \pi_t^{URV}$, then $\pi_{i,t}^{URV} = 0$. That is, if the CR$ inflation of an item equals the inflation of the weekly weighted average of the CR$/URV, then its converted URV inflation should be zero. However, given that we convert prices to URV using the corresponding weekly CR$/URV, it is unlikely that the ratio $P_{i,t}^{URV}/P_{i,t-4}^{URV}$ would remain stable across two consecutive periods. In other words, one would necessarily find evidence of a URV price change (which is represented by an indicator function $I_{i,t}^{URV} = 1$). As a result, we represent a URV price variation by setting an inflation threshold $\bar{\pi}_{i,t}^{URV}$ instead of checking for $P_{i,t}^{URV} \neq P_{i,t-4}^{URV}$.[17]

$$I_{i,t}^{URV} = \begin{cases} 1, \; if \; \left|\pi_{i,t}^{URV}\right| > \bar{\pi}_{i,t}^{URV} \; \leftrightarrow \; \pi_{i,t}^{URV} > \bar{\pi}_{i,t}^{URV} \; or \; \pi_{i,t}^{URV} < -\bar{\pi}_{i,t}^{URV} \\ 0, \; if \; \left|\pi_{i,t}^{URV}\right| \leq \bar{\pi}_{i,t}^{URV} \; \leftrightarrow \; -\bar{\pi}_{i,t}^{URV} \leq \pi_{i,t}^{URV} \leq \bar{\pi}_{i,t}^{URV} \end{cases}, \quad (2)$$

$$\text{for } t = 1, \; \dots, \; T \text{ and } i = 1, \; \dots, \; I$$

Consequently, our baseline results consider $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ for all i and t to define a threshold $\bar{\pi}_{i,t}^{URV}$ that varies according to π_t^{URV} (see Equation 1). As a result, if the CR$ inflation of an item exceeds the CR$/URV inflation by more than (less than or equal to) 1 p.p. in absolute terms, then the converted URV inflation of an item is higher than (less than or equal to) the threshold so that there is (not) a price change. To illustrate the rationale, given the Table 4 official CR$/URV monthly rates of 43.3% (March) and 46.6% (June), the corresponding thresholds are 0.7% and 0.68%, respectively, which are roughly in line with the small impreciseness of our estimative (V_t^{URV}) suggested by the exercises above.

[17] Specifically, if the indicator function of URV-converted price change had been defined in terms of $P_{i,t}^{URV} \neq P_{i,t-4}^{URV}$, it would be unlikely price stability across two consecutive periods given the data characteristics.

$$I_{i,t}^{URV} = \begin{cases} 1, \; if \; P_{i,t}^{URV} \neq P_{i,t-4}^{URV} \; \leftrightarrow \; \pi_{i,t}^{URV} \neq 0 \\ 0, \; if \; P_{i,t}^{URV} = P_{i,t-4}^{URV} \; \leftrightarrow \; \pi_{i,t}^{URV} = 0 \end{cases}$$

1.3. Price-Setting Statistics

In this section, we review, compute, and discuss a series of pricing statistics before, during, and after the URV phase of the Real Plan using the FIPE microdata and the URV official information under two price conversion approaches – the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750.000) and the weekly weighted average of the CR$/URV.

1.3.1. Statistics Review

Following Araujo and Rodrigues (2018), we briefly summarize the relevant pricing statistics while adjusting for the URV particularities.

The most disaggregated unit in our analysis is item *i*. It represents a unit of a brand of a product *j* sold at a certain outlet *n* in a particular week *t*. Specifically, each product is the aggregation of different brands, which are composed of different items (as exemplified in Table 2). For each *i*, we have the respective price information $P_{i,n,t}$.

Consider an indicator function ($I_{i,n,t}$) that takes a unit value in case there is a price change through comparison between the same weeks in consecutive months. As initially discussed in Section 2, we differentiate price changes in CR$ ($\pi_{i,t}^{CR\$} \neq 0$) from the ones in URV ($\left|\pi_{i,t}^{URV}\right| > \bar{\pi}_{i,t}^{URV}$) to account for estimates of the weekly URV that prevailed when each item was surveyed. Nevertheless, one can generally represent a price change function as:

$$I_{i,n,t} = \begin{cases} 1, \; if \; P_{i,n,t} \neq P_{i,n,t-4} \\ 0, \; if \; P_{i,n,t} = P_{i,n,t-4} \end{cases}, \tag{3}$$

$$\text{for } i = 1, \; \dots, \; I, n = 1, \dots, N, \text{ and } t = 1, \dots, T$$

Let $\pi_{i,n,t}$ be the inflation of an item *i* sold at outlet *n* at time *t* and $\pi_{i,t}$ the average inflation of the brand (of *i*) across its *n* sellers. Then, we obtain the average inflation of product *j* across all of its brands ($\pi_{j,t}$), where $card(s_{j,t})$ is the number of brands within

j. Subsequently, we can calculate a measure of aggregate inflation, that is, the overall weighted inflation (π_t) using the product-specific FIPE CPI weights ($w_{j,t}$).[18]

$$\pi_t = \sum_{j=1}^{J} w_{j,t}\, \pi_{j,t}, \tag{4}$$
$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,\ \pi_{j,t} = \frac{1}{card(s_{j,t})} \sum_{i \in s_{j,t}} \pi_{i,t},$$
$$\pi_{i,t} = \frac{1}{n_{i,t}} \sum_n \pi_{i,n,t}, \text{ and } \pi_{i,n,t} = \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}}$$
$$\text{for all } i, n, t, \text{ and } j$$

The frequency of price change or extensive margin (f_t) indicates the proportion of products that changed prices relative to the same week of the previous month. To compute it, we aggregate i into j through simple means to find the average frequency of price change for j ($f_{j,t}$), where $card(s_{j,t})$ is the number of non-missing brands that composes j.[19] Afterward, we obtain the overall f_t using the official CPI weights.

$$f_t = \sum_{j=1}^{J} w_{j,t}\, f_{j,t}, \tag{5}$$
$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,\ f_{j,t} = \frac{1}{card(s_{j,t})} \sum_{i \in s_{j,t}} I_{i,t}, \text{ and } I_{i,t} = \frac{1}{n_{i,t}} \sum_n I_{i,n,t}$$
$$\text{for all } i, n, t, \text{ and } j$$

The average absolute size of a price change or intensive margin (ΔP_t) appraises situations in which there exist price changes (that is, $I_{i,n,t} = 1$). It shows how much prices have changed in absolute terms over consecutive periods (correspondent weeks in two consecutive months). Once again, we aggregate brands into products ($\Delta P_{j,t}$) using a simple means, where $card(s_{j,t}^{*})$ represents the number of non-zero price changes of i within j, before computing the general ΔP_t.

$$\Delta P_t = \sum_{j=1}^{J} w_{j,t}\, \Delta P_{j,t}, \tag{6}$$
$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,\ \Delta P_{j,t} = \frac{1}{card(s_{j,t}^{*})} \sum_{i \in s_{j,t}^{*}} \Delta P_{i,t},$$

[18] The product weights are recalculated within each group or category so that it adds up to unity every FIPE week while preserving the household budget structure.
[19] Before aggregating brands into products, we aggregate all of the units that compose a brand using sample mean.

$$\Delta P_{i,t} = \frac{1}{n_{i,t}} \sum_n \Delta P_{i,n,t}, \text{ and } \Delta P_{i,n,t} = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} \neq 0 \right|$$

for all $i, n, t,$ and j

In short, these indicators provide an overview of the proportion of products that changed prices over two consecutive months and, for those that altered their price, the amount of that change. Furthermore, we decompose f_t and ΔP_t into the frequency of price increases (f_t^+) and decreases (f_t^-), and the size of price increases (Δp_t^+) and decreases (Δp_t^-), respectively.[20, 21]

$$f_t = f_t^+ + f_t^- \text{ such that} \tag{7}$$

$$f_t^+ = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^+)} \sum_{i \in s_{j,t}^+} I_{i,t}^+ \right),$$

$$\text{where } I_{i,t}^+ = \frac{1}{n_{i,t}} \sum_n I_{i,n,t}^+, \text{ and } I_{i,n,t}^+ = \begin{cases} 1, \ if \ P_{i,n,t} > P_{i,n,t-4} \\ 0, \ if \ P_{i,nt} \leq P_{i,n,t-4} \end{cases}$$

$$f_t^- = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^-)} \sum_{i \in s_{j,t}^-} I_{i,t}^- \right),$$

$$\text{where } I_{i,t}^- = \frac{1}{n_{i,t}} \sum_n I_{i,n,t}^-, \text{ and } I_{i,n,t}^- = \begin{cases} 1, \ if \ P_{i,n,t} < P_{i,n,t-4} \\ 0, \ if \ P_{i,nt} \geq P_{i,n,t-4} \end{cases}$$

for all $i, n, t,$ and j

$$\Delta P_t = \frac{f_t^+}{f_t} \Delta P_t^+ + \frac{f_t^-}{f_t} \Delta P_t^- \text{ such that} \tag{8}$$

$$\Delta P_t^+ = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^{+*})} \sum_{i \in s_{j,t}^{+*}} \Delta P_{i,t}^+ \right),$$

$$\text{where } \Delta P_{i,t}^+ = \frac{1}{n_{i,t}} \sum_n \Delta P_{i,n,t}^+, \text{ and } \Delta P_{i,n,t}^+ = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} > 0 \right|$$

$$\Delta P_t^- = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^{-*})} \sum_{i \in s_{j,t}^{-*}} \Delta P_{i,t}^- \right),$$

---

[20] For the frequency decomposition using URV prices, we adapt Equation 7 so that:

$$I_{i,n,t}^+ = \begin{cases} 1, \ if \ P_{i,n,t} > P_{i,n,t-4} \\ 0, \ if \ P_{i,n,t} \leq P_{i,n,t-4} \end{cases} \rightarrow I_{i,n,t}^{URV+} = \begin{cases} 1, \ if \ \pi_{i,n,t}^{URV} > \bar{\pi}_{i,n,t}^{URV} \\ 0, \ if \ \pi_{i,n,t}^{URV} \leq \bar{\pi}_{i,n,t}^{URV} \end{cases}$$

$$I_{i,nt}^- = \begin{cases} 1, \ if \ P_{i,n,t} < P_{i,n,t-4} \\ 0, \ if \ P_{i,n,t} \geq P_{i,n,t-4} \end{cases} \rightarrow I_{i,t}^{URV-} = \begin{cases} 1, \ if \ \pi_{i,n,t}^{URV} < -\bar{\pi}_{i,n,t}^{URV} \\ 0, \ if \ \pi_{i,n,t}^{URV} \geq -\bar{\pi}_{i,n,t}^{URV} \end{cases}$$

[21] For the absolute size decomposition using URV prices, we adapt Equation 8 so that:

$$\Delta P_{i,n,t}^+ = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} > 0 \right| \rightarrow \Delta P_{i,n,t}^{URV+} = \left| \frac{P_{i,n,t}^{URV} - P_{i,n,t-4}^{URV}}{P_{i,n,t-4}^{URV}} > \bar{\pi}_{i,n,t}^{URV} \right|$$

$$\Delta P_{i,n,t}^- = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} < 0 \right| \rightarrow \Delta P_{i,n,t}^{URV-} = \left| \frac{P_{i,n,t}^{URV} - P_{i,n,t-4}^{URV}}{P_{i,n,t-4}^{URV}} < -\bar{\pi}_{i,n,t}^{URV} \right|$$

$$\text{where } \Delta P^{-}_{i,t} = \frac{1}{n_{i,t}} \sum_n \Delta P^{-}_{i,n,t}, \text{ and } \Delta P^{-}_{i,n,t} = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} < 0 \right|$$

$$\text{for all } i, n, t, \text{ and } j$$

In parallel, we use the intra-market relative price variability (RPV) indicators to verify the degree of price synchronization. It evaluates the dispersion of price adjustments among all outlets selling the same brand. There are two commonly used RPV statistics in the literature – the standard deviation of price changes across outlets ($SDP_{i,t}$) and the coefficient of variation of price levels across stores ($CV_{i,t}$). While the SDP computes deviations of the inflation of a given brand to its average across all of its sellers, the CV measures discrepancies between the price of a given brand and its average market price. Similar to the previous price statistics, we compute it for items and then aggregate it into products to find an aggregate measure. The SDP is defined as:

$$SDP_t = \sum_{j=1}^{J} w_{j,t} \, SDP_{j,t}, \tag{9}$$

$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,$$

$$SDP_{j,t} = \frac{1}{card(s_{j,t})} \sum_{i \in s_{j,t}} SDP_{i,t}, \text{ and } SDP_{i,t} = \left[\frac{1}{n_{i,t}} \sum_n \left(\pi_{i,n,t} - \pi_{i,t} \right)^2 \right]^{1/2}$$

$$\text{for all } i, n, t, \text{ and } j$$

Finally, given that $P_{i,t}$ is the average price level of *i* at *t*, then the CV is:

$$CV_t = \sum_{j=1}^{J} w_{j,t} \, CV_{j,t}, \tag{10}$$

$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,$$

$$CV_{j,t} = \frac{1}{card(s_{j,t})} \sum_{i \in s_{j,t}} CV_{i,t}, \text{ and } CV_{i,t} = \left[\frac{1}{n_{i,t}} \sum_n \left(\frac{P_{i,n,t} - P_{i,t}}{P_{i,t}} \right)^2 \right]^{1/2}$$

$$\text{for all } i, n, t, \text{ and } j$$

1.3.2. Price-Setting Evidence

We introduced two different approaches to converting prices to R$ – the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750.000) and the weekly weighted average of the CR$/URV – and reviewed price-setting statistics. Now, we compute these pricing indicators using both conversion methods to document the characteristics of the microdata on prices at the time of the Real Plan stabilization program that ended hyperinflation in Brazil. Specifically, our interest lies primarily in verifying if the URV policies affected prices, which should exhibit an increased level of synchronization besides being relatively closer to zero and more symmetric.

First, we calculate the price-setting statistics of inflation, frequency, absolute size of price changes, their respective separation into positive and negative changes, and their heterogeneous behavior across product categories. Second, we illustrate movements in the distribution of monthly price changes across time through kernel density estimation (KDE). Third, we evaluate the price synchronization degree via brand-level intra-market relative price variability (RPV) indicators, namely the coefficient of variation of price levels (CV) and the standard deviation of price changes (SDP). All statistics are initially computed at the most disaggregated level (i.e., item), then successively aggregated to obtain overall indexes using the product-specific FIPE CPI weights.

Table 5 summarizes the main results of inflation, frequency, and absolute size of price changes using both conversion approaches for the pre-URV, URV, and post-URV periods. The statistics are computed monthly by taking the weekly statistics' average within a month before obtaining the respective means for each phase. Importantly, as discussed in Section 2, we arbitrarily set $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ in combination with a varying π_t^{URV} to define price change thresholds for each t ($\bar{\pi}_{i,t}^{URV}$) specifically for assessing the frequency and absolute size of price changes in URV. In the Appendix, Table A3 illustrates the sensibility of these statistics for different levels of $\pi_{i,t}^{CR\$} - \pi_t^{URV}$ during the URV stabilization phase as a robustness check.

Additionally, Figures 6-10 display the time series of the inflation, frequency, and absolute size of price change statistics under both price conversion approaches while exploring possible heterogeneities among groups of products – food at home, industrial goods, services, and regulated prices. In the Appendix, Figures A4-A6 complement the analysis by computing these pricing statistics under the FIPE classification, which separates products into food, housing, transportation, personal expenses, apparel, healthcare, and education categories, while Figures A7-A9 evaluate the statistics after separating products into tradables, non-tradables, and regulated prices.

Table 5 - Inflation, Frequency, and Absolute Size of Price Change (Fixed and CR$/URV Conversion)

Currency	CR$							URV | R$						
Statistics	**Infl.**	**Size**	**Size +**	**Size -**	**Freq.**	**Freq. +**	**Freq. -**	**Infl.**	**Size**	**Size +**	**Size -**	**Freq.**	**Freq. +**	**Freq. -**
Pre-URV	**39.1**	**41.3**	**42.0**	**8.9**	**96.8**	**94.1**	**2.7**	**2.6**	**13.5**	**15.6**	**11.1**	**92.8**	**49.2**	**43.6**
Aug-93	36.6	38.9	39.5	8.0	96.1	93.9	2.2	4.5	13.8	15.6	11.5	87.2	54.7	32.4
Sep-93	35.4	38.1	38.6	8.8	95.5	93.0	2.5	1.6	13.2	14.7	11.5	88.8	50.3	38.5
Oct-93	37.2	39.2	39.7	7.7	97.0	95.0	2.0	0.9	12.1	15.2	10.4	94.1	40.1	54.0
Nov-93	38.1	40.0	40.5	9.0	97.0	95.2	1.7	3.4	12.1	13.6	11.1	95.2	62.4	32.8
Dec-93	44.2	46.3	47.1	6.8	96.8	94.4	2.5	5.3	15.2	17.7	10.6	92.4	56.3	36.1
Jan-94	41.3	43.3	44.7	8.7	98.1	92.6	5.5	1.8	14.9	17.5	11.2	95.5	41.3	54.2
Feb-94	41.2	43.3	43.9	13.6	97.4	94.9	2.5	0.4	13.3	15.0	11.2	96.5	39.6	56.9
URV	**49.1**	**50.7**	**51.5**	**9.6**	**98.5**	**96.1**	**2.4**	**4.5**	**14.6**	**16.4**	**11.3**	**95.1**	**56.0**	**39.1**
Mar-94	45.1	47.0	47.8	8.7	98.3	95.5	2.8	3.7	14.7	16.3	12.0	94.1	55.8	38.4
Apr-94	49.3	51.1	52.1	11.4	98.4	95.6	2.8	3.5	14.5	16.6	11.7	95.8	50.0	45.8
May-94	48.4	50.0	50.7	9.8	98.6	96.3	2.3	4.8	15.3	17.2	11.2	94.2	56.0	38.2
Jun-94	53.4	54.7	55.3	8.8	98.8	97.1	1.7	6.0	14.0	15.6	10.1	96.2	62.1	34.1
Post-URV	**-**	**-**	**-**	**-**	**-**	**-**	**-**	**3.5**	**14.5**	**17.1**	**11.0**	**64.5**	**36.8**	**27.7**
Jul-94	-	-	-	-	-	-	-	9.9	16.0	18.1	7.9	100.0	66.7	33.3
Aug-94	-	-	-	-	-	-	-	2.3	14.8	17.2	11.6	49.7	28.4	21.3
Sep-94	-	-	-	-	-	-	-	0.7	12.9	15.6	11.0	58.9	24.9	34.0
Oct-94	-	-	-	-	-	-	-	3.3	13.7	19.2	10.5	67.8	32.6	35.3
Nov-94	-	-	-	-	-	-	-	3.5	14.8	16.7	11.8	57.9	37.0	20.9
Dec-94	-	-	-	-	-	-	-	1.5	14.5	15.6	12.9	52.5	31.4	21.1

Notes: The table provides monthly price setting indicators (in %) alongside its averages for the pre-URV, URV, and post-URV periods. It uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) and the weekly weighted average of the CR$/URV to convert prices to R$. The CR$/URV price change thresholds ($\bar{\pi}_{i,t}^{URV}$) assume $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ (while being time-variant dependent from π_t^{URV}) and affect the URV statistics of frequency and absolute size of price change.

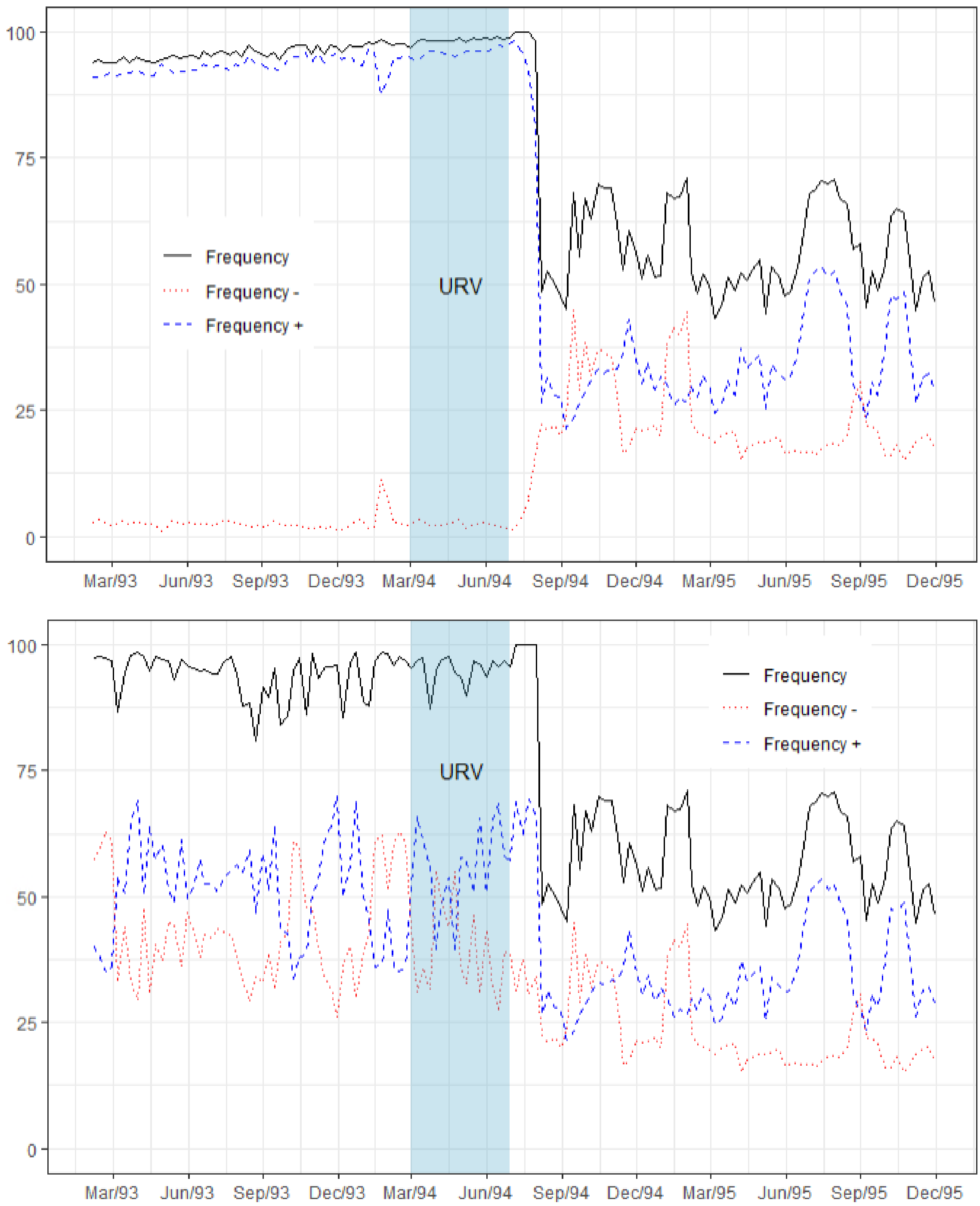
In short, our results reinforce that the Real Plan significantly altered price-setting characteristics in Brazil, but most of the stabilization effects emerged only after introducing the R$. Accordingly, prices changed less frequently, by a lower amount, and in a more evenly distributed way following the URV period. Nevertheless, the monetary change marks a more subtle change in URV compared to CR$, which suggests that URV was relatively efficient as a stabilization tool.

Figure 6 sets out the CR$ and URV inflation in the top and bottom graphs, respectively. In a context of greater uncertainty, especially during the URV phase of the Real Plan, there is evidence that CR$ monthly inflation rose steadily. In parallel, the URV inflation was different from zero, but relatively smaller, indicating that URV-converted prices were unstable and varied more than the official URV rates. Moreover, there are heterogeneities among groups, especially in the regulated prices category whose behavior differs substantially from the market ones.

Figure 6 - Inflation (Fixed and CR$/URV Conversion)

Notes: The top graph uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) while the bottom one uses the weekly weighted average of the CR$/URV to convert prices to R$. The highlighted area represents the Real Plan's URV phase.

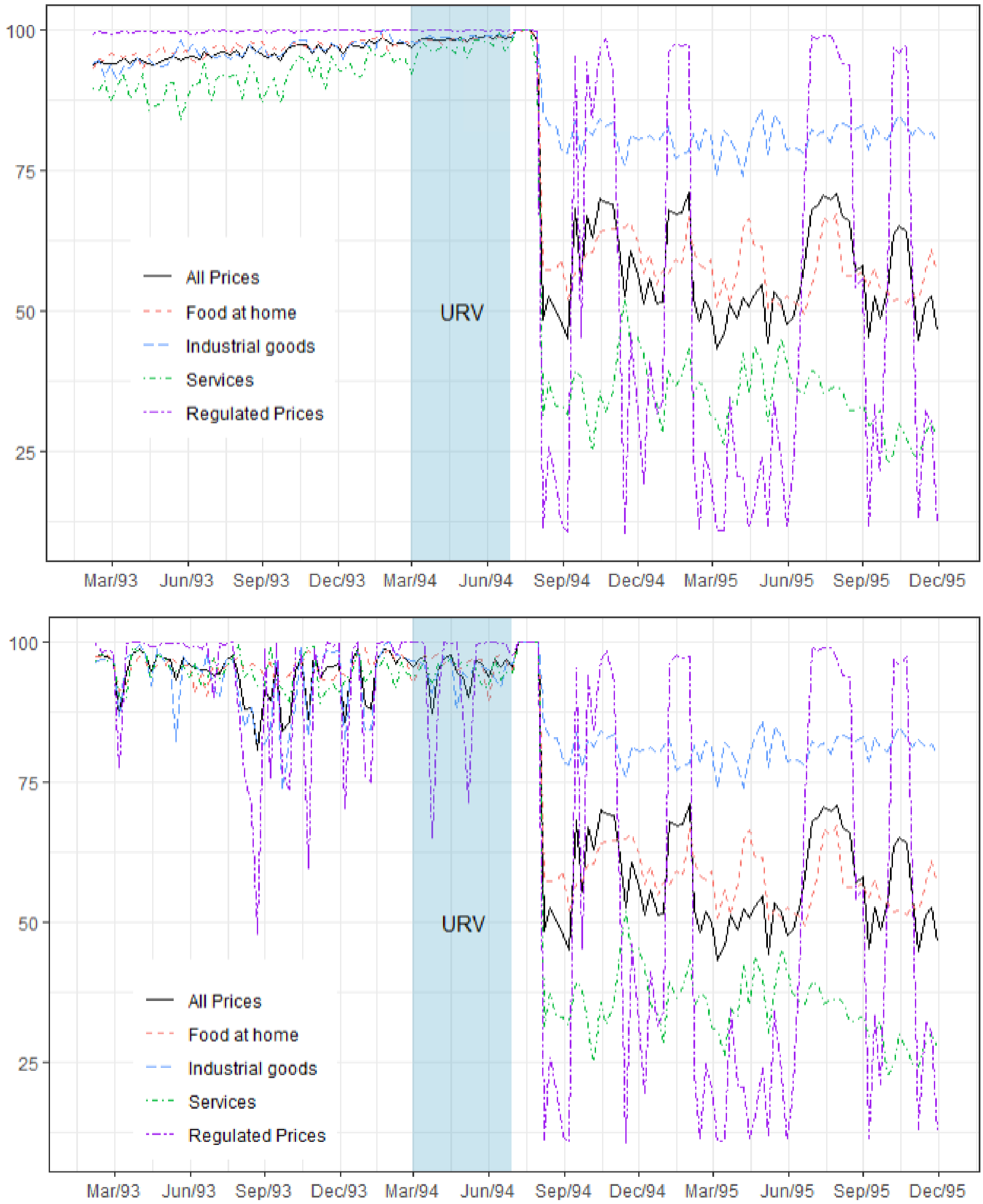
Figure 7 - Frequency of Price Change (Fixed and CR$/URV Conversion)



Notes: See explanations in Figure 6. The CR$/URV price change thresholds ($\bar{\pi}_{i,t}^{URV}$) assume $\pi_{i,t}^{CR\$} - \pi_t^{URV} =$ 1% while being time-variant dependent from π_t^{URV}.

The frequency of price adjustment informs the proportion of products that changed prices from the same week of the previous month. Figure 7 shows prices in CR$ changed very frequently during the URV phase, but the frequency is highly dominated by positive changes in such a hyperinflationary scenario (see graph above). On the other hand, the frequency in URV, despite being similarly elevated, is much more even with a relevant proportion of negative price changes before the monetary change (see graph below). Afterward, both the positive and negative frequency of changes reduce so that the overall frequency declines remarkably. Figure 8 adds on the frequency disaggregation among product categories.

Figure 8 - Frequency of Groups of Products (Fixed and CR$/URV Conversion)



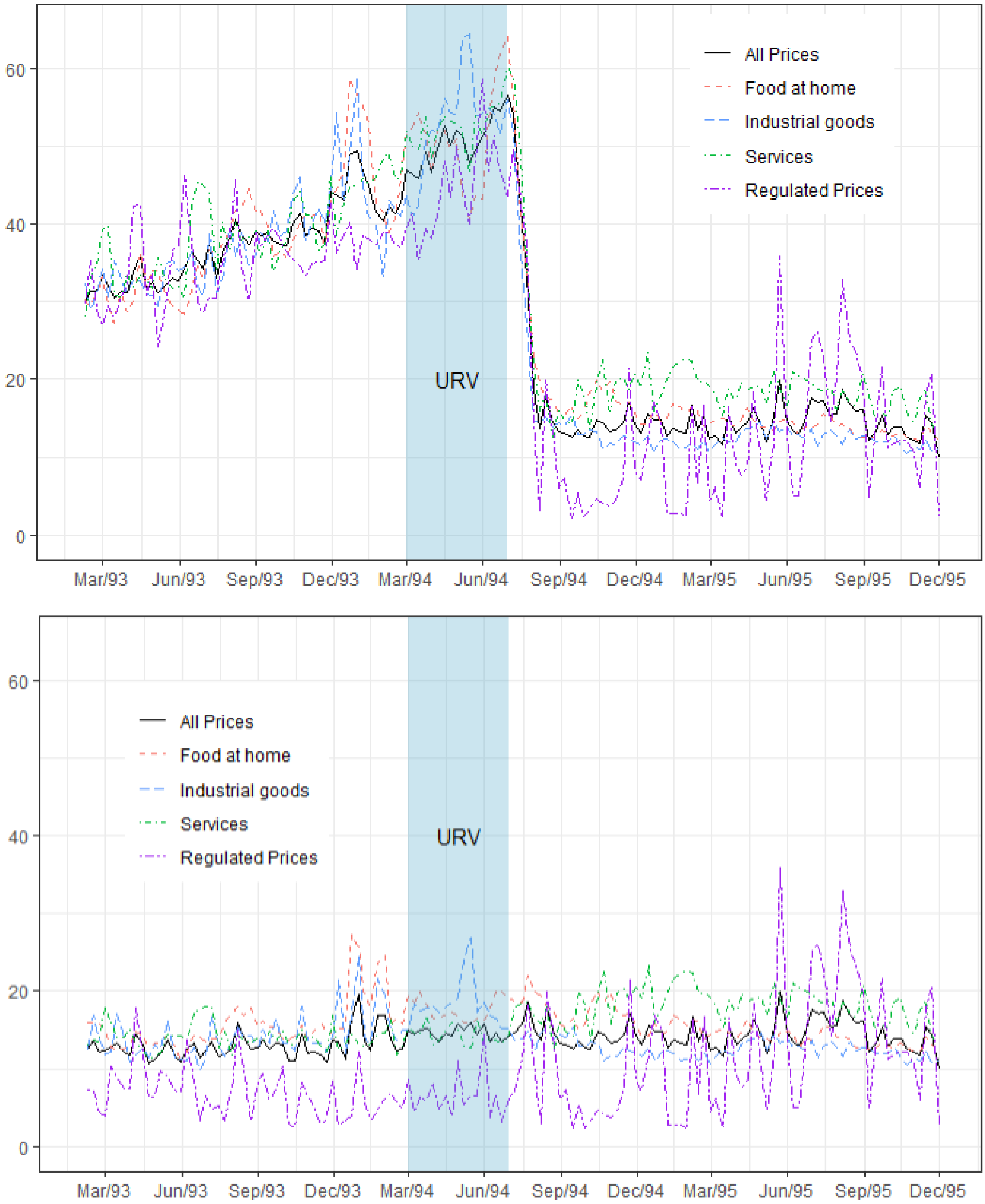
Notes: See explanations in Figure 7.

