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The end of Brazilian big inflation: lessons to monetary policy from a standard New Keynesian model

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Abstract The paper analyzes economic stabilization in Brazil in the context of a New Keynesian model estimated with Bayesian techniques. Dataset covers the period 1975–2012. Our methodology is based on tests for multiple structural breaks at unknown dates and counterfactual exercises. The results show that inflation and output volatility present an inverted U-shape pattern, peaking at the 1985–1994 sample. Changes in the monetary policy stance and milder shocks accounted for the reduced inflationary volatility (about 50% each, in some specifications). However, some assumptions indicated that a sharp decline in the Phillips curve slope was also important for controlling inflation. Concerning to output, the sole explanation for its volatility fall seemed to be smaller shocks. Therefore, we conclude that a mix of the "good luck" and "good policy" hypotheses mainly originated the current period of increased stability in the country.

Keywords Monetary policy · New Keynesian model · Bayesian estimation · Brazil

JEL Classification C52 · E32 · E52

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1 Introduction

By the middle of the 2000s, macroeconomics seemed to have reached its maturity. A bulk of the empirical literature pointed toward a substantial decline in the volatility of economic activity in several industrialized countries, and there was a widespread view that improved monetary policy conduction could explain such phenomenon in a significant extent (see, e.g., Clarida et al. 2000; Lucas 2003; Boivin and Giannoni 2006; Giannone et al. 2008).

The beginning of the period of increased stability, i.e., the Great Moderation, is usually dated at the middle of the 1980s, with 1984 being a good point estimate (Kim and Nelson 1999; McConnell and Perez-Quiros 2000; Stock and Watson 2002). Additionally, the assumption relating inflation and output stabilization with central banks policies have come to be known as the good policy hypothesis (henceforth GPH).

However, following the Great Recession in late 2007, consensus toward GPH has weakened. Barnett and Chauvet (2011), for instance, posit that monetary policy had any connection with the Great Moderation in the U.S.; instead, they argue, FED's interventions tended to fuel business cycles fluctuations due to the problem of faulty monetary aggregates. Moreover, according to Taylor (2007, 2011), during 2003–2005 FED kept interest rates below the level implied by sound monetary principles for too long. Such policy overheated economic activity, stimulated risk taking in housing finance, and ultimately led to the bubble boom and bust.

Thus, economic studies have also tried to associate the Great Moderation with other elements, unrelated to monetary policy. In this sense, McConnell and Perez-Quiros (2000), Blanchard and Simon (2001) and Barnett and Chauvet (2008, 2011) found that it resulted from changes in the private sector, such as improvements in inventory management and information technology, higher productivity growth, and innovations in financial markets. In the present paper, we address to this interpretation as the private sector hypothesis, or PSH. By its turn, another strand of literature claims that the moderate business cycles were originated by milder exogenous shocks hitting the economies, that is, the good luck hypothesis, or GLH. In the latter, the volatility drop, especially output variability, is associated with a favorable macroeconomic environment. Stock and Watson (2002), Moreno (2004), Primiceri (2005) and Canova (2009) present results in this direction.

As can be seen, the Great Moderation, its sources, and the events that led to the Great Recession are still prolific research topics, especially when it comes to the case of developing countries, where econometric evidence in the subject is relatively scarce. Therefore, the present paper aims to investigate the Brazilian stabilization in light of a New Keynesian (NK) DSGE model, which encompasses, in a standard framework, parameters related to each one of the previously presented hypotheses. Specifically, it is intended to delimitate the effects of the 1994 monetary reform[1] on the reduction of inflation and real gross domestic product (GDP).

[1] In July 1994, Brazilian government launched a monetary plan in order to fight soaring inflation rates, the so-called Real Plan.

Brazilian recent past of high inflation imposes serious barriers to empirical testing in the monetary policy field, whenever longer periods are considered. Nevertheless, important results are found in the literature. Pastore (1997), for example, shows that monetary policy was essentially passive from 1975 to 1983, and real exchange rates were kept fairly constant over time, due to import substitution programs. Minella (2003) investigated monetary policy and basic macroeconomic relationships involving output, inflation, interest, and money. Based on vector autoregression models (VAR), the author compared three different periods, that is, 1975–85, when inflation was moderately increasing, 1985–94, with high inflation, and 1994–2000, characterized by low inflation. His main results showed that monetary policy shocks had significant effects on output and that monetary policy shocks did not lead to reduced inflation rates in the first two samples, but they appeared to have gained power to affect prices after 1994.

Regarding exclusively the Real Plan period, Barbosa-Filho (2008) found that inflation targeting managed to reduce inflation in Brazil after the 1999 and 2002 currency crises. By their turn, Mello and Moccero (2011) estimated a conventional NK model in a VAR context for Brazil, Chile, Colombia and Mexico. Their main findings are that inflation targeting has been associated with greater responsiveness of the monetary authority to changes in expected inflation in Brazil and Chile; the lower interest rate volatility after 1999 owes more to a favorable economic environment; and, the change in the monetary regime has not yet resulted in reduced output volatility in these countries.

Therefore, there is room for further research on the Brazilian stabilization, which is the focus of the present work. Its main goal is to provide answers to questions as the following: Has a phenomenon related to the Great Moderation happened in the country, that is, has the volatility of inflation and GDP dropped in a statistical sense? When? And, finally, what are the main determinants of this process?

We believe, however, that experience would lead most professionals to suggest that the answer would be the Real Plan period. It is natural to think that way, since output gap[2] standard deviation fell from 3.3% per quarter between 1975 and 1994, to 1.6% during 1995–2012. Moreover, changes in inflation were even more pronounced, since average quarterly inflation fell from about 39 to 2%, and its volatility diminished from 38 to 2.2% in the same period.[3] Thus, our work also addresses to additional research topics, namely: How important was the Real Plan and its monetary policy for these reductions?

Our paper analyzes the period 1975–2012, using quarterly data on inflation, output gap, and interest rate. The methodological approach in this paper is described as follows. First, data are sorted in subsamples according to a test of (possibly multiple) breaks at unknown dates, applied on a VAR system. We then characterize the most significant variations in the macroeconomic aggregates across the samples, identifying periods of higher or lower volatility. Second, we estimate the theoretical model in each subsample, using Bayesian techniques, and highlight any systematic changes in the

[2] Measured by the cyclical component of HP-filter decomposition ($\lambda = 1600$; Hodrick and Prescott 1997). GDP data are provided by Bonelli and Rodrigues (2012).

[3] General Price Index—Internal Availability (IGP–DI), published by Getúlio Vargas Foundation.

parameters associated with private sector, monetary policy, and exogenous shocks over time. Finally, data are divided in two periods: “pre-” and “post”-Real Plan. Then, we investigate which set of parameters explains the stabilization in the country, using a counterfactual method based on Canova (2009).

The relevance of the paper lies on the evidence that the welfare losses due to output fluctuations are significant in Brazil, reaching up to 10% of aggregate consumption (Cunha and Ferreira 2004). Besides, Fischer (1981) and Lucas (2000) found that high and volatile inflation is socially costly. A 10% annual rate, for example, can produce losses of around 3% of the real GDP, via misallocation of savings and investments, or losses in the value of real balances (Moreno 2004). In countries where inflation has a history of values far higher than that, such as Brazil,[4] inflation is even more problematic. Thus, as pointed out by Moreno (2004), the current period of low inflation constitutes a major macroeconomic achievement, so that understanding the forces behind it can greatly improve monetary policy effectiveness.

The remainder of this paper is organized as follows. Section 2 provides an overview of the macroeconomic setting throughout the analyzed period and describes the NK theoretical model. Section 3 presents the econometric methodology, the description of the data, and the structural breaks results. Sections 4 and 5 provide our main results, and Sect. 6 concludes the paper.

2 Macroeconomic setting review and theoretical model

This section presents an overview of the Brazilian macroeconomic setting between 1975 and 2012, giving special attention to the context in which monetary policy has been implemented. Such information will provide grounds for the econometric findings we present later in the paper. Next, we describe the theoretical model chosen to describe the economy in the period.

2.1 Macroeconomics and monetary policy in Brazil, with some selected empirical evidence

In order to understand the evolution of policy-making in the analyzed period, it is interesting to take a step back to 1964, year when the Brazilian Central Bank (BCB) was created.[5] BCB was born out of a far-reaching monetary and fiscal reform, aimed to control a raising rate of inflation and to stimulate the economy. From 1961 to 1964, for instance, prices increased[6] 70% p.a. on average, while real GDP growth was only 0.6% in 1963. In fact, in the same year industrial production varied −0.2%. In this context, the Brazilian economic team opted for a policy of gradual adjustment, which was translated in an increasing degree of indexation. By issuing indexed bonds, the

4 From 1988 to 1993, annual inflation reached triple-digit figures in every year, except in 1991.

5 BCB started to operate in 1965.

6 As measured by the general prices index.

central government managed to reduce monetary expansion due to fiscal deficits, while keeping its expansionary investment policies (Pastore et al. 2015).

Initially, the reform met its targets reasonably, bringing mean inflation rates to about 20% p.a. between 1967 and 1973. Moreover, the Brazilian economy grew at unusually high rates in the same period; GDP, for example, increased 10% p.a. on average; employment, 4%; fixed capital investments were steady at 21% of the GDP; and, industrial production increased 10, and 14% per year between 1967–1970 and 1971–1973, respectively. That was the "Brazilian Miracle."

