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Empirical estimates for the Brazilian total imports equation using quarterly national accounts data (1996–2010)☆

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Abstract

This paper presents econometric estimates for the Brazilian aggregate imports over the period 1996–2010. To the best of our knowledge, this is the first paper that uses the Brazilian quarterly national accounts with this goal in mind. Besides estimating a demand equation (canonical model), as it is usual in the literature, we also explore the co-movements among total imports, gross fixed capital formation and household consumption (alternative model). The results underscore the role played by domestic income, which is the main determinant of total imports. The limited domestic supply of capital goods makes the allocation of domestic income also relevant to the imports dynamics. It should be noted, however, that this specification needs a better theoretical grounding. Evidence of nonlinearities has been found by different estimation techniques. Unfortunately its precise characterization turns out to be difficult due to the diversity of the results obtained. The out-of-sample assessment (one-step-ahead forecast) shows a good performance of the long-run vectors of the alternative models in predicting aggregate imports, meanwhile the best performance is obtained by the error correction representations of the canonical model. We believe those differences emerge from the different speeds of adjustment toward the long-run solutions.

JEL classification: C22; C52; F1; F17

Keywords: Quarterly Brazilian imports; Cointegration models with structural breaks; Markov-switching models; State-space models

Resumo

O presente artigo apresenta estimativas econométricas das importações agregadas brasileiras. A primeira inovação reside na utilização dos dados de importações das contas nacionais trimestrais. A segunda contribuição se refere aos modelos empíricos, pois, além da equação de demanda (modelo canônico), exploramos o comovimento das importações com a formação bruta de capital fixo, além do consumo das famílias (alternativo). Em terceiro lugar, a instabilidade dos parâmetros foi examinada por diferentes técnicas. Em relação aos resultados, vale destacar o papel da renda interna, que é o principal determinante das importações. A limitada oferta doméstica de bens de capital faz com que a alocação da renda interna seja relevante, cabendo notar que essa especificação necessita de uma fundamentação teórica melhor. Embora tenham sido encontradas evidências de não linearidades, a diversidade de resultados torna difícil sua caracterização exata. Na avaliação fora da amostra (previsões um passo à frente), os vetores de longo prazo apresentaram bom desempenho para o modelo alternativo, enquanto as representações de correção de erros do modelo canônico

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obtiveram os melhores resultados, estando essa diferença possivelmente relacionada às distintas velocidades de ajustamento na direção da solução de longo prazo.

Palavras-chave: Importações brasileiras trimestrais; Cointegração com quebras estruturais; Modelos de alternância de regimes markovianos; Modelos de espaço-estado

1. Introduction

There have been several studies discussing the long-run and quarterly dynamics of Brazilian imports. We aim to contribute to this literature in three innovative ways. First, given the recent current account scenario, there is a growing interest in predicting the official data published by the Brazilian Statistical Office (Institute of Geography and Statistics – IBGE). Therefore, our motivations are similar to the ones found in macroeconometric models of projection (Reis et al., 1999). This fact led us to use imports data measured at constant prices from the quarterly national accounts meanwhile the majority of the previous papers used quantity indexes.

Second, we investigate possible nonlinearities in the data using three different techniques: cointegration, Markov switching and state-space models. This emphasis on nonlinearities comes from the perception that there might have occurred many structural breaks in the Brazilian time series over the period under study. There were many episodes of external crises, the exchange rate regime changed from fixed to floating with different degrees of central bank interventions, and the output growth rate varied substantially from 1996 to 2010. However, it is not possible to determine a priori using only economic theory how these facts may have changed the relation between total imports and its determinants.

In addition to changes in macroeconomic environment, it is also possible to note a smooth modification in the imports composition. The high growth rates of the imports of consumption goods in the period is probably related to the process of income distribution started in 1995, which expanded the mass consumption market in the second part of the sample (Brasil, 2003). There is also evidence that the higher the investment growth rate the higher the imports of capital goods due to bottlenecks in the productive sector (Freitas and Dweck, 2010). Indeed, as we will show later, there is a strong correlation between the Brazilian series of investment and total imports.

Finally, one of the innovative aspects of this study is the comparison between two different specifications. The canonical one, which is frequently found in the literature, is a demand equation that includes domestic income and relative price as imports determinants (details in Section 2). The alternative specification substitutes domestic income by household consumption expenditures and gross capital formation, since their behavior in the period boosted imports of goods and services in different ways. The results suggest, in accordance with previous studies, that domestic income is the main determinant of the imports dynamics, but they also show the relevance of the composition of demand. Although the real exchange rate plays a less important role in determining Brazilian imports, it must be considered in the estimations in order to avoid omitted variable bias. It is important to realize that our estimates are not directly comparable with the previous literature and, therefore, should be interpreted with caution.