Figure 9 - Absolute Size of Price Change (Fixed and CR$/URV Conversion)

Notes: See explanations in Figure 7.

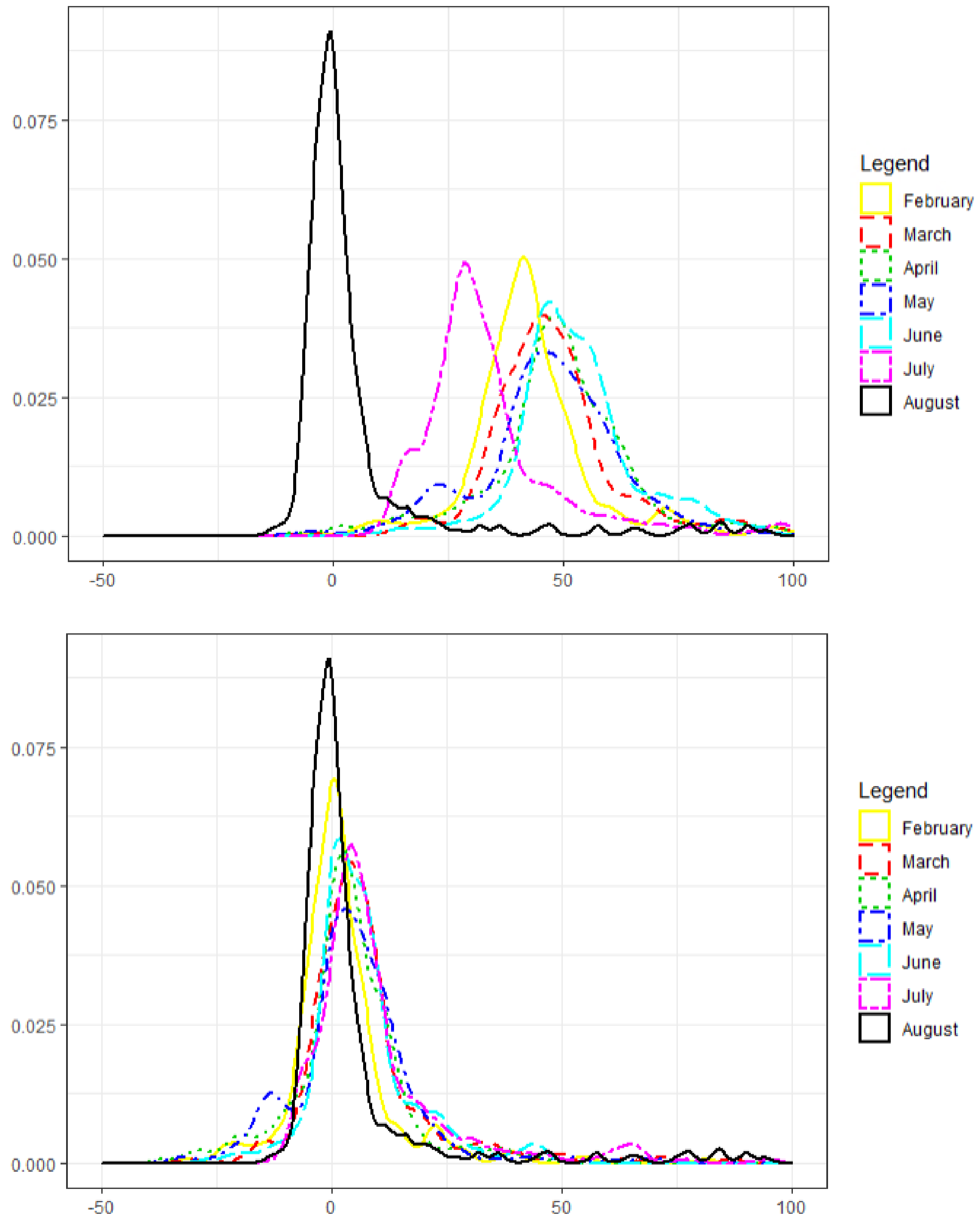
Figures 9 and 10 display the absolute size of price change and its breakdown into different groups, respectively. The absolute size of CR$ changes is quite similar to the CR$ inflation given the high frequency of price changes (notably increases) under that hyperinflationary environment. Besides, in terms of the URV, the absolute size of changes is higher than inflation because the latter considers a relevant share of price decreases within the frequency of changes.

Figure 10 - Absolute Size of Groups of Products (Fixed and CR$/URV Conversion)



Notes: See explanations in Figure 7.

Figure 11 - Distribution of Price Changes (Fixed and CR$/URV Conversion)



Notes: These figures illustrate the distributions of the average monthly inflation of products from February to August 1994. The top graph uses the official fixed conversion rate (R$ 1 = CR$ 2,750) while the bottom one uses the weekly weighted average of the CR$/URV to convert prices to R$.

Second, Figure 11 gives the product distribution of price changes across the stabilization months through KDE. Interestingly, the monthly distribution of changes in

CR$ started moving toward zero only in July of 1994 and then became more symmetric and concentrated around zero (see graph above). Hence, there is no meaningful adjustment when the URV coexisted with CR$. In contrast, the URV monthly distribution reinforces the role of the URV as a mechanism of price alignment to eliminate distribution asymmetry, as suggested by Almeida and Bonomo (2002). Specifically, after converting prices to URV, the distribution of inflation shows an immediate shift in terms of symmetry and concentration, even though there is still a marked difference compared to the higher levels of August 1994 (see graph below).

Third, the brand-level intra-market relative price variability indicators of the CV and SDP are depicted in Figures 12 and 13, respectively. Using either metric, the price synchronization degree does not improve significantly during the URV period irrespective of the conversion method (see Equation 10). At best, the RPV indicators started stabilizing around June 1994. Subsequently, with the monetary change, these statistics move to an enhanced level of synchronization.

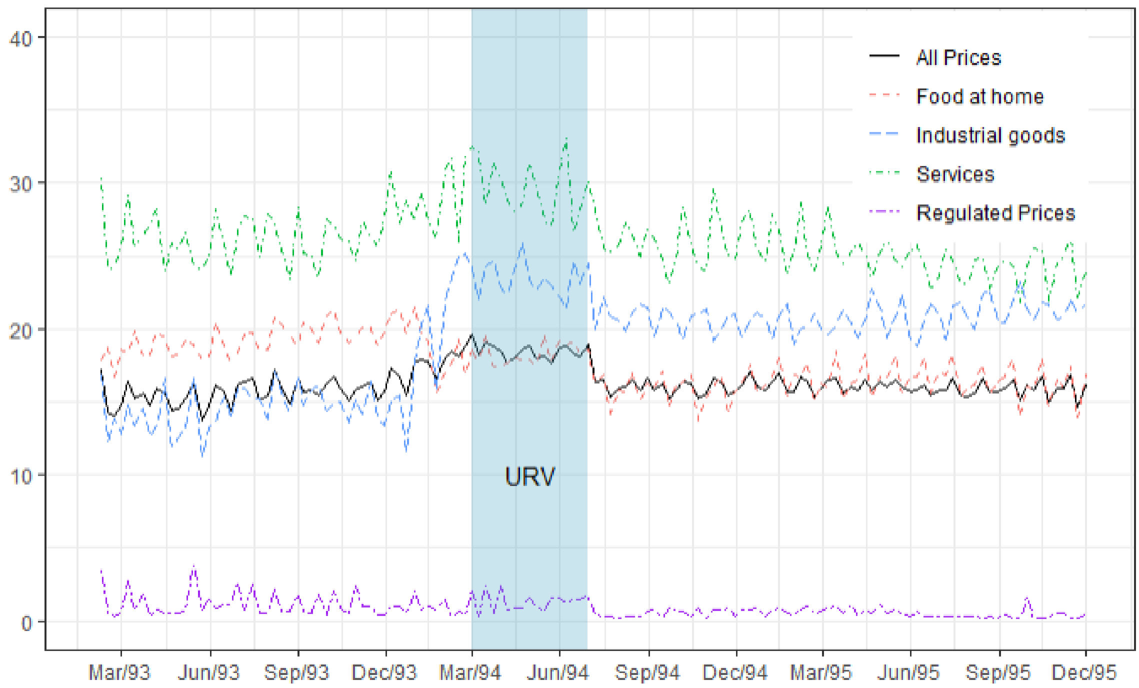
To conclude, we combine microdata of prices at the store level at the time of the Real Plan stabilization program with the URV official information so that our findings hinge on the underlying micro price adjustments. Our results indicate that the Real Plan promoted a major turnaround in the Brazilian price-setting statistics. Overall, prices changed less frequently, by a lower amount, and in a more evenly distributed way due to the program. Nevertheless, the pricing characteristics at the time of the URV (which supposedly coincides with the stabilization period) contain small evidence of that transformation, which manifests itself only after the monetary reform that launched the R$. If the URV were acting as expected, we should observe an immediate shift in the distribution of URV-converted prices accompanied by an increasing degree of price synchronization, but there is no evidence of the latter effect. Therefore, based on micro price adjustments, our results downplay the claim that the URV had a decisive role in altering the pricing dynamics and attaining sustainable price stability.

A few explanations contribute to finding muted supportive evidence of stabilization effects during the URV period. First, although the microdata structure is weekly-based, it is a monthly dataset since the quotation of a specific unit of analysis is usually once a month, so it possesses a relatively low frequency during the 4-month URV period. Second, as a rule, URV price quotations effectively started in May while price surveys reflected inflation with a delay of 2-4 weeks, thus reducing the chances of finding evidence during that spell. Third, a positive association between rising inflation and greater RPV, especially before the introduction of the R$ in a context of greater uncertainty, is also a factor that weakens stabilization signals.

Figure 12 - Standard Deviation of Price Changes (Fixed and CR$/URV Conversion)

Notes: See explanations in Figure 6.

Figure 13 - Coefficient of Variation of Price Levels (Fixed and CR$/URV Conversion)



Notes: The coefficient of variation of price levels is the same irrespective of the conversion method.

As a result, given these constraints, we cannot infer that URV did not matter. We shed light on how the mechanism enabled a milder transition to a new currency while promoting an immediate shift of the distribution of prices towards symmetry around zero. The URV policies also guaranteed real wage purchasing power during the stabilization reform, which certainly contributed to the URV's broad and voluntary acceptance – a feature that micro price adjustment cannot account for.

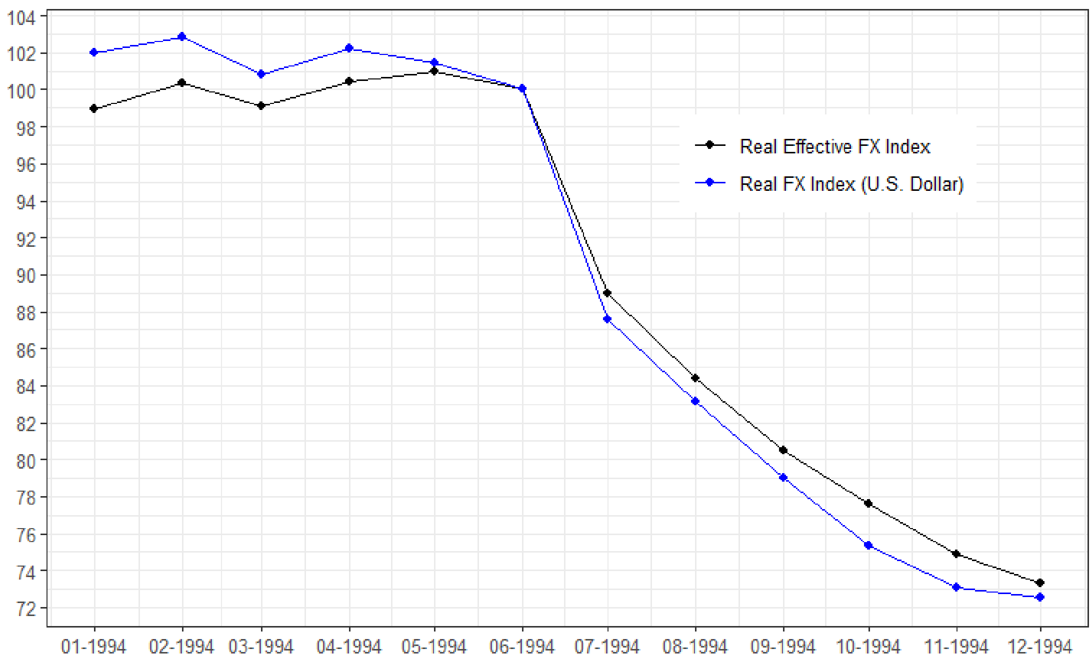
1.3.3. Foreign Exchange Rate Evidence

Following Brazil's successful transition out of the hyperinflationary era through the Real Plan stabilization program, a central question arises: why did the disinflationary effects become markedly more pronounced only after the monetary reform that introduced the R$? Empirical evidence suggests that the exchange rate policy adopted during this period played a pivotal role in shaping the observed outcomes.

During the implementation of the Real Plan, Brazil operated under a managed exchange rate regime, whereby the BCB actively intervened in foreign exchange markets to stabilize the value of the CR$ relative to the U.S. dollar. Although the URV was not a transactable currency, its valuation was tied to the dollar, thereby influencing the exchange rate framework. As explained in Section 2, the URV was introduced with an initial parity of approximately URV 1 = USD 1, and its daily quotation in CR$ was adjusted by the BCB based on prevailing inflation indices. This approach ensured that the URV retained a relatively stable nominal value vis-à-vis the dollar. Consequently, the URV functioned as a nominal anchor, signaling the government's commitment to exchange rate stability in the transition to the new currency.

Figure 14 illustrates two monthly indexes calculated by the BCB using the IPCA: the real effective exchange rate (REER) – which measures the Brazilian currency's value against a trade-weighted basket of foreign currencies – and the bilateral real exchange rate vis-à-vis the U.S. dollar. An upward trend in these indices indicates a real depreciation of the Brazilian currency, which can improve the trade balance by boosting export attractiveness, but it may simultaneously increase the domestic cost of imports, thereby exerting inflationary pressures. Conversely, a downward trend denotes a real appreciation of the national currency, which tends to facilitate imports and exert downward pressure on domestic prices.

Figure 14 - Monthly Real Foreign Exchange Rate Indexes (Jun/1994=100)



Notes: The real effective exchange rate (REER) and the real exchange rate to the U.S dollar indexes are monthly computed by the BCB using the IPCA (series 11752 and 11753, respectively).

Specifically, the real exchange rates remained relatively stable during the first half of 1994, including the URV phase of the Real Plan. Subsequently, from July 1994 onwards, given that interventions by the BCB limited nominal depreciations to maintain the exchange rate near parity with the U.S. dollar, the Brazilian Real experienced a real appreciation, as domestic inflation continued to exceed inflation in the United States. This appreciation facilitated imports and was consistent with the macroeconomic strategy of exerting downward pressure on domestic prices to support disinflation.

To assess whether this exchange rate policy evidence is supported by the FIPE microdata, we conduct an additional analysis of the pricing dynamics of tradable goods, non-tradables, and regulated prices. Both tradables and non-tradables are set by free market mechanisms, but tradable prices are directly influenced by exchange rate fluctuations, whereas non-tradable prices – including services – tend to be more sensitive to wage dynamics. In contrast, regulated prices are subject to certain limits imposed by regulatory authorities.

In the Appendix, Figures A7-A11 present statistics on inflation, frequency, absolute size, and standard deviation of price changes, as well as the coefficient of variation of price levels. Once again, we find no conclusive evidence of significant shifts in price dynamics during the URV phase, suggesting that stabilization effects primarily took hold with the monetary reform. Consequently, prices began to change less frequently, by smaller amounts, and in a more synchronized fashion. Also, the distribution of price adjustments became more symmetric and clustered around zero after the introduction of the new currency. Nevertheless, the patterns observed in these figures are consistent with the behavior of real exchange rates during that period, lending further support to the hypothesis that the exchange rate policy played a central role in the disinflationary process associated with the Real Plan.

1.4. Concluding Remarks

In this paper, we revisited the Real Plan – the economic stabilization program that put an end to hyperinflation in Brazil in 1994 – using unique micro data on prices at the establishment level of that period.

The main feature of the Real Plan is the URV, a virtual currency that preceded the introduction of a new monetary unit (R$) serving only as a temporary unit of account (from March to June 1994) while the cruzeiro real (CR$) was the official currency. The objective was to stimulate the synchronization of price adjustments in the economy so that it would reduce inertial inflation and achieve disinflation.

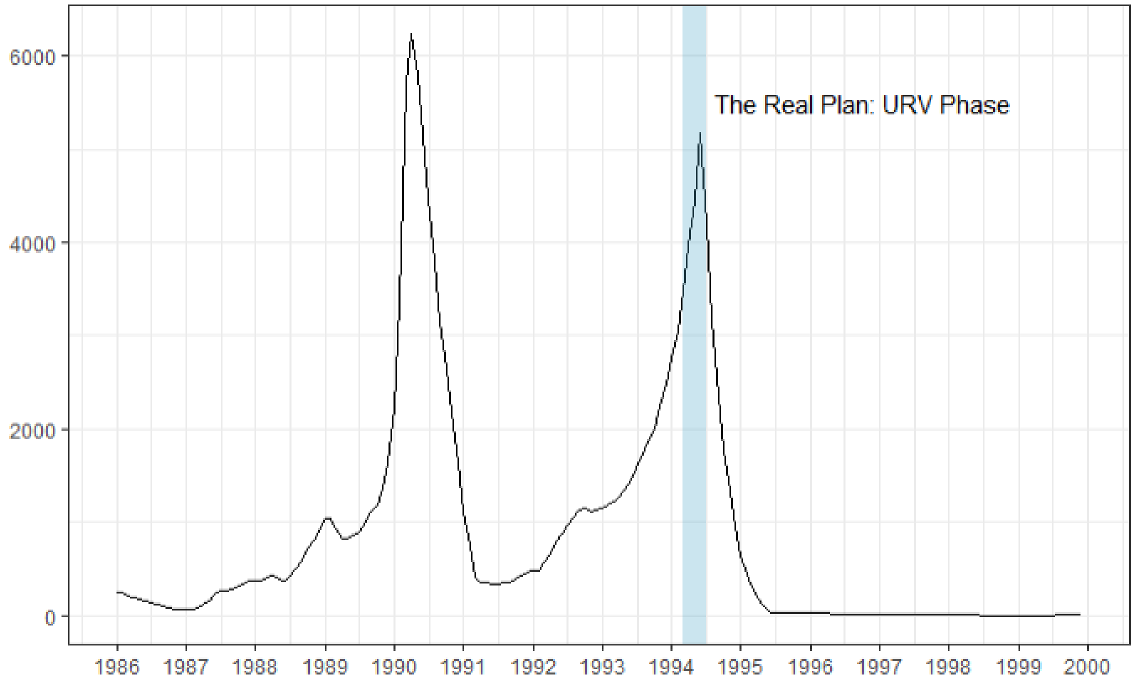
We characterized price setting during the URV phase of the Real Plan and compared it with the preceding and succeeding periods. Based on micro price adjustments, our results reinforced that the Real Plan changed Brazil's price-setting features. Nevertheless, we found that the stabilization effects emerged only after introducing the R$. Accordingly, prices changed less frequently, by a lower amount, and in a more synchronized way compared to previous periods, and the distribution of price adjustments became much more symmetric and concentrated around zero only after the monetary substitution. Besides, we provided evidence that the exchange rate policy contributed to this resulting outcome through a major real appreciation of the domestic currency following the monetary change.

As a result, notwithstanding some caveats related to the microdata and the URV period, we presented empirical evidence that downplays the role of the URV in altering the pricing dynamics and attaining sustainable price stability. In particular, there is muted evidence of price dynamic changes during the URV phase. However, these findings do not imply that the URV did not matter because the mechanism allowed for a gradual switch to the new currency while promoting a quick shift in the product’s inflation distribution, in addition to ensuring real wage purchasing power during the stabilization period – a feature that micro price adjustment cannot account for.

To conclude, despite the singularity of the Brazilian stabilization program, the puzzling results for Latin America’s largest economy in terms of disinflation may offer lessons and insights that could carry over elsewhere, especially in studies about the economic stabilization processes in developing countries.

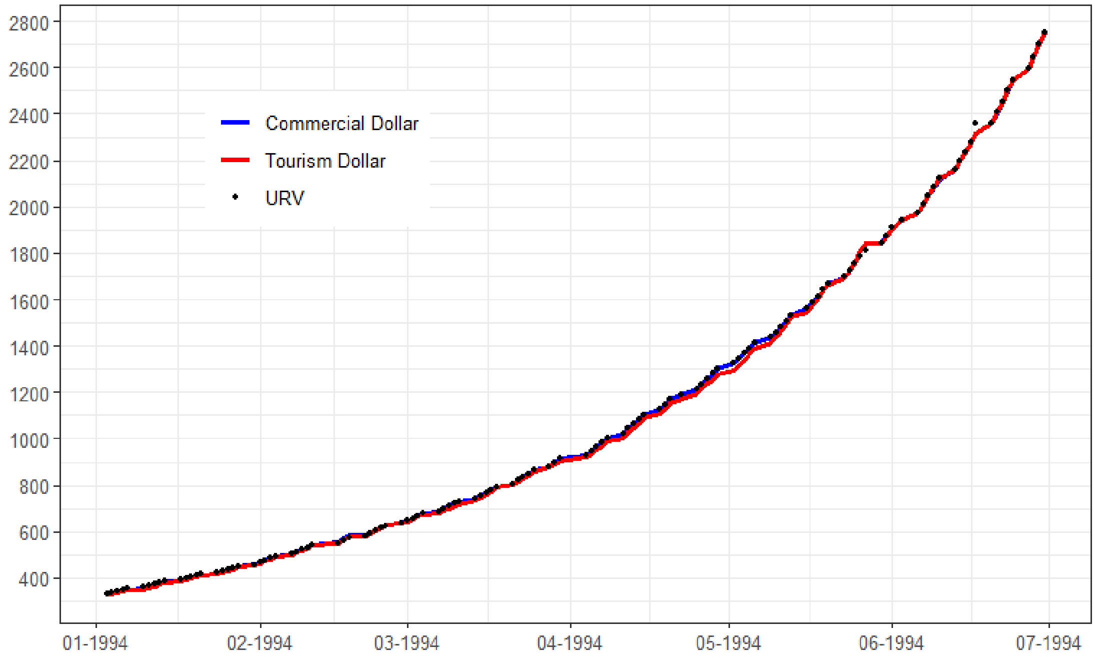
1.5. Appendix

Figure A1 - FIPE CPI: Yearly Inflation (%)



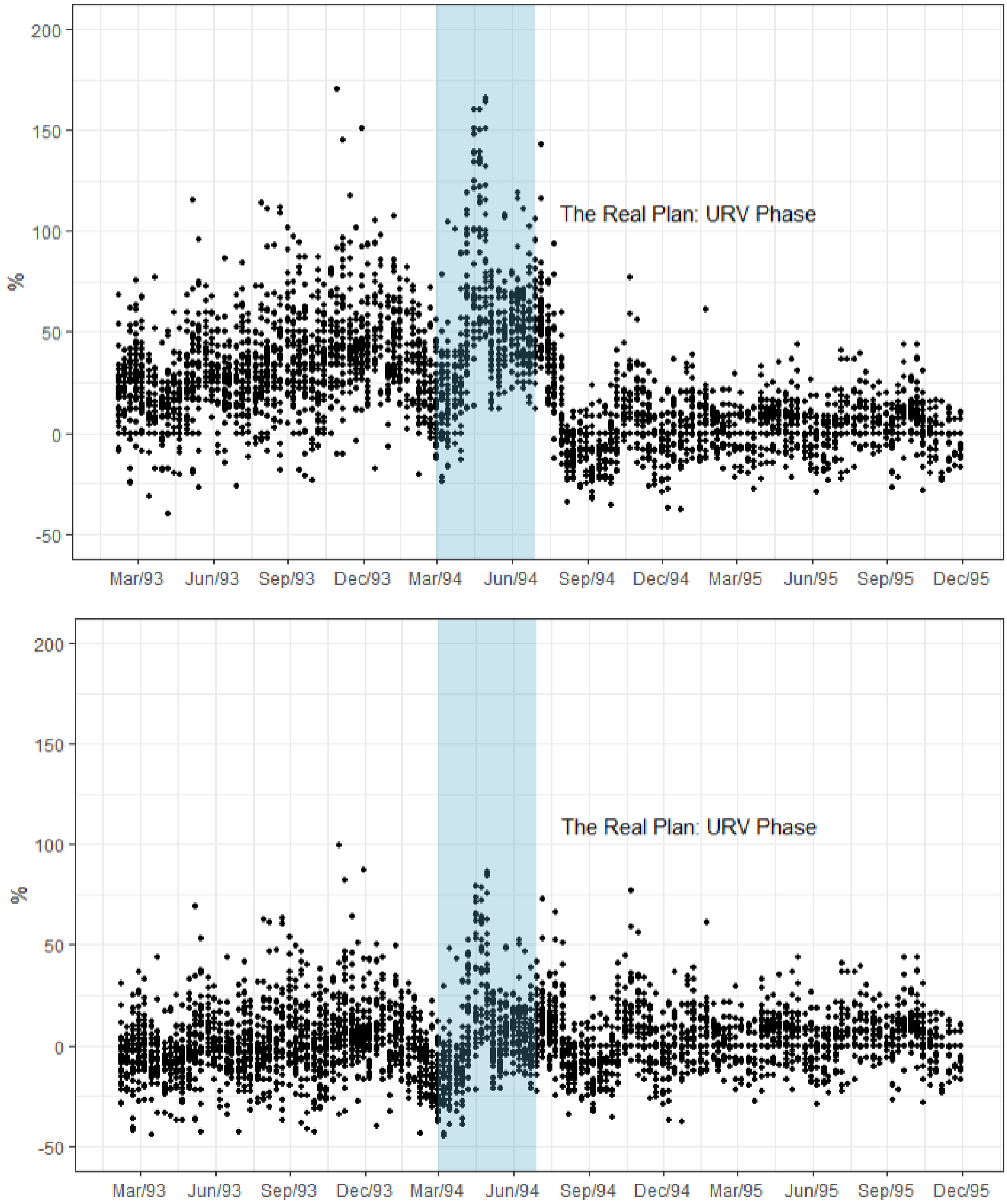
Notes: The highlighted area represents the Real Plan's unit of real value (URV) phase.

Figure A2 - Daily URV and Nominal Foreign Exchange Rate (CR$/U.S. Dollar)



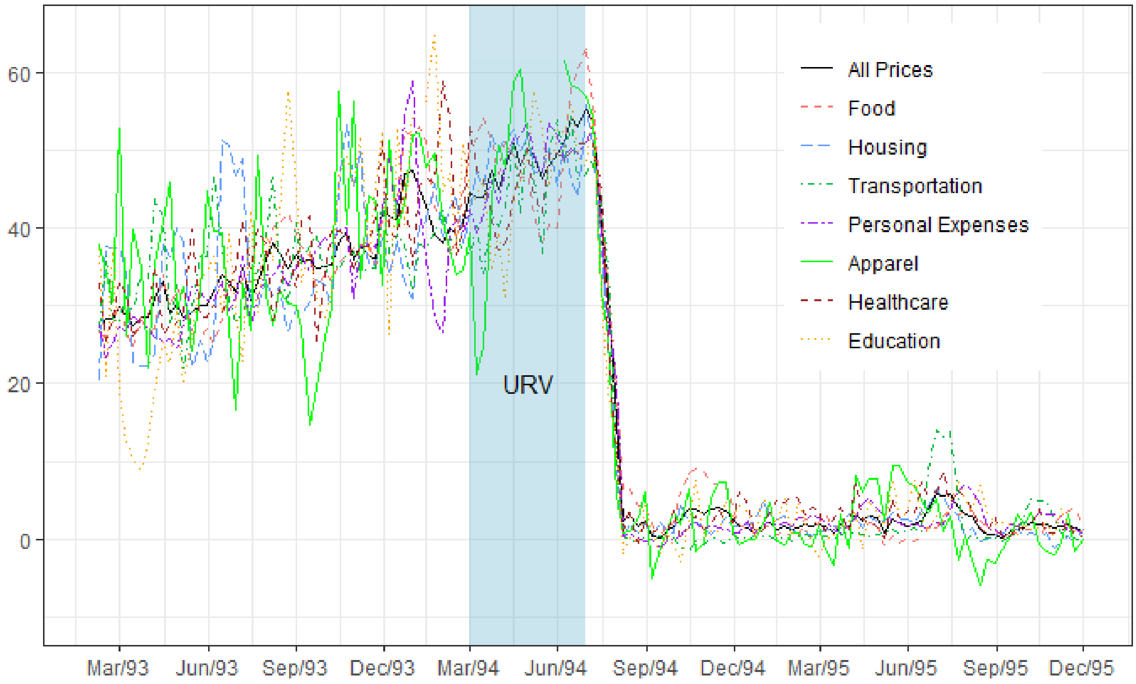
Notes: The URV was pegged to the U.S. dollar at parity and daily updated to the CR$. The commercial dollar, also known as PTAX, and tourism dollar coexisted from March 1993 until January 1999. These series are available by the BCB (series 13, 1, and 3, respectively).

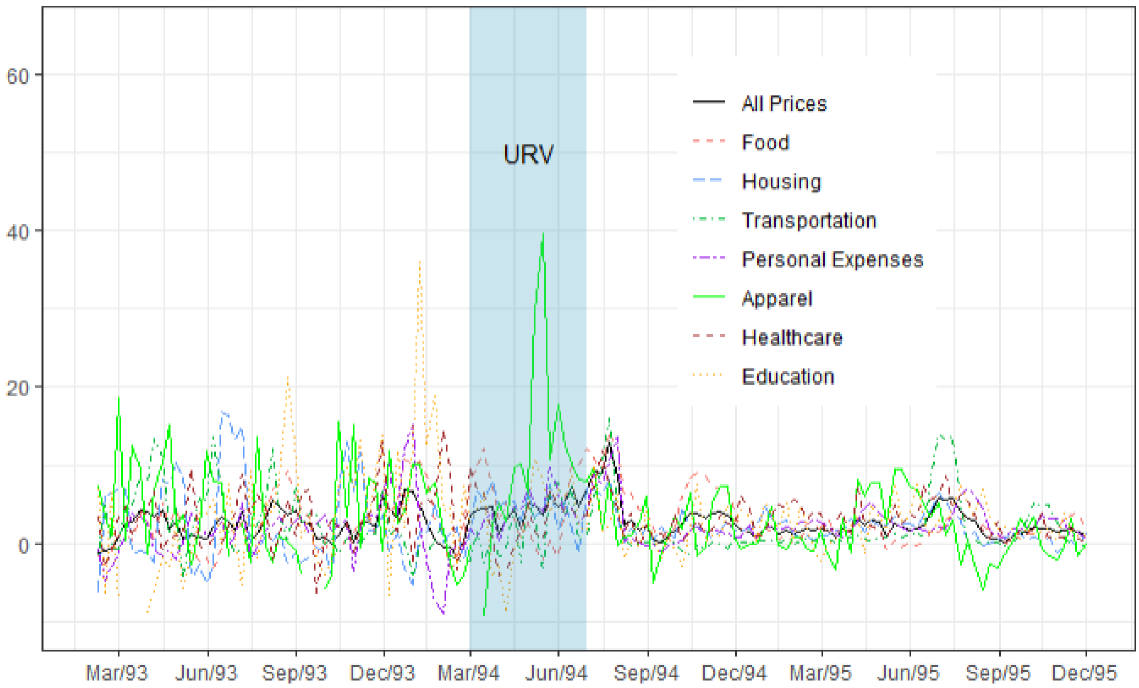
Figure A3 - Inflation Trajectory under the Fixed and the CR$/URV Conversion



Notes: The graphs illustrate the monthly inflation using the brand's price trajectories (see Figure 3) exemplified in Table 2 whereby each point represents a monthly rate quoted in a specific outlet in a given week. The top graph uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000). The bottom graph uses the weekly weighted average of the CR$/URV to convert prices to R$.