By 1973–1974, however, the economy was flooded with external capital loans, and industrial capacity was close to a maximum. Inflation rate had returned to an accelerating path, and reached 35% in 1974, 29% in 1975, and 51% on average during 1976–1979 (Carneiro 1989). Not surprisingly, in the beginning of the 1980s the Brazilian economy was suffering the effects of high levels of external debt, unbalanced public finances, and an upward trending inflation. Meanwhile, inertial inflation concepts have gained force among economists and politicians (see, e.g., Lopes 1984; Resende 1985). In such class of models, in chronic inflation environments, economic agents are assumed to adopt a defensive, backward-looking stance on pricing. In such a situation, stopping high inflation with conventional restrictive policies would be socially costly and ineffective. Thus, in February 1986 it was launched the first attempt to stabilize the economy based on the inertial inflation concepts, the so-called Cruzado Plan.

With the new plan, prices, wages and exchange rates were indeterminately frozen at their February 1986 level. Additionally, indexation was regulated in long-term contracts and not allowed in short-term ones (maturity up to one year). There was not, however, a clear position for the monetary and fiscal policies. In fact, real interest rate was usually negative and below international levels in the period (see, Moraes 1990).

The Cruzado Plan was initially successful; real GDP growth rate was about 7.5%, and monthly inflation was stable at 1% from March to October 1986. However, a consumption bubble originated from appreciated salaries and negative interest rates had quickly pressed on prices (Marques 1988). Commodities disappeared from stores, and goods such as gasoline were under rationing (Cati et al. 1999). Extra premiums over the price of many products were commonly asked to consumers in black markets. In April 1987, monthly inflation rate had returned to 20%, staying at this level until June of the same year, when the economic team gave up on the plan.

In January 1989, Cruzados' accumulated inflation was about 7000%, the highest ever seen in the Brazilian history. After that, another four short-lived monetary plans were attempted,[7] which one of them utilizing a mix of actions that comprised the introduction of a new currency, restrictions on indexation, price freeze, and announcements of contractionary policies. Nonetheless, in practice, monetary policy was very accommodative. For instance, during the entire 1980 decade, inflation tax, as measured by Rocha and Saldanha (1995), was steady at 20% of the central government revenues. Additionally, some evidence indicates that BCB was collecting seigniorage in order to keep public deficits stable (Rocha 1997; Issler and Lima 2000).

[7] The so-called Bresser, Summer, Collor I, and Collor II plans.

Economic activity fell sharply in the period; industrial production and industrial employment reduced 40 and 5% in 1990, respectively. Real GDP growth rate was −4.5% in this year. Overall, annual growth rates in Brazil swung around −0.14% between 1988 and 1992. Nevertheless, inflation was above 1000% in 1992, with monthly rates reaching up to 30 or 40% between January 1993 and July 1994, when the Real Plan was launched.

The transition to Real Plan was announced with three phases. The first one tried to balance the budget, predicting cuts of about 20% in 1994 and 1995 fiscal years; the second one introduced a parallel unit of account, aiming to adjust relative prices; and the third one established the conversion between this unit of account and the new currency, the Real (Bacha 1997).

The idea was first to promote the full indexation of the economy to a single and trustable unit to, only then, replace the old currency by the new one. In this sense, in March 1994 the government introduced a device known as "Real-Value Unit," or URV, a daily-adjusted index, which had a maximum parity of one-to-one with the US Dollar, that would serve as benchmark for the economy, readjusting prices, wages, and contracts. URV was replaced in July 1994. In the period, monthly inflation rate fell dramatically from 40% on average (January to July 1994) to 2% (August to December 1994).[8] Later, in 1995, BCB formally adopted a system of crawling peg, keeping exchange rates oscillating within narrow bands with the aid of very high real interest rates.

The exchange rate policy was an important instrument during the earlier years of the Real Plan. However, its overvaluation generated persistent trade and current account deficits, which by 1998 became critical (Goldfajn and Minella 2005). Additionally, in 1997 the economy was hit by an external financial crisis that also affected the Asian Tigers, Mexico, and Russia. Thus, the macroeconomic panorama requested a change in the exchange rate policy, which occurred in January 1999, with the implementation of a system of dirty fluctuations.

In fact, alongside with the changes in the exchange policy, monetary authorities introduced a regime of inflation targeting (IT) for anchoring agents' expectations, which also rely on more sound results in the fiscal area (Carvalho and Minella 2012). Regarding the policy conduction, medium-run targets are defined by the National Monetary Council (CMN), and pursued by the Central Bank mainly via interest rate changes.[9]

Inflation and its volatility have been smaller since the adoption of inflation targeting, but how much of this phenomenon is due to the policy itself, or to favorable shocks is still an open research topic. In this regard, Barbosa-Filho (2008) argues that IT managed to reduce inflation in Brazil after the 1999 and 2002 currency crises and that the economic growth was slower and less volatile under this regime than during exchange targeting. By their turn, Mello and Moccero (2011) posits that the lower

[8] In addition, real GDP growth was around 3.8% p.a. on average between 1994 and 1997.

[9] CMN is somehow related to the Federal Reserve Board of Governors, while the BCB is its executive branch. Nowadays, three members, the Minister of Finance (chair), the Minister of Planning and Budget, and the President of the Central Bank constitute CMN.

volatility of interest rate and output in the post-1999 regime owes more to a benign economic environment than to the policy change.

Carvalho and Minella (2012) posit that credibility of IT has improved in Brazil recently. Their econometric evidence shows that BCB targets tended to guide inflation expectations in a larger extent after the serious confidence crises of 2002–2003. Additionally, private agents' interest rate forecasts indicate that they perceive BCB as following a "Taylor principle," i.e., reacting more than proportionally to deviations of inflation from its target. This result is in line with Sánchez-Fung (2011).

In this sense, the effect of the monetary policy on the economic performance of the country is still a fruitful line of research, whether considering the period before or after the Real Plan. The present paper provides additional information regarding such issues by estimating a standard NK model, presented in the next section, with Bayesian techniques.

2.2 A standard three-equation New Keynesian model for monetary policy evaluation

The NK model contains three equations, namely an aggregate demand, or IS-curve; an aggregate supply, or Phillips curve; and a Taylor-type policy rule. The first two equations are behavioral, derived from optimizing and forward-looking agents, thus addressing the Lucas critique. Versions of this model have been largely employed; see, for example, Silveira (2008), Canova (2009), Mello and Moccero (2011) and Creel et al. (2013), among others.

The IS-curve is described in Eq. (1). It presents endogenous output persistence, which can be derived from models with some form of adjustment costs (Clarida et al. 1999), or habit formation (Fuhrer 2000). However, the main reason to include endogenous persistence is empirical, as demonstrated by Fuhrer (1997).

$$\text{IS-curve:} y_t = -\delta(i_t - E_t\{\pi_{t+1}\}) + \theta_g y_{t-1} + (1-\theta_g)E_t\{y_{t+1}\} + v_{1,t}, \quad (1)$$

where y refers to the output gap; i, the interest rate; π, the inflation rate; t is the time index; it δ is the inverse of the consumption elasticity parameter; $0 \leq \theta_g \leq 1$ measures output persistence, indexing the effect of lagged versus expected future output gap on its current values; E_t is the expectation operator, dependent on the information set available at time t; and $v_{1,t}$ is an $i.i.d.$ disturbance related to measurement errors and/or exogenous shocks. An important feature of Eq. (1) is that it implies an inverse relationship between expected *ex ante* real interest rate and output gap.

The economy side supply is represented by Eq. (2), the Phillips curve, which associates inflation and output gap. It is possible to find this type of formulation by using Calvo (1983) approach with a slightly modified price-setting scheme (Christiano et al. 2005).

$$\text{Phillips curve:} \pi_t = \kappa y_t + \theta_i \pi_{t-1} + (1-\theta_i)\beta E_t\{\pi_{t+1}\} + v_{2,t}, \quad (2)$$

where it κ is the Phillips curve inclination, which represents real price rigidities; $0 < \beta < 1$ is a discount factor; $0 \leq \theta_i \leq 1$ index the effect of lagged versus future inflation on its current value, i.e., inflation inertia; and we added an $i.i.d.$ disturbance capturing measurement errors and/or exogenous shocks on the aggregate supply, $\upsilon_{2,t}$.

Now, in order to close the model, it is necessary to describe our monetary policy rule, which determines the nominal interest rate. In this sense, we follow Clarida et al. (1999), where monetary policy has two components, explicitly,

$$i_t = \rho i_{t-1} + (1-\rho) i_t^*, \tag{3}$$
$$i_t^* = \bar{i}^* + \gamma_i (E_t \pi_{t+1} - \bar{\pi}) + \gamma_g y_t. \tag{4}$$

where $\bar{i}^*$ is the desired interest rate, and $\bar{\pi}$ is the long-term desired inflation. Equation (3) models the tendency of central banks toward smoothing interest rates, measured by ρ, $0 \leq \rho \leq 1$. In its turn, Eq. (4) is a Taylor rule, in which the Central Bank reacts to deviations of expected inflation from its target, weighted by γ_i, and to deviations of output from its equilibrium level, weighted by γ_g.

By plugging (3) into (4), and adding a $i.i.d.$ monetary shock, $\upsilon_{3,t}$, we write the final equation of the model, the monetary policy (MP) rule:

$$\text{MP-rule} : i_t = \bar{c} + \rho i_{t-1} + (1-\rho)(\gamma_i E_t\{\pi_{t+1}\} + \gamma_g y_t) + \upsilon_{3,t}, \tag{5}$$

where $\bar{c} = (1-\rho)(\bar{i}^* - \gamma_i \bar{\pi})$.