Since the use of nonlinear models result in the loss of many degrees of freedom, we carried out one-step-ahead predictions to discard overparameterized models that perform bad in the out-of-sample assessment. The alternative specification presented good results in forecasting four and eight quarters ahead. The high speed of adjustment found in the cointegration models might be an explanation for this fact. Nonetheless, the error correction models of the canonical specification were the ones that showed the best results in terms of prediction. In our view this implies that the short-run dynamics of total imports should not be ignored when building scenarios for the current account for one or two years ahead.

The remainder of this paper is organized in five sections. Section 2 presents theoretical and empirical literature. Section 3 describes the data and shows some interesting correlations. The econometric techniques and the main results are presented in Sections 4 and 5, respectively. Finally, in Section 6 we present some concluding remarks.

2. Literature review

Total imports is certainly one of the most studied macroeconomic aggregates of the Brazilian economy. Despite this fact, results regarding its main determinants are mixed: the size of the income elasticity and the impact of the real exchange rate, or relative price of imports, vary according to the period under study, data source, frequency of time series, empirical models and econometric techniques. In the following we discuss the common hypotheses and main results found in the literature.

As a starting point, it should be noted that the hypothesis of imperfect substitution between local and foreign goods is commonplace in the Brazilian macroeconometric literature. Regarding demand and supply equations, they match the consumer and producer theories. The demand for any product depends positively on the domestic income and the local price of substitutes, and negatively on the price of imports (including taxes). The supply varies positively according to the price of imports (added exports incentives, such as producer subsides), and negatively with the price of the same product in the destination country (Carneiro, 2013). These relationships can be represented as

$$M^D = f\left(\overset{+}{Y}, \overset{-}{P_M}, \overset{+}{P}\right), \tag{1}$$

$$M^S = g\left(P_M^*(\overset{+}{1} + \delta), \overset{-}{P^*}\right), \tag{2}$$

where M^D and M^S mean demand and supply quantities; Y is the domestic income; PM is the price of imports in local currency; and PM* is the exports price in international currency; δ is the rate of exports incentives in the rest of the world; and P and $P*$ are the price level of local and international goods and services.

Most papers about the Brazilian total imports estimate the demand equation using a single-equation approach. It is important to clarify, however, that the results will be valid only when the supply elasticity is infinity, i.e. the supply curve is horizontal. This implies assuming the so-called small-country hypothesis, which means that whenever there is an increase in Brazilian demand for imports, the rest of the world will supply goods and services to meet it without an increase in prices. In some papers the empirical models include also a proxy for the level of capacity utilization (variable U with positive expected sign) to deal with restrictions not related to prices (for example, non-tariff barriers).

Studies previous to Portugal (1992) did not take into consideration the integration order of the series (Abreu, 1986; Fachada, 2013). Portugal used cointegration techniques to analyze quarterly Brazilian imports data of total quantum, intermediary goods and capital goods over the period 1975/1976–1988. The results of the long-run equations (cointegration vectors) and short-run dynamics (error correction models) for intermediary goods were in line with previous studies. For capital goods, however, the results were unsatisfactory, showing even a wrong sign (negative) for the income elasticity. In the case of total imports, the income elasticity was estimated in 0.3, smaller than the ones found in the literature (parameter close to unity). The author argued that such result may have been influenced by the "misbehavior" of the imports of capital goods and indicates the need to deal with possible parametric instability, which might be another source of estimation bias.

Castro and Cavalcanti (1998) used annual data over the period 1955–1995 to estimate import and export demand equations. They examined total imports and exports functions as well as their disaggregation by principal end-use category (imports) and aggregate factor (exports). This study was part of the external sector module of the IPEA macroeconometric model (Reis et al., 1999) developed by GAMMA (*Grupo de Análise e Modelagem Macroeconômica*). Therefore, the authors decided to use imports and exports data measured in dollars instead of indexes of quantum since this enabled them to direct compare their predictions with actual trade balance data. Moreover, it was possible to extend sample size using such data. Compared to Portugal (1992), Castro and Cavalcanti (1998) found very different results. The estimates for aggregate imports and capital goods were satisfactory: along with strong evidence of cointegration, the authors obtained plausible coefficients for the long-run vectors and the error correction models. On the other hand, imports of intermediary goods (exclusive oil) did not even cointegrate.

Muinhos et al. (2003) developed specifications of total imports and exports for the Brazilian Central Bank macroeconometric model (structural model with external sector). This paper was the starting point to a more ambitious work, the medium-size macroeconomic model for the Brazilian economy (Muinhos and Alves, 2003). In spite of that, the authors decided to use quarterly indexes of quantum over the period 1988–2001. The empirical exercise did not fully explore the short-run dynamics as the authors chose to work with a partial adjustment equation. Outliers and structural breaks were incorporated into the models using impulse and level dummy variables. The estimated coefficients suggest that the major influence on the quarterly dynamics of imports is the domestic income (impact elasticity of 1.2 and long-run multiplier of 2.7), since the exchange rate coefficients imply an impact elasticity of −0.2 and a long-run multiplier of -0.4.