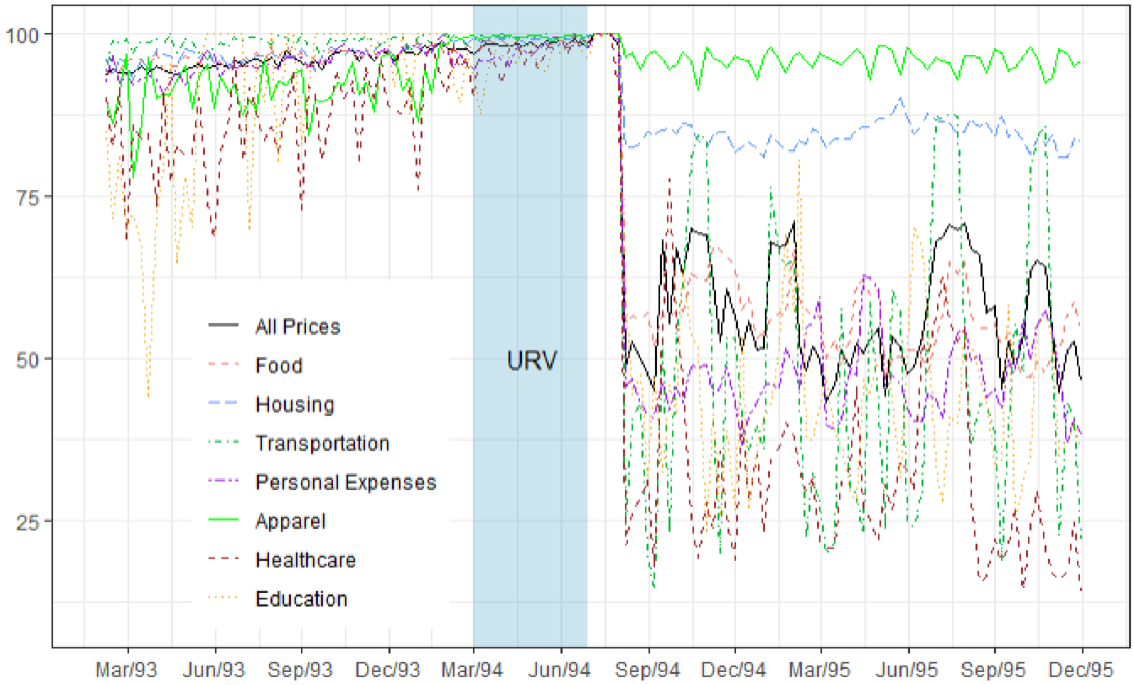
Figure A4 - Inflation of the FIPE Groups of Products (Fixed and CR$/URV Conversion)

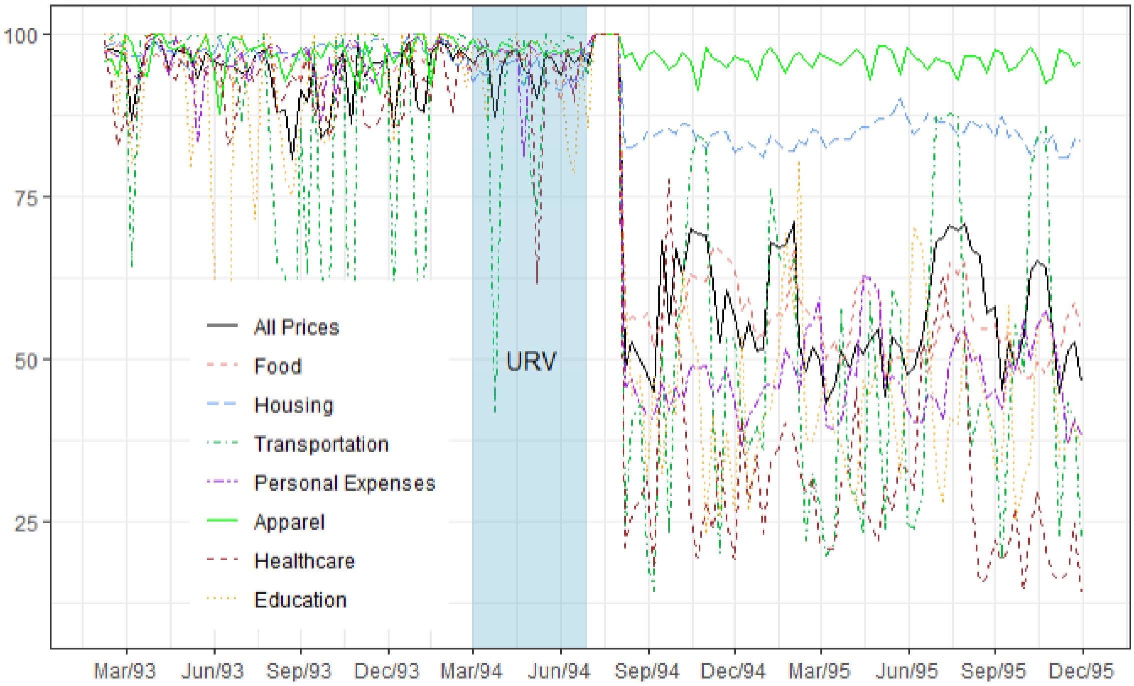




Notes: The top graph uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) while the bottom one uses the weekly weighted average of the CR$/URV to convert prices to R$. The highlighted area represents the Real Plan's URV phase.

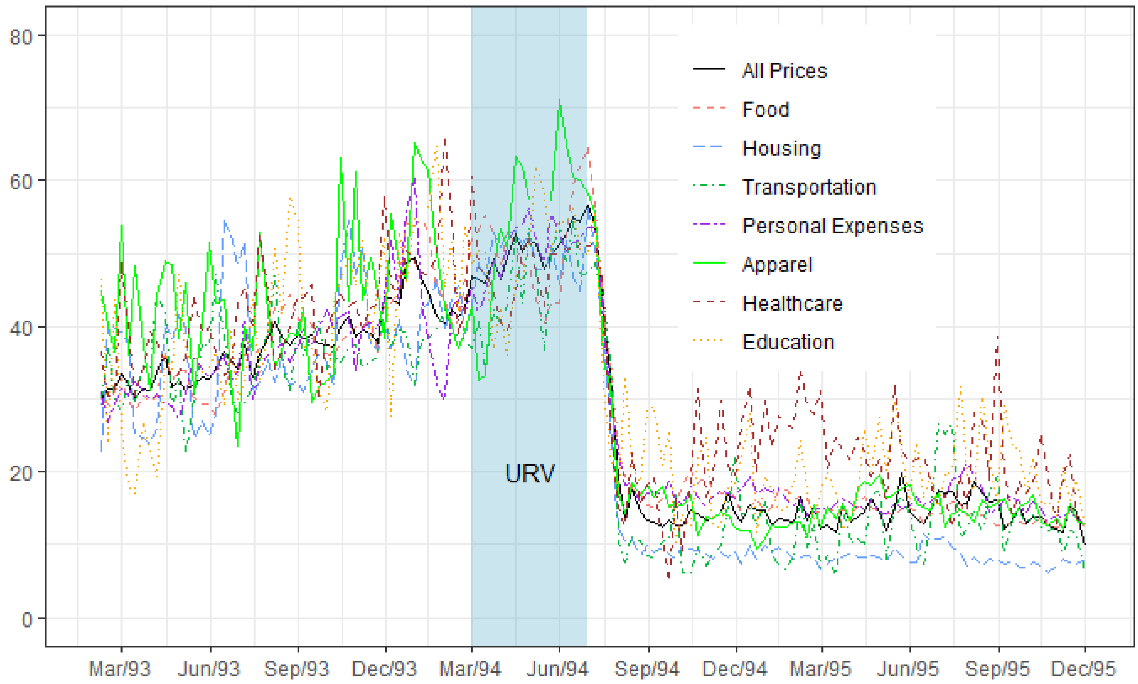
Figure A5 - Frequency of the FIPE Groups of Products (Fixed and CR$/URV Conversion)

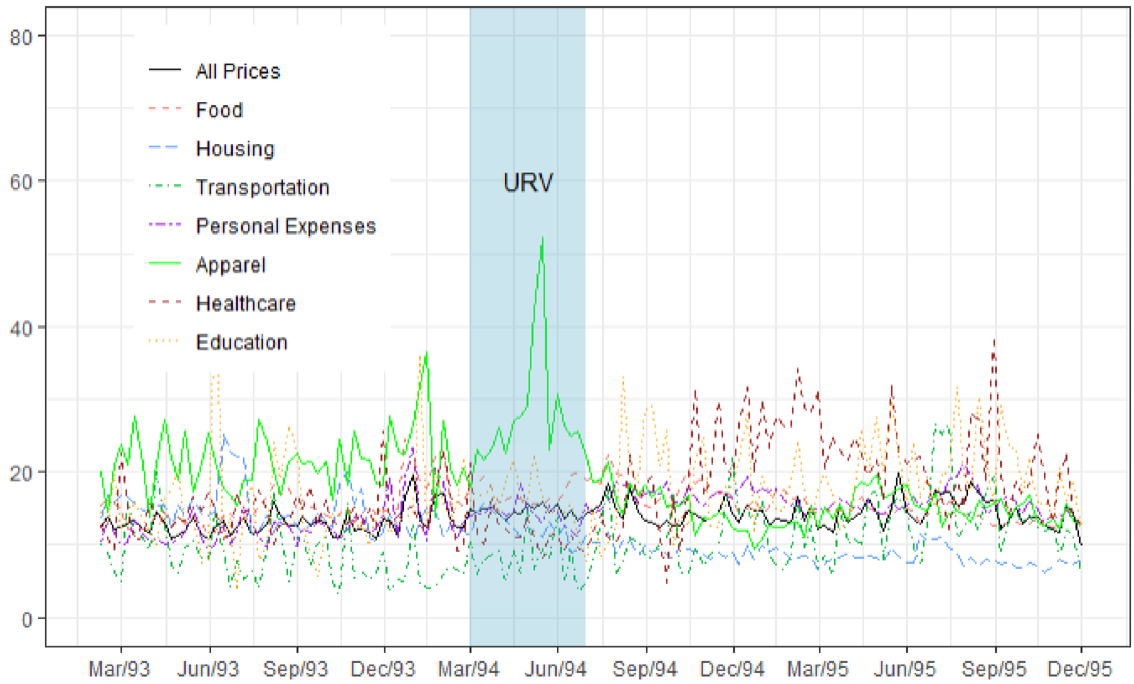




Notes: See explanations in Figure A4. The CR$/URV price change thresholds ($\bar{\pi}_{i,t}^{URV}$) assume $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ while being time-variant dependent from π_t^{URV}.

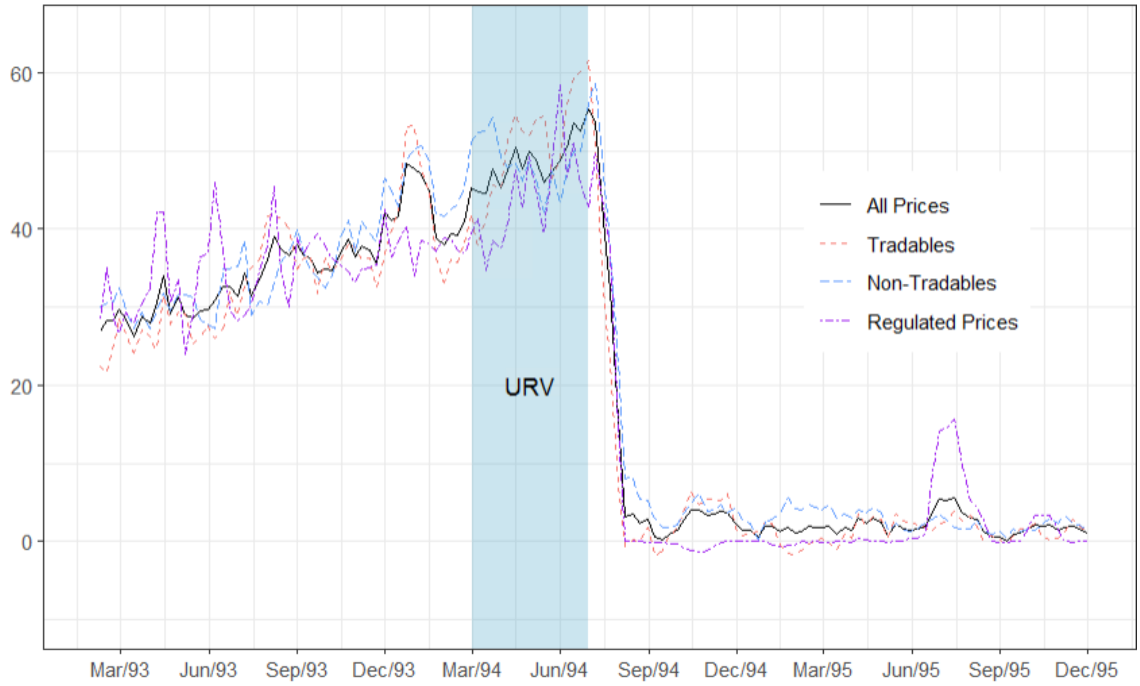
Figure A6 - Absolute Size of the FIPE Groups of Products (Fixed and CR$/URV Conversion)

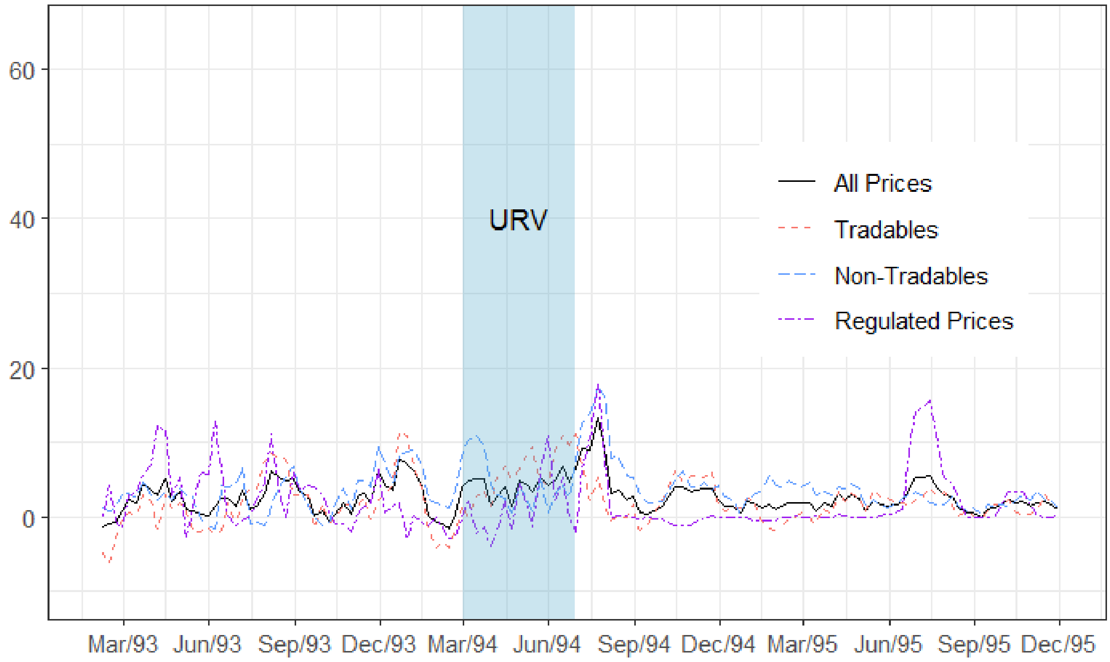




Notes: See explanations in Figure A4.

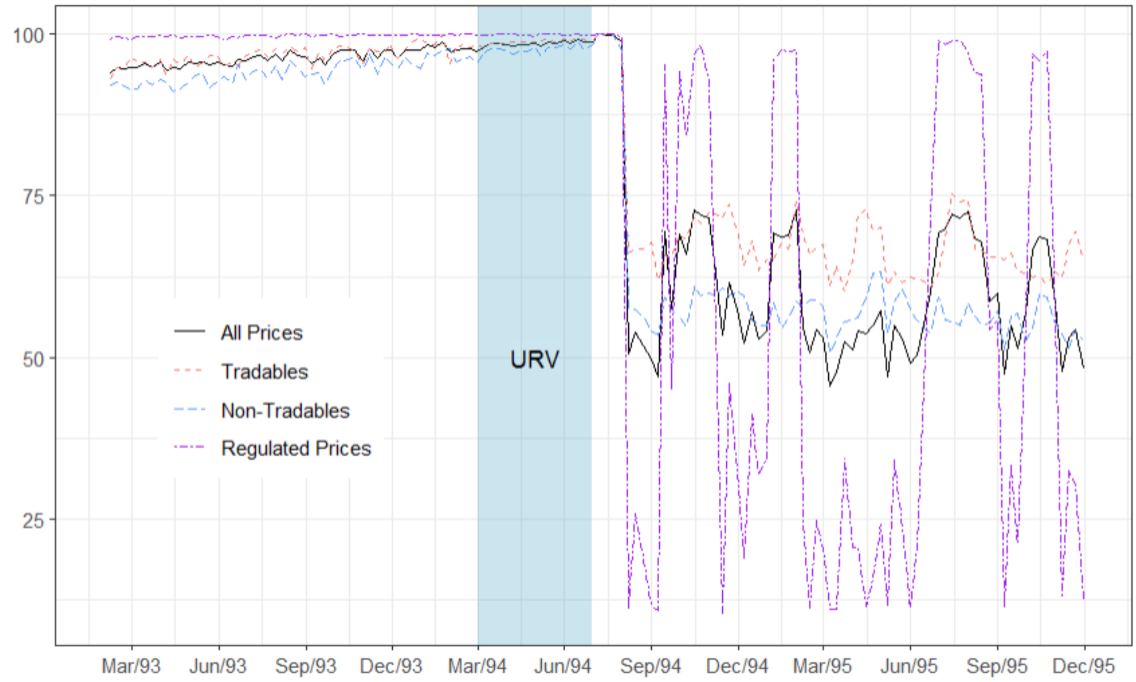
Figure A7 - Inflation of Tradables, Non-tradables, and Regulated prices (Fixed and CR$/URV Conversion)

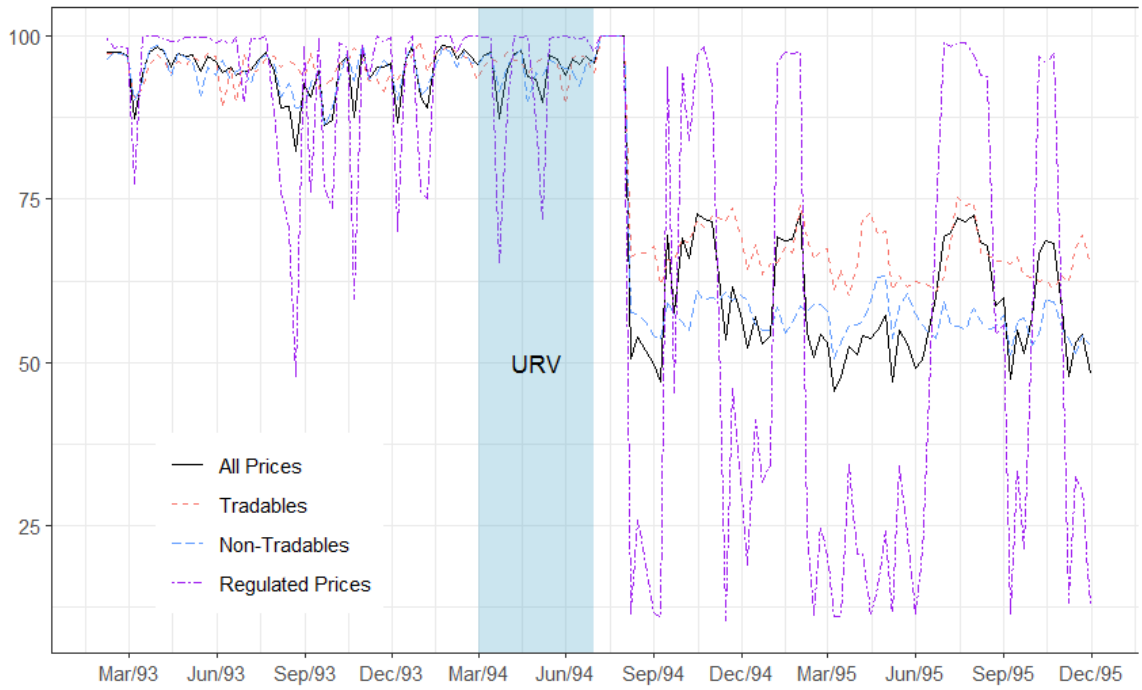




Notes: The top graph uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) while the bottom one uses the weekly weighted average of the CR$/URV to convert prices to R$. The highlighted area represents the Real Plan's URV phase.

Figure A8 - Frequency of Tradables, Non-tradables, and Regulated prices (Fixed and CR$/URV Conversion)



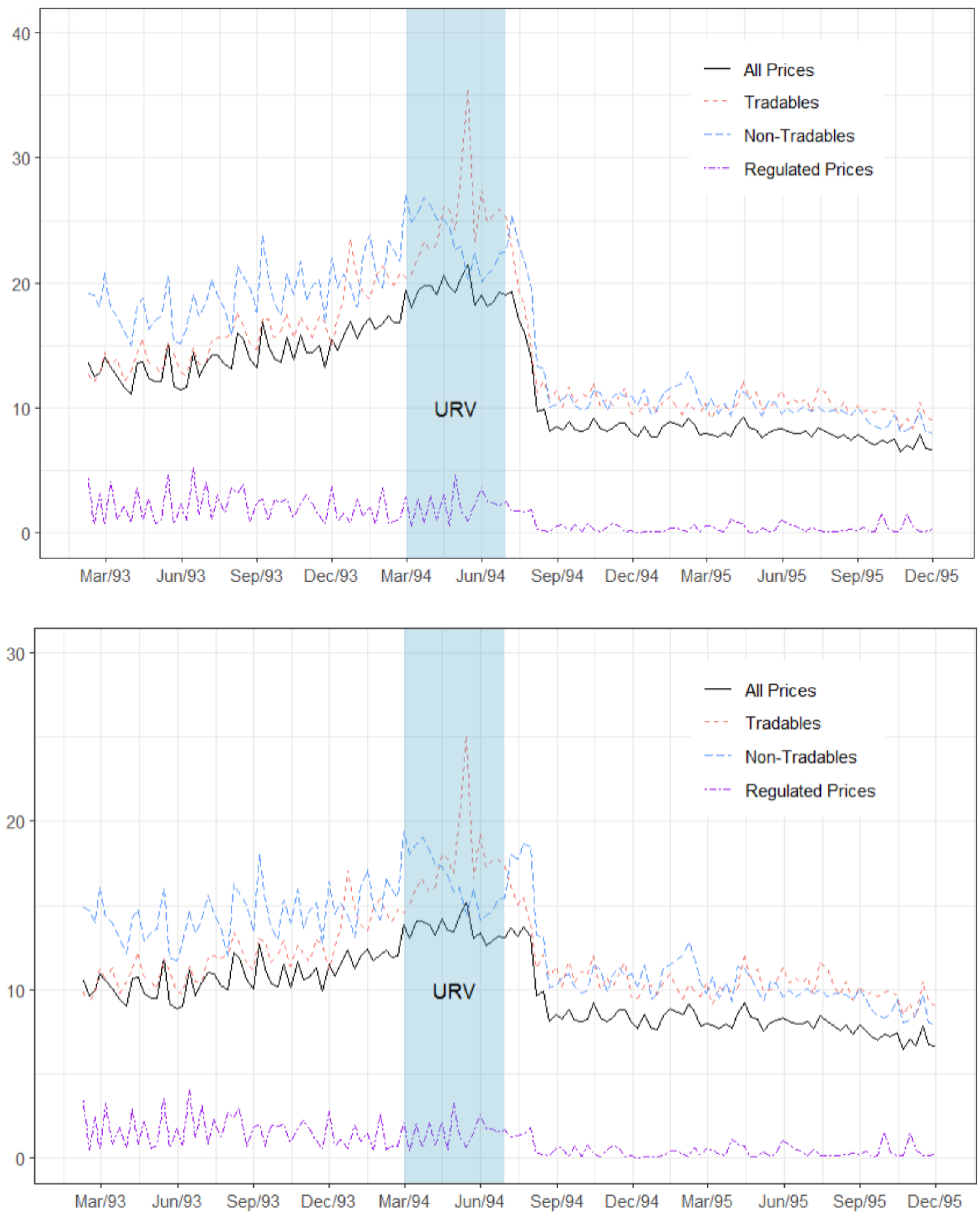


Notes: See explanations in Figure A4. The CR$/URV price change thresholds ($\bar{\pi}_{i,t}^{URV}$) assume $\pi_{i,t}^{CR\$} - \pi_{t}^{URV} = 1\%$ while being time-variant dependent from π_{t}^{URV}.

Figure A9 - Absolute Size of Tradables, Non-tradables, and Regulated prices (Fixed and CR$/URV Conversion)

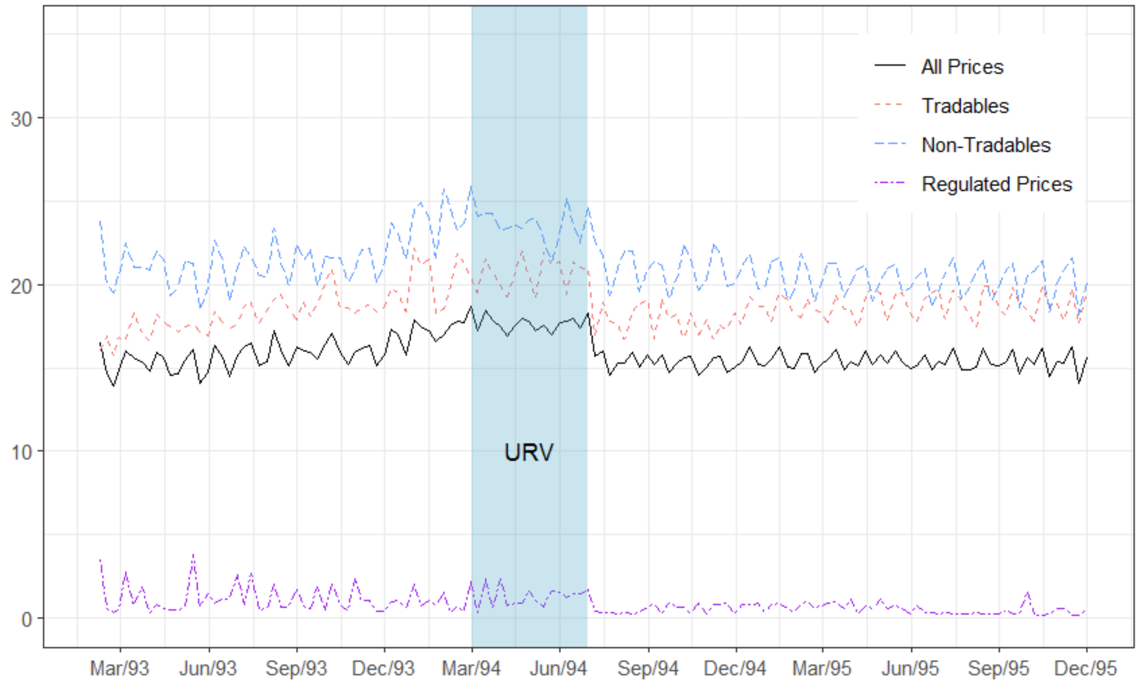
Notes: See explanations in Figure A7.

Figure A10 - Standard Deviation of Price Changes (Fixed and CR$/URV Conversion)



Notes: The top graph uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) while the bottom one uses the weekly weighted average of the CR$/URV to convert prices to R$. The highlighted area represents the Real Plan's URV phase.

Figure A11 - Coefficient of Variation of Price Levels (Fixed and CR$/URV Conversion)



Notes: The coefficient of variation of price levels is the same irrespective of the conversion method.

Table A1 - FIPE CPI: Weeks of Price Survey

Months	**Days**	**FIPE Weeks (# Days)**			
		1	**2**	**3**	**4**
January	31	7	8	8	8
February	28	7	7	7	7
March	31	7	8	8	8
April	30	7	8	7	8
May	31	7	8	8	8
June	30	7	8	7	8
July	31	7	8	8	8
August	31	7	8	8	8
September	30	7	8	7	8
October	31	7	8	8	8
November	30	7	8	7	8
December	31	7	8	8	8

Notes: The table summarizes information on the number of days in each (FIPE) week and month during the non-leap years of 1993-1995. Each month equals four periods of 7 to 8 days corresponding to four index monthly updates.

Table A2 - FIPE CPI: Groups and Weights (%)

Groups of Products	**POF 81-82 1989-1993**	**POF 90-91 1994-1999**
General Index	100.0	100.0
Food	37.7	30.8
Housing	18.3	26.5
Transportation	10.5	13.0
Personal Expenses	19.6	12.5
Apparel	8.1	8.7
Healthcare	3.8	4.6
Education	2.0	3.9

Notes: The table summarizes the official FIPE weighting scheme across groups of products. During the 1993 to 1995 window, the FIPE changed the index methodology once (at the end of 1993) to incorporate changes in consumption behavior. For the sake of consistency, we use the weights based on the Consumer Expenditure Survey (POF) of 1990-1991.

Table A3 - Sensibility of the Inflation, Frequency, and Absolute Size of Price Change for different levels of $\pi_{i,t}^{CR\$} - \pi_t^{URV}$ (CR$/URV Conversion)

Currency	**URV | R$**						
Statistics	**Infl.**	**Size**	**Size +**	**Size -**	**Freq.**	**Freq. +**	**Freq. -**
1%	**4.5**	**14.6**	**16.4**	**11.3**	**95.1**	**56.0**	**39.1**
Mar-94	3.7	14.7	16.3	12.0	94.1	55.8	38.4
Apr-94	3.5	14.5	16.6	11.7	95.8	50.0	45.8
May-94	4.8	15.3	17.2	11.2	94.2	56.0	38.2
Jun-94	6.0	14.0	15.6	10.1	96.2	62.1	34.1
5%	**4.5**	**17.5**	**19.2**	**14.0**	**79.4**	**47.8**	**31.6**
Mar-94	3.7	18.2	19.9	14.6	82.5	45.3	37.2
Apr-94	3.5	17.4	19.3	14.9	78.6	46.8	31.8
May-94	4.8	17.5	19.5	13.6	79.8	48.7	31.0
Jun-94	6.0	16.9	18.2	12.8	76.7	50.4	26.3
10%	**4.5**	**21.3**	**23.9**	**17.5**	**67.3**	**39.6**	**27.8**
Mar-94	3.7	21.5	25.1	17.1	70.0	36.1	33.9
Apr-94	3.5	21.0	22.9	18.5	70.3	41.9	28.5
May-94	4.8	21.5	25.1	16.9	63.4	36.1	27.3
Jun-94	6.0	21.1	22.7	17.5	65.6	44.3	21.4

Notes: The table provides monthly price setting indicators (in %) alongside its averages for different levels of $\pi_{i,t}^{CR\$} - \pi_t^{URV} = \{1\%, 5\%, 10\%\}$, which affect CR$/URV price change thresholds ($\bar{\pi}_{i,t}^{URV}$), during the URV period. It uses the weekly weighted average of the CR$/URV to convert prices to R$.