The next section deals with the estimation method. However, before we proceed, it is noteworthy that we would rather use a model without exchange rates in this paper. This variable was used before as a policy instrument, stimulating investments with the import substitution programs (between 1975 and 1984), and controlling inflation, as the nominal anchor in the early years of the Real plan (1994–1998). Thus, its value had been controlled by BCB for several periods, which hindered its utilization in our regression methods. Besides, Mello and Moccero (2011), while estimating a model similar to ours, found that changes in the nominal exchange rate were not statistically significant at 5% in the Brazilian Central Bank's reaction function. In the same line, Sánchez-Fung (2011), estimating a Taylor rule for Brazil in a data-rich environment, showed that BCB behaves as following a policy rule in which the relevant variables are inflation deviations and output gap. Exchange rate was not selected in the final reaction function.[10]

3 Methodology: estimation, series, and samples

In this section, we first describe the Bayesian method of estimation; then, we present the time series employed here; and in the third part, we sort the data into subsamples of estimation, according to a test developed by Qu and Perron (2007).

[10] In addition, our model does not explicitly consider fiscal variables. While exchange rate changes and fiscal policy are relevant for inflation determination, we would rather keep our model parsimonious, leaving changes in these as exogenous disturbances that affect BCB behavior. By doing this, we are also able to focus on the possible effects of monetary policies on the Brazilian stabilization.

3.1 Estimation method

The theoretical model employed in this research has 12 parameters, including nine with structural interpretation (δ, θ_g, κ, θ_i, β, $\bar{c}$, ρ, γ_i, γ_g), and three auxiliary ones ($\sigma_1^2, \sigma_2^2, \sigma_3^2$). The first set of coefficients is related to the PSH (δ, κ, β), to persistence (θ_g, θ_i), and to the GPH ($\bar{c}$, ρ, γ_i, γ_g). The second set represents the variance of the disturbances in each equation, which are related to the GLH. Estimation is carried out by Bayesian methods, as described below.

Zellner (1971) demonstrates that Bayesian estimation connects calibration exercises, through the specification of priors, to methods of estimation, via maximum likelihood function, thus confronting theory and empirics. In this sense, we follow Fernández-Villaverde (2010) and An and Schorfheide (2007) defining $p(\boldsymbol{\alpha})$ as a density function describing a prior belief, where $\boldsymbol{\alpha}$ is a vector containing all the model's parameters; and $p(\cdot)$ stands for a general probability density function (pdf). The likelihood function is written in a state space format, so that it can be calculated by means of the Kalman filter and the prediction error decomposition (Canova 2009).

Our goal is to use $p(\boldsymbol{\alpha})$ and the likelihood function to estimate the parameters posterior density. When doing so, we employ the Random-Walk Metropolis-Hastings (RWMH) algorithm,[11] as suggested by Schorfheide (2000). Specifically, we set the scale parameter so that acceptance rate lies between 20 and 50%. Thus, the algorithm is allowed to visit the tails of the parameter distribution. Moreover, in all samples of estimation, we draw five MH-chains of 40,000 iterations each, keeping 14,000 draws in order to construct histograms. Convergence is addressed by the Brooks and Gelman (1998) method.

Parameters are assumed to have independent distributions. When selecting the shape of the priors distributions, we proceed as suggested by Canova (2009), setting: inverse gamma distributions for parameters that must be positive (δ, κ, σ_1, σ_2, σ_3); gamma distributions for parameters that are truncated below a particular value ($\bar{c} \geq 0$, $\gamma_i \geq 1$); beta distributions for parameters lying in an interval (θ_g, θ_i, β, ρ); and normal distributions for the remaining parameters (γ_g). By assuming $\gamma_i \geq 1$, we assure a single solution for the theoretical system, given reasonable values of the other parameters (see, for example, Clarida et al. 2000).

Means of the priors are set around typical calibrated values, while standard deviations are set loosely, in order to minimize the effect of subjective beliefs, thus allowing the algorithm to move away from the priors if data provides useful information. Table 1 shows the first two moments of each pdf.

3.2 Data and procedures

The time series used in this work are quarterly and, when necessary, seasonally adjusted with the X-12 additive methodology. The sample ranges from the first quarter of 1975 to the last one of 2012.

[11] Dynare 4.3.3 is the utilized software.

Table 1 Priors' shapes, means and standard deviations

Parameter	Shape	Prior mean	Prior standard deviation	Range
δ	Inverse gamma	2.00	0.25	Strictly positive
κ	Inverse gamma	2.00	1.00	Strictly positive
β	Beta	0.98	0.01	[0, 1]
θ_g	Beta	0.80	0.25	[0, 1]
θ_i	Beta	0.80	0.25	[0, 1]
$\bar{c}$	Gamma	2.00	0.25	$[0, \infty)$
ρ	Beta	0.80	0.25	[0, 1]
γ_i	Gamma	1.50	0.25	$[1, \infty)$
γ_g	Normal	0.50	0.25	$(-\infty, \infty)$
σ_1	Inverse gamma	0.01	0.50	Strictly positive
σ_2	Inverse gamma	0.01	0.50	Strictly positive
σ_3	Inverse gamma	0.01	0.50	Strictly positive

Quarterly inflation rate, π_t, is approximated by the average of the monthly rate within each quarter.[12] The latter is measured as variations of the General Price Index—Internal Availability (IGP-DI), published by Getúlio Vargas Foundation. Instead of focusing on consumer price indexes, this work deals with the general price, provided that it has a longer time series, including data between 1975 and 1980, essential for our comparative purposes.

In order to keep the database conformity, the quarterly nominal interest rate, i_t, is proxied by the average of the monthly Selic rates within each quarter. Selic is the Brazilian overnight rate, being published by BCB.[13]

When it comes to the output gap, y_t, we follow Orphanides (2004), who suggests a flexible way to compute it. Thus, we consider two measures, namely: (a) deviations from a HP-filter with smoothing parameter of 1600 (Hodrick and Prescott 1997) and (b) the real log-GDP growth rate. GDP data are provided by Bonelli and Rodrigues (2012).

Concerning whether inflation and interest rates are stationary, results presented in the literature so far are conflicting. Regular ADF tests tend to conclude that inflation and interest rates are $I(0)$, but this feature is probably due to the presence of structural breaks and/or abrupt governmental interventions. Cati et al. (1999), for example, built a test that accounts for such interventions, obtaining pieces of evidence that support the unit root hypothesis for inflation and interest rates in Brazil between 1974 and 1993. In the present paper, as an alternative of thoroughly investigating this issue, we assume both specifications for the variables entering into our systems, that is, one with

12 This proxy allows for a better fit in the Bayesian estimation. Besides, as long as target inflation goes to the intercept term in Eq. (5), we do not necessarily need to use annualized quarterly inflation rates in our estimations.

13 Inflation and Selic data are available at https://www.ipeadata.gov.br.

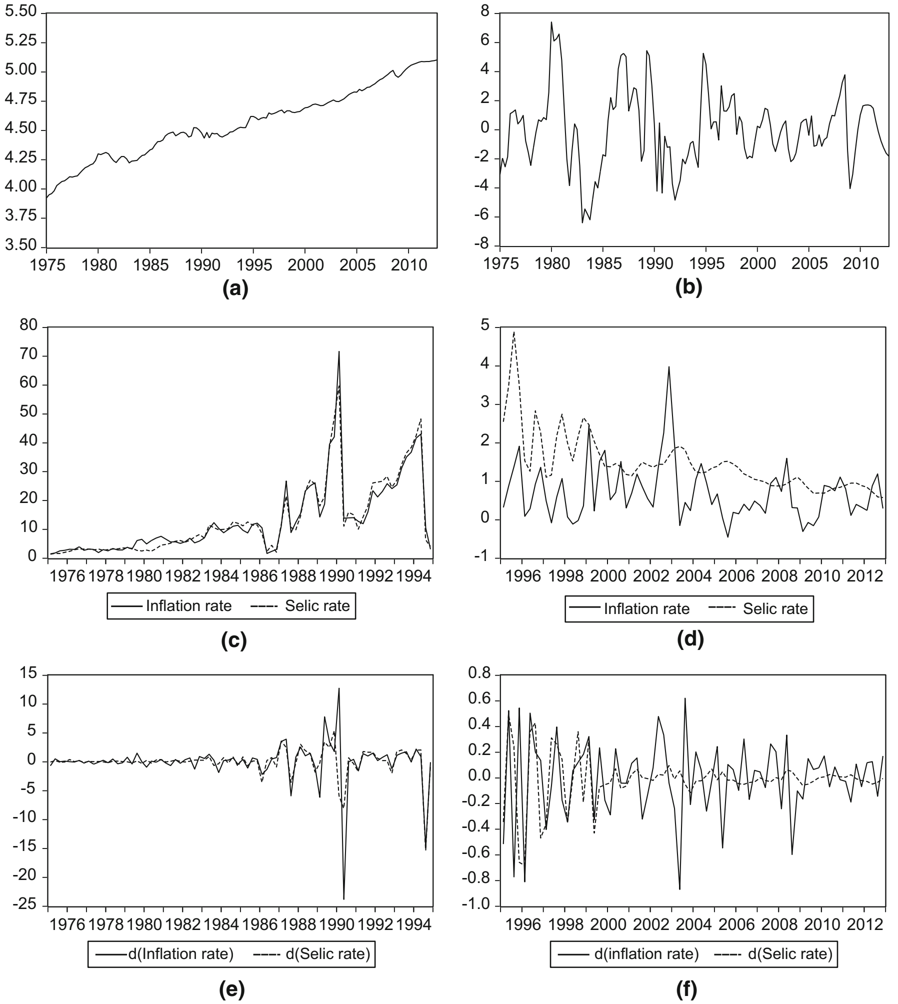

Fig. 1 Time series. **a** Real log-GDP in levels; **b** HP-filtered output gap; **c** IGP-DI inflation and Selic rate, 1975–1994; **d** IGP-DI inflation and Selic rate, 1975–2012; **e** inflation and Selic, first differences, 1975–1994; and, **f** inflation and Selic, first differences, 1995–2012

inflation, output gap, and interest rates in levels; and another, where we take the first difference of inflation and interest rates. Figure 1 depicts the analyzed time series.