To our knowledge, Ferreira (1994) is the first paper to delve deep into the investigation of structural breaks. The estimations were performed using quarterly data of total imports (quantum index) from 1973 to 1989. A partial adjustment mechanism was introduced into the estimated regression. Stability tests suggested a structural break in 1981. The estimated coefficients for the structural break model indicated a rise in the price elasticity and a reduction in the income elasticity of imports during the 1980s.

Table 1 summarizes the point estimates for total import in the studies reviewed so far. It is clear that there is a great variability among these estimates. The studies are different in many respects, such as the data set used, the empirical model and the econometric technique, among others. In the presence of many confounding factors, it turns out to be difficult to find a definitive reason for the differences in the main parameter estimates. Nonetheless, there is an indication that different samples are the preeminent factor behind such discrepancies. In fact, the comparison among different studies, as well as for different periods in the same paper, indicates that the income elasticity of imports was higher in the 1990s than in other decades.

Resende (2001) argued that, due to imperfect financial markets, imports were restricted in many periods by the level of international reserves. The author, then, defended the inclusion of a proxy for the capacity to import in estimations of the Brazilian import demand function. Resende applied cointegration and error correction models to estimate total imports and end-use category import equations using quarterly indexes of quantum from 1978 to 1998. Stability tests and structural break regressions were also performed. The results suggest that omitted variable bias occurs when the proxy for capacity to import is not included in models that cover periods before the Real plan. Moreover, the estimations indicate an increase in the income elasticity of imports during the 1990s.

Cavalcanti and Frischtak (2001) was the first paper in the Brazilian literature that employed cointegration tests with endogenous structural break to detect parametric instability. Analyzing the quantum of total imports and end-use category over 1980 and 2000 (quarterly data), the authors found evidence in line with the previous literature that a structural break happened in the long-run relationship among imports and its major determinants during the 1990s. According to the estimations, there was an important increase in the income elasticity of imports.

Azevedo and Portugal (1998) studied the period over 1980 and 1994 using quarterly data of total imports. The authors found clear evidence of parametric instability in the beginning of the 1990s. Structural break regressions showed that domestic income, meanwhile not significant in the first part of the sample (1980s), became an important determinant of imports during the 1990s (with an estimated income elasticity of 2.1). The capacity utilization, however, lost its predictive power at the end of the sample. This fact is related to the non-tariff policies used during the 1980s. According to the authors, the trade liberalization in the 1990s changed the relative importance of those variables in terms of predictive power.

Finally, it is important to cite two other studies that provided important innovations to deal with parametric changes. Silva et al. (2001) studied the period from 1978 to 1999 using the quarterly indexes of quantum of total imports and intermediary goods. Employing artificial neural networks technique, the authors found ruptures in the data behavior in 1989 and 1994. The main results were that the capacity utilization became irrelevant to explain the evolution of total imports and there was an increase in the income elasticity of imports. Morais and Portugal (2005) analyzed the evolution of total imports using an annual Laspeyres index from 1947 to 2002 and a quarterly index of quantum from 1978 and 2002. The authors estimated error correction models with regime changes using markov-switching regressions. The reported annual model has three regimes with changes only in the intercept. The changes are associated with periods of external sector adjustments, periods of closed economy and trade liberalization reforms. The quarterly model in its turn has only two regimes, but it allows for changes in all parameters, including the variance. However, the characterization of each regime depends only on the intercepts.

[1] The dataset is available under request.

Table 1
Point estimates of the relevant long-run elasticities found in the literature regarding the Brazilian total imports.