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Chapter 2

Micro Evidence on the Real Plan Stabilization Program in Brazil: State-Dependent Menu-Cost Model[1]

2.1. Introduction

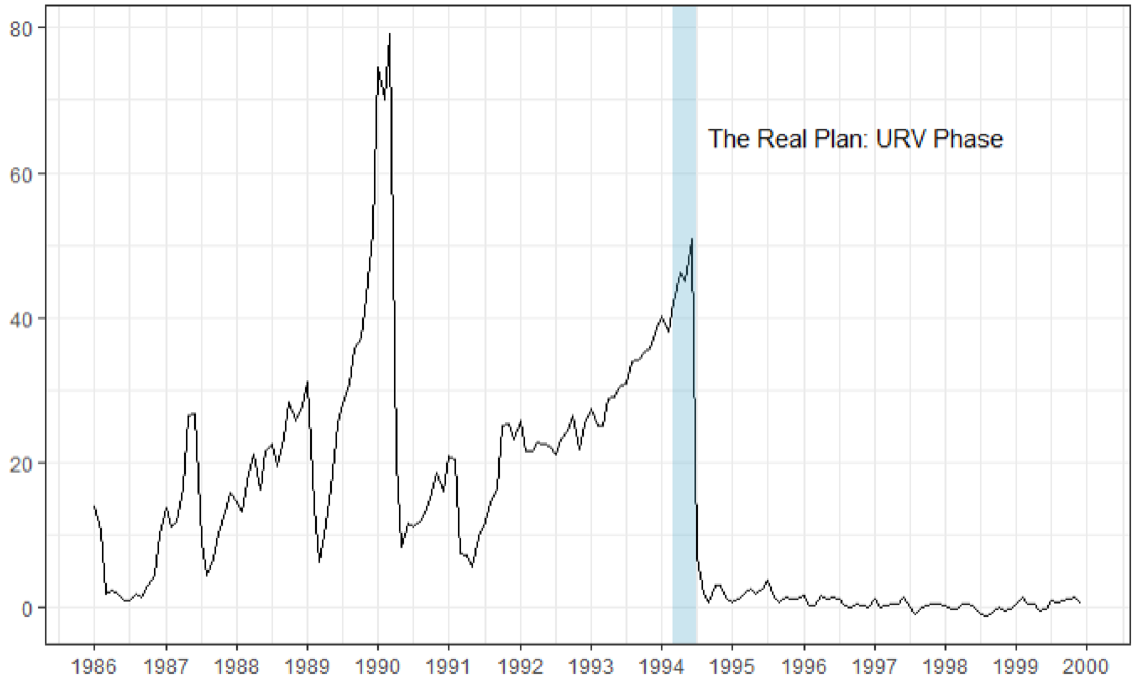
The Real Plan, implemented in Brazil in 1994, was a crucial economic reform aimed at stabilizing the country's economy, which had been plagued by hyperinflation throughout the 1980s and early 1990s. The plan achieved significant success, bringing inflation under control and laying the groundwork for economic growth. This paper revisits the Brazilian hyperinflationary experience and provides micro evidence by examining it through the lens of a menu cost model with state-dependent pricing rules.

Amidst fiscal and monetary instability, Brazil faced extraordinarily high inflation during the late 1980s and early 1990s, with none of the stabilization attempts successfully reducing inflation to sustainable low levels. The persistence of Brazilian inflation had both structural and inertial components: the former was driven by fiscal deficits, while the latter was linked to the unsynchronized indexation of contracts. Figure 1 presents the monthly inflation rate, as measured by the Consumer Price Index (CPI) of the Economic Research Institute Foundation (FIPE), from 1985 through the end of 1999, and highlights the URV phase of the Real Plan, during which the stabilization process occurred. As shown in Figure 1, the Real Plan achieved substantial disinflation and established long-term price stability. This illustrates the Plan's success through a pre-announced strategy,

[1] This is joint work with Julia Araujo, Marco Bonomo, and Mauro Rodrigues.

avoiding the confiscation of household savings or the imposition of price freezes, which had characterized previous stabilization attempts.[2]

Figure 1 - FIPE CPI: Monthly Inflation (%)



Notes: The highlighted area represents the Real Plan's unit of real value (URV) phase.

The Real Plan has its conceptual origins in the Larida Plan, an economic stabilization proposal developed by Arida and Lara-Resende (1985), in which the authors proposed the creation of an indexed currency, the Readjusted Obligations of the National Treasury (ORTN) pro-rata, to synchronize price adjustments throughout the economy. This idea resurfaced in the Real Plan in the form of the Units of Real Valor (URV), drawing inspiration from the ORTN. The URV served as an alternative unit of account and was instrumental in the Plan's approach to stabilization. Almeida and Bonomo (2002) argue that the Real Plan illustrates how rapid disinflation can be achieved with relatively low costs, even in a context of limited credibility, as long as a price alignment mechanism – such as the URV – is employed to correct distributional asymmetries.

[2] During the presidency of José Sarney (1985–1989), four stabilization plans were implemented: Cruzado I (Feb/1986), Cruzado II (Nov/1986), Bresser (Jun/1987), and Verão (Jan/1989). His successor, Fernando Collor de Mello (1990–1992), introduced two additional programs: Collor I (Mar/1990) and Collor II (Jan/1991). Common to all these plans was the attempt to control hyperinflation through price and wage freezes, though Collor I also included the controversial measure of seizing household savings.

The Real Plan can be divided into three distinct phases. The preliminary phase began with the introduction of a new currency, the cruzeiro real (CR$), on August 1, 1993. This phase was marked by fiscal adjustments through austerity measures, such as the Immediate Action Program (PAI) in July 1993 and the creation of the Emergency Social Fund (FSE) in February 1994, signaling a firm commitment to addressing the structural component of inflation by curbing fiscal deficits.

The second phase, representing the core of the Real Plan, spanned four months, from March 1 to June 30, 1994, during which a temporary unit of account, the URV, was introduced and coexisted with the CR$. Although both the CR$ and URV functioned as units of account, the URV was specifically designed to offer greater price stability. While the CR$ absorbed the effects of hyperinflation, the URV was pegged to the U.S. dollar at a one-to-one exchange rate and adjusted daily against the CR$ based on inflation indices. Consequently, the CR$/URV exchange rate reflected the depreciation of the CR$, leading to corresponding increases in prices and wages in CR$. However, these would remain relatively stable when expressed in URV, theoretically encouraging its adoption as the primary unit of account.[3] The rationale behind the URV was that, as individuals observed the stabilization of purchasing power, they would begin to reference prices in URV rather than CR$, reducing inflationary expectations. By implementing daily indexation through the URV, the Real Plan aimed to synchronize price adjustments across the economy, promoting alignment in relative prices and thus addressing the inertial component of inflation. This alignment would be crucial to breaking the cycle of continuous price increases that had plagued the Brazilian economy for years.

In the final phase of the Real Plan, the URV was converted into the Brazilian real (R$), transitioning from a mere unit of account to a full-fledged medium of exchange. On July 1, 1994, the real replaced the cruzeiro real (CR$) at a fixed conversion rate of CR$ 2,750 to R$ 1. From that point onward, the real (R$) became Brazil's new official currency, marking the conclusion of the stabilization process and the formal establishment of the monetary reform.

In this paper, we analyze price adjustment patterns during the implementation of the Real Plan stabilization program in Brazil, which represented a transitional phase from

[3] Initially, prices of goods and services for immediate payment were exclusively quoted in CR$ until the end of April, to prevent the reduction of menu costs associated with daily price adjustments. After this period, URV quotations were allowed, enabling people to acclimate to stable prices. Simultaneously, salaries were first adjusted based on their average value over the preceding four months, then indexed daily and paid monthly in CR$, ensuring the preservation of purchasing power during the transition.

hyperinflation to relatively low inflation, by examining it through the lens of a menu cost model with state-dependent pricing rules. Our analysis is based on non-publicly disclosed FIPE CPI microdata, which provides detailed price information at the store level. This dataset is particularly valuable as it is the only available source of price data for the period of hyperinflation in Brazil, offering unique insights into the dynamics of price-setting behavior during this critical economic transition.

Given the price statistics calculated from the FIPE CPI microdata by Araujo, Bonomo, Kudamatsu, and Rodrigues (2024), from now on referred as ABKR (2024), we apply the state-dependent menu-cost model of Almeida and Bonomo (2002) calibrated with Brazilian data to track the behavior of prices throughout the stabilization process. Specifically, we compare price settings in Brazil across different periods: pre-URV (from August 1993 until February 1994, with prices in CR$), URV (from March until June 1994, with prices in both CR$ and URV), and post-URV (from July until December 1994, with prices in R$). For the URV period, we also contrast the estimated costs associated with the CR$ and URV conversions to describe the theoretical incentives for using either metric. In this way, our analysis provides a theoretical foundation to characterize this transitional economic period of the Real Plan and offers additional insights into the mechanisms underlying disinflation during the Real Plan.

Our paper contributes to the literature on price setting using microdata and explores how pricing behavior varies across different inflationary environments. Klenow and Malin (2010) and Nakamura and Steinsson (2013) provide comprehensive reviews of studies that shed light on the features of micro-level price adjustments and how the sluggishness of these adjustments influences the impact of demand and monetary shocks on short-run output.[4] The context of high inflation, especially hyperinflation, is typical of developing economies, as well as some developed nations, and provides unique opportunities to study the dynamics of price adjustments. Relevant studies include Lach and Tsiddon (1992) for Israel, Konieczny and Skrzypacz (2005) for Poland, Gagnon (2009) for Mexico, Wulfsberg (2016) for Norway, Álvarez et al. (2018) for Argentina.

In Brazil, some studies have focused on the country's price-setting mechanisms following economic stabilization, including Gouvea (2007), Lopes (2008), and Barros et

[4] Firms optimally change prices subject to menu costs in state-dependent models while the timing of price changes is exogenous in time-dependent sticky-price models either because prices are maintained for a given number of periods (Taylor, 1999) or because only a fraction of total firms can adjust prices each period (Calvo, 1983). As a result, price changes are not synchronized among firms, and monetary policy can affect economic activity in the short run.

al. (2009). Angelis (2012), Araujo and Rodrigues (2018), Araujo (2019), and ABKR (2024) show that the Real Plan fundamentally altered Brazil's price-setting characteristics. With the exemption of ABKR (2024), these studies do not explore the specific role of the URV in the stabilization process. Moreover, none of them employ theoretical models to illustrate the disinflation period. This paper complements the detailed analysis by ABKR (2024) on micro-level price adjustments, which explicitly investigates the relationship between store-level price microdata and URV information during the transition period, by employing a state-dependent menu cost model calibrated with data from that era.

We emphasize that the Real Plan led to significant changes in Brazil's price-setting mechanisms, though the stabilization effects only became pronounced after the introduction of the Real (R$). Consequently, the empirical evidence suggests that the URV's role in reshaping price dynamics and securing long-term price stability was less impactful than initially assumed. Specifically, there is limited evidence of substantial changes in pricing behavior during the URV phase, despite theoretical incentives for URV-based price conversions to the detriment of the official currency, which were relatively cost-effective during the coexistence of both units of account – not only the estimated menu cost in URV was lower than the one in CR$, but also its expected expenses. However, these findings do not imply that the URV was without importance. Rather, the URV acted as a transitional tool that facilitated the gradual shift to the new currency, accelerated the realignment of inflation rates across products, and preserved real wage purchasing power during the stabilization period—an aspect that is not fully reflected in the micro-level price adjustment data.

The rest of the paper is organized as follows: data, price-setting statistics evidence, menu-cost model, and concluding remarks.

2.2. Data

Our analysis focuses on the period extending from January 1993 to December 1995. Consequently, the timeframe under consideration encompasses the pre-stabilization phase, the URV phase of the Real Plan (March to June 1994), and the post-implementation period following the introduction of the Brazilian Real (R$) as the official currency in July 1994. To examine the distinctions between these periods and assess the impact of the URV on price-setting dynamics in Brazil, we utilize the non-publicly available microdata from the FIPE Consumer Price Index (CPI) and official URV information.

The FIPE CPI computes price change in the municipality of Sao Paulo.[5] The FIPE microdata is composed of price surveys that provide detailed information on price quotes for consumer goods and services at the store level. This dataset is particularly unique as it is the only one available covering the years of hyperinflation in Brazil.[6] The dataset's most granular unit of observation is a specific brand of a product, with its price collected during a particular week at a specific store (henceforth referred to as an "item"). Moreover, although the data are collected on a weekly basis, a given item is typically recorded only once per month.[7]

The URV data goes from January 1993 to June 1994. The Central Bank of Brazil (BCB) established distinct methodologies for the pre-URV period (January 1993 to February 1994) and the URV phase of the Real Plan (March to June 1994) through Act 1066, enacted in February 1994. At the time the guidelines were issued, the government retroactively published all historical daily Cr$/URV and CR$/URV exchange rates to facilitate the understanding and adoption of the URV as a unit of account for updating contracts. Furthermore, these publications outlined the methodology that would be employed for future calculations of the URV.

During the pre-URV period, the CR$/URV monthly rate was determined retrospectively as the arithmetic average of three existing monthly inflation indices: the FIPE CPI (measured at the third week), the IBGE IPCA-E, and the FGV IGP-M. As such, the CR$/URV rate captured the average depreciation of the CR$ in terms of purchasing power. In the URV period, the BCB would announce the CR$/URV exchange rate for the following business day, based on price surveys derived from those three inflation indicators, which reflected inflation with a delay of 2 to 4 weeks. Moreover, the CR$/URV monthly rate was required to fall within the range established by the highest and lowest monthly inflation rates of these indices.[8] Following this period, the

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[5] The Broad National Consumer Price Index (IPCA), the main inflation index in Brazil, is not available for the early 1990s. Despite lacking the national coverage of the IPCA, the FIPE CPI shares a similar structure, enabling both indices to capture parallel inflationary trends over time.

[6] Lopes (2008), Angelis (2012), Araujo and Rodrigues (2018), Araujo (2019), and ABKR (2024) also use the FIPE microdata in their price-setting studies.

[7] Each item is identified by a specific code that carries information on its product, brand, outlet, and week that price was surveyed.

[8] The initial value of the URV was set at CR$ 647.5 on March 1, and it was updated daily, reaching CR$ 2,750 by June 30.

government not only abolished the CR$ but also converted the URV into the Brazilian Real (R$), maintaining parity with the U.S. dollar.[9] Table 1 presents a comparison between the monthly CR$/URV inflation rate and the behavior of its underlying components. In short, the monthly variation in the URV closely tracked the overall inflationary trends in Brazil.

Table 1 - Monthly Inflation (%): CR$/URV and its Components

Period	CR$/URV	Avg. Infl.	FIPE-3	IPCA-E	IGP-M
Pre-URV					
Jan-93	27.71	27.71	27.83	29.47	25.83
Feb-93	26.73	26.73	25.06	26.72	28.41
Mar-93	25.74	25.74	25.02	25.96	26.25
Apr-93	27.98	27.98	27.78	27.34	28.83
May-93	29.19	29.19	29.26	28.61	29.70
Jun-93	29.80	29.80	30.28	27.61	31.50
Jul-93	30.71	30.71	30.21	30.67	31.25
Aug-93	32.39	32.39	33.40	31.99	31.79
Sep-93	34.55	34.55	34.00	34.38	35.28
Oct-93	34.90	34.90	34.48	35.17	35.04
Nov-93	35.19	35.19	35.51	33.90	36.15
Dec-93	37.59	37.59	37.76	36.69	38.32
Jan-94	39.73	39.73	40.94	39.17	39.07
Feb-94	39.17	39.17	37.04	39.70	40.78
URV					
Mar-94	43.26	43.55	41.31	43.63	45.71
Apr-94	42.60	42.53	45.43	41.25	40.91
May-94	44.00	43.82	44.66	44.21	42.58
Jun-94	46.60	46.28	48.97	44.65	45.21

Notes: The table provides the monthly inflation (%) of the CR$/URV and its underlying components, that is, the FIPE-3 CPI (measured in the 3rd week), the IBGE IPCA-E, and the FGV IGP-M for both the pre-URV and the URV periods. Because the Cr$ was the Brazilian official currency until July 1993, Cr$/URV quotes were converted to CR$/URV at $CR\$1 = Cr\$1000$.

Given our focus on understanding the mechanics of the Real Plan stabilization program, we combine FIPE microdata – weekly price quotes expressed in the official

[9] In the transition from the URV to the real, the government established that inflation in July of 1994 would be computed through the comparison of prices in R$, collected at that month, with prices in URV of the previous month (after being converted from CR$).

monetary units (Cr$, CR$, or R$) – with daily URV data. In this context, the timing of price collection is of critical importance. However, since the FIPE did not record the exact dates of the price surveys in its files, it is necessary to estimate the CR$/URV exchange rates that were in effect at the time of each item's price surveys.

Following ABKR (2024), we convert prices into URV (or R$) using a weekly weighted average of the CR$/URV exchange rate, assigning a higher probability that prices were collected on one of the first three working days.[10] Additionally, we represent a URV price variation through an indicator function $I_{i,t}^{URV}$ that relies on an inflation threshold $\bar{\pi}_{i,t}^{URV}$ (see Equation 1). Finally, the baseline results assume $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ for all *i* and *t* to define a threshold $\bar{\pi}_{i,t}^{URV}$ that varies according to π_t^{URV}.[11]

$$I_{i,t}^{URV} = \begin{cases} 1, \; if \; \left|\pi_{i,t}^{URV}\right| > \bar{\pi}_{i,t}^{URV} \; \leftrightarrow \; \pi_{i,t}^{URV} > \bar{\pi}_{i,t}^{URV} \; or \; \pi_{i,t}^{URV} < -\bar{\pi}_{i,t}^{URV} \\ 0, \; if \; \left|\pi_{i,t}^{URV}\right| \leq \bar{\pi}_{i,t}^{URV} \; \leftrightarrow \; -\bar{\pi}_{i,t}^{URV} \leq \pi_{i,t}^{URV} \leq \bar{\pi}_{i,t}^{URV} \end{cases}, \quad (1)$$
$$\text{for } t = 1, \; \dots, T \text{ and } i = 1, \; \dots, I$$

2.3. Price-Setting Statistics Evidence

In this section, we briefly discuss a series of pricing statistics results computed by ABKR (2024) for the periods around the URV phase of the Real Plan, utilizing the FIPE microdata and official URV data. This analysis is conducted under two distinct price conversion methodologies: the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) and the weekly weighted average of the CR$/URV exchange rate. The primary objective is to ascertain whether the URV policies influenced price behavior.

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[10] As discussed by ABKR (2024), there are two sources of potential measurement error: uncertainty regarding the exact date of the price survey and the precision of the point estimate itself.

[11] Consider $P_{i,t}^{CR\$}$ and $\pi_{i,t}^{CR\$}$ the CR$ price and inflation of item *i* at time *t*, V_t^{URV} and π_t^{URV} the weekly weighted average of the CR$/URV and its resulting inflation at *t*, and $P_{i,t}^{URV}$ and $\pi_{i,t}^{URV}$ the converted URV price and inflation of *i* at *t*. Then, a threshold $\bar{\pi}_{i,t}^{URV}$ is calculated as:

$$\pi_{i,t}^{URV} = \frac{P_{i,t}^{URV}}{P_{i,t-4}^{URV}} - 1 = \frac{P_{i,t}^{CR\$} / V_t^{URV}}{P_{i,t-4}^{CR\$} / V_{t-4}^{URV}} - 1 = \frac{P_{i,t}^{CR\$} / P_{i,t-4}^{CR\$}}{V_t^{URV} / V_{t-4}^{URV}} - 1 = \frac{1 + \pi_{i,t}^{CR\$}}{1 + \pi_t^{URV}} - 1 = \frac{\pi_{i,t}^{CR\$} - \pi_t^{URV}}{1 + \pi_t^{URV}}$$

Table 2 - Inflation, Frequency, and Absolute Size of Price Change (Fixed and CR$/URV Conversion)

Currency	**CR$**							**URV | R$**						
Statistics	**Infl.**	**Size**	**Size +**	**Size -**	**Freq.**	**Freq. +**	**Freq. -**	**Infl.**	**Size**	**Size +**	**Size -**	**Freq.**	**Freq. +**	**Freq. -**
Pre-URV	**39.1**	**41.3**	**42.0**	**8.9**	**96.8**	**94.1**	**2.7**	**2.6**	**13.5**	**15.6**	**11.1**	**92.8**	**49.2**	**43.6**
Aug-93	36.6	38.9	39.5	8.0	96.1	93.9	2.2	4.5	13.8	15.6	11.5	87.2	54.7	32.4
Sep-93	35.4	38.1	38.6	8.8	95.5	93.0	2.5	1.6	13.2	14.7	11.5	88.8	50.3	38.5
Oct-93	37.2	39.2	39.7	7.7	97.0	95.0	2.0	0.9	12.1	15.2	10.4	94.1	40.1	54.0
Nov-93	38.1	40.0	40.5	9.0	97.0	95.2	1.7	3.4	12.1	13.6	11.1	95.2	62.4	32.8
Dec-93	44.2	46.3	47.1	6.8	96.8	94.4	2.5	5.3	15.2	17.7	10.6	92.4	56.3	36.1
Jan-94	41.3	43.3	44.7	8.7	98.1	92.6	5.5	1.8	14.9	17.5	11.2	95.5	41.3	54.2
Feb-94	41.2	43.3	43.9	13.6	97.4	94.9	2.5	0.4	13.3	15.0	11.2	96.5	39.6	56.9
URV	**49.1**	**50.7**	**51.5**	**9.6**	**98.5**	**96.1**	**2.4**	**4.5**	**14.6**	**16.4**	**11.3**	**95.1**	**56.0**	**39.1**
Mar-94	45.1	47.0	47.8	8.7	98.3	95.5	2.8	3.7	14.7	16.3	12.0	94.1	55.8	38.4
Apr-94	49.3	51.1	52.1	11.4	98.4	95.6	2.8	3.5	14.5	16.6	11.7	95.8	50.0	45.8
May-94	48.4	50.0	50.7	9.8	98.6	96.3	2.3	4.8	15.3	17.2	11.2	94.2	56.0	38.2
Jun-94	53.4	54.7	55.3	8.8	98.8	97.1	1.7	6.0	14.0	15.6	10.1	96.2	62.1	34.1
Post-URV	**-**	**-**	**-**	**-**	**-**	**-**	**-**	**3.5**	**14.5**	**17.1**	**11.0**	**64.5**	**36.8**	**27.7**
Jul-94	-	-	-	-	-	-	-	9.9	16.0	18.1	7.9	100.0	66.7	33.3
Aug-94	-	-	-	-	-	-	-	2.3	14.8	17.2	11.6	49.7	28.4	21.3
Sep-94	-	-	-	-	-	-	-	0.7	12.9	15.6	11.0	58.9	24.9	34.0
Oct-94	-	-	-	-	-	-	-	3.3	13.7	19.2	10.5	67.8	32.6	35.3
Nov-94	-	-	-	-	-	-	-	3.5	14.8	16.7	11.8	57.9	37.0	20.9
Dec-94	-	-	-	-	-	-	-	1.5	14.5	15.6	12.9	52.5	31.4	21.1

Notes: The table provides monthly price setting indicators (in %) alongside its averages for the pre-URV, URV, and post-URV periods. It uses the official fixed conversion rates (R$ 1 = CR$ 2,750 = Cr$ 2,750,000) and the weekly weighted average of the CR$/URV to convert prices to R$. The CR$/URV price change thresholds ($\bar{\pi}_{i,t}^{URV}$) assume $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ (while being time-variant dependent from π_t^{URV}) and affect the URV statistics of frequency and absolute size of price change.

Table 2 summarizes the price-setting statistics of inflation, frequency, absolute size of price changes, and their respective separation into positive and negative changes using both conversion approaches for the pre-URV, URV, and post-URV periods.[12] The statistics are calculated on a monthly basis by averaging the weekly statistics within each month, followed by computing the respective means for each phase. Notably, as outlined in Section 2, the results consider the assumption of $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$, in conjunction with a varying π_t^{URV}, to establish price change thresholds for each t ($\bar{\pi}_{i,t}^{URV}$), specifically to evaluate the frequency and absolute magnitude of URV price changes.[13]

In summary, the findings indicate that the Real Plan substantially modified price-setting behavior in Brazil. However, the majority of stabilization effects only materialized

[12] The statistics are computed at the most disaggregated level, and then successively aggregated to obtain overall indexes using the product-specific FIPE CPI weights.

[13] See ABKR (2024) for the sensitivity of these price-setting statistics across different levels of $\pi_{i,t}^{CR\$} - \pi_t^{URV} = 1\%$ during the URV stabilization phase.

subsequent to the introduction of the Real (R$) currency. Specifically, prices adjusted with lower frequency, reduced magnitude and exhibited a more evenly distribution following the URV period. Consequently, based on the analysis of micro-level price adjustments, the results suggest that the role of the URV in fundamentally altering pricing behavior and achieving durable price stability was less decisive than expected.[14]

On the other hand, the monetary transition from URV to R$ represents a more nuanced shift compared to the preceding CR$ phase, suggesting that the URV functioned relatively effectively as a stabilization instrument. The policies associated with the URV also ensured the preservation of real wage purchasing power throughout the stabilization reform, a factor that undoubtedly facilitated its widespread and voluntary adoption—a dimension that micro-level price adjustment data are unable to capture fully.

2.4. State-Dependent Menu-Cost Model

This section applies a state-dependent sticky-price model to reconcile the Brazilian price-setting characteristics across the Real Plan stabilization program with the related literature.

We model nominal rigidity as resulting from a fixed cost of changing prices. Specifically, firms face menu costs when optimally setting prices through asymmetric and two-sided state-dependent pricing rules.[15] As opposed to time-dependent models (wherein prices are fixed for a preset amount of time), optimal adjustments are infrequent so that price rigidity remains as long as the optimal price is not driven too far away from the current price.

In what follows, we build on Almeida and Bonomo's (2002) inflationary economy setting to compare the price setting in Brazil across three different periods – before the URV (from August 1993 until February 1994), during the URV (from March until June 1994), and after the URV (from July until December 1994). Accordingly, we consider

[14] According to ABKR (2024), some factors that attenuate potential stabilization signals during the URV period like the underlying weekly structure of the microdata, the evidence that URV-based price quotations predominantly began in May, and the positive relationship between increasing inflation and higher relative price variability (RPV) preceding the implementation of the Real (R$) amidst heightened uncertainty.

[15] Bonomo (2000) shows that optimal two-sided rules are in practice similar to suboptimal one-sided ones because one of the bounds is not reached frequently.

CR$ prices for the pre-URV, both CR$ and URV prices for the URV period, and R$ prices afterward.

Almeida and Bonomo (2002) characterize an inflationary economy as an asymmetric distribution of deviations of individual prices from optimal ones that precede a disinflation policy. In that scenario of positive and constant inflation, there are more firms with prices lower than optimal compared to those with relatively higher prices. As policymakers reduce nominal money growth to zero, this price asymmetry interacts with symmetric idiosyncratic shocks (which trigger more upward than downward adjustments) to produce inflation persistence, thus increasing the disinflation costs. The authors argue this mechanism of inflationary inertia justifies the policymaker's concerns with the alignment of prices (concerning the frictionless optimal levels) in high-inflation economies, such as Brazil at the time of the stabilization plan.[16] Besides, they claim that the Real Plan is an example that, even in a context of low credibility, a fast disinflation without high costs is possible conditional on a mechanism of price alignment – the URV – to eliminate distribution asymmetry. As discussed in Section 2, the Real Plan used the URV as a price referential with a stable real value so that prices would henceforth be theoretically closer to the frictionless optimal levels before altering the currency to the R$.

Firstly, we revisit the menu cost model, including individual agents' state-dependent optimal pricing rule and the aggregate equilibrium arising from constant nominal aggregate demand growth and inflation, and the expected time of hitting a specific barrier (thus triggering price adjustments) for the first time. Then, we calibrate the model for each stabilization phase (before, during, and after the URV) using price quotes of the FIPE CPI microdata, and proceed with the discussion of the results.

2.4.1. Menu Cost Model Review

[16] The authors compute the costs of a money-based disinflation under different hypotheses about the credibility of the policy change. The assumptions range from full credibility (in which agents choose their optimal rules by forecasting the monetary policy as announced) to lack credibility scenario (in which agents do not change their pricing rules because they do not believe that there will be any monetary policy change).

We start revisiting the optimal pricing rule and the inflationary steady state as in Almeida and Bonomo (2002). In the following menu cost model, all the variables are in the log.

From the utility maximization, the optimal price (p_i^*) relative to the average level of prices (p) depends on the aggregate demand (y) and shocks specific to the firm (e_i) so that:[17]

$$p_i^* - p = vy + e_i \tag{2}$$

Given the nominal aggregate demand ($y + p$) equals the money supply (m), and assuming that v is one (which rules out strategic complementarities in prices and simplifies aggregation), the previous equation yields that the individual optimal price is the quantity of money plus an idiosyncratic shock.

$$p_i^* = m + e_i \tag{3}$$

However, it is costly to maintain individual prices aligned to their optimal levels because of fixed adjustment costs (k). Consequently, there is inaction at the microeconomic level, and some firms will have prices that are different from the frictionless optimal. In addition, as the individual price drifts away from its optimal, profit reduces at a rate $l(p_i - p_i^*)^2$, whereby l is also assumed to equal one.

Each firm solves for the optimal state-dependent pricing rule by taking the stochastic process for the p_i^* as given. Consider that e_i follows a Brownian motion without drift and m has a deterministic constant growth rate π (inflation rate). The frictionless optimal price is a Brownian motion with a drift given by the π and a Wiener process w_i so that:

$$dp_i^* = \pi dt + \sigma dw_i \tag{4}$$

Rewriting it in terms of the dynamics of price deviation $z_i = p_i - p_i^*$:

$$dz_i = -\pi dt + \sigma dw_i' \tag{5}$$

[17] See the utility maximization as in Ball and Romer (1989).

To find the optimal rule, whenever the price deviation z_i is inside the region where it is optimal not to adjust, the value function (C) – obtained through the cost minimization problem – should satisfy the following:[18]

$$\rho C(z_i)dt = z_i^2 dt + E_t(dC) \tag{6}$$

Subsequently, we find the ordinary differential equation for C by applying Ito's Lemma in the $E_t(dC)$, that is, the conditional expectation given price deviation at time t. It suggests the ensuing general format:[19]

$$C(x) = Ae^{\alpha x} + Be^{\beta x} + \frac{x^2}{\rho} - \frac{2\pi x}{\rho^2} + \frac{\sigma_i^2}{\rho^2} + 2\frac{\pi^2}{\rho^3}\text{, where} \tag{7}$$

$$\alpha = \frac{\pi}{\sigma_i^2} - \sqrt{\frac{\pi^2}{\sigma_i^4} + 2\frac{\rho}{\sigma_i^2}} \text{ and } \beta = \frac{\pi}{\sigma_i^2} + \sqrt{\frac{\pi^2}{\sigma_i^4} + 2\frac{\rho}{\sigma_i^2}}$$

Substituting the value function (7) into the value matching and smooth pasting conditions (8), one can obtain a system with five equations and five unknowns (A, B, L, c, U) that can be solved numerically.[20]

$$\begin{gathered} C(L) = C(c) + k \\ C(U) = C(c) + k \\ C'(L) = C'(U) = C'(c) = 0 \end{gathered} \tag{8}$$

In other words, α, β, and the constants A and B are jointly determined with the optimal rule parameters (L, c, U), where c is the target level for adjustments, and L and U are the levels of price deviation which trigger upward and downward adjustments, respectively.

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[18] Intuitively, this equation indicates that the required addition to the value function C should be equal to the current flow cost plus the stock increase in C.

[19] The ordinary differential equation for C is: $\frac{\sigma_i^2}{2}C''(x) - \pi C'(x) - \rho C(x) + x^2 = 0$.

[20] See Dixit (1993) for the intuition and derivation of the value matching and smooth pasting conditions. Specifically, while the value matching conditions state that the value function at a trigger level should be equal to the value function at the target level plus the adjustment cost, the smooth pasting conditions are optimality conditions so that the derivative of the value function at the optimal trigger and target levels should be zero.

In this setup, it is important to highlight that c is greater than zero and increases with π because price-setters adjust their prices at a level higher than the optimal one in an inflationary economy (to account for the expectation that prices might depreciate rapidly). Additionally, the size of the upward adjustments (c – L) also grows with π to prevent too high a frequency of adjustments (and the associated costs).[21]

In parallel, we can obtain that p is a function of both m and the average price deviation z as in equation (9). Besides, since $y + p = m$, the output level y is symmetric to z, that is, $y = -z$.

$$p = \int p_i di = \int (p_i^* + z_i) di = \int (m + e_i + z_i) di = m + z \qquad (9)$$

In this context, an inflationary steady state is a situation in which the rate of change of the money supply is unaltered for a reasonable amount of time so that the distribution of the price deviations becomes ergodic and the inflation rate equals the money supply growth rate. As a result, each inflationary steady state is associated with an ergodic distribution of price deviations through a pricing rule; even so, firm frictionless optimal price varies with time (due to the existence of idiosyncratic shocks).[22] Additionally, the distribution of price deviations is highly sensitive to (positive) inflation and asymmetric. Since optimal price tends to follow inflation, a higher fraction of firms is closer to the lower barrier L in comparison to the upper barrier U, which results in much more frequent upward than downward price changes.[23]

In parallel, Bonomo (2000) developed an explicit formula for the expected time of hitting a specific barrier for the first time. It follows similar steps employed to find the

[21] Almeida and Bonomo (2002) show how different values of π impact the optimal rule parameters (L, c, U) ceteris paribus.