This figure shows that from 1975 to the mid-1990s Brazilian inflation and interest rates were highly volatile, with periods of rapid growth and sudden stops. Besides, they became remarkably more stable after 1995. Such feature makes the estimation for the entire 1975–2012 period unmanageable, with inconsistently estimated parameters. Regarding this problem, we classify our data into specific subsamples, according to a three-equation test for multiples structural breaks that occur at unknown dates (Qu and Perron 2007).

Table 2 Dates of multivariate structural changes

Specification[a]	SupLR. statistic[b]	Break number 1 [90% CI]	Break number 2 [90% CI]	Break number 3 [90% CI]
Levels (at most 1 break)	482*	–	–	**1994.I**
				[1993.III, 1994.II]
Levels (at most 2 breaks)	685*	–	**1985.I**	**1994.II**
			[1984.IV, 1986.IV]	[1993.III, 1994.II]
Levels (at most 3 breaks)	880*	**1985.II**	**1994.I**	**2002.III**
		[1984.IV, 1986.IV]	[1993.III, 1994.II]	[2001.IV, 2003.IV]
Differences (at most 1 break)	560*	–	–	**1995.IV**
				[1989.II, 1996.I]
Differences (at most 2 breaks)	786*	–	**1985.II**	**1995.IV**
			[1985.I, 1992.IV]	[1987.III, 1996.I]
Differences (at most 3 breaks)	897*	**1985.II**	**1994.II**	**2003.III**
		[1985.I, 1993.II]	[1992.II, 1994.IV]	[2002.II, 2003.IV]

The bold entries represent estimated breakpoint dates

[a] 18 parameters changing across regimes

[b] *, ** and *** denote statistical significance at 1, 5 and 10%, respectively. The null tests zero against one, two or three breaks, depending on the case

3.3 Structural breaks and subsamples of estimation

Hess and Iwata (1997) showed that an ARIMA(1, 1, 0) model describes well the duration and amplitude of real log-GDP in the USA. Following this result, we estimated our breakpoint dates assuming a VAR(1) system with all variables in levels, and taking the first difference of inflation and interest rates. We allow for breaks in every parameter, including intercepts, autoregressive coefficients, and the error covariance matrix. Moreover, given the length of the data, we allow for three breaks at most, that is, four different regimes, and we exclude breaks that could appear at the first and final 7.5% of the series (setting the trimming parameter at 0.15). Estimation is carried out via an iterative feasible generalized least squares procedure (details in Qu and Perron 2007). Test results are described in Table 2.

As observed in Table 2, for the assumption of one structural break, in both specifications, that is, levels or first differences, a breakpoint is estimated around 1994/1995. The latter is the most dominant structural change in our system, since it is the first highlighted by the data. It is also statistically significant, with tight confidence intervals.

Considering the possibility of two structural changes, specification in levels suggests the first quarters of 1985 and 1994 as breakpoints, while with the first differences, the second quarter of 1985 and the fourth quarter of 1995 are the estimated breakpoints. Again, the dates of changes are similar in these frameworks, but confidence intervals are narrower when variables are in levels. The last significant break was estimated around 2003, with the sequential likelihood test suggesting three as the number of structural changes in our estimated VAR. Statistics are 195 and 102 for inflation and nominal rates in levels and in differences, respectively. Whatever the case, significance is high and therefore rejects the null of two breaks against the alternative of three breaks.

Based on the results of Table 2, we are able to define our four subsamples: the first one ranging from 1975.I to 1984.IV; the second one, from 1985.I to 1994.IV; the third one, from 1995.I to 2004.IV; and the last one from 2005 on. When doing so, we delimit our different samples according to the structural breaks test, while keeping each one of them, except to the last one, with the same length, i.e., 40 time series points. This procedure is important for minimizing the differences in the parameter estimation that would emerge from changing the precision of the estimates (Canova 2009).

In Table 3, we depict mean, standard deviation, and persistence of each variable across the samples. Persistence is defined as the first autocorrelation coefficient, as often referred to in the business cycle literature. The analysis of this table reveals some interesting facts.

First, mean and standard deviation of the output gap have been falling over the years. For instance, the standard deviation after 1995 is half of its value in the previous years. Moreover, we cannot reject the null of equal means and variances between the samples of 1995–2004 and 2005–2012. Conversely, output gap persistence is quite stable across the samples.

Second, inflation rate in levels shows a remarkable decrease in its mean and standard deviation after 1995. During the 1985–1994 period, it is observed a peak of these statistics, when seasonally adjusted inflation reached an average value of 840% per year, approximately.

In a comparison of the last two samples regarding inflation mean and variance, t and F equality tests lead to the rejection of the null hypothesis at 1% of statistical significance. Thus, the reduction in the mean inflation rate and its instability after 2005 is significant, which agrees with our structural break tests. Nominal interests rates have a pattern similar to the one described for inflation.

Additionally, figures for the first difference of inflation and interest rate have also decreased over the years. Thus, in the last sample, BCB has been controlling inflation with milder Selic variations. These results may reflect an improved credibility of Brazilian Central Bank in keeping inflation stable some years after the adoption of inflation targeting, as found by Carvalho and Minella (2012).

According to the results in Table 3, the Brazilian economy has clearly entered in a period of increased stability after 1995. As mentioned before, in developed countries this phenomenon is usually connected to changes in the private sector, monetary policy and/or in the shocks hitting the economy. In the present paper, all these assumptions are addressed in a simple way as the following. First, we estimate the New Keynesian model across the samples, delimitating systematic variations in the estimated parame-

Table 3 Descriptive statistics of subsamples

Stat.[a]/Sample	1975–1984	1985–1994	1995–2004	2005–2012	Full sample
Output gap (HP-filter)					
Mean	−0.52	0.24	0.21	0.09	0.00
Std. Dev.	3.46	3.06	1.53	1.74	2.62
Persistence	0.87	0.69	0.65	0.71	0.77
Inflation					
Mean	5.45	20.56	0.91	0.46	7.18
Std. Dev.	2.94	14.1	0.84	0.47	11.03
Persistence	0.96	0.55	0.29	0.42	0.81
Nominal rates—Selic					
Mean	5.02	21.14	1.86	0.97	7.58
Std. Dev.	3.20	13.75	0.79	0.25	10.96
Persistence	1.00	0.63	0.60	0.96	0.85
Δ(Inflation)					
Mean	0.06	−0.07	−0.01	0.01	−0.004
Std. Dev.	0.63	5.45	0.37	0.21	2.81
Persistence	−0.24[NS]	−0.26[NS]	−0.45	−0.42	−0.26
Δ(Selic)					
Mean	0.08	−0.06	−0.02	−0.01	−0.002
Std. Dev.	0.33	3.51	0.27	0.03	1.80
Persistence	−0.45	0.04[NS]	0.01[NS]	0.38	0.04[NS]
N. obs.	40	40	40	32	152

[a] NS stands for a not significant parameter at 10%

ters that could match the patterns in Table 3. Second, the model is estimated before and after the Real Plan launch, when we take notes of the simulated variances and parameter values. Finally, we re-estimate over the Real Plan sample while fixing a specific set of parameters at their pre-Real Plan values. Therefore, the question is whether the restricted model reproduces the observed fall in the volatility of output and inflation.

4 Volatility drop and parameter changes

As previously shown, the Brazilian economy has passed by quite a few structural breaks between 1975 and 2012. The most remarkable change across these subsamples is the volatility reduction: real GDP standard deviation fell 50% after 1995, while the instability of inflation and nominal interest rate presented even sharper reductions (see Table 3).

In this sense, we first verify the adequacy of the estimated models in reproducing this feature, which is done in Tables 4 and 5. In Table 4, we depict the estimated posterior standard deviation when output gap is derived from a HP-filter, and inflation and interest rates are in levels in Panel (a). By its turn, in Panel (b) inflation and Selic are

Table 4 Estimated posterior standard deviations of the endogenous variables for the HP-filter specification

Variable/sample	1975–1984	1985–1994	1995–2004	2005–2012
Panel (a): levels				
Inflation	8.41	41.26	1.98	1.45
	[5.5, 10.5]	[27.6, 53.5]	[1.5, 2.4]	[1.2, 1.6]
Gap (HP-filter)	3.77	5.01	1.83	2.02
	[3.1, 4.3]	[4.2, 5.8]	[1.3, 2.1]	[1.6, 2.4]
Selic rate	8.02	39.87	2.69	1.76
	[5.2, 10.1]	[22.8, 51.4]	[2.2, 3.2]	[1.3, 2.2]
Panel (b): differences				
ΔInflation	1.95	18.36	1.87	1.03
	[1.7, 2.2]	[14.8, 21.8]	[1.5, 2.1]	[0.9, 1.2]
Gap (HP-filter)	2.80	5.11	1.63	1.59
	[2.3, 3.2]	[4.1, 6.6]	[1.3, 1.9]	[1.3, 1.8]
ΔSelic rate	1.67	12.81	1.33	0.82
	[1.3, 1.9]	[9.5, 16.6]	[1.0, 1.5]	[0.6, 0.9]

90% confidence interval is between brackets

in first differences. Ninety percent confidence intervals are presented inside brackets. Table 5 has the same structure, except that, in this case, output gap is measured by the first difference of real log-GDP.

Table 4 shows that the volatility drop is present. In Panel (a), for example, the upper limit of the variables standard deviation after 1995 is smaller than the lower limit in the previous samples, which denotes a significant structural break in the country instability. In Panel (b), although results show a reduced volatility, significant breaks occur only after 2005, since the 1995–2004 statistics are not statistically different from those obtained in the first sample, considering a 90% confidence interval.

In Table 5, the reduction in the standard deviation of the variables is significant in 1995. Therefore, whichever the case depicted in Tables 4 and 5, a clear inverted U-shape pattern emerges from the estimated standard deviation of the variables across the samples, with peaks observed in the 1985–1994 sample, reproducing the facts we report in Table 3.[14]

We now address to systematic changes in the model parameters. Results are presented graphically, in order to facilitate the exhibition of a large set of information.[15] Dataset and subsamples are informative for all parameters: estimated coefficients present significant variations across time, several times away from the prior mean;

[14] Nevertheless, it is noteworthy that the posterior standard deviations implied by the model tended to overestimate the volatility of inflation and interest rate. Interestingly, Canova (2009) obtained a similar result. We guess that it could be due to a characteristic of the theoretical model, which is relatively simple, or even to difficulties found in DSGE models when matching some moments of aggregate prices, as Caballero (2010) pointed it.

[15] A complete list of the results is available in "Appendix," Tables 9, 10 and 11.

Table 5 Estimated posterior standard deviations of the endogenous variables for the output growth specification

Variable/sample	1975–1984	1985–1994	1995–2004	2005–2012
Panel (a): levels				
Inflation	5.39	39.51	1.49	0.95
	[3.5, 6.6]	[24.9, 49.9]	[1.2, 1.7]	[0.7, 1.1]
Gap (ΔGDP)	1.55	2.58	0.62	0.53
	[1.2, 1.8]	[2.1, 3.0]	[0.5, 0.7]	[0.4, 0.6]
Selic rate	5.70	38.50	2.11	1.27
	[3.6, 7.0]	[23.0, 48.7]	[1.7, 2.4]	[1.0, 1.4]
Panel (b): differences				
ΔInflation	1.74	16.95	1.14	0.59
	[1.5, 1.8]	[14.7, 18.5]	[1.0, 1.2]	[0.57, 0.64]
Gap (ΔGDP)	0.70	2.63	0.37	0.28
	[0.6, 0.8]	[2.2, 3.0]	[0.3, 0.4]	[0.2, 0.3]
ΔSelic rate	1.30	13.18	0.73	0.22
	[1.1, 1.5]	[10.7, 14.6]	[0.6, 0.8]	[0.17, 0.22]

90% confidence interval is between brackets

and in most of the cases, posterior standard deviations of the estimated parameters are smaller than the prior standard deviations.

We begin by analyzing private sector and inertia coefficients, depicted in Fig. 2. The estimated parameters in this figure are presented separately in four parts, *i* to *iv*. In each one of these, letters *a*, *b*, *c* and *d* stand for the models that consider, respectively: (a) HP-filter output gap, inflation and Selic in levels; (b) HP-filter output gap, inflation and Selic in differences; (c) first difference output gap, inflation and Selic in levels; and (d) first differences of output gap, inflation and Selic.

The parameter controlling the IS-curve inclination, δ (delta), which measures contemporaneous dependence of output gap on changes in *ex ante* real interest rate is fairly constant over time in all four models. It is estimated at 1.5 or 2.0 (see Fig. 2, part *i*), a value closer to the prior. Notwithstanding, the posterior standard deviations of the parameter were smaller than the prior belief in all but two cases.[16] This leads to the conclusion that monetary policy impacts real economic activity in Brazil more than proportionally. Similar values for the IS inclination were found by Mello and Moccero (2011), while using a VAR methodology for the periods 1996–1998 and 1999-Feb/2006, with monthly data.

The influence of output gap on inflation, measured by κ in the Phillips curve, has an interesting dynamic over time. Usually, this parameter reaches a peak in the 1985–1994 sample (the only exception is depicted in Fig. 2, part *ii*, model *c*). Besides, considering models where inflation and interest rate are in differences, this effect is even more remarkable (see models *b* and *d*). This dynamics is probably reflecting

[16] Specifically, in model *b*, samples 1975–84 and 1985–94.

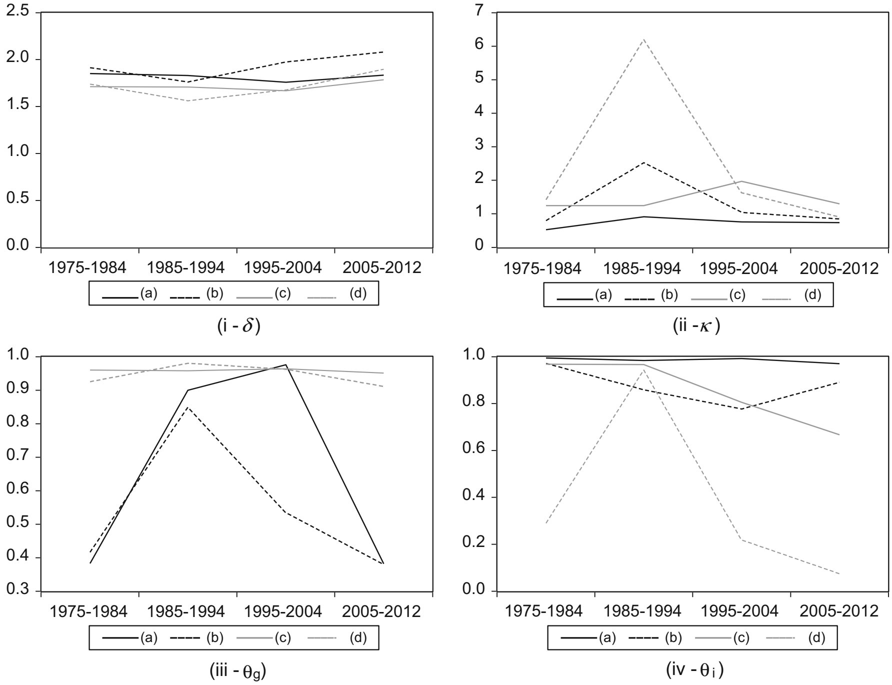

Fig. 2 Changes in private sector and inertia parameters across subsamples in different models Notes: **a** inflation and Selic rate in levels, HP-filter output gap; **b** inflation and Selic rate in differences, HP-filter output gap; **c** inflation and Selic rate in levels, log-GDP differences output gap; and **d** inflation and Selic rate in differences, log-GDP differences output gap

the effects of consumption bubbles generated right after the price-freezing policies. Following the Real Plan implementation, the output gap coefficient in the Phillips curve inflation has been smaller than one, around 0.75 and 0.9.

Endogenous output persistence is captured by θ_g. Whenever the output gap is measured by the HP-filter, this parameter has tended to be smaller in the recent years of our dataset. However, if one considers the first difference of GDP as a proxy for the output gap, θ_g is estimated around 1.0 in most cases, which approximates the IS-curve to the backward-looking framework proposed by Svensson (1997). In this sense, based on our estimations, drawing general conclusions about the persistence of the aggregate demand is troublesome. Table 3 may be more informative in this regard. It shows that the autoregressive coefficient of output gap oscillates around 0.75 over the years.

The θ_i parameter is in Fig. 2, part *iv*. It captures inflation inertia and is, without any doubt, one of the most important sources of information in our estimations.[17] For the case illustrated in model *a*, where variables are in levels, and output gap is derived from the HP-filter, our results show that inflation inertia has slowly decreased over time, with 90% confidence interval of [0.98, 1.00] in the first sample, and

[17] Indexation has regularly attracted a great deal of attention in Brazilian policy-making.

[0.92, 1.00] in the last one. In model *c* of Fig. 2, where we measure the output gap as the first difference of real log-GDP, this result is strengthened, with θ_i point estimates starting at 0.97 ([0.91, 1.00]) between 1975 and 1984, and 0.97 ([0.89, 1.00]), 0.81 ([0.66, 1.00]), and 0.67 ([0.50, 0.81]) in the next three samples, respectively.

In models *c* and *d* of Fig. 2, part *iv*, where inflation and Selic are presented in growth rates, both estimations indicate reduced inflationary inertia. In model *c*, the smallest value for θ_i is found in the 1994–2005 sample, that is, 0.78. In model *d*, our results are rather interesting, with θ_i being estimated at 0.29 in the first sample, then 0.94, 0.21, and only 0.07 in the next three samples. Thus, according to model *d*, inflation inertia has greatly reduced after 1995, with a peak found between 1985 and 1994.

Overall, we can conclude that inflationary inertia has decreased in Brazil over the years, but this reduction is extremely more apparent in the specifications with variables in growth rates. Therefore, even in the years following 1995, the inflationary process (its mean value) still tends to be persistent, but it does not have an explosive dynamics, as it used to have before this year. Our results, regarding the cases where inflation is in levels, are related to the findings of Figueiredo and Marques (2009), who used ARFIMA-GARCH methodology and monthly IGP-DI data from Aug/1994 to Jan/2008, showing that inflation is a stochastic process with long memory.