[22] The ergodic distribution of price deviations has a density function in the format (see Bertola and Caballero, 1990):

$$f(z) = \begin{cases} Me^{\tau z} + N, & L \leq z \leq c \\ Pe^{\tau z} + Q, & c \leq z \leq U \\ 0, & otherwise \end{cases}$$

where $\tau = -{}^{2\pi}/_{\sigma^2}$

Since the density must be zero at the extremes, that is, $f(L) = 0 = f(U)$, continuous at c, that is, $f(c)^+ = f(c)$, and the integral of the density function over the appropriate range is one, then the constants (M, N, P, Q) are jointly determined by the following equations:

$$Me^{\tau L} + N = 0 = Pe^{\tau U} + Q$$
$$Me^{\tau c} + N = Pe^{\tau c} + Q$$
$$\frac{M}{\tau}(e^{\tau c} - e^{\tau L}) + \frac{P}{\tau}(e^{\tau U} - e^{\tau c}) + N(c - L) + Q(U - c) = 1$$

[23] Almeida and Bonomo (2002) illustrate the ergodic density of price deviations considering situations with zero and positive inflation.

optimal rule parameters (L, c, U). This statistic is convenient in situations wherein downward adjustments do not happen often. In this case, the expected time until an upward price adjustment (starting from price reset position c) becomes a good approximation for the time between general adjustments.

Starting from a position x, let $\theta_U(x)$ and $\theta_L(x)$ be the expected time until the respective barriers U and L are hit, then:

$$\theta_U(x) = \frac{(c-L)e^{\frac{2\pi x}{\sigma^2}}}{\pi\left(e^{\frac{2\pi L}{\sigma^2}} - e^{\frac{2\pi c}{\sigma^2}}\right)} + \frac{x}{\pi} - \frac{U}{\pi} - \frac{(c-L)e^{\frac{2\pi U}{\sigma^2}}}{\pi\left(e^{\frac{2\pi L}{\sigma^2}} - e^{\frac{2\pi c}{\sigma^2}}\right)} \tag{10}$$

$$\theta_L(x) = \frac{(c-U)e^{\frac{2\pi x}{\sigma^2}}}{\pi\left(e^{\frac{2\pi U}{\sigma^2}} - e^{\frac{2\pi c}{\sigma^2}}\right)} + \frac{x}{\pi} - \frac{L}{\pi} - \frac{(c-U)e^{\frac{2\pi L}{\sigma^2}}}{\pi\left(e^{\frac{2\pi U}{\sigma^2}} - e^{\frac{2\pi c}{\sigma^2}}\right)}$$

2.4.2. Model Calibration and Results

We solve the model for the equilibrium to characterize the inflationary steady state in Brazil before the URV (from August 1993 until February 1994) when prices were quoted in CR$, during the URV (from March until June) with both the CR$ and URV prices and after the URV (from July 1994 onwards) with the introduction of the R$. For each stabilization stage, we find the distribution of price deviations corresponding to a certain inflation rate by aggregating optimal individual pricing rules assuming that inflation is constant and positive, and all uncertainty is idiosyncratic.

Initially, the model has four parameters (ρ,π,k,σ) to be calibrated.

The time-discount rate parameter ρ is fixed at 2.5% annually following Bonomo (2000). However, given the monthly structure of the dataset, we use $\rho = 2.5\%/12$. Additionally, we set the monthly inflation parameter π according to the average monthly inflation over the corresponding period (see Table 2). Amid heightened uncertainty, particularly during the URV phase of the Real Plan, evidence suggests that monthly inflation of the cruzeiro real (CR$) exhibited a steady upward trend. Concurrently, URV inflation, while not negligible, remained comparatively lower, indicating that prices converted to URV were more volatile and exhibited greater fluctuations than the official URV indices.

Consequently, for each stabilization stage, we only need to calibrate two model parameters – the menu cost parameter k and the standard deviation of the idiosyncratic shock σ – based on two moments of the microdata – the monthly frequency f and absolute size of price change ΔP.

The frequency of price adjustment indicates the proportion of products whose prices changed compared to the corresponding week in the previous month. According to Table 2, prices denominated in CR$ experienced frequent adjustments throughout the URV phase, predominantly positive changes, as would be expected in a hyperinflationary environment. In contrast, the frequency of URV-based price adjustments, while similarly elevated, is more balanced, with a substantial proportion of negative price changes observed prior to the monetary transition. Following the implementation of the new currency, both positive and negative price adjustment frequencies decline significantly, resulting in a marked reduction in overall adjustment frequency.

The absolute magnitude of CR$ price changes closely aligns with CR$ inflation due to the high frequency of price adjustments, predominantly upward, under hyperinflationary conditions. Conversely, in the case of URV-denominated prices, the absolute size of changes exceeds the rate of URV inflation, as the latter metric incorporates a considerable share of price decreases within the overall frequency of changes.

Table 3 compares the targeted data moments with those obtained through model calibration and reports parameter estimates for every period. In addition to fitting for these targeted moments, we consider both the positive and negative frequency and size of price changes as untargeted ones (as shown in the numerical simulations of Table 4). In this case, once inverting $\theta_L(c)$ and $\theta_U(c)$ – the expected time of hitting a specific barrier L and U (which triggers price increases and decreases) starting from c – we obtain estimates for the average frequency of price increases and decreases, respectively, whereas $(c - L)$ and $(U - c)$ represent the absolute size of the upward and downward adjustments, respectively. In either case, the calibration results are satisfactory as seen by the data comparison (see Table 2).

Table 3 - Menu Cost Model Calibration

Periods	**Pre-URV**		**URV**				**Post-URV**	
Statistics & Parameters	**CR$**		**CR$**		**URV | R$**		**URV | R$**	
	Data	**Model**	**Data**	**Model**	**Data**	**Model**	**Data**	**Model**
Inflation rate	39.1	39.1	49.1	49.1	4.5	4.5	3.5	3.5
Frequency of change	96.8	94.7	98.5	96.3	95.1	92.6	64.5	65.0
Absolute size of change	41.3	41.3	50.7	51.0	14.6	14.9	14.5	15.3
Idiosyncratic shock		8.0		9.0		14.0		12.0
Menu cost		3.0		4.5		0.4		0.6

Notes: This table presents the monthly average (in %) of the targeted data moments (frequency and absolute size of price change) and the parameter estimates – menu cost and standard deviation of the idiosyncratic shock – across different periods. The pre-URV period goes from August 1993 until February 1994 with CR$ prices. The URV period goes from March until June 1994 with both the CR$ and URV prices. The post-URV goes from July until December 1994 with R$ prices. The time-discount rate parameter ρ is fixed at $2.5\%/12$.

Regarding the calibration of the parameters in Table 3, it is noticeable the menu costs of price adjustments during the URV phase of the stabilization. When both units of account coexisted, the estimated cost of changing prices was substantially higher in CR$ (4.5%) than in URV (0.4%) indicating that there were economic incentives to adhere to URV price conversion. Additionally, given the rising inflation rate context, the CR$ menu costs increased from 3% before the URV to 4.5% afterward.

In Table 4, we perform numerical simulations using the estimated parameters to find the optimal pricing rules (L, c, U). Throughout the URV period, the optimal rules (L, c, U) are (-26.3%, 24.7%, 51%) and (-14.1%, 1.2%, 15.4%) for prices in CR$ and URV, respectively. In other words, when the actual CR$ (URV) price becomes 26.3% (14.1%) lower or 51% (15.4%) higher than the optimal one, it is reset to a value of 24.7% (1.2%) higher than the optimum. The estimated rule also provides the absolute amount of that change through the size of price increases $(c - L)$ and decreases $(U - c)$, which are 51% (15.3%) and 26.3% (14.3%) for CR$ (URV) prices.

Table 4 - Numerical Simulations

Periods		Parameters			Optimal Pricing Rules					Frequency		Value
		π	σ	k	L	c	U	c-L	U-c	$\frac{1}{\theta_L(c)}$	$\frac{1}{\theta_U(c)}$	C(0)
Pre-URV		39.1	8.0	3.0	-21.5	19.8	41.3	41.3	21.5	94.7	0.0	20.5
URV	**CR$**	49.1	9.0	4.5	-26.3	24.7	51.0	51.0	26.3	96.3	0.0	31.2
	URV | R$	4.5	14.0	0.4	-14.1	1.2	15.4	15.3	14.3	60.0	32.6	3.5
Post-URV		3.5	12.0	0.6	-14.4	1.3	15.9	15.7	14.6	43	22.0	3.7

Notes: This table exhibits the model parameters' monthly average estimates (in %), the optimal pricing rules (L, c, U), the inverse of the expected time of hitting a specific barrier L and U starting from c $[\theta_L(c), \theta_U(c)]$, and the value function at $x = 0$ $[C(0)]$ across different periods. The pre-URV period goes from August 1993 until February 1994 with CR$ prices. The URV period goes from March until June 1994 with both the CR$ and URV prices. The post-URV goes from July until December 1994 with R$ prices. The time-discount rate parameter ρ is fixed at $2.5\%/12$.

Furthermore, we compute the expected time of hitting a specific barrier $[\theta_L(c), \theta_U(c)]$ and invert it to find the average frequencies related to it (f^+, f^-). The expected time to trigger a CR$ price adjustment shows that a positive variation $\theta_L(c)$ occurs practically every month while a negative one $\theta_U(c)$ is unlikely to happen. Contrastingly, the expected time of a URV price change is less asymmetric although positive adjustments are still significantly more likely. At the URV phase, for the CR$ (URV), the average frequencies of price increase and decrease are 96.3% (60%) and 0% (32.6%), respectively.

Finally, Table 4 contrasts the expected cost starting from $x = 0$ in CR$ and URV to verify the theoretical incentives for using either metric during the URV period. The increase in cost due to the usage of CR$ rather than URV is substantial (31.2% vs 3.5%). Hence, it strengthens the URV conversion appeal (contrary to the CR$) throughout the stabilization since it possesses not only a reduced adjustment cost but also lower projected expenses associated with it.

2.5. Concluding Remarks

In this paper, we revisited the Real Plan – the economic stabilization program that put an end to hyperinflation in Brazil in 1994 – by examining it through the lens of a menu cost model with state-dependent pricing rules, calibrated using unique micro data on prices at the establishment level of that period.

The central component of the Real Plan was the URV, a virtual currency introduced as a temporary unit of account between March and June 1994, prior to the implementation of the new monetary unit, the Real (R$), while the cruzeiro real (CR$) remained the official currency. The URV was designed to promote the synchronization of price adjustments across the economy, thereby mitigating inertial inflation and facilitating the disinflation process.

Given the Almeida and Bonomo (2002) menu cost model and the ABKR (2024) price-setting shreds of evidence around the Real Plan, we characterize price-setting behaviors during the URV phase and compare them with the periods preceding and succeeding the stabilization program. We reinforce that the Real Plan brought significant changes in Brazil's price-setting mechanisms, but the stabilization effects became prominent only after introducing the R$. Hence, the empirical evidence suggests that the URV's role in transforming price dynamics and achieving long-term price stability was less influential than initially hypothesized. Specifically, there is limited evidence of any substantial alteration in pricing behavior during the URV phase, despite theoretical incentives for URV-based price conversions (which were relatively cost-effective when both units of account coexisted).

Nonetheless, these results do not suggest that the URV was inconsequential. Rather, the URV served as a mechanism that facilitated a gradual transition to the new currency, enabled a rapid adjustment in the distribution of inflation rates across products, and safeguarded real wage purchasing power during the stabilization period – an effect not fully captured by micro-level price adjustment data.

2.6. Appendix: Pricing Statistics Review

Following ABKR (2024), we provide a brief summary of the relevant pricing statistics, taking into account the specific adjustments related to the URV.

The most granular unit in our analysis is item *i*, which represents a specific brand of product *j* sold at outlet *n* during week *t*. Each product is an aggregation of various brands, with each brand comprising multiple items. For each item i, we have its corresponding price information $P_{i,n,t}$.

We define an indicator function $I_{i,n,t}$, which takes a value of 1 if a price change occurs when comparing the same weeks across consecutive months. It can generally be represented as:

$$I_{i,n,t} = \begin{cases} 1, \ if \ P_{i,n,t} \neq P_{i,n,t-4} \\ 0, \ if \ P_{i,n,t} = P_{i,n,t-4} \end{cases}, \tag{11}$$

$$\text{for } i = 1, \ \dots, I, n = 1, \dots, N, \text{ and } t = 1, \dots, T$$

Let $\pi_{i,n,t}$ represent the inflation of item *i* sold at outlet *n* at time *t*, and $\pi_{i,t}$ denote the average inflation of the brand associated with the item *i* across all its outlets *n*. From this, we derive the average inflation of product *j* across all its brands, denoted $\pi_{j,t}$, where $card(s_{j,t})$ represents the number of brands within product *j*. Finally, we calculate the aggregate inflation measure, or overall weighted inflation (π_t), using the product-specific FIPE CPI weights $w_{j,t}$.

$$\pi_t = \sum_{j=1}^{J} w_{j,t} \, \pi_{j,t}, \tag{12}$$

$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1, \pi_{j,t} = \frac{1}{card(s_{j,t})} \sum_{i \in s_{j,t}} \pi_{i,t},$$

$$\pi_{i,t} = \frac{1}{n_{i,t}} \sum_n \pi_{i,n,t}, \text{ and } \pi_{i,n,t} = \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}}$$

$$\text{for all } i, n, t, \text{ and } j$$

The frequency of price changes, or extensive margin (f_t), reflects the proportion of products that experienced a price change compared to the same week in the previous month. To calculate this, we aggregate items *i* into products *j* by taking the simple mean to determine the average frequency of price changes for j ($f_{j,t}$), where $card(s_{j,t})$ represents the number of non-missing brands that comprise j. We then calculate the overall f_t using the official CPI weights.

$$f_t = \sum_{j=1}^{J} w_{j,t}\, f_{j,t}, \tag{13}$$

$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,\ f_{j,t} = \frac{1}{card(s_{j,t})}\sum_{i\in s_{j,t}} I_{i,t}, \text{ and } I_{i,t} = \frac{1}{n_{i,t}}\sum_n I_{i,n,t}$$

$$\text{for all } i, n, t, \text{ and } j$$

The average absolute size of a price change, or intensive margin (ΔP_t), evaluates instances where price changes occur (i.e., $I_{i,n,t} = 1$). It measures the magnitude of price changes in absolute terms between corresponding weeks of consecutive months. As before, we aggregate brands into products (ΔP_t) using simple means, where $card(s_{j,t}^{*})$ represents the number of non-zero price changes for items i within product j, before calculating the overall ΔP_t.

$$\Delta P_t = \sum_{j=1}^{J} w_{j,t}\, \Delta P_{j,t}, \tag{14}$$

$$\text{where } \sum_{j=1}^{J} w_{j,t} = 1,\ \Delta P_{j,t} = \frac{1}{card(s_{j,t}^{*})}\sum_{i\in s_{j,t}^{*}} \Delta P_{i,t},$$

$$\Delta P_{i,t} = \frac{1}{n_{i,t}}\sum_n \Delta P_{i,n,t}, \text{ and } \Delta P_{i,n,t} = \left|\frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} \neq 0\right|$$

$$\text{for all } i, n, t, \text{ and } j$$

In summary, these indicators offer a view of the proportion of products that experienced price changes over two consecutive months, and for those that did, the magnitude of the change. Additionally, we decompose f_t and ΔP_t into the frequency of price increases (f_t^{+}) and decreases (f_t^{-}), as well as the size of price increases (Δp_t^{+}) and decreases (Δp_t^{-}).

$$f_t = f_t^{+} + f_t^{-} \text{ such that} \tag{15}$$

$$f_t^{+} = \sum_{j=1}^{J} w_{j,t}\left(\frac{1}{card(s_{j,t}^{+})}\sum_{i\in s_{j,t}^{+}} I_{i,t}^{+}\right),$$

$$\text{where } I_{i,t}^{+} = \frac{1}{n_{i,t}}\sum_n I_{i,n,t}^{+}, \text{ and } I_{i,n,t}^{+} = \begin{cases} 1, & if\ P_{i,n,t} > P_{i,n,t-4} \\ 0, & if\ P_{i,nt} \leq P_{i,n,t-4} \end{cases}$$

$$f_t^- = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^-)} \sum_{i \in s_{j,t}^-} I_{i,t}^- \right),$$

$$\text{where } I_{i,t}^- = \frac{1}{n_{i,t}} \sum_n I_{i,n,t}^-, \text{ and } I_{i,n,t}^- = \begin{cases} 1, \; if \; P_{i,n,t} < P_{i,n,t-4} \\ 0, \; if \; P_{i,nt} \geq P_{i,n,t-4} \end{cases}$$

$$\text{for all } i, n, t, \text{ and } j$$

$$\Delta P_t = \frac{f_t^+}{f_t} \Delta P_t^+ + \frac{f_t^-}{f_t} \Delta P_t^- \text{ such that} \tag{16}$$

$$\Delta P_t^+ = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^{+*})} \sum_{i \in s_{j,t}^{+*}} \Delta P_{i,t}^+ \right),$$

$$\text{where } \Delta P_{i,t}^+ = \frac{1}{n_{i,t}} \sum_n \Delta P_{i,n,t}^+, \text{ and } \Delta P_{i,n,t}^+ = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} > 0 \right|$$

$$\Delta P_t^- = \sum_{j=1}^{J} w_{j,t} \left(\frac{1}{card(s_{j,t}^{-*})} \sum_{i \in s_{j,t}^{-*}} \Delta P_{i,t}^- \right),$$

$$\text{where } \Delta P_{i,t}^- = \frac{1}{n_{i,t}} \sum_n \Delta P_{i,n,t}^-, \text{ and } \Delta P_{i,n,t}^- = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} < 0 \right|$$

$$\text{for all } i, n, t, \text{ and } j$$

To compute price changes in URV, it is necessary to slightly modify these pricing statistics. Crucially, price changes in CR\$ ($\pi_{i,t}^{CR\$} \neq 0$) differ from the ones in URV ($\left|\pi_{i,t}^{URV}\right| > \bar{\pi}_{i,t}^{URV}$) because the latter depends on estimates of the weekly URV that prevailed when each item was surveyed. As a result, the frequency and the absolute size of price changes in URV can be decomposed, respectively, as:

$$I_{i,n,t}^+ = \begin{cases} 1, \; if \; P_{i,n,t} > P_{i,n,t-4} \\ 0, \; if \; P_{i,n,t} \leq P_{i,n,t-4} \end{cases} \rightarrow I_{i,n,t}^{URV+} = \begin{cases} 1, \; if \; \pi_{i,n,t}^{URV} > \bar{\pi}_{i,n,t}^{URV} \\ 0, \; if \; \pi_{i,n,t}^{URV} \leq \bar{\pi}_{i,n,t}^{URV} \end{cases} \tag{17}$$

$$I_{i,nt}^- = \begin{cases} 1, \; if \; P_{i,n,t} < P_{i,n,t-4} \\ 0, \; if \; P_{i,n,t} \geq P_{i,n,t-4} \end{cases} \rightarrow I_{i,t}^{URV-} = \begin{cases} 1, \; if \; \pi_{i,n,t}^{URV} < -\bar{\pi}_{i,n,t}^{URV} \\ 0, \; if \; \pi_{i,n,t}^{URV} \geq -\bar{\pi}_{i,n,t}^{URV} \end{cases}$$

$$\Delta P_{i,n,t}^+ = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} > 0 \right| \rightarrow \Delta P_{i,n,t}^{URV+} = \left| \frac{P_{i,n,t}^{URV} - P_{i,n,t-4}^{URV}}{P_{i,n,t-4}^{URV}} > \bar{\pi}_{i,n,t}^{URV} \right| \tag{18}$$

$$\Delta P_{i,n,t}^- = \left| \frac{P_{i,n,t} - P_{i,n,t-4}}{P_{i,n,t-4}} < 0 \right| \rightarrow \Delta P_{i,n,t}^{URV-} = \left| \frac{P_{i,n,t}^{URV} - P_{i,n,t-4}^{URV}}{P_{i,n,t-4}^{URV}} < -\bar{\pi}_{i,n,t}^{URV} \right|$$

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Chapter 3

The Effects of High-Frequency Monetary Policy Shocks on Financial Assets in Brazil[1]

3.1. Introduction

Monetary policy is a central tool for influencing economic activity, ensuring financial stability, and guiding market expectations. Recent advances in high-frequency data availability have promoted a shift in the literature, enabling the identification and measurement of monetary policy shocks with greater precision. This paper explores the effects of different types of high-frequency monetary policy shocks on financial assets in Brazil.

Over the last decades, several studies have employed changes in interest rates around central bank announcements to measure monetary policy effects on financial markets. Seminal contributions by Kuttner (2001) and Gurkaynak et al. (2005) have highlighted the immediate impact of monetary policy on asset prices, while Bernanke and Kuttner (2005) and Wright (2012) advanced the analysis by focusing on broader financial market implications. More recently, Swanson (2021) provided a comprehensive overview of how financial markets respond to monetary policy in low-interest environments.

Beyond financial markets, the macroeconomic effects of monetary policy have been extensively studied. Cochrane and Piazzesi (2002) examined the relationship between policy rates and macroeconomic indicators, while Faust et al. (2003, 2004) and Stock and Watson (2012) provided critical insights into the transmission mechanisms of monetary policy shocks. Gertler and Karadi (2015) and Ramey (2016) further explored how these shocks affect macroeconomic variables, emphasizing the importance of identifying credible shock measures.

[1] This is joint work with Marco Bonomo and Miguel Bandeira.

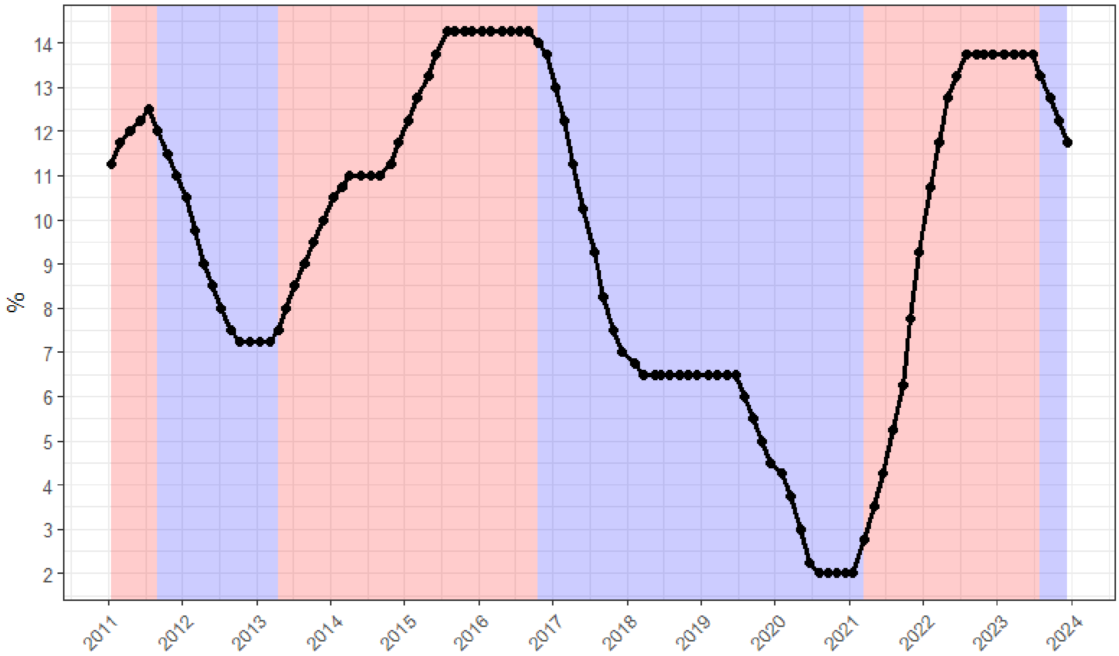
Recent developments have introduced novel high-frequency shock measures, particularly for identifying monetary policy shocks. Swanson and Jayawickrema (2024) developed a comprehensive database for the U.S., while Altavilla et al. (2019) and Braun et al. (2024) compiled comparable data to the Euro Area and the U.K., respectively. However, the application of high-frequency shock measures to emerging markets, such as Brazil, remains limited.

This paper seeks to address this gap by compiling a novel dataset of Brazilian monetary policy events spanning 1996 to 2023. This comprehensive dataset includes traditional monetary announcements, namely the Monetary Policy Committee (MPC) statements, minutes, and quarterly inflation reports. It also includes the inflation target announcement setting by the National Monetary Council (NMC), the open letters by the Central Bank of Brazil (BCB)'s Governor to the chairman of the NMC (in case the inflation target is missed in the previous year), and, importantly, the high-frequency information on the agendas of BCB officials. As a result, the dataset represents a relevant contribution to the literature as it provides a foundation for future research on Brazilian monetary policy by enabling a nuanced analysis of the effects of different types of monetary communication.

The Brazilian economy offers a unique context for such an investigation, given its history of inflationary pressures, structural reforms, and the BCB's evolving monetary policy framework. Figure 1 illustrates the historical evolution of the Selic interest rate, which has been a central tool of monetary policy. Over the period from 2011 to 2023, the Selic rate experienced significant fluctuations, with periods of monetary tightening (highlighted in red) and easing (highlighted in blue). These shifts reflect the Central Bank's response to varying macroeconomic conditions, including inflationary pressures and economic downturns.

Within the Brazilian context, numerous studies have contributed to the understanding of monetary policy transmission mechanisms and effects of BCB's communication. Gonçalves and Guimarães (2011) examined the relationship between monetary policy, sovereign default risk, and exchange rate dynamics, highlighting how policy decisions can affect external vulnerability and currency valuations. Carvalho et al. (2013) and Machado (2014) conducted a quantitative analysis of the content and tone of the MPC statements, providing evidence that the informational content exerts meaningful effects on the term structure of interest rates. Cabral and Guimarães (2015) analyzed the impact of MPC statements on domestic financial markets, demonstrating that such communications exert significant influence not only across the yield curve and the stock market index, but also enhance the predictability of the content subsequently disclosed in the MPC minutes.

Figure 1 - Selic Interest Rate (%): 2011-2023



Notes: The black dots indicate the Selic rate for each MPC statement. The red (blue) sections represent periods of monetary tightening (easing).

More recently, Da Silva et al. (2022) applied a bivariate VAR-GARCH model to jointly model the dynamic interactions between news related to the inflation targeting regime and financial variables. Fasolo et al. (2022) introduced a novel hierarchical topic model approach to classify and evaluate MPC statements and minutes, offering a more detailed and data-driven characterization of their informational structure. Feibert (2022) estimated the impact of monetary policy surprises on multiple asset classes in Brazil controlling for fiscal and political turmoil obtained through natural language processing (NLP) techniques on news articles.

While these studies have advanced the empirical characterization of monetary policy transmission in Brazil, most analyses have focused on MPC statements and minutes. In contrast, this study expands the scope of investigation by integrating a comprehensive set of monetary policy communications – including MPC statements, minutes, inflation reports, inflation target announcements, open letters, and the high-frequency agendas of BCB authorities – and by assessing their impact across a broader range of financial assets. As a result, it enables a more granular understanding of how central bank communication affects the domestic financial market dynamics.

First, we combine the Brazilian monetary policy events database with daily financial data from the Broadcast Terminal, which is available since 2011 to compare the influence of different monetary shocks on financial assets. We collected price quotes (open, close, maximum, and minimum) of the following assets: Ibovespa stock price index, USDBRL foreign exchange rate, and the DI-PRE swap yields of different maturities ranging from 1 month to 5 years. Specifically, since there is information on the open and close price quotes, each asset is observed twice daily, allowing us to calculate overnight returns (close-to-open price changes) for events such as MPC statements, MPC minutes, and inflation reports, and intraday returns (open-to-close price changes) for events such as speeches, inflation target settings, and open letters.

The findings indicate that MPC statements and the public agenda of BCB authorities exert considerable influence on financial assets, with these engagements emerging as a particularly influential source of shocks during trading hours compared to other forms of monetary policy communication.

Second, we estimate a series of Structural Vector Autoregressive (SVAR) models with identification through heteroskedasticity. Rigobon (2003) and Rigobon and Sack (2003, 2004, 2005) pioneered the use of heteroskedasticity-based identification strategies to disentangle monetary shocks from market responses, while Wright (2012) expanded these methodologies to high-frequency data. Under the assumption that periods with monetary events exhibit higher conditional volatility than non-event ones, this method allows to recover structural monetary policy shocks from high-frequency data using selected financial variables, including DI-PRE swap yields of various maturities, the Ibovespa stock index, and the USDBRL exchange rate. Then we estimate impulse-response functions (IRFs) via Local Projections (LP), which regress those financial variables at each horizon on the estimated shocks to trace their dynamic responses over time in a specification-robust and model-flexible manner (Jordà, 2005). In other words, the SVAR identified through heteroskedasticity delivers the exogenous high-frequency shock series, while the LP framework provides a robust method to estimate the corresponding IRFs. Moreover, our LP estimations explicitly account for asymmetries in the responses during periods of monetary tightening versus easing, by allowing impulse responses to differ across these policy stances.

We estimate four different models. The first model isolates the effects of monetary policy shocks by focusing exclusively on MPC statement announcements. The second model accounts for the participation of BCB officials in various events. The third model expands the analysis to include all forms of monetary announcements, encompassing MPC statements, MPC minutes, quarterly inflation reports, inflation target settings, open letters from the BCB Governor, and the agendas of BCB Governors. The fourth model refines this approach by restricting the analysis to the most significant shocks, identified by aligning monetary announcements with periods of extreme changes in the DI-PRE swap rates for 1-year and 2-year maturities.

The results underscore the critical role of monetary policy announcements in shaping financial market dynamics. Across all specifications, the linear estimated IRFs reveal positive, statistically significant, and persistent responses of interest rates to a contractionary monetary policy shock. Overall, the shocks induce a stronger upward repricing of short- and medium-term rates, with a more moderate pass-through to longer maturities. Notably, the MPC President's public agenda is associated with a comparatively higher immediate response on longer-term maturities, highlighting the importance of forward guidance in shaping investor expectations for future policy directions. Equity markets exhibit pronounced and statistically significant negative reactions with varying degrees of late reversal. In contrast, the evidence for the Brazilian real is mixed: while most impulse responses indicate short-lived effects, MPC agenda stands out for producing a sharp and persistent depreciation of the BRL on impact.

Additionally, the analysis demonstrates that the monetary policy stance significantly conditions the transmission of monetary shocks, as revealed by the non-linear estimated IRFs. For interest rates, shocks produce persistent and statistically significant DI-PRE yield responses that exhibit clear state dependence. Under tightening, yield increases are more durable and sustained, whereas under easing, responses follow a hump-shaped pattern, building through intermediate horizons before reversing. This configuration implies that easing actions exert stronger effects on the term structure in the short to medium term (up to roughly six months), while tightening effects dominate over longer horizons, particularly around one year.

Other risk-sensitive assets also display regime dependence, with the most pronounced effects occurring during tightening phases. Equity prices show persistent declines in response to tightening shocks, while easing episodes typically generate initial gains that later reverse or lose statistical significance. This pattern aligns with higher discount rates and weaker growth expectations during restrictive monetary cycles. In foreign exchange markets, easing responses are mixed, whereas the BRL depreciates following tightening shocks, potentially reflecting structural vulnerabilities in Brazil's macroeconomic fundamentals and heightened perceptions of country-specific risk.

We also assess the effects of the identified monetary policy shocks on macroeconomic variables, specifically the Gross Domestic Product (GDP) and the Consumer Price Index (CPI), through linear LP. Consistent with expectations, the estimated tightening shocks lower the price level at medium horizons while imposing comparatively small output costs, irrespective of the shock source.

From an economic perspective, these results align with theoretical expectations about the transmission mechanisms of monetary policy. Short-term interest rates act as immediate benchmarks for financial instruments, while longer-term rates capture expectations about future policy and macroeconomic conditions. Equity and currency markets, in turn, reflect the broader implications of monetary policy, encompassing corporate profitability, capital flows, and risk perceptions.

In parallel, from a policy perspective, the findings emphasize the importance of clear, consistent, and transparent communication strategies. Effective monetary policy announcements not only manage market expectations but also mitigate volatility, thereby supporting the central bank's objectives of financial stability and economic growth.

The remainder of the paper is organized as follows. Section II details the construction and characteristics of the monetary policy announcement database. Section III provides a comparative analysis of the influence of different monetary shocks on financial assets. Section IV reviews the SVAR and the LP models. Section V presents the IRFs and discusses the results. Section VI concludes.

3.2. Brazilian Monetary Policy Events Database

In this section, we present the features of our Brazilian monetary policy events database. The data contains public information that was collected from the Central Bank of Brazil (BCB) official website. For each event, there is information regarding the type of monetary announcement, a brief description, and the date.[2]

Table 1 provides a summary of the monetary policy events in Brazil from 1996 to 2023. The table categorizes monetary events into six distinct types, providing information on the total number of occurrences, the earliest available data for each category, and their typical release periods.