We now analyze changes in monetary policy parameters. Figure 3 brings information about the coefficients governing interest rate smoothness, ρ, and the weights of inflation and output gap in the Taylor rule, γ_i and γ_g, respectively. Beginning with ρ, in part *i*, it tends to be small and alike in all estimation setups. Its values are around 0.05, ranging from 0.01 in 1985–1994, when all variables are in differences (Fig. 3, model *d*), to 0.12 in 2005–2012, when inflation and interest rate are in differences, and output gap is measured by the HP-filter (Fig. 3, model *b*). Besides, 90% confidence interval includes 0.00 most of the times. In this sense, our results support that the Brazilian Central Bank has not been pursuing gradual monetary policy conduction over time; instead, interest rate swings tended to be sharp, especially before 2003. This is an intriguing feature of the Brazilian economy, since some authors, including Moreno (2004) and Canova (2009), found higher values for ρ (usually above 0.8), which reflects the FED smoothing behavior.

Concerning to the weights of the policy rule in parts *ii* and *iii* of Fig. 3, when inflation and interest are in levels, one can see that the models present some interesting features, such as: (i) before the Real Plan, monetary authority imputed higher weight on output stabilization and (ii) BCB has become more adverse and responsive to elevations in inflation expectations since 1995, especially after the adoption of inflation targeting.

Another interesting information arises from the examination of the policy rule in the first two samples. As Table 10 shows, when inflation and interest rates are in levels, as in Panels *a* and *c*, inflation deviations had a smaller weight than output deviations. The inverse situation is seen when these variables are in first differences (Panels *b* and *d*). Such results are indicating that monetary measures in Brazil would

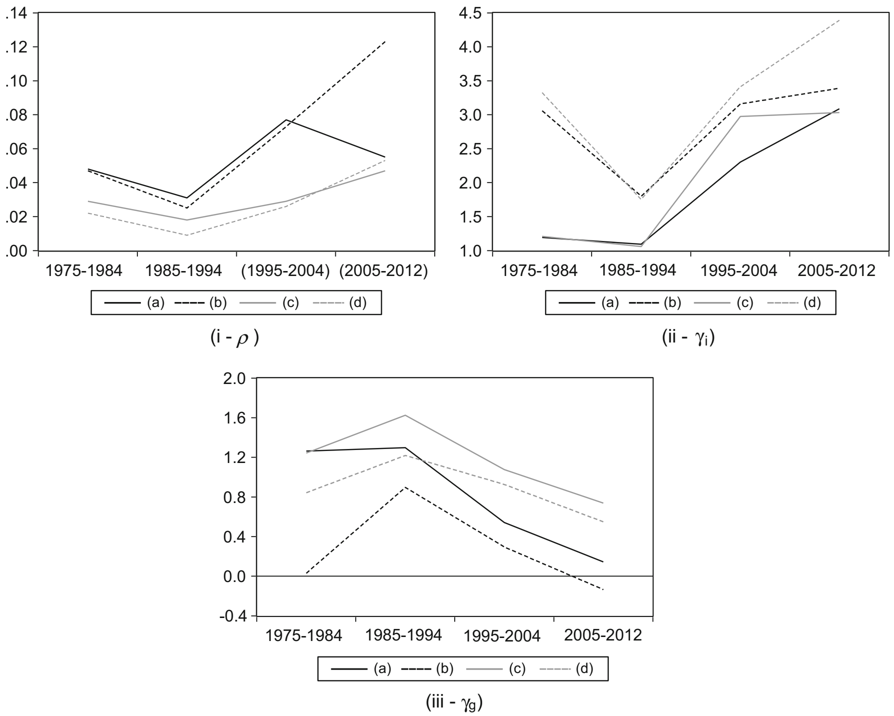

Fig. 3 Changes in policy parameters across subsamples in different models. *Notes* **a** inflation and Selic rate in levels, HP-filter output gap; **b** inflation and Selic rate in differences, HP-filter output gap; **c** inflation and Selic rate in levels, log-GDP differences output gap; and, **d** inflation and Selic rate in differences, log-GDP differences output gap

provide inflation tax in order to accommodate government fiscal and credit policies, whenever the level of inflation was close to some kind of a tolerable value. It would become more reactive, however, if inflation started to gain momentum, accelerating more rapidly. This view is able to fit the increasing rates of seigniorage as a share of GDP in the period, which went from 2% between 1975 and 1984, to 4% between 1985 and 1994 (Pastore et al. 2015).[18] After 1995, this pattern is not present anymore, once the coefficient of inflation in the Taylor rule is higher than the one estimated for the output gap, regardless the assumed specification, reflecting a change in the Central Bank behavior.

Figure 4 presents the last set of estimated parameters, referring to the exogenous instabilities in IS and Phillips curves, and in the Taylor rule, respectively, σ_1, σ_2, and σ_3. Two facts emerge from this figure. First, the shocks faced by the Brazilian economy in 1985–1994 were much greater than those observed in the USA. Estimates for that country show that the standard deviation of the disturbances reaches a maximum value

[18] The fact that a Taylor Principle is satisfied for the 1985–1994 period, when inflation was very high (see Table 10), could also suggest that BCB had targets for the inflation tax (as a share of GDP). With prices growing quickly and eroding the real stock of money, financing public deficits with seigniorage required increasing monetary expansions and, therefore, higher inflation rates.

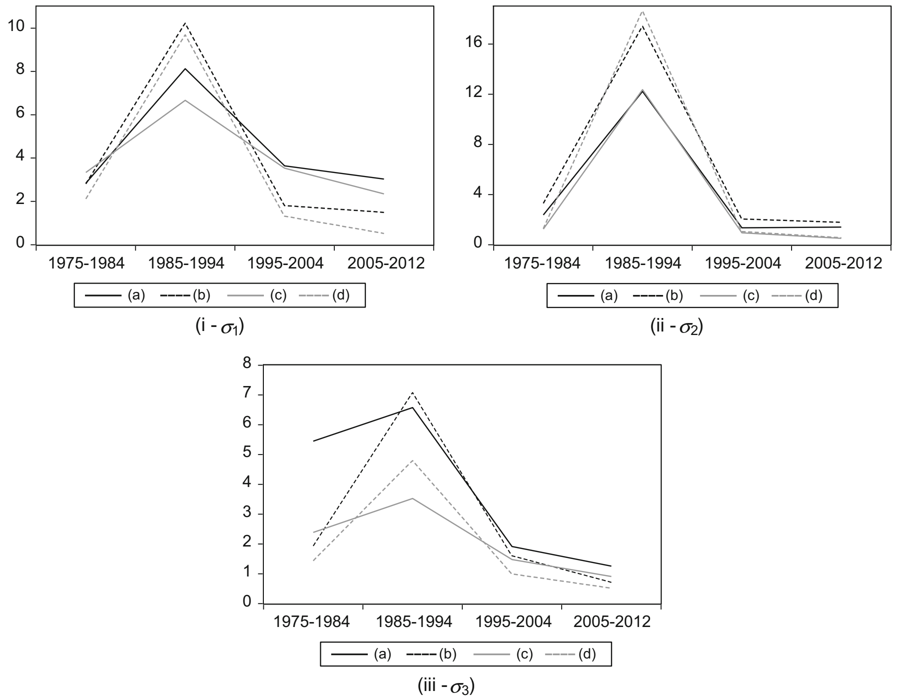

Fig. 4 Changes in shock parameters across subsamples in different models. *Notes* **a** Inflation and Selic rate in levels, HP-filter output gap; **b** inflation and Selic rate in differences, HP-filter output gap; **c** inflation and Selic rate in levels, log-GDP differences output gap; and, **d** inflation and Selic rate in differences, log-GDP differences output gap

of 3.0 in systems similar to the one employed here (see, e.g., Stock and Watson 2002; Moreno 2004; Canova 2009).

Since the early and mid-1990s, however, shocks in Brazil have also become milder and were estimated at 2.2 (σ_1), 1.2 (σ_2), and 1.2 (σ_3). Besides, all three graphs in Fig. 4 display an inverted U-shape pattern, comparable to that observed in the raw data (see Table 3). This result indicates that milder shocks are probably able to explain part of the stabilization process in Brazil. This issue is addressed in the next section.

5 Explanations for the Brazilian stabilization

Now, we seek to explain the causes of the drop in the volatility of output and inflation in Brazil. To analyze this issue, we estimate the theoretical model using two samples, the first one over 1975 and 1993, named "pre-Real Plan" and the second one over 1995 and 2012, named "post-Real Plan".[19] Next, we perform the estimation again over the post-Real Plan period, while restricting some parameters at their 1975–1993 estimated values. According to Canova (2009), this procedure allows us to investigate

[19] We discard the year of 1994, because it can be seen as a transition period.

Table 6 Posterior moments: restricted and unrestricted specifications, 1995–2012 sample

Variable	Unrestricted stand. error, post-real plan [90% CI]	Stand. errors restricting parameters (at their pre-real plan values)		
		Private and inertia	Policy	Shocks
Panel a: inflation and Selic in levels, HP-filter gap				
Inflation	1.6 [1.4, 1.6]	1.4	8.2	5.6
Output gap	1.9 [1.7, 2.1]	1.8	1.6	3.1
Panel b: inflation and Selic in differences, HP-filter gap				
Inflation	1.2 [1.1, 1.3]	2.5	2.0	7.5
Output gap	1.3 [1.1, 1.5]	1.3	1.6	3.5
Panel c: inflation and Selic in levels, ΔGDP gap				
Inflation	1.3 [1.1, 1.4]	1.3	5.3	5.2
Output gap	0.6 [0.4, 0.6]	0.7	0.7	1.8
Panel d: inflation and Selic in differences, ΔGDP gap				
Inflation	1.0 [0.9, 1.0]	2.6	1.1	6.4
Output gap	0.4 [0.4, 0.8]	0.3	0.3	1.8

whether the restricted models are able to reproduce the volatility fall observed in the data. Results are presented in Table 6.

The evidence in Table 6 is clear. First, whenever we consider inflation and interest rate in levels, that is, specifications *a* and *c*, if the policy parameters had remained at their pre-Real Plan values, inflation instability would have been about 5 times higher (compare, for example, the values 1.6 and 8.2 in Panel *a*). Output gap volatility, however, is rather unresponsive to changes in the policy rule in all estimations. By their turn, exogenous shocks soundly affect the volatility of inflation and output gap. In some estimation, these variables would have been four or five times higher if the shocks had stayed at their 1975–1993 values. Thus, the improved stability of the economic environment has also been a key element of the stabilization in the country.

Private sector as a whole had no statistically significant effect on the volatility of inflation and output gap, whenever one considers variables in levels (Panels *a* and *c*). However, in specifications *b* and *d*, where inflation and interest rates are in differences, this situation changes, and private parameters affect inflation significantly.

In order to filter which parameter of this set matters the most, we have implemented our comparative exercise and restricted, in each turn, the Phillips curve inclination and the inflationary inertia to their pre-Real plan values.[20] Table 7 brings the results.

If inflationary inertia is restricted, the effects on the volatility of inflation and output gap are almost irrelevant. For example, considering the specifications *b* and *d* of

[20] Consumption elasticity and output persistence are quite stable along the estimation samples and are disregarded at this point of our analysis.

Table 7 Posterior moments when some private parameters are restricted, models b and d, 1995–2012 sample

Variable	Unrestricted stand. error, post-real plan [90% CI]	Stand. errors restricting (pre-real plan values)	
		Inflationary inertia (θ_i)	Phillips curve inclination (κ)
Panel b: inflation and Selic in differences, HP-filter gap			
Inflation	1.2 [1.1, 1.3]	1.3	2.3
Output gap	1.3 [1.1, 1.5]	1.5	1.4
Panel d: inflation and Selic in differences, ΔGDP gap			
Inflation	1.0 [0.9, 1.0]	1.3	2.6
Output gap	0.4 [0.4, 0.8]	0.5	0.4

Table 7, when we restrict θ_i to their pre-Real Plan values, inflation and output volatilities would be, respectively, 1.3 and 1.5 (Panel *b*), and 1.3 and 0.5 (Panel *d*). Thus, the reduced inflationary inertia does little to explain the regime of low volatility in the country.

The inflation inertia issue deserves an in-depth look, given its importance for the formulation of recent monetary plans. Our results show that reduction in this parameter, i.e., in the importance of past inflation on its future values, is not a decisive factor for improving macroeconomic stability in the country. In other words, even if the inertia had been eliminated, while monetary policy and exogenous shocks had remained at their pre-Real Plan levels, inflation would have continued to increase and, almost certainly, new indexing mechanisms would have been created. Moreover, GDP volatility is not affected by the inflationary inertia degree.

On the other hand, if the Phillips curve inclination were restricted, inflation volatility would be twice as higher after 1995. A possible explanation for this result is very interesting indeed. Such parameter, κ, stands for real rigidities in price setting, and it reached a peak in the sample from 1985 to 1994.[21] During these years, monetary policy often froze prices, thus leading price setters to readjust them in every possible situation in order to restore relative prices and profits. Moreover, it is easy to think that menu costs were negligible in this hyperinflationary context. Our paper corroborates, therefore, the vision that an important feature of the Real Plan was the URV mechanism, which allowed for a gradual realignment of prices in the economy and stabilized the behavior of price setters after its implementation.

There has been a long controversy between Thomas Sargent (see Sargent 1982) and Rudi Dornbusch (sided with Stanley Fischer and M. H. Simonsen, see, for instance, Dornbusch 1982; Dornbusch et al. 1986) about the importance of inflation inertia in hyperinflation process. According to Dornbusch (1982), inflation inertia, as defined

[21] A higher value for κ denotes prices that are more flexible.

by a persistent transmission of past inflation through time, brings prohibitively high costs of eradicating inflation. As long as firms and workers form their expectations using past information, conventional restrictive monetary policy causes lower output and employment, while prices keep their growing pace. In other words, aggregate demand would have to be kept at lower levels for long periods until expectations could change.

The idea of conventional monetary policy ineffectiveness in hyperinflation contexts was extremely appellative during the 1980s in Brazil and other Latin American countries, spurring experiments with heterodox-type stabilization plans. However, considering that inflation rate was reaching about 1000% per year between 1990 and 1993, it is understandable that Brazilian policy makers were giving up on heterodox-plans by the time, in favor to an alternative view, in line with Sargent (1982) positioning.

According to this author, firms and workers are rational and forward-looking agents bargaining in an inflationary environment. Thus, inflation now comes to be high exactly because people anticipate government expansionary policies, such as budget deficits, loosen monetary conditions, and price freezing into their decision-making process; that is, government's current and prospective policies warrant people's inflationary expectations. In this sense, in order to stop inflation it would be required (Sargent 1982, p. 42): *a change in the policy regime (…)*.

Our findings point toward Thomas Sargent's view in many ways. First, as Table 7 shows, reduced nominal indexation had no significant statistical effect on inflationary volatility. Second, we find that monetary policy did change after 1994, and this new regime has had a central focus on controlling inflation, i.e., there was a policy regime shift. Finally, as our regressions show, Phillips curve inclination was steeper in the 1985–1994 sample, when inflation reached its higher historical values, indicating that agents adjusted prices aggressively under hyperinflation due to an anticipation of future conditions.[22]

As the Real Plan did not rely on artificial price freeze, and managed to implement sound monetary policies, it built credibility with the public, which according to Végh (1992) is a key factor for a successful, low cost, disinflation program. Thus, our results are also in line with Végh (1992) findings, who studied monetary reforms in several countries, including Argentina, Bolivia, Brazil, Israel and Uruguay. Additionally, Wolf (2001), studying the Latvian successful stabilization in 1922, argued that its main determinants were a mix of monetary and fiscal policies committed with stopping inflation, a gradual realignment of prices by the introduction of a device much like URV, and a bit of luck; i.e., some of the same elements we highlight here.

Table 8 brings additional information on the volatility fall in Brazil. In this table, the "pre-Real Plan" column denotes the posterior standard error of the variables in the 1975–1993 period; by its turn, the "Restricted" column shows the same statistics for the post-Real plan period, while the exogenous shocks and the policy rule

[22] Not coincidentally, we titled our paper honoring Sargent (1982) work.

Table 8 Comparisons between pre- and restricted post-Real Plan volatilities

Variable/std. Deviation	Pre-real	Post-real (restricting policy + shocks)	Percentage	Pre-real	Post-real (restricting policy + shocks)	Percentage
	Panel a			Panel b		
Inflation	41.2	38.5	93.4	10.6	8.6	81.1
Output gap	3.8	3.9	102.6	4.1	4.5	109.8
Selic	40.4	37.8	93.6	8.1	5.1	63.0
	Panel c			Panel d		
Inflation	37.8	38.5	101.9	8.8	7	79.5
Output gap	1.8	1.8	100.0	1.1	1.4	127.3
Selic	36.8	37.8	102.7	6.7	4	59.7

parameters are fixed at their first sample values. Therefore, we intend to investigate whether the volatility in the economy after 1995 would be similar to that observed during 1975–1993 if shocks and monetary policy had remained unchanged over time.

Table 8 shows that the restricted models tended to replicate a significant percentage of the 1975–1993 instability. Especially in the levels assumptions (Panels *a*, and *c*), all ratios are very close to 100% and, sometimes, they are even higher. Thus, the exercises in Table 8 show that both sources of explanations, that is, the good policy and good luck hypotheses, can account for the stabilization in the country.

6 Conclusions

The present paper investigated the occurrence of a volatility break in Brazil, in a process similar to those occurred in several OECD countries. For such, we estimate a small-scale New Keynesian DSGE model, using time series spanning from 1975 to 2012, and Bayesian techniques. Our main conclusions are outlined below.

It is widely known that output and inflation have been presenting milder fluctuations since 1995 in Brazil. The present paper highlighted parts of the very mechanisms driving this structural change. Some figures we obtained came to confirm what previous studies have already found, while some new information came to expand our understanding about the accomplishments of Real Plan.

Regarding inflation, the main explanations its reduced volatility are smaller shocks, related to the good luck hypothesis, and changes in the monetary policy parameters. Over time, the weight of inflation in the Taylor rule has greatly increased, in opposition to the weight of output gap, which tended to reduce. Thus, between 1995 and 2012 the Brazilian Central Bank has focused on controlling prices changes, rather than constantly exploring the tradeoff between inflation and unemployment, an important policy for maintaining inflation on track.

Concerning to output, we found that the sole reason for its volatility fall was the reduction of the shocks hitting the economy. Thus, output stabilization was not due to improved monetary policy, but rather to changes in parameters or variables that are exogenous to our monetary model. In this sense, our results are in line with some papers, which focused on the US case, including Stock and Watson (2002), Moreno (2004), Primiceri (2005), Sims and Zha (2006) and Canova (2009).

Changes in private sector parameters, such as consumer preferences, output persistence and inflationary inertia, cannot explain the Brazilian stabilization. Thus, decreased indexation, which has been regarded as a key element for controlling inflation in the country, was not significant in our estimations. On the other hand, we found evidence suggesting that a more sticky price setting has contributed to reduce inflation volatility (the Phillips curve has become flatter after 1995). That is, if agents increasingly understand aggregate demand shifts as generalized pressures on the economy, they will pursue price adjustments more aggressively. As we argued before, this result strongly discourages the adoption of price-freezing policies, since long periods of price misalignments can only magnify this price-setting behavior.