Table 1 - Monetary Policy Communication in Brazil (1996-2023)

Monetary Policy Events	#	Earliest Available Data	Release Period
MPC Statements	259	June-96	Post-market
MPC Minutes	247	May-98	Pre-market
Quarterly Inflation Report	99	June-99	Pre-market
Inflation target setting	26	June-99	Irregular
Open letter by the BCB's Governor	7	January-02	Irregular
Agenda of BCB Authorities	1997	November-08	Irregular

The Monetary Policy Committee (MPC) statements comprise 259 recorded events, covering both scheduled and unscheduled monetary policy decisions. The first MPC meeting dates back to June 1996. These statements, typically released post-market following the meetings, are crucial in communicating monetary policy decisions to the market.

The MPC minutes comprise 247 instances since May 1998 (the MPC Meeting #21). These documents are released pre-market, providing a detailed account of the discussions within the MPC that led to policy decisions.

The Quarterly Inflation Report, introduced in June 1999, includes 99 reports released pre-market and offer a comprehensive analysis of the macroeconomic environment and the BCB's projections for inflation and economic activity.

The inflation target setting, irregularly released by the National Monetary Council (NMC) since June 1999, accounts for 26 events. These announcements typically occur

[2] Starting from February 19, 2018, detailed information regarding the expected time interval for each monetary policy event is also available. Prior to this date, the scheduling of events in the agenda was limited to broader time categories, specifying only whether the occasion was planned for the morning, afternoon, or evening.

once a year to define the inflation targets for the following years, which serve as a key anchor for monetary policy.

The dataset also includes 7 open letters from the BCB Governor to the NMC chairman, starting in January 2002. These letters are issued irregularly and only when inflation deviates significantly from the set target, exceeding the upper limit of the inflation target interval, explaining the reasons for the deviation and the measures to address it.

Finally, the dataset contains the BCB Governor' agenda, which includes public and private engagements, dating back to November 2008.[3] This represents the most frequent type of monetary announcement in the dataset, with 1,997 filtered event days. This frequency underscores their significance as a source of monetary policy signals. Before any filtering, there were 4,139 days entries in the dataset. However, entries for weekends, holidays, vacations, blackout periods, internal activities, and transportation-only activities were excluded since they were unlikely to contain episodes relevant to monetary policy.

As a result, our comprehensive dataset captures the diversity and richness of monetary policy communications in Brazil, providing a robust foundation for analyzing the impact of high-frequency monetary policy shocks on financial assets.

3.3. Monetary Policy Shocks and Financial Asset Prices

In this section, we combine our Brazilian monetary policy events database with financial data to compare the influence of different monetary shocks on financial assets.

From the Broadcast Terminal, we collected price quotes of the following assets: Ibovespa stock price index, USDBRL foreign exchange rate, and the DI-PRE swap yields of different maturities ranging from 1 month to 5 years.[4] Daily information is available from 2011 onwards for each asset, including the open, close, maximum, and minimum price quotes. In Appendix A, Figures A1 to A4 illustrate the evolution of these assets over time.

[3] We also collected the agenda of two of the most important directors of the Board of Governors – the Monetary Policy and the Economic Policy Directors. Information dates back to May 2012.

[4] The DI PRE swap is a financial derivative in which one party exchanges a fixed interest rate (PRE) for a floating rate based on the DI (the overnight interbank deposit rate). By doing so, one is able to hedge against fluctuations in interest rates.

Given the open and close price quotes for the analyzed assets, we compute series of intraday (open-to-close) and the overnight (close-to-open) log-returns. Intraday returns apply for speeches, inflation target setting, and open letter, which tend to occur during market hours. On the other hand, overnight returns correspond to MPC statements, MPC minutes, and quarterly inflation report. The statements are typically released after market whereas the minutes and inflation report are usually published before the financial market opens.

Table 2 presents the mean absolute changes (log-returns) in financial assets per monetary policy announcement from 2011 to 2023. This table shows the quantities of each type of monetary event and brings the magnitude associated to it across different asset classes. In addition, we test the null hypothesis that the mean absolute return observed during periods without monetary shocks equals the mean observed during periods with any monetary shock, employing the unequal variance t-test (Welch's t-test).

Table 2 reports 848 periods (which last about 12 hours each) featuring a monetary announcement and 4,270 periods without such shocks. Given the large number of engagements on the BCB President's agenda, we retain only those entries that coincide with market reactions outside the interquartile range of returns on 1- and 2-year swap rates – that is, reactions sufficiently pronounced to plausibly convey monetary-policy information. Under this criterion, 606 announcement windows include at least one engagement by the BCB President.

The results indicate that the absolute changes in financial assets are markedly higher during periods with monetary policy events compared to periods without such events (see columns 2 and 3). All of these differences are statistically significant at the 1% level. This evidence underscores the pronounced impact of monetary policy announcements on market volatility.

Additionally, Table 2 highlights the pronounced influence of MPC statements and the engagements of BCB authorities on financial assets, with absolute changes significantly exceeding those observed during periods without monetary shocks. This evidence highlights their role as a critical source of market-relevant information. Specifically, the MPC statements are associated with the largest changes on DI-PRE swap yields across most maturities (ranging from 3 months to 1 year) while the BCB Governor's engagements tend to have a significant impact even on longer-term maturities, the Ibovespa stock price index, and USDBRL exchange rate. It is worth noticing that the

Ibovespa index and the exchange rate exhibit relatively muted responses to MPC statements, likely due to their release outside of regular trading hours.

The MPC minutes and the quarterly inflation reports tend to show less significant effects on interest rates compared to other monetary event types. Since the minutes and the inflation reports are released outside of regular trading hours, their effect on equities and FX markets are relatively muted as well.

Table 2 - Mean Absolute Change (log-return) per Announcement (in %): 2011-2023

	No Shock	Any Shock	Agenda of BCB Authorities	MPC Statements	MPC Minutes	Quarterly Inflation Report	Inflation target setting	Open letter by the BCB's Governor
#	4270	848	606	85	98	51	13	4
DI PRE - 1mo	0.58	0.75***	0.63	1.02**	0.84	1.41**	**1.53**	0.62
DI PRE - 3mo	0.68	0.83***	0.89***	**0.92***	0.58	0.54	0.35**	0.21**
DI PRE - 6mo	0.56	0.88***	0.92***	**1.13*****	0.58	0.7	0.38**	0.26**
DI PRE - 1y	0.56	1.11***	1.17***	**1.36*****	0.78**	0.7	0.68	0.43
DI PRE - 2y	0.63	1.25***	**1.39*****	1.24***	0.75	0.64	0.93	0.74
DI PRE - 3y	0.63	1.15***	**1.35*****	0.79**	0.58	0.58	0.99	0.78
DI PRE - 4y	0.61	1.08***	**1.25*****	0.79	0.62	0.53	0.87	0.85
DI PRE - 5y	0.65	1.03***	**1.19*****	0.74	0.62	0.53	0.91	0.89
Ibovespa	0.40	0.86***	**1.2*****	0.01***	0.01***	0.01***	0.9***	1.1
USDBRL	0.27	0.6***	**0.82*****	0.04***	0.04***	0.07***	0.74**	0.82

Notes: This table presents the mean absolute log-returns of various financial assets (listed in rows) during periods associated with a specific monetary shock, any monetary shock, or no monetary shock, spanning the years 2011 to 2023. Log-returns are calculated by comparing the close-to-open price for overnight returns and the open-to-close price for intraday returns. Intraday returns apply to events such as speeches, inflation target settings, and open letters, while overnight returns correspond to statements, minutes, and inflation reports. For each asset, the largest change is highlighted in bold. To evaluate whether the mean return in periods without monetary shocks differs significantly from the mean return in periods with any monetary shock (as specified in the columns), we conduct an unequal variance (Welch) t-test. * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

The inflation target settings and the open letters by the BCB Governor, though less frequent, also produce relevant market reactions. The influence of open letters can be attributed to the exceptional circumstances under which they are issued – when the IPCA inflation deviates from upper limit of the inflation target in the preceding year – since these periods tend to be associated with higher volatility. On the other hand, the magnitude of changes associated with inflation target settings indicates their importance in signaling shifts in long-term monetary policy frameworks.

In the Appendix, Table A1 shows the median absolute changes (log-returns) in financial assets per monetary policy announcement from 2011 to 2023.[5] In general, Table A1 confirms the findings of Table 2, showing that monetary shocks lead to consistent market movements, albeit with slightly less variability than the mean changes suggest.

[5] While mean changes may be influenced by outliers, median changes provide a more stable measure of typical market reactions.

In summary, the evidence suggests that periods with monetary policy events tend to exhibit a relatively higher volatility across a range of financial assets in Brazil. Moreover, high-frequency monetary policy shocks are marked by heterogeneous effects across asset classes, with engagements by BCB authorities emerging as a particularly influential source of market disturbances during trading hours. These findings underscore the critical role of monetary policy communication in shaping financial market dynamics and highlight the importance of incorporating a broad set of events in analyzing monetary policy impacts.

3.4. Structural Vector Autoregressive (SVAR) Model with Identification through Heteroskedasticity and Local Projections (LP) Model Review

In this section, we briefly revisit the Structural Vector Autoregressive (SVAR) model with identification through heteroskedasticity and the Local Projections (LP) model. The former is used to obtain high-frequency series of monetary policy shocks while the latter is employed to trace out impulse response functions using the estimated shocks.

3.4.1. SVAR Model with Identification through Heteroskedasticity

Following the methodologies outlined by Rigobon (2003) and Rigobon and Sack (2003, 2004, 2005), as well as the adaptation by Wright (2012), this identification strategy assumes that monetary policy event periods exhibit higher volatility compared to non-event periods. In other words, the SVAR model relies on changes in the covariance matrix of asset returns during high-volatility periods to achieve identification, that is, to disentangle structural shocks from the reduced-form residuals.

Given the reduced-form VAR representation for a vector Y_t that contains p financial variables, the reduced-form errors ε_t can be expressed as a series of structural shocks $\eta_{i,t}$ independent of each other and over time:

$$A(L)Y_t = \mu + \varepsilon_t \text{ , where } \varepsilon_t = \sum_{i=1}^{p} R_i \eta_{i,t} \quad (1)$$

Assuming $\eta_{1,t}$ refers to the monetary shock, which has mean zero and variance σ_1 on periods with monetary events and σ_0 otherwise (periods without monetary announcements). Hence, the identification assumption is $\sigma_1 \neq \sigma_0$. In turn, the remaining structural stocks have mean zero and variance one, that is, there is no difference in volatility between those periods. It is important to highlight that our interest lies in identifying only the monetary policy shock, not the other structural shocks.

We estimate the VAR, compose the sample variance-covariance matrix of residuals on periods with and without monetary events $\hat{\Sigma}_1$ and $\hat{\Sigma}_0$, respectively, so that the vector R_1 can be estimated according to:[6]

$$\hat{R}_1 = \arg min_{R_1} \left[vech\left(\hat{\Sigma}_1 - \hat{\Sigma}_0\right) - vech(R_1 R_1')\right]' \left[\hat{V}_0 + \hat{V}_1\right]^{-1}$$
$$\left[vech\left(\hat{\Sigma}_1 - \hat{\Sigma}_0\right) - vech(R_1 R_1')\right], \quad (2)$$

where Σ_1 and Σ_0 are the variance-covariance matrix of reduced-form errors on events and non-events periods, respectively, such that $\Sigma_1 - \Sigma_0 = R_1 R_1' \sigma_1^2 - R_1 R_1' \sigma_0^2 = R_1 R_1'(\sigma_1^2 - \sigma_0^2)$,

As Wright (2012) demonstrated, bootstrap inference can be used to test the identification condition hypothesis that periods with monetary announcements are no different that periods without it, that is, $\Sigma_1 = \Sigma_0$.[7] Moreover, bootstrap also allows for the estimation of confidence intervals for the impulse response functions.

In our application, we construct separate volatility regimes for periods with and without monetary policy events. By combining the SVAR model with high-frequency data, we estimate series of monetary policy shocks to evaluate their effects on financial asset prices.

3.4.2. LPs Model

[6] The vector R_1 can be identified using only $\Sigma_1 - \Sigma_0 = R_1 R_1' \sigma_1^2 - R_1 R_1' \sigma_0^2 = R_1 R_1'(\sigma_1^2 - \sigma_0^2)$, where $\sigma_1^2 - \sigma_0^2 = 1$.

[7] The test statistic for the identification condition is: $\left[vech\left(\hat{\Sigma}_1 - \hat{\Sigma}_0\right)\right]' \left[\hat{V}_0 + \hat{V}_1\right]^{-1} \left[vech\left(\hat{\Sigma}_1 - \hat{\Sigma}_0\right)\right]$.

To estimate the dynamic responses of financial variables to monetary shocks, this study employs the Local Projections method introduced by Jordà (2005). Unlike the traditional VAR approach, which derives impulse response functions through recursive substitution of the estimated system, the LPs rely on direct estimation of the response at each forecast horizon via horizon-specific regressions.[8]

Formally, for each forecast horizon $h = 0, \dots, H$, the LP framework estimates a regression of the form:

$$y_{t+h} = \alpha_h + \beta_h x_t + \sum_{j=1}^{p} \gamma_{h,j} y_{t-j} + \varepsilon_{t+h} \tag{3}$$

where y_{t+h} is the future value of the dependent variable of interest observed h periods ahead of time t, α_h is the horizon-specific intercept, β_h is the response coefficient that captures the effect of the shock x_t on the future value y_{t+h}, $\gamma_{h,j}$ are the coefficients on the lagged dependent variables y_{t-j} for $j = 1, \dots, p$ lags, and ε_{t+h} is the horizon-specific error term.

Each horizon-specific regression in (3) can be estimated via Ordinary Least Squares (OLS), provided that the error terms are conditionally homoskedastic and serially uncorrelated. To account for potential serial correlation in the forecast errors, especially at longer horizons, HAC (Heteroskedasticity and Autocorrelation Consistent) robust standard errors (e.g., Newey and West) are employed.

The sequence of coefficients $\{\beta_h\}_{h=0}^{H}$ traces out the impulse response function (IRF) of y_t to a unit change in x_t. In our application, we evaluate the response of different asset classes to the estimated monetary shocks using IRF that capture both the magnitude and persistence of these shocks.

3.5. Impulse-Response Functions (IRFs) Results

In this section, we estimate distinct models using he SVAR and the LP frameworks – the heteroskedasticity-identified SVAR employed to obtain the shock series, and the LP specification used to estimate impulse-response functions conditional

[8] By estimating one regression per horizon, the error terms are not constrained to follow any particular structure, which contrasts with the VAR framework.

on those shocks. First, we present the linear estimated IRFs. Second, we evaluate potential asymmetries in the responses during periods of monetary tightening and easing through non-linear estimated IRFs. Finally, we also compute the effects of those shocks on selected macroeconomic variables.

3.5.1. Linear Estimated IRFs

Following the methodology proposed by Wright (2012), we employ four distinct SVAR models with identification through heteroskedasticity to obtain high-frequency monetary policy shocks. It is assumed that periods characterized by monetary policy announcements exhibit greater volatility relative to non-event periods. As discussed in Section 3.3, this assumption is supported by the evidence according to Table 2.

The first model isolates the effects of monetary policy shocks by focusing exclusively on MPC statement announcements. The second model accounts for the participation of BCB officials in various public and private events. The third model expands the analysis to include all forms of monetary announcements, encompassing MPC statements, MPC minutes, quarterly inflation reports, inflation target settings, open letters from the BCB Governor, and the public agenda of BCB Governors. The fourth model refines this approach by restricting the analysis to the most significant shocks, identified by aligning monetary announcements with periods of extreme changes in the DI-PRE swap rates for 1-year and 2-year maturities.[9]

To implement these models, we utilize daily data on six endogenous variables: DI-PRE swap yields with maturities of 1 month, 1 year, 2 years, and 5 years, the Ibovespa stock price index, and the USDBRL foreign exchange rate. Additionally, log-transformations are applied to the Ibovespa index and the USDBRL rate to account for potential nonlinearities and ensure consistent scaling. Each financial asset is observed twice daily, allowing us to calculate overnight returns (close-to-open price changes) and intraday returns (open-to-close price changes) associated with the analyzed monetary events, that is, the MPC statements, MPC minutes, inflation reports, MPC agenda, inflation target settings, and open letters.

[9] Given the changes in the swap interest rates for the 1-year and 2-year maturities individually, the thresholds for extreme returns were defined as the 5th and 95th percentiles. Subsequently, periods containing extreme returns in either of these assets were selected for analysis.

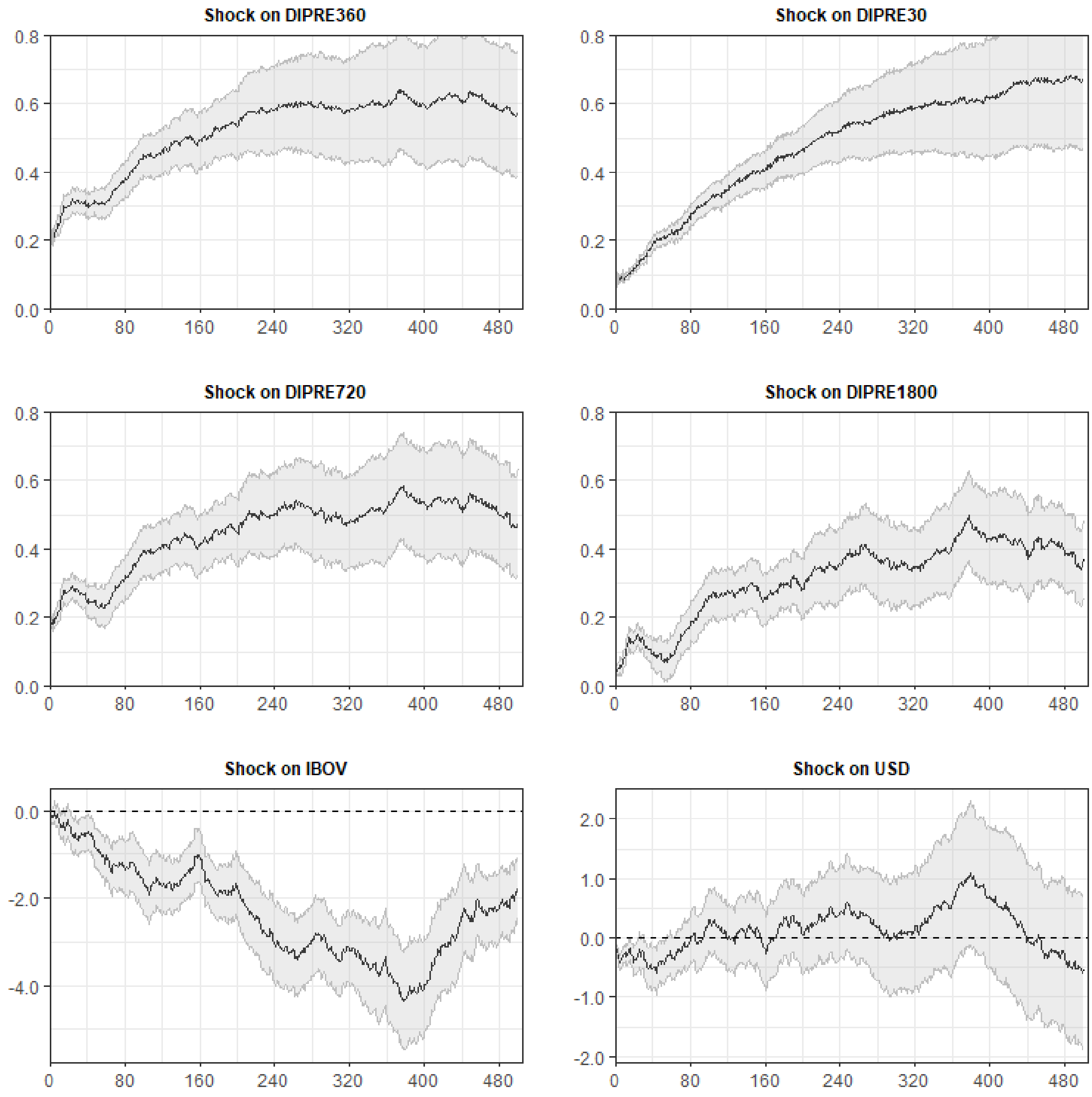
About the SVAR modeling, it is important to clarify three key points. First, the null hypothesis that periods with monetary announcements are no different than otherwise was rejected in all SVAR model estimations (with p-value 0.000). Second, the lag length of the SVAR models was determined using the Bayesian Information Criterion (BIC), which prioritizes model simplicity and parsimony. Accordingly, the four models were estimated with lag length 2, 3, 3, and 2, respectively. These lag lengths are also used in the LP models. Third, to ensure comparability across SVAR models, the identified monetary policy shock is normalized to induce an immediate increase of 25 basis points in the 1-year DI-PRE yield. The Generalized Least Squares (GLS) projection along the identified column R_1 (even though the other columns of R are unknown) delivers the true structural monetary policy shock η_1 (as seen in Equation 1) that is used as the shock regressor in the LP estimations.

By integrating high-frequency data with the SVAR and LP modeling, this framework also enables the generation of IRFs that quantify both the magnitude and persistence of monetary policy shocks across various asset classes. The resulting IRFs are computed with a 68% bootstrap confidence interval, spanning 0 to 500 periods (approximately one year). Figures 2–5 present the estimated LP impulse responses for, respectively, MPC decisions, the MPC agenda, the full set of monetary policy events, and the filtered monetary announcements. The vertical axis is expressed in percentage points (p.p.) for DI-PRE swap rates and in percent changes for both the Ibovespa index and the spot exchange rate. In the Appendix, Figures A5 to A8 provide the corresponding estimated SVAR IRFs, whose behaviors are fairly similar to the LP IRFs.

The estimated IRFs for MPC decisions, as depicted in Figure 2, yield important insights into the economic impact of these monetary policy announcements on financial markets. The results indicate a positive, significant, and persistent response of interest rates to a positive monetary policy shock (scaled to an immediate 25-bp increase in the 1-year DI-PRE yield) with shorter-term rates responding more strongly than longer-term maturities over the horizon. Intuitively, the amplified responses of short-term rates suggest that these announcements are perceived as direct signals of immediate monetary adjustments, while the reactions of longer-term rates reflect adjustments in expectations for future policy paths and economic conditions.

The responses of the Ibovespa stock index and the USDBRL exchange rate are also noteworthy. Equity prices fall slightly on impact and continue to decline for several months before a late, incomplete reversion, a pattern consistent with higher discount rates and weaker expected activity, signaling a statistically meaningful contractionary effect. The exchange rate response points to a short-lived statistically significant appreciation of the Brazilian real (a decline in USD/BRL) following the tightening, which signals the sensitivity of the foreign exchange market to shifts in interest rate differentials, as higher domestic interest rates attract foreign capital, increasing demand for the currency. Nevertheless, the effects dissipate and become non-significant after about a month.

Figure 2 - Estimated LP Impulse-Responses: MPC Decisions



Notes: Linear LP impulse responses using shocks identified from structural VAR models with heteroskedasticity-based identification. The monetary policy shock corresponds to an immediate increase of 25 basis points in the 1-year DI-PRE yield. The IRFs are computed with a 68% bootstrap confidence interval (grey area), spanning 0 to 500 periods (approximately one year).

Overall, the impulse responses depict conventional transmission of tighter MPC decisions, which short-term yields rise more than long-term yields, equity prices decline, and the currency appreciates temporarily, with magnitudes and persistence consistent with the economic theory.

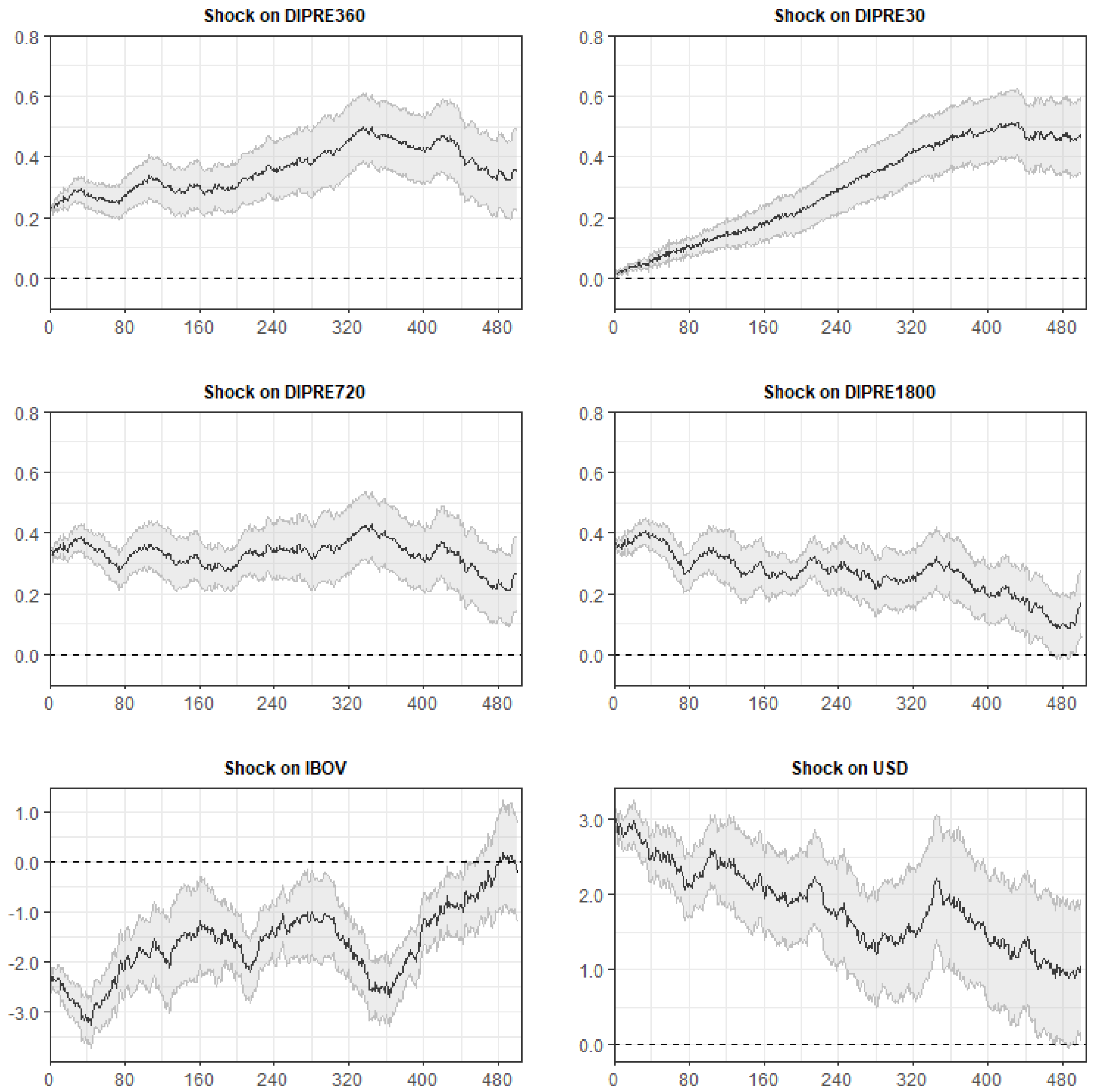
As presented in Figure 3, the IRFs for the MPC agenda demonstrate meaningful impacts on financial assets, highlighting the importance of verbal communication by central bank officials in shaping market dynamics. Speeches can provide an opportunity for central bank officials to address market concerns, clarify policy stances, or provide forward guidance, which can significantly shape market sentiment. For instance, a speech emphasizing future tightening or concerns over inflationary pressures would likely prompt a rise in longer-term yields as market participants adjust their expectations for future interest rates.

The DI-PRE swap rates across all maturities respond positively and persistently to monetary shocks, with longer maturities (720-day and 1800-day) exhibiting the most pronounced immediate reactions. This result reflects the market's interpretation of speeches as providing critical guidance on the central bank's medium- to long-term policy trajectory, influencing expectations beyond immediate rate changes. However, the transmission pattern is generally weaker on the yield curve than the one obtained for MPC decisions.

In parallel, the response of the Ibovespa stock index to MPC presidential agenda appearances is sharp and negative, with a relevant decline immediately after the monetary shock which remains economically large and statistically meaningful for extended horizons before a late reversal. By contrast, the Brazilian real exhibits a pronounced depreciation on impact that persists throughout the sample window, potentially reflecting underlying structural vulnerabilities in the domestic economy and heightened risk perceptions.

Consequently, the IRFs associated with the MPC agenda underscore its role as a central-bank communication instrument capable of influencing short- and medium-term interest rates, equity valuations, and the exchange rate. Relative to formal MPC decision events, however, agenda-driven communications generate more limited repricing at long maturities, consistent with limited effects on long-run expectations. These findings reinforce the importance of transparent and consistent verbal guidance for shaping market expectations and supporting the transmission of monetary policy.

Figure 3 – Estimated LP Impulse-Responses: MPC Agenda



Notes: See Figure 2.

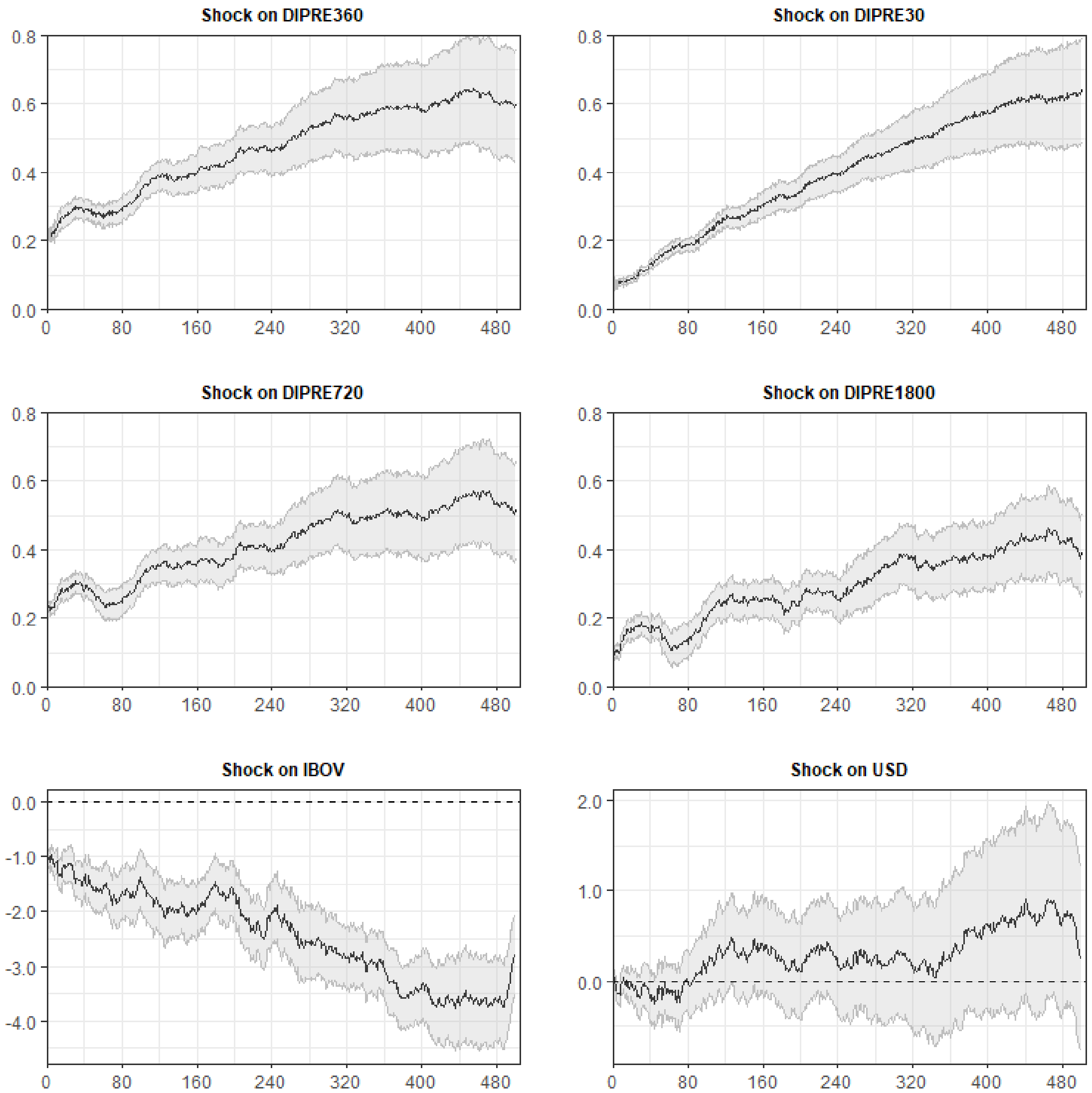
Figure 4 presents the IRFs for all monetary events (MPC statements, minutes, and inflation reports, inflation-target settings, open letters, and MPC agenda), which provide important insights into the impacts of overall monetary policy on financial markets. As seen in Table 2, the MPC presidential agenda account for most of the monetary policy events considered in our analysis.