In sum, the main finding of the paper is that the Real Plan and its subsequent monetary policy were successful in reducing inflationary instability in the country, but they are not the sole determinant of this structural change. Milder shocks hitting the economy and changes in the price-setting behavior have also been contributing to it. Besides, output volatility did not seem to respond to the improved monetary regime, at least in the context we have considered here. For that reason, Brazilian Central Bank should keep its focus on stabilizing prices, assuming that the continuity of a low inflation regime rests on sound monetary policies. It should be clear that depending on the size of the exogenous perturbations, inflation might rise again. However, by using the correct instruments, the Central Bank will be able to mitigate their effects throughout the economy.

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Appendix

See Tables 9, 10, and 11.

Table 9 Private and inertia parameters estimates in each subsample

Coefficient/sample	Priors	1975–1984			1985–1994			1995–2004			2005–2012		
		Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%
Panel a													
Delta (δ)	2.00	1.851	1.503	2.190	1.831	1.435	2.174	1.759	1.439	2.002	1.834	1.496	2.174
Kappa (κ)	2.00	0.527	0.452	0.585	0.910	0.728	1.177	0.759	0.585	0.914	0.735	0.555	0.911
Beta (β)	0.98	0.980	0.965	0.995	0.979	0.964	0.993	0.983	0.974	0.994	0.980	0.964	0.996
Theta g (θ_g)	0.80	0.383	0.202	0.556	0.900	0.706	1.000	0.976	0.921	1.000	0.382	0.001	0.840
Theta i (θ_i)	0.80	0.994	0.984	1.000	0.984	0.948	1.000	0.992	0.977	1.000	0.970	0.915	1.000
Panel b													
Delta (δ)	2.00	1.913	1.607	2.290	1.760	1.483	2.062	1.974	1.633	2.313	2.080	1.724	2.484
Kappa (κ)	2.00	0.797	0.624	0.956	2.526	1.426	3.795	1.040	0.726	1.330	0.845	0.658	1.047
Beta (β)	0.98	0.980	0.966	0.996	0.979	0.963	0.995	0.980	0.966	0.995	0.981	0.966	0.995
Theta g (θ_g)	0.80	0.416	0.268	0.599	0.849	0.580	1.000	0.534	0.293	0.782	0.379	0.166	0.583
Theta i (θ_i)	0.80	0.971	0.911	1.000	0.858	0.653	1.000	0.777	0.503	1.000	0.891	0.682	1.000
Panel c													
Delta (δ)	2.00	1.712	1.425	1.995	1.707	1.416	2.005	1.668	1.374	1.946	1.785	1.454	2.114
Kappa (κ)	2.00	1.247	0.929	1.573	1.245	0.858	1.622	1.968	1.458	2.508	1.297	0.975	1.593
Beta (β)	0.98	0.982	0.970	0.996	0.980	0.966	0.996	0.980	0.964	0.995	0.979	0.960	0.992
Theta g (θ_g)	0.80	0.960	0.881	1.000	0.958	0.868	1.000	0.964	0.886	1.000	0.951	0.841	1.000
Theta i (θ_i)	0.80	0.968	0.906	1.000	0.966	0.893	1.000	0.805	0.658	1.000	0.667	0.504	0.810

Table 9 continued

Coefficient/sample	Priors	1975–1984			1985–1994			1995–2004			2005–2012		
		Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%
Panel d													
Delta (δ)	2.00	1.739	1.449	2.015	1.560	1.348	1.748	1.676	1.396	1.914	1.897	1.578	2.228
Kappa (κ)	2.00	1.420	1.025	1.788	6.194	4.393	7.775	1.624	1.162	2.025	0.899	0.681	1.100
Beta (β)	0.98	0.979	0.963	0.994	0.979	0.969	0.995	0.981	0.967	0.996	0.979	0.962	0.993
Theta g (θ_g)	0.80	0.925	0.783	1.000	0.980	0.934	1.000	0.962	0.885	1.000	0.911	0.738	1.000
Theta i (θ_i)	0.80	0.290	0.159	0.419	0.943	0.791	1.000	0.218	0.114	0.315	0.074	0.007	0.136

CI10% and CI90% stand for the lower and upper limits of the 90% confidence interval, respectively

Table 10 Policy parameters estimates in each subsample

Coefficient/sample	Priors	1975–1984			1985–1994			1995–2004			2005–2012		
		Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%
Panel a													
Constant ($\bar{c}$)	2.0	1.506	1.208	1.849	1.831	1.498	2.181	1.790	1.531	2.069	1.591	1.239	1.946
Rho (ρ)	0.8	0.048	0.000	0.100	0.031	0.000	0.066	0.077	0.000	0.155	0.055	0.001	0.118
Gamma i (γ_i)	1.5	1.194	1.082	1.299	1.092	1.036	1.146	2.302	1.848	2.716	3.086	2.523	3.727
Gamma g (γ_g)	0.5	1.264	0.994	1.531	1.298	1.025	1.566	0.541	0.296	1.031	0.144	−0.146	0.438
Panel b													
Constant ($\bar{c}$)	2.0	1.654	1.380	1.977	1.965	1.634	2.378	1.520	1.213	1.896	1.328	0.996	1.633
Rho (ρ)	0.8	0.047	0.001	0.104	0.025	0.000	0.050	0.073	0.001	0.151	0.123	0.012	0.218
Gamma i (γ_i)	1.5	3.059	2.442	3.593	1.798	1.398	2.116	3.157	2.519	3.894	3.390	2.720	4.040
Gamma g (γ_g)	0.5	0.029	−0.197	0.251	0.897	0.599	1.256	0.293	0.009	0.606	−0.136	−0.374	0.104
Panel c													
Constant ($\bar{c}$)	2.0	1.461	1.145	1.722	1.700	1.321	2.052	1.717	1.365	2.043	1.584	1.231	1.868
Rho (ρ)	0.8	0.029	0.000	0.058	0.018	0.000	0.038	0.029	0.000	0.062	0.047	0.000	0.100
Gamma i (γ_i)	1.5	1.208	1.103	1.331	1.060	1.026	1.096	2.974	2.432	3.504	3.032	2.367	3.671
Gamma g (γ_g)	0.5	1.243	0.924	1.578	1.625	1.320	1.967	1.077	0.742	1.416	0.738	0.394	1.083
Panel d													
Constant ($\bar{c}$)	2.0	1.627	1.306	1.960	1.950	1.538	2.372	1.435	1.105	1.703	1.010	0.800	1.192
Rho (ρ)	0.8	0.022	0.000	0.049	0.009	0.000	0.019	0.026	0.000	0.051	0.053	0.000	0.107
Gamma i (γ_i)	1.5	3.326	2.667	3.949	1.754	1.572	1.938	3.409	2.847	3.953	4.391	3.929	4.959
Gamma g (γ_g)	0.5	0.842	0.494	1.184	1.219	0.985	1.548	0.925	0.593	1.275	0.548	0.213	0.894

CI10% and CI90% stand for the lower and upper limits of the 90% confidence interval, respectively

Table 11 Exogenous shocks estimates in each subsample

Coefficient/sample	Priors	1975–1984			1985–1994			1995–2004			2005–2012		
		Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%	Posterior	CI10%	CI90%
Panel a													
Sigma 1 (σ_1)	0.01	2.839	2.110	3.678	8.124	5.996	10.390	3.643	2.789	4.767	3.028	2.047	3.935
Sigma 2 (σ_2)	0.01	2.366	1.910	2.895	12.233	10.366	14.499	1.343	1.078	1.606	1.405	0.973	1.853
Sigma 3 (σ_3)	0.01	5.450	3.977	6.600	6.578	5.299	7.907	1.914	1.458	2.574	1.255	0.665	1.941
Panel b													
Sigma 1 (σ_1)	0.01	2.807	2.112	3.518	10.230	7.445	12.818	1.802	1.363	2.278	1.485	1.008	1.999
Sigma 2 (σ_2)	0.01	3.297	2.514	4.046	17.425	12.112	21.292	2.042	1.438	2.564	1.771	1.196	2.288
Sigma 3 (σ_3)	0.01	1.934	1.350	2.428	7.079	5.363	8.579	1.604	1.072	2.095	0.711	0.425	0.985
Panel c													
Sigma 1 (σ_1)	0.01	3.334	2.298	4.330	6.665	4.353	8.512	3.534	2.689	4.407	2.344	1.704	3.067
Sigma 2 (σ_2)	0.01	1.248	1.012	1.450	12.365	10.029	14.549	0.952	0.725	1.171	0.511	0.386	0.629
Sigma 3 (σ_3)	0.01	2.386	1.907	2.962	3.524	2.866	4.382	1.475	1.195	1.746	0.907	0.712	1.134
Panel d													
Sigma 1 (σ_1)	0.01	2.113	1.554	2.599	9.692	7.349	11.320	1.319	0.994	1.618	0.515	0.359	0.675
Sigma 2 (σ_2)	0.01	1.286	1.030	1.547	18.662	14.507	24.187	1.044	0.766	1.225	0.553	0.436	0.668
Sigma 3 (σ_3)	0.01	1.435	1.109	1.740	4.797	4.096	5.518	0.993	0.776	1.216	0.511	0.370	0.675

CI10% and CI90% stand for the lower and upper limits of the 90% confidence interval, respectively

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