The DI-PRE swap rates rise gradually and near-monotonically, remaining statistically above zero across horizons. The peak response is smaller at the long end (DI-PRE1800), suggesting that announcements meaningfully reprice the near-term policy path while only partially shifting long-run expectations.

Equity prices fall markedly and persistently – the Ibovespa declines on impact and continues to drift downward over the year – highlighting a contractionary risk-pricing channel consistent with higher discount rates and weaker growth prospects. In contrast, the exchange-rate response is limited and non-statistically significant.

Overall, the effect of monetary announcement shocks produces conventional tightening signatures with broad upward pressure on yields and strong negative equity effects, whose persistence is comparable to the responses to formal MPC decisions.

Figure 4 - Estimated LP Impulse-Responses: All Monetary Announcements



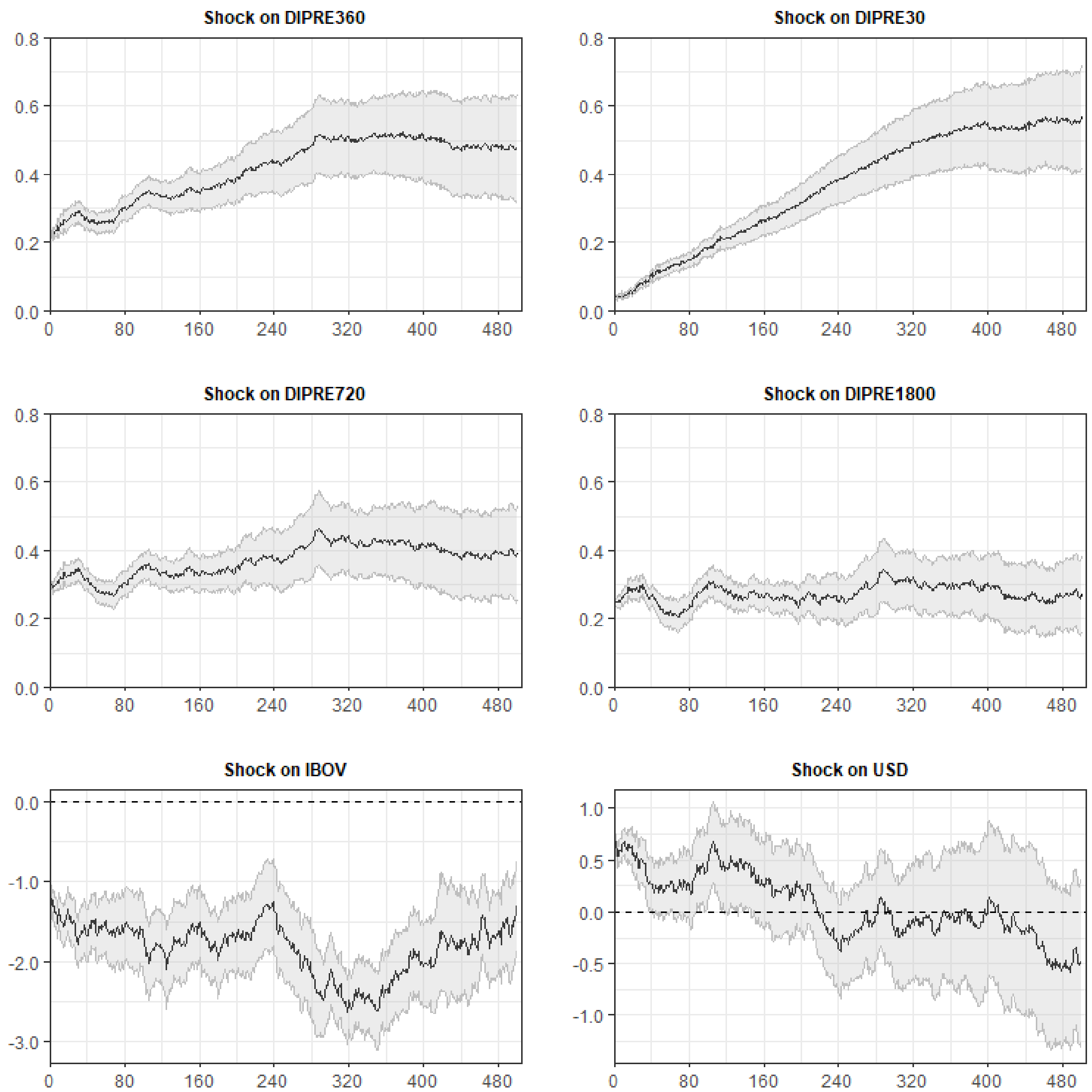
Notes: See Figure 2.

Finally, Figure 5 illustrates the IRFs for filtered monetary announcements based on the extreme returns in the DI-PRE swap rates for 1-year and 2-year maturities. This filtering process focuses on monetary events associated with pronounced market responses, thereby enhancing the robustness of the identification strategy.

The DI-PRE swap rates across all maturities exhibit significant responses, with a notable upward shift in the 30-day and 360-day maturities, which rises almost monotonically. The longer-term maturities (720-day and 1800-day DI-PRE rates) show a more measured and persistent adjustment with flatter dynamics.

The Ibovespa stock index displays an immediate and significant decline in response to the monetary shock, consistent with the tightening of monetary conditions that typically accompanies such events, while the USDBRL exchange rate exhibits an initial depreciation, but the movement fades away over time.

Figure 5 - Estimated LP Impulse-Responses: Filtered Monetary Announcements



Notes: See Figure 2.

Accordingly, the IRFs based on filtered monetary announcements also reflect the conventional transmission of tightening shocks. However, relative to the impulse responses that pool all monetary events (Figure 4), longer-maturity yields exhibit flatter dynamics suggesting a stable repricing of long-run expectations.

Taken together, the sets of LP impulse responses show a coherent and robust monetary-policy transmission of a tightening shock in Brazilian asset markets with a few differences by event type and impact. The distinct responses across asset classes reinforce the central role of monetary policy announcements in managing market expectations and achieving macroeconomic objectives. All the IRFs indicate a positive, statistically significant, and persistent response of interest rates to a positive monetary policy shock. In general, the shocks deliver a larger upward re-pricing of short- and medium-term DI-PRE rates, and a moderate pass-through to the long end. Nevertheless, the MPC president's public agenda is associated with a relatively higher immediate response on longer-term maturities. Moreover, all the IRFs show pronounced and significant negative reactions of equities with distinct degrees of late reversal. On the other hand, the evidence for the Brazilian real is mixed. While most impulse responses display only short-lived effects, the MPC engagements exhibit a strong depreciation on impact that persists over time.

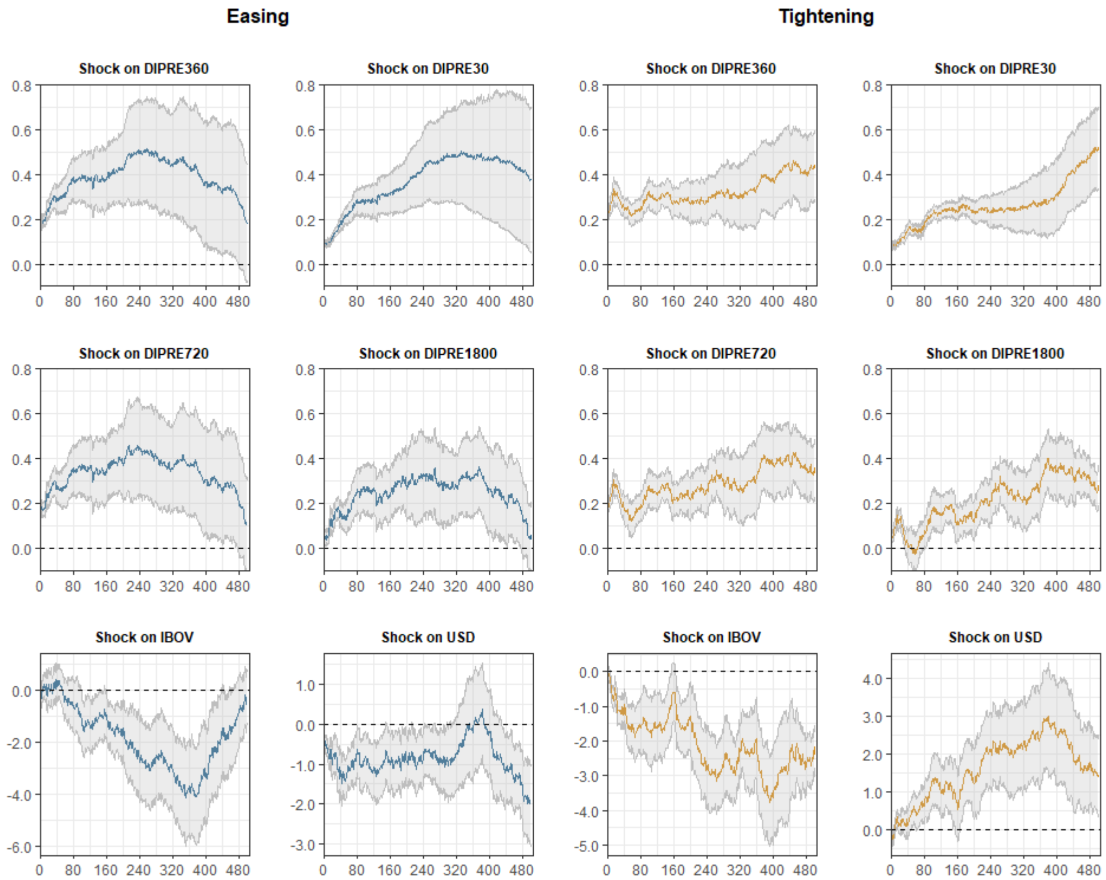
3.5.2. Non-Linear Estimated IRFs: Tightening and Easing

We next examine potential asymmetries in the responses of financial variables to the estimated monetary shock across periods of monetary tightening and easing, with policy stance indicated in Figure 1, using non-linear LP IRFs. The goal is to assess whether financial assets react differently when the BCB undertakes distinct phases monetary policy via the Selic rate.

Figures 6 to 9 illustrate the estimated non-linear IRFs for the MPC decisions, MPC agenda, all monetary policy events, and filtered monetary announcements, respectively, separately for tightening and easing periods. The vertical axis reports effects in percentage points (p.p.) for DI-PRE swap rates and in percent changes for both the Ibovespa index and the spot exchange rate. For comparability, signs are defined relative to monetary tightening; accordingly, the sign is inverted when interpreting responses during easing episodes.

Based on Figure 6, the identified 25 bp monetary shock originated from MPC decisions generates positive, persistent, and generally statistically significant responses of DI-PRE yields across most horizons during tightening episodes, with short-maturity rates reacting more strongly than longer maturities. On the other hand, during easing periods, the responses are more hump-shaped, building through intermediate horizons before a late reversal.

Figure 6 - Non-Linear LP IRFs (Tightening vs Easing): MPC Decisions



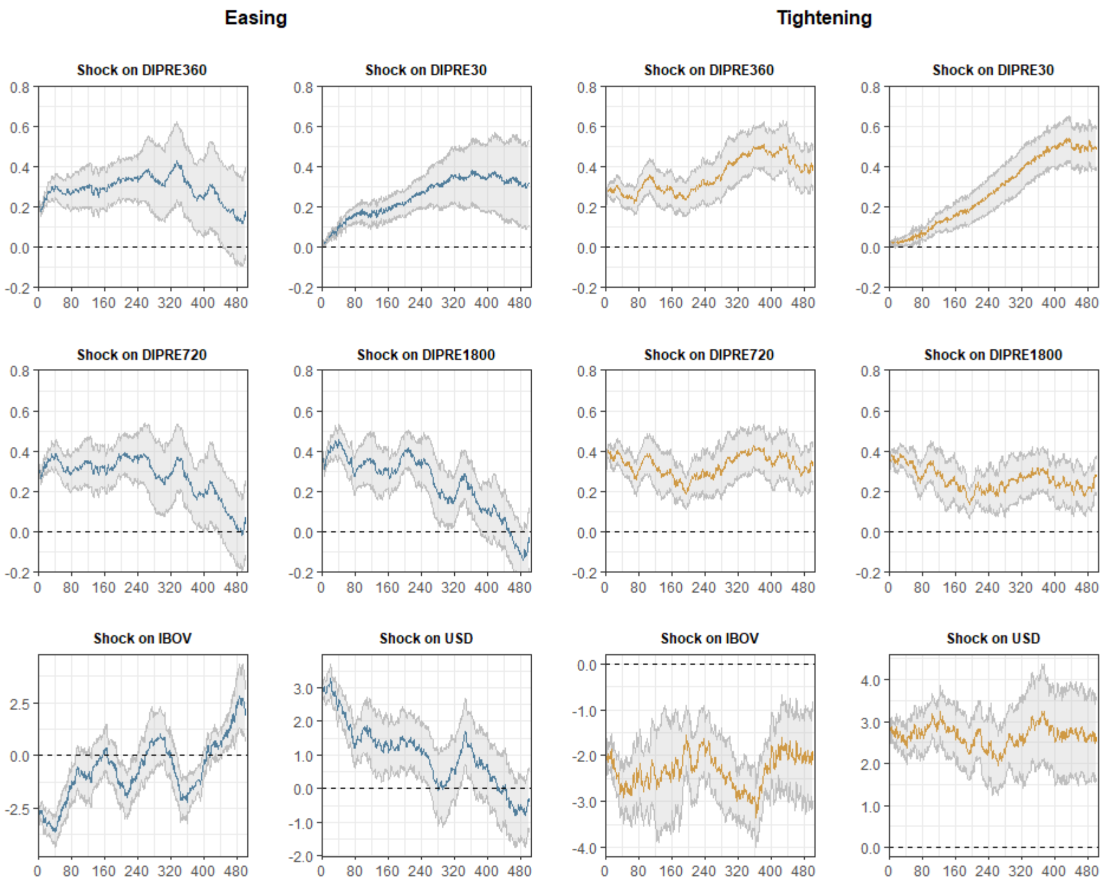
Notes: Non-linear LP impulse responses using shocks identified from structural VAR models with heteroskedasticity-based identification. The monetary policy shock corresponds to an immediate increase of 25 basis points in the 1-year DI-PRE yield. The IRFs are computed with a 68% bootstrap confidence interval (grey area), spanning 0 to 500 periods (approximately one year). The vertical axis reports effects in percentage points (p.p.) for DI-PRE swap rates and in percent changes for both the Ibovespa index and the spot exchange rate. Axis signs are defined relative to monetary tightening so that the signs are inverted when interpreting responses during easing episodes.

The Ibovespa declines on impact and continue to fall over subsequent quarters in tightening cycles, whereas equity prices initially rise before reversing at longer horizons under easing periods. Moreover, the BRL depreciates in both policy states.

Taken together, these patterns indicate state-dependent sensitivity of interest rates and equities to policy shocks, which appears more pronounced during tightening.

Figure 7 presents the non-linear IRFs for the MPC agenda. The identified monetary shock generates positive, persistent, and statistically significant DI-PRE yield responses during tightening episodes, with longer maturities reacting strongly already at short horizons. By contrast, during easing episodes, the yield profiles are hump-shaped, with effects building through intermediate horizons and subsequently dissipating, particularly at the long end.

Figure 7 - Non-Linear LP IRFs (Tightening vs Easing): MPC Agenda

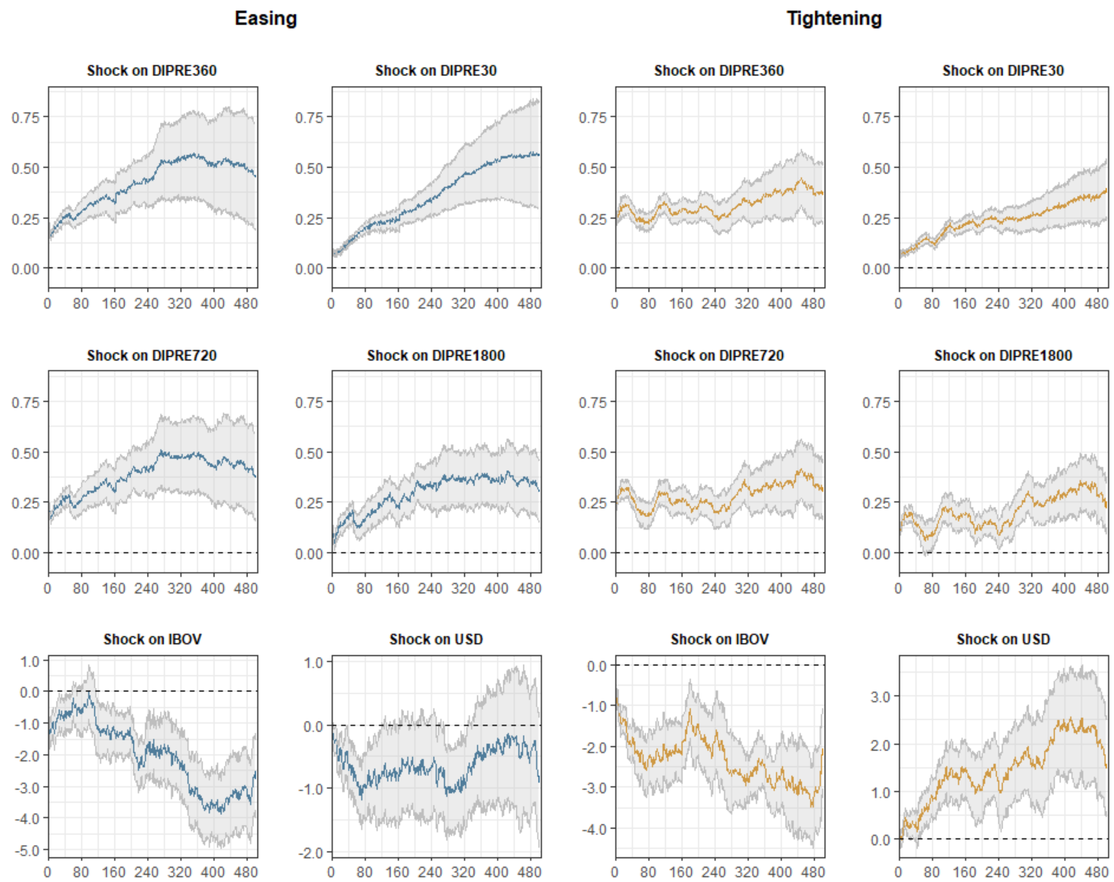


Notes: See Figure 6.

Risk assets also display state dependence. In tightening phases, the Ibovespa declines persistently and the USD appreciates (implying BRL depreciation). In easing phases, the stock index initially rises and the BRL appreciates, after which both series alternate between gains and losses, yielding less one-sided responses.

In sum, the MPC agenda shocks exhibit state dependence: the term-structure pass-through is stronger and more durable under tightening, while in easing the yield responses peak and then unwind, especially for longer maturities. Equity and exchange-rate reactions are also more pronounced and distinguishable under tightening. These patterns imply that agenda-related communications and appearances by MPC members impact financial assets differently across policy cycles.

Figure 8 - Non-Linear LP IRFs (Tightening vs Easing): All Monetary Announcements



Notes: See Figure 6.

Aggregating all monetary announcements – MPC statements, minutes, and inflation reports, inflation-target settings, open letters, and MPC agenda – Figure 8 illustrates the respective non-linear IRFs. For interest rates, the identified shock induces positive, persistent, and statistically significant responses across DI-PRE maturities under both policy stances, with a stronger pass-through during easing. In contrast, the increases in yields are more gradual and of smaller magnitude, indicating weaker transmission during tightening.

Equity responses switch sign across stances: the Ibovespa declines under tightening, whereas it rises under easing, with broadly comparable amplitudes. The exchange rate moves in the same direction across regimes – the USD appreciates (BRL depreciates) – but the response is statistically stronger during tightening.

Overall, considering all monetary events, the IRFs exhibit state-dependent transmission that is more pronounced for yields in easing phases and for the exchange rate in tightening periods.

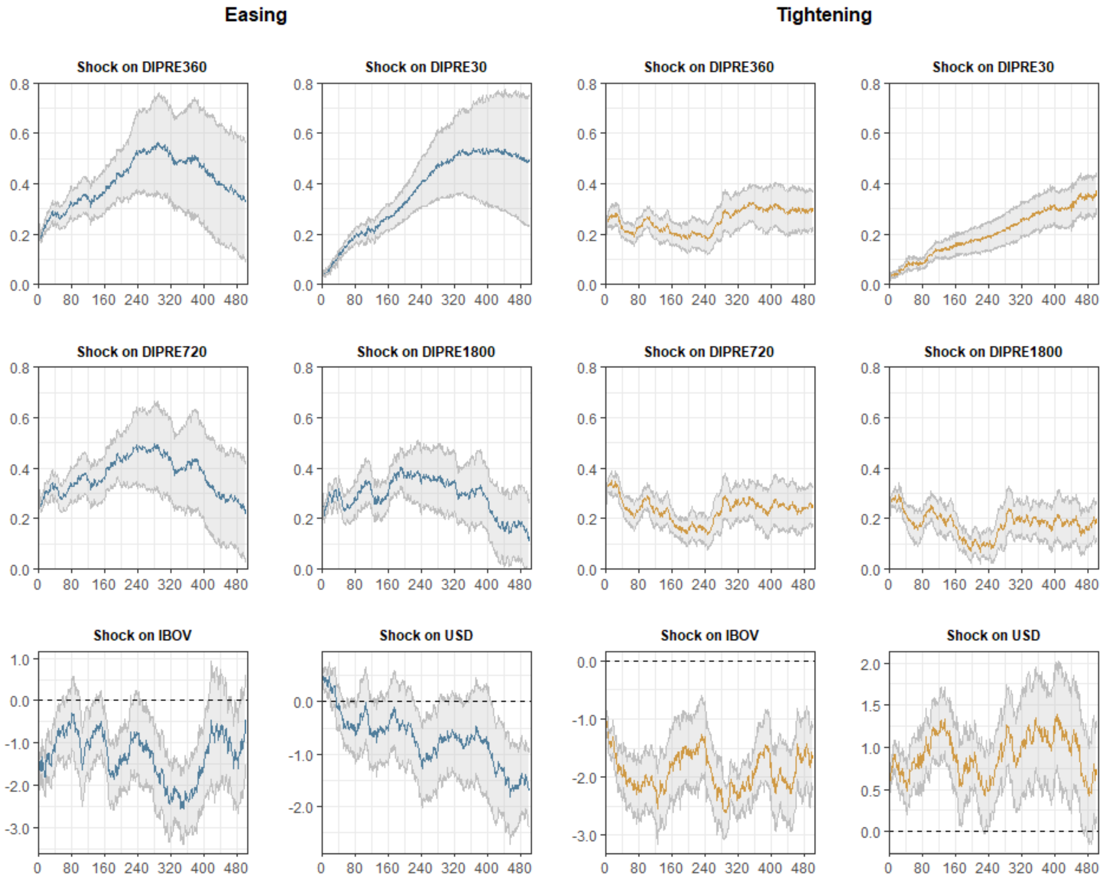
Conditioning on high-impact announcements, identified by extreme 1Y–2Y DI-PRE returns, sharpens the IRFs signals (Figure 9). During easing episodes, DI-PRE yields increase rapidly in a pronounced hump-shaped pattern, reach their peak at intermediate horizons, and then partially retrace, especially at longer maturities. Under tightening, the pass-through is flatter so that responses remain positive but accumulate more gradually. The confidence bands exclude zero, indicating precise estimates.

For equities, tightening shocks produce a persistent decline in the Ibovespa, whereas in easing the stock-price response is positive, albeit with some intervals that are not statistically significant. The exchange rate also behaves differently across regimes: in tightening, the USD appreciates (BRL depreciates) with sizable and persistent effects; in easing, the BRL initially appreciates but subsequently depreciates over the horizon.

Overall, filtering announcements by extreme DI-PRE moves yields another state-dependent transmission. The term-structure channel is hump-shaped with partial reversion in easing, while it is more linear in tightening. The FX impact reverses on impact across regimes (BRL down in tightening, up in easing), although BRL depreciation ultimately prevails over time. These results support the view that high-impact monetary communications transmit differently across policy cycles and that restricting attention to these events contributes to the identification.

In conclusion, the evidence indicates that policy stance decisively shapes the size, timing, and persistence of monetary-shock transmission, with patterns varying by the information set. For interest rates, shocks generate persistent and statistically significant DI-PRE yield responses with clear state dependence. Overall, shorter-term yields rise more than long-term yields over the horizon. Under tightening, yield increases are more durable and sustained. Under easing, responses are characteristically hump-shaped, building at intermediate horizons and subsequently reversing. This configuration implies that easing actions exert a stronger influence on the term structure for roughly the first six months, whereas tightening effects dominate over a one-year horizon.

Figure 9 - Non-Linear LP IRFs (Tightening vs Easing): Filtered Monetary Announcements



Notes: See Figure 6.

Other risk assets also exhibit regime dependence, with the sharpest effects under tightening. For equities, all IRFs display persistent declines following tightening shocks, whereas easing typically produces initial gains that later reverse or lose statistical precision, yielding a more decisive cumulative response of the Ibovespa under tightening – consistent with higher discount rates and weaker growth prospects. For the exchange-rate, the IRFs show BRL depreciation in tightening, while easing responses are mixed. The former likely reflects the BRL's longer-run depreciation tendency (as in Figure A4), suggestive of structural vulnerabilities in the domestic economy and heightened risk perceptions.

3.5.3. Linear Estimated IRFs: Macroeconomic Variables

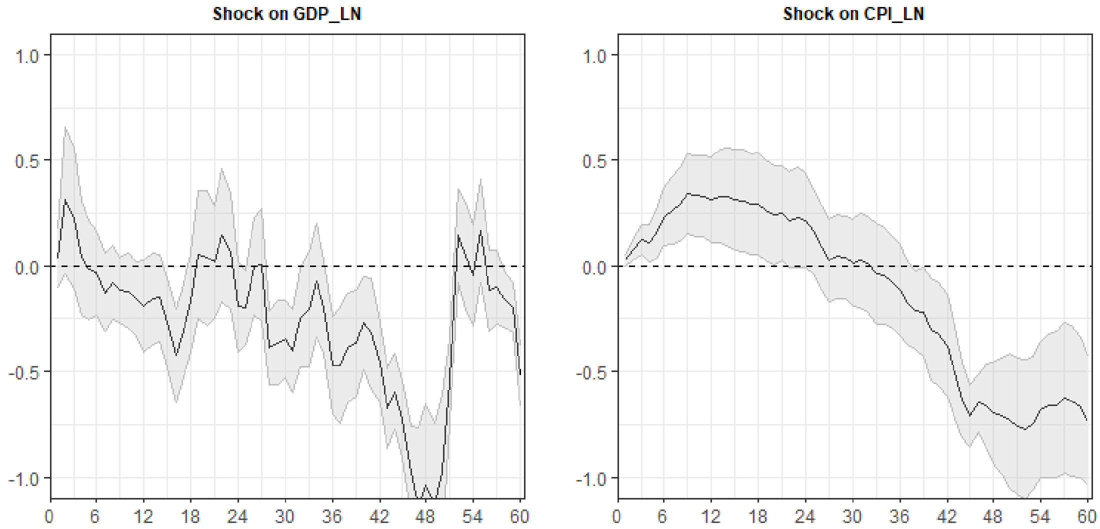
In this section, we examine the effects of the identified monetary shocks on selected macroeconomic variables using linear LP. Our objective is to assess whether the shocks that move financial markets also propagate to economic indicators.

First, we aggregate daily shocks – recovered from structural VAR models identified through heteroskedasticity and normalized to produce an immediate 25 basis points increase in the 1-year DI-PRE yield – into monthly series by summing within each month. We then standardize the series to have zero mean and unit variance. Subsequently, we estimate linear LP impulse responses to a tightening shock for two Brazilian monthly indexes – the Gross Domestic Product (GDP) proxy calculated by the Getulio Vargas Foundation (IBRE-FGV) and the Consumer Price Index (CPI) measured by the Broad National Consumer Price Index (IPCA) of the Brazilian Institute of Geography and Statistics (IBGE). The lag length of the LP models is set to 2 based on the Bayesian Information Criterion (BIC). In the Appendix, Figures A9 and A10 display the time-series behavior of these macroeconomic series.

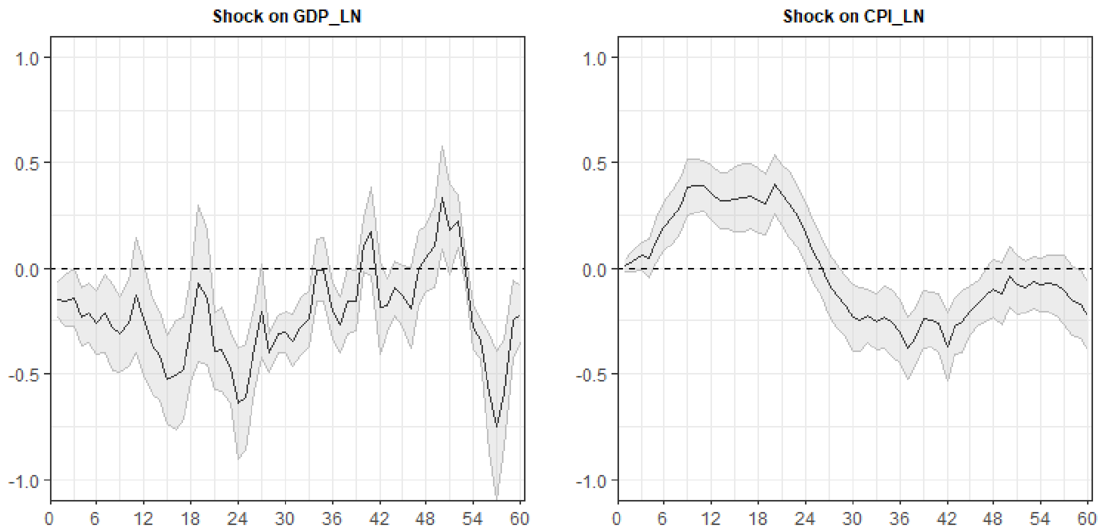
Figure 10 reports the cumulative LP impulse responses of the log-transformed GDP and IPCA indexes to a one standard-deviation change in the estimated tightening shocks derived from four different sources: (i) MPC statements; (ii) the public agenda of BCB Governors; (iii) the full set of MPC-related announcements (MPC statements, MPC minutes, quarterly inflation reports, inflation target settings, open letters from the BCB Governor, and the Governors' public agenda); and (iv) the subset of most significant MPC shocks. These monetary shock series are the same as those underlying the IRFs presented in Sections 3.5.1 and 3.5.2. The resulting IRFs are computed with a 68% bootstrap confidence interval, spanning 0 to 60 periods (approximately five years). The vertical axis is expressed in cumulative percent changes.

Across the CPI impulse responses, a one standard-deviation tightening shock produces an initial, transitory increase in the price level that turns decisively disinflationary at medium horizons (roughly after 24 months). An initial increase in the CPI is consistent with a policy reaction to persistent, higher-than-expected inflation, whereby the monetary authority raises interest rates to curb price pressures. Because the transmission of tighter policy is gradual, inflation does not fall immediately; consequently, prices continue to rise for a time before the disinflationary effects materialize. This pattern holds for all IRFs. The main exception is the shock originated from the MPC agenda, for which the medium-to-long-run disinflation is smaller and often not statistically distinguishable from zero near the end of the horizon.

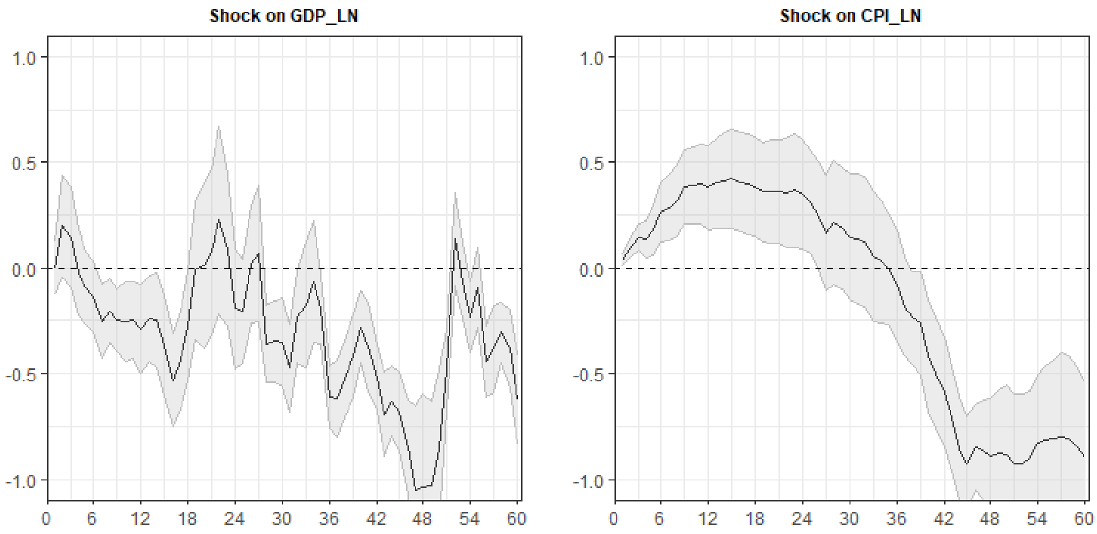
Figure 10 - Estimated LP Cumulative Impulse-Responses: GDP and CPI



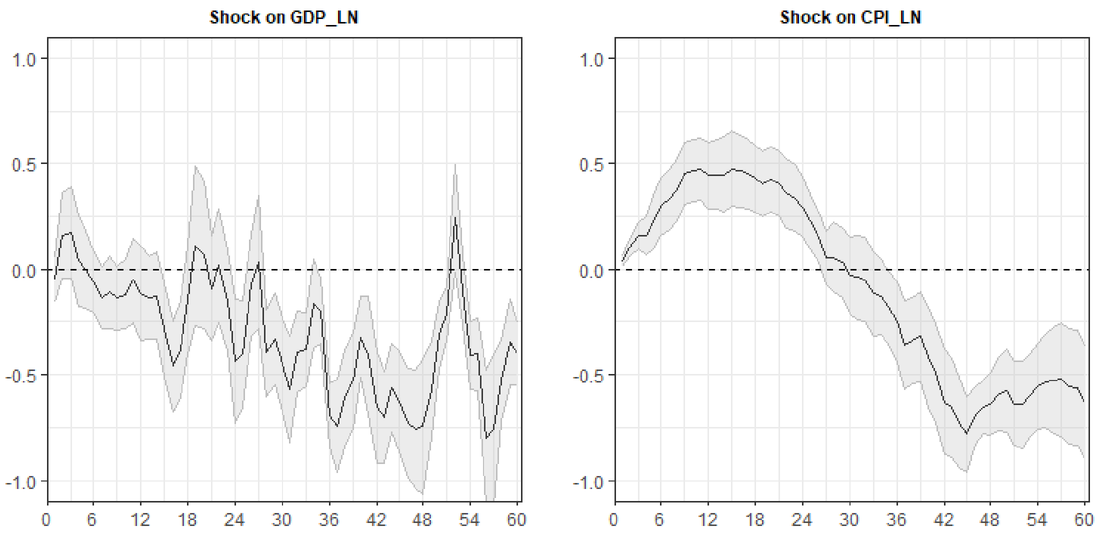
(i) MPC Decisions



(ii) MPC Agenda



(iii) All Monetary Announcements



(iv) Filtered Monetary Announcements

Notes: Linear LP cumulative impulse responses to a one standard-deviation change in the estimated tightening shock, which is identified from structural VAR models with heteroskedasticity-based identification. The IRFs are computed for 60 periods (approximately five years). The vertical axis reports effects in cumulative percent changes.

About the GDP, the cumulative IRFs are broadly contractionary but comparatively small and noisy. The responses generally exhibit a mild downward trajectory with intermittent declines and limited statistical precision, as reflected in frequent oscillations around zero. Consequently, the real effects are modest relative to the disinflationary response.

In the Appendix, Figure A11 presents the responses of GDP and CPI to alternative identified monetary shocks based on Gomes et al. (2024), estimated over a comparable window.[10] The results are broadly consistent with those obtained using our various shock sources, thereby reinforcing the validity of our estimates.

In sum, the results suggest our estimated tightening shocks reduce the price level over medium horizons while imposing, at most, moderate output costs. This pattern is consistent with effective transmission to inflation and suggests that the real costs associated with these shocks are limited.

[10] Gomes et al. (2024) define the monetary policy surprise by the variation in the nearest-maturity DI futures contract from the closing price before the announcement to the opening price on the next day.

3.6. Concluding Remarks

This paper evaluated the effects of high-frequency monetary policy shocks on financial asset prices in Brazil. We estimated SVAR models with identification through heteroskedasticity to obtain monetary policy shocks from high-frequency data under the assumption that periods with monetary events exhibit an elevated volatility compared to non-event periods. Then we computed impulse-response functions through LPs by multistep regressions for different types of events while also explicitly accounting for asymmetries in the responses during periods of monetary tightening and easing, and their effects on macroeconomic variables.

For this purpose, we constructed an extensive database containing information on monetary policy events in Brazil from 1996 until 2023. This dataset incorporates not only traditional monetary occasions such as the MPC statements and the associated MPC meeting minutes, but, crucially, the agendas of BCB officials. Hence, it represents a significant contribution to the literature as it provides a foundation for future research on Brazilian monetary policy.

The findings indicate that MPC statements and the public agenda of BCB authorities exert considerable influence on financial assets, with these engagements emerging as a particularly influential source of shocks during trading hours, compared to other sources of monetary communication. LP estimates reveal that monetary policy shocks have significant and persistent impacts across asset classes, especially yields of different maturities and stock markets. Moreover, the transmission of these shocks is strongly conditioned by the policy stance, with more sustained effects during tightening phases. Finally, we show that the identified tightening shocks also affect macroeconomic indicators, reducing inflation over medium horizons while imposing comparatively small output costs, irrespective of the shock source.

In conclusion, this study contributes to the understanding of how monetary policy shocks affect financial markets in Brazil. It highlights the heterogeneous effects of different types of monetary policy communication, offering insights may provide valuable guidance for policymakers seeking to optimize monetary policy, ensuring its effectiveness in achieving macroeconomic objectives.

3.7. Appendix

Figure A1 - Market Interest Rates (%): DI x PRE Swap with Shorter-Term Maturities

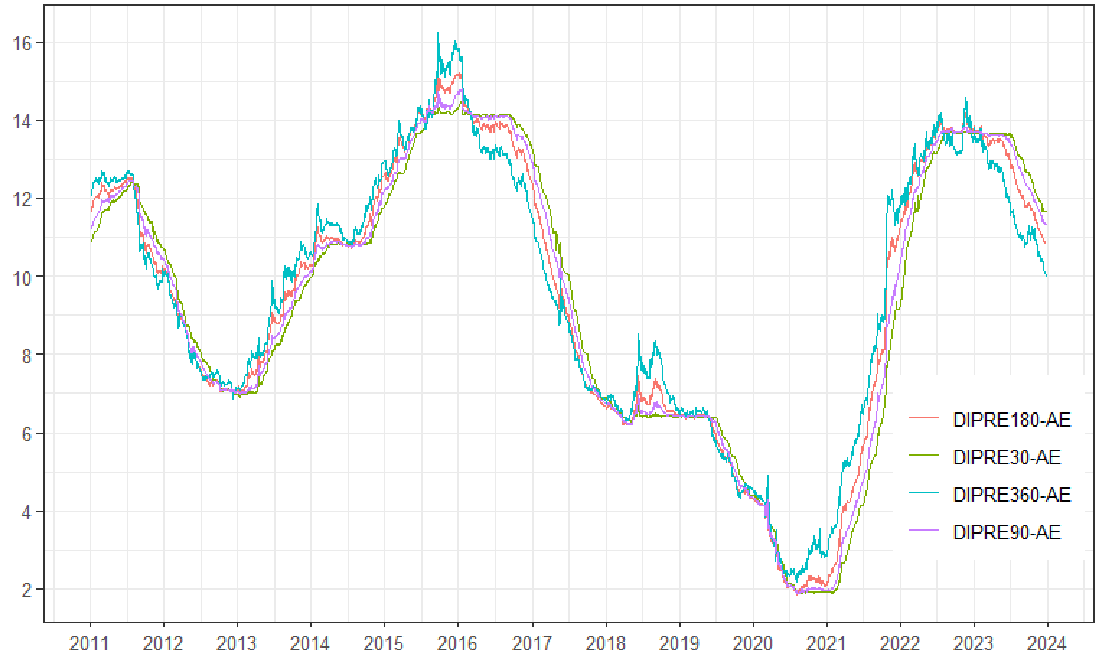


Figure A2 - Market Interest Rates (%): DI x PRE Swap with Longer-Term Maturities

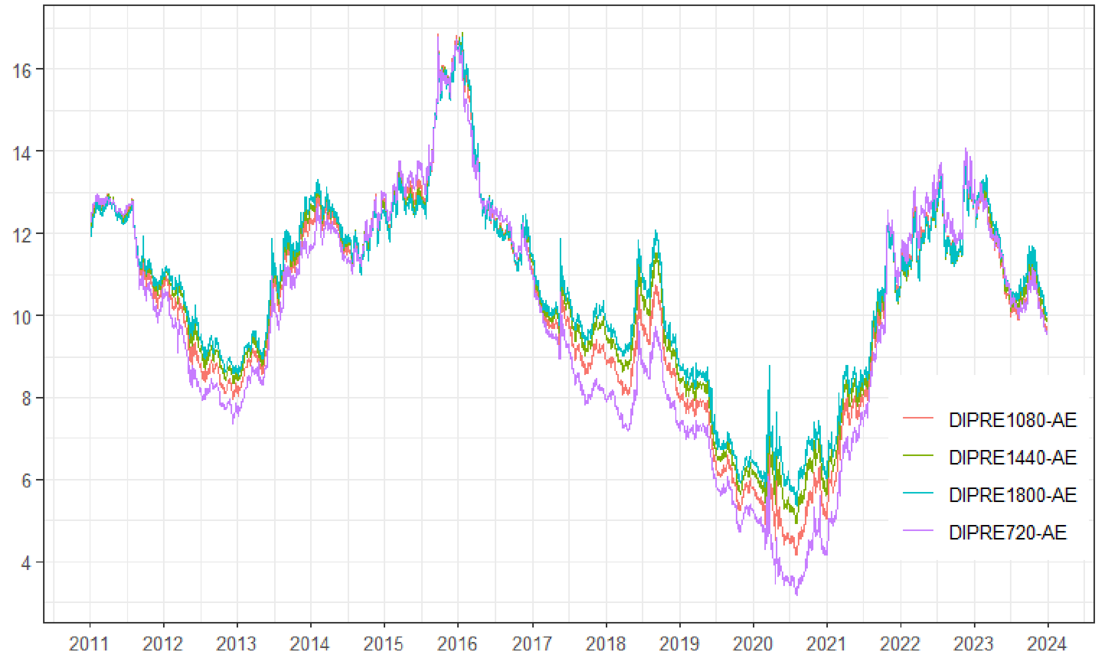
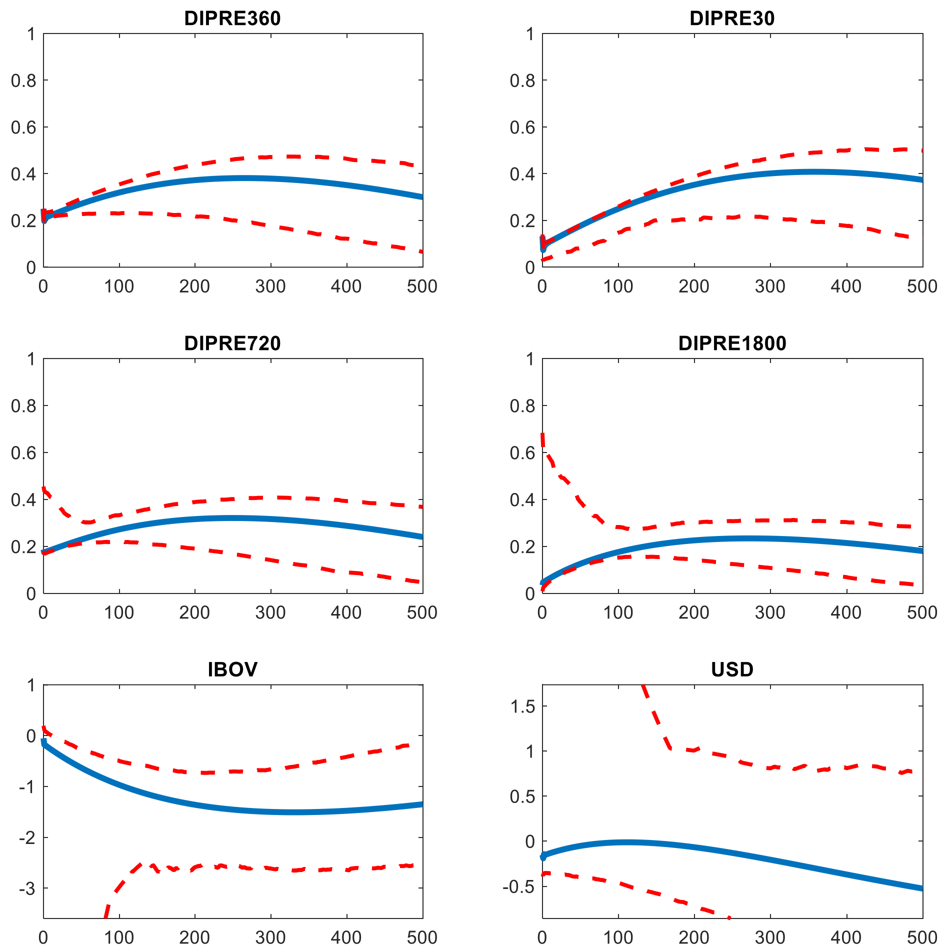


Figure A3 - Ibovespa Stock Market Index (Close Price)

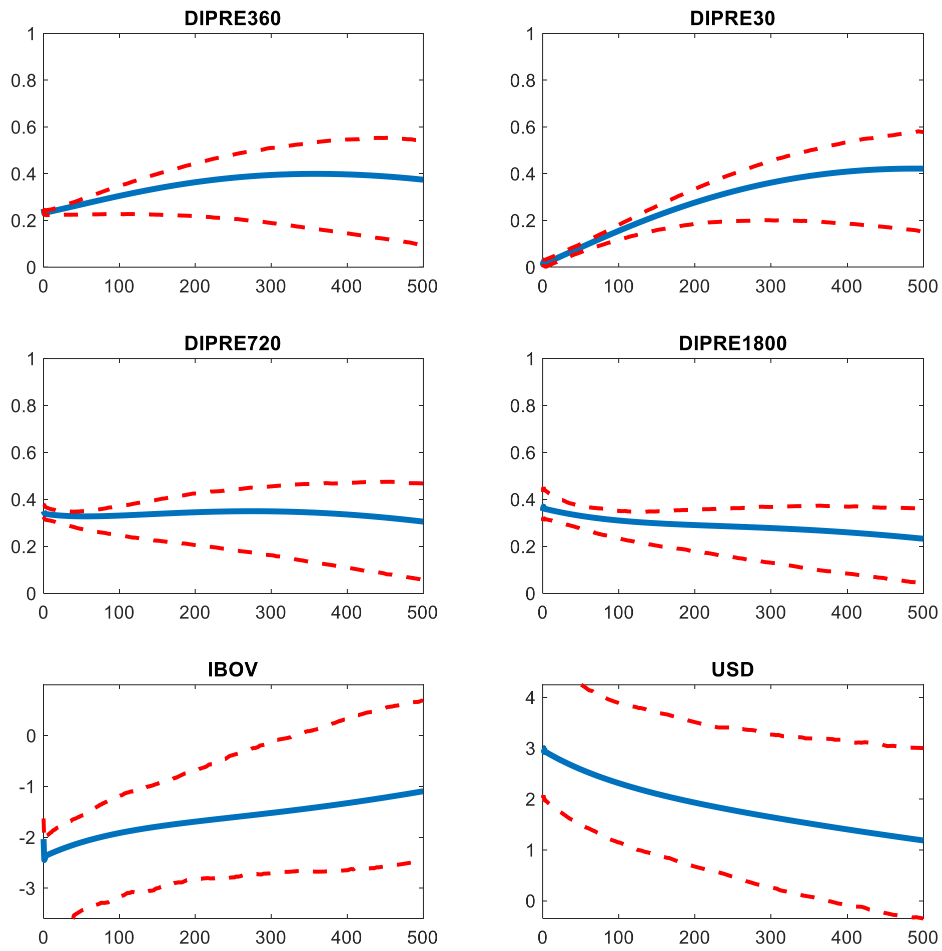
Figure A4 - U.S. Dollar (USD) to Brazilian Real (BRL) Exchange Rate

Figure A5 – Estimated SVAR Impulse-Responses: MPC Decisions



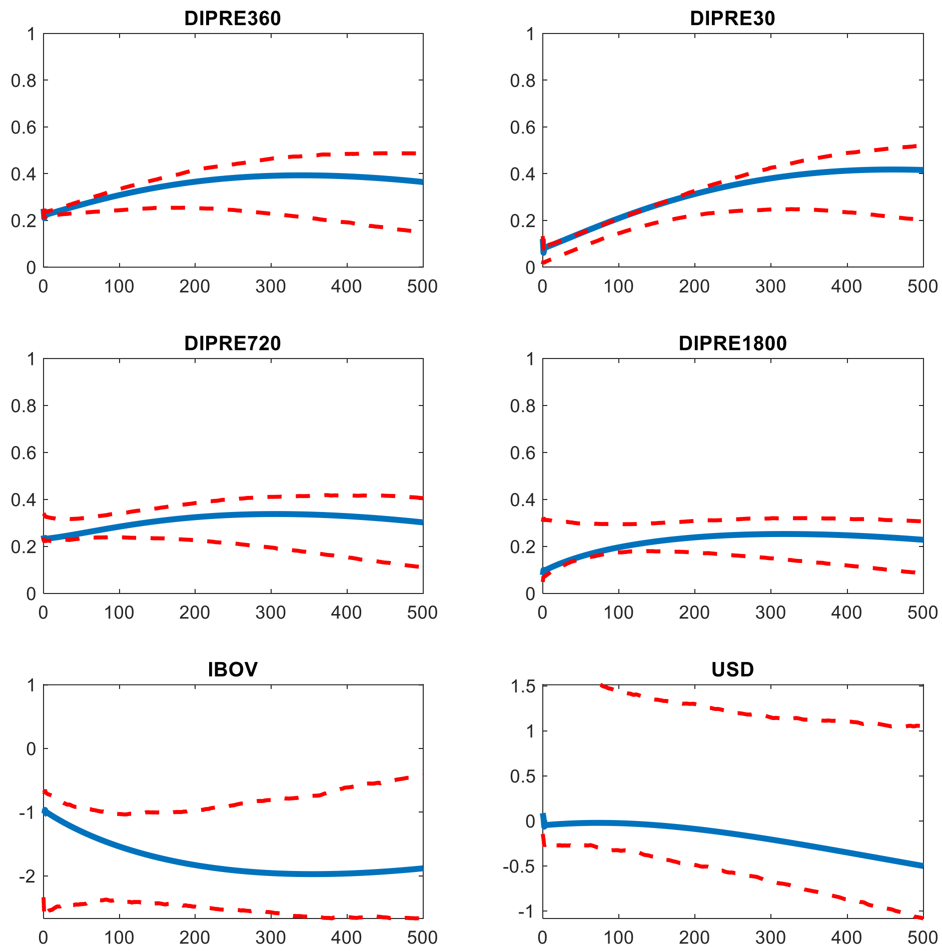
Notes: The monetary policy shock is normalized to induce an immediate increase of 25 basis points in the 1-year DI-PRE yield. The IRFs are computed with a 68% bootstrap confidence interval (in red dashed lines), spanning 0 to 500 periods (approximately one year).

Figure A6 - Estimated SVAR Impulse-Responses: MPC Agenda



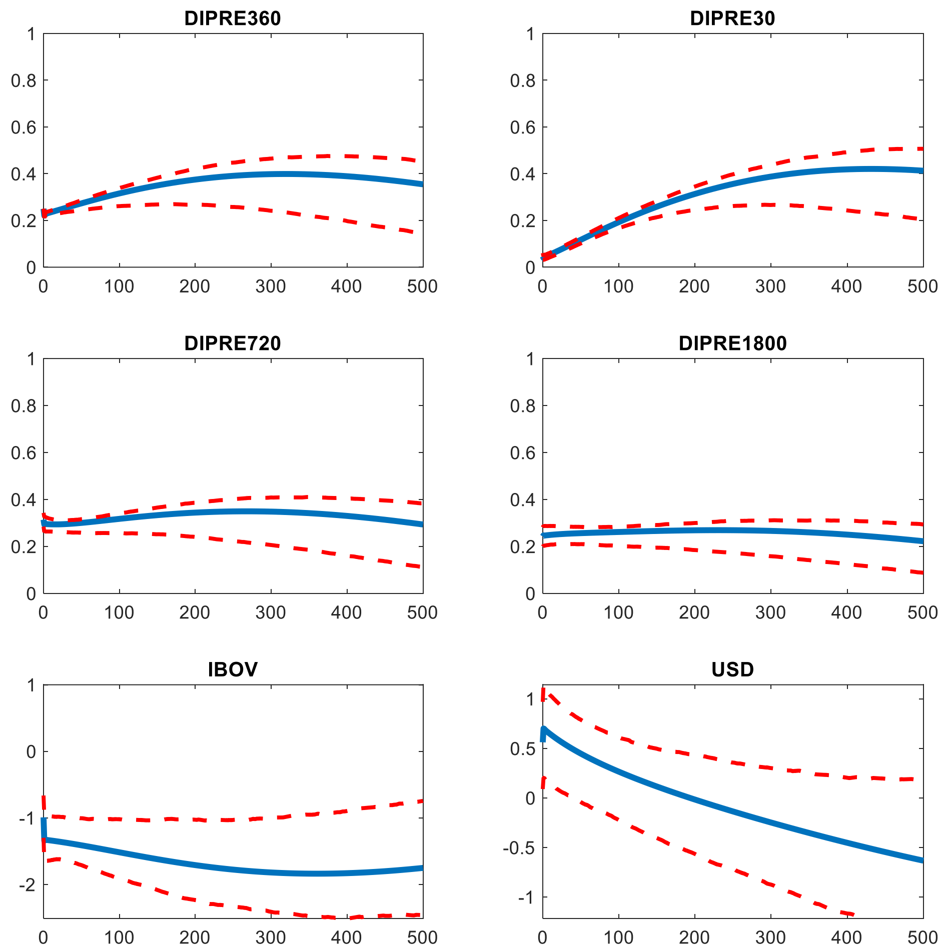
Notes: See Figure 2.

Figure A7 – Estimated SVAR Impulse-Responses: All Monetary Announcements



Notes: See Figure 2.

Figure A8 – Estimated SVAR Impulse-Responses: Filtered Monetary Announcements



Notes: See Figure 2.

Figure A9 - IBRE-FGV Monthly Gross Domestic Product (GDP) Seasonal Adjusted Volume Index

Figure A10 - IBGE Monthly Consumer Price Index (CPI) – IPCA

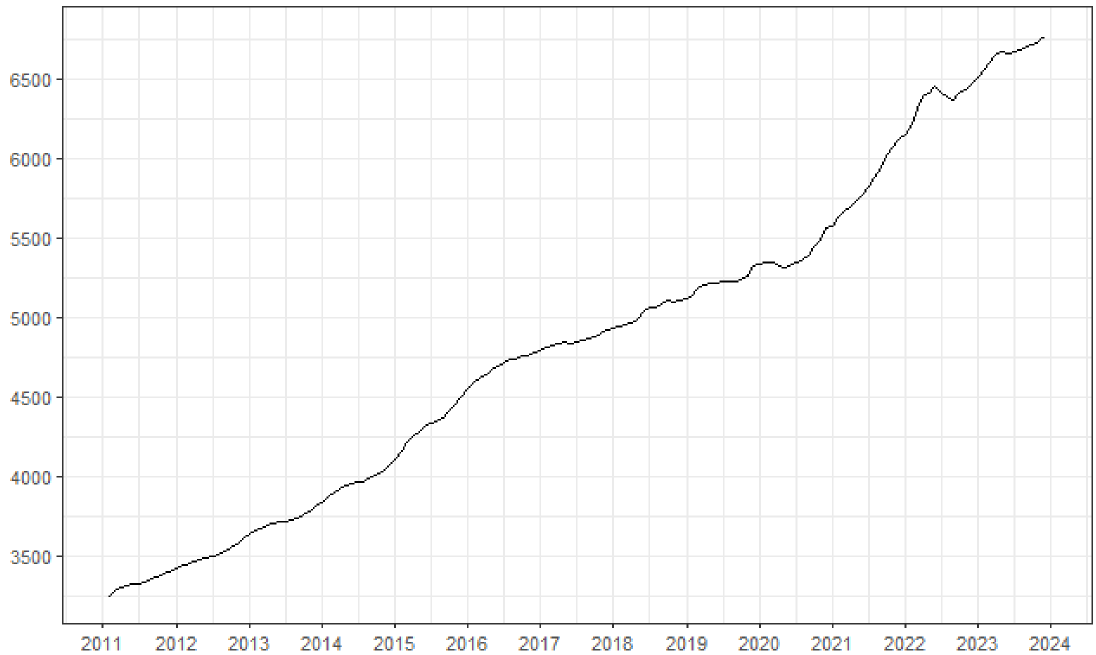
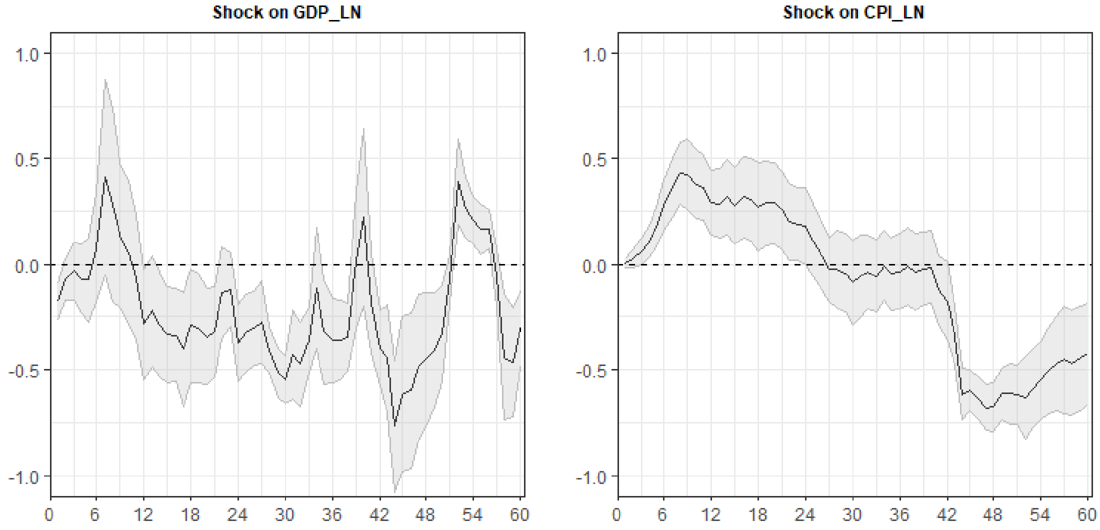
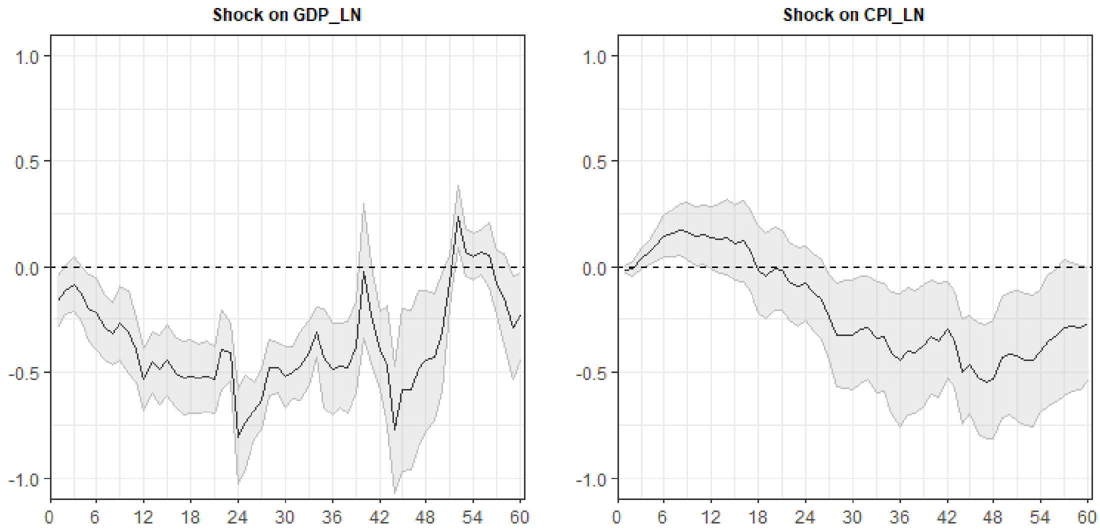


Figure A11 - Estimated LP Cumulative Impulse-Responses Using Other Monetary Policy Shocks based on Gomes et al. (2024)



(i) Shortest-maturity DI



(ii) 1-month DI Interpolation

Notes: Linear LP cumulative impulse responses to a one standard-deviation monetary shock, computed over 60 periods (approximately five years). The vertical axis reports effects in cumulative percent changes. Gomes et al. (2024) measure monetary policy surprises as the change in the shortest maturity DI future contract between the opening price on the trading day following the announcement and the pre-announcement closing price. The IRFs display the effect of two different monetary policy shocks, one based on the shortest-maturity DI contract available and the other on an interpolated 1-month DI contract, corresponding to panels (i) and (ii), respectively..

Table A1 - Median Absolute Change (log-return) per Announcement (in %): 2011-2023

	No Shock	Any Shock	Agenda of BCB Authorities	MPC Statements	MPC Minutes	Quarterly Inflation Report	Inflation target setting	Open letter by the BCB's Governor
#	4270	848	606	85	98	51	13	4
DI PRE - 1mo	0.14	0.14	0.12	0.37	0.14	0.19	0.25	**0.63**
DI PRE - 3mo	0.27	0.31	0.35	**0.38**	0.28	0.25	0.17	0.11
DI PRE - 6mo	0.26	0.41	0.43	**0.58**	0.32	0.29	0.27	0.32
DI PRE - 1y	0.32	0.75	0.80	**0.91**	0.47	0.51	0.54	0.40
DI PRE - 2y	0.40	0.93	**1.06**	0.88	0.52	0.48	0.48	0.52
DI PRE - 3y	0.43	0.92	**1.10**	0.58	0.40	0.38	0.85	0.56
DI PRE - 4y	0.42	0.86	**1.05**	0.63	0.48	0.40	0.63	0.72
DI PRE - 5y	0.44	0.82	**0.99**	0.56	0.43	0.33	0.69	0.86
Ibovespa	0.01	0.50	0.87	0.00	0.00	0.00	0.86	**1.19**
USDBRL	0.00	0.35	0.58	0.00	0.00	0.00	0.66	**0.81**

Notes: This table presents the median absolute log-returns of various financial assets (listed in rows) during periods associated with a specific monetary shock, any monetary shock, or no monetary shock, spanning the years 2011 to 2023. Log-returns are calculated by comparing the close-to-open price for overnight returns and the open-to-close price for intraday returns. Intraday returns apply to events such as speeches, inflation target settings, and open letters, while overnight returns correspond to statements, minutes, and inflation reports. For each asset (listed in rows), the largest change is highlighted in bold, while shades of blue within the cells visually depict the relative magnitude of changes across selected columns, ranging from no monetary shocks to the quarterly inflation report.

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