	Domestic income	Relative price/exchange rate	Capacity utilization	Sample	Source
Portugal (1992)	0.34	−0.91	3.87	1976:1–1988:4	FGV (trade indexes and prices), IBGE (economic activity) and Central Bank (exchange rate)
Ferreira (1994)	2.00/−0.21	−0.11/−1.32	2.21	1973:1–1981:3/1981:4–1989:4	FGV (price indexes, exchange rate, exchange rate premium, import tax revenues, unit value of imports and import quantum index), Central Bank (unit value of imports, import quantum index), IBGE and others (GDP)
Azevedo and Portugal (1998)	NA/2.11	−0.58	4.55/−2.01	1980:1–1989:4/1990:1–1994:4	Banco Central and Secex (import value and price indexes), FGV (price indexes, real exchange rate and capacity utilization), FIESP (capacity utilization), IBGE (quarterly GDP)
Castro and Cavalcanti (1998)	0.73	−2.23	NA	1955–1995 (dados anuais)	Central Bank (import value and price indexes), IMF (real exchange rate) and IBGE (GDP)
Resende (2001)	0.54/3.31/3.31	NA/NA/−1.39	NA	1978:3–1989:4/1990:1–1994:2/1994:3–1998:4	FUNCEX (Import quantum and price indexes), Central Bank (exchange rate), FGV (price indexes) and IBGE (GDP)
Cavalcanti and Frischtak (2001)	0.45/5.53	−0.54/−0.65	NA	1980:1–1991:3/1991:4–2000:4	FUNCEX (Import quantum and price indexes), FGV (price indexes) and IBGE (GDP)
Silva et al. (2001)	−0.01/0.18/1.23	−0.23/−0.91/−1.18	0.05/0.04/0.29	1978:1–1989:3/1989:4–1994:2/1994:3–1999:4	FUNCEX and IPEA
Muinhos et al. (2003)	2.71	−0.44	NA	1988:1–2001:2	Central Bank (imports and exchange rate), IBGE (GDP)
Morais and Portugal (2005)	−0.69/0.82	−0.94/−0.91	2.62	1947–2002 (dados anuais)/1978:1–2002:2 (trimestrais)	FUNCEX (import quantum index), IMF and others (Laspeyre import index) FGV (GDP, exchange rate and capacity utilization), Federal Revenue Service and MDIC (import tax rate), Central Bank (exchange rate) and IBGE (GDP)

Authors' elaboration.
Note: "NA" means "do not apply".

Table 2
Average percent variation of the quantum imported (1996–2010).

End-use category	1996/98	1999/2001	2002/04	2005/07	2008/10
Consumption goods	−1.81	−5.52	−7.54	15.60	18.32
Durable goods	0.88	−8.46	−1.21	35.98	26.31
Non-durable goods	−1.55	−4.47	−8.90	9.18	13.92
Raw materials and intermediary goods	3.60	2.31	0.01	10.37	8.63
Capital goods	19.74	3.55	−4.53	38.29	29.37
Fuel and lubricants	−0.36	0.78	1.98	5.38	3.83
Total	1.04	0.98	0.50	8.14	7.04

Source: FUNCEX.
Growth rates were not weighed by the share of their sub-groups and, consequently, do not represent the percentual variation in the index of volume.

Table 3
Average share of principal end-use groups in total imports (1996–2010).

End-use category	1996/98	1999/2001	2002/04	2005/07	2008/10
Consumption goods	18.51	13.60	11.61	12.63	15.72
Durable goods	9.06	6.27	5.13	6.28	8.89
Non-durable goods	9.45	7.32	6.48	6.35	6.83
Raw materials and intermediary goods	45.96	49.77	52.16	50.06	47.00
Capital goods	26.34	26.19	21.80	20.82	22.20
Fuel and lubricants	9.19	10.44	14.43	16.50	15.08
Total	100.0	100.0	100.0	100.0	100.0

Source: FUNCEX.

3. Data[1]

In this section we present the dataset used in the empirical exercises along with some of their interesting correlations. As discussed before, the main determinants of imports in the canonical specification are the domestic income and the relative price of imports. In the studies reviewed, the domestic income variable is proxied by the Gross Domestic Product (GDP). It is important to highlight that household consumption expenditure is the major component of GDP from the demand side, representing nearly 60% of the aggregate. Two important developments in its recent dynamics are worth mentioning. First, the expansion of credit markets allowed previously credit-constrained consumers to increase their consumption expenditures. Second, consumption raised significantly in the poor deciles of the income distribution due to reduction in income inequality. It is not clear, however, how these changes may have impacted imports of consumption goods since household expenditures are concentrated in domestic goods and services.

The gross capital formation in turn is responsible for nearly 17% of GDP on average in our sample period and it has a stronger correlation with imports than does household consumption expenditures. Furthermore, anytime large companies decide to implement their investment plans, what generally happens when domestic income is growing, imports rise faster as domestic supply of capital goods is very limited.

Table 2 shows average percent variation in the quantum imported by principal end-use category over the sample period for consecutive three-year intervals. Imports of consumption goods reduced until 2002–2004, but they have grown very fast starting in 2005–2007. The growth rate of imports of durable goods in particular was much higher than the growth rate of total imports. Imports of capital goods in turn grew fast during the three-year period of 1996–1998 and from 2005–2007 onwards. In 2005–2007 and 2008–2010, the growth rates of imports of capital and durable goods were very similar. Also, it is not surprising that the growth rates of raw materials and intermediary goods increased after 2005–2007 (see Table 2), given the complementarity with capital goods. Table 3 presents the share of each group in total imports (in US$), showing that the capital goods weight is much higher than the durable consumer goods (23.1% and 7.4%, respectively). Its share is only lower than that for raw materials and intermediary goods (48.8% on average).

Given the weight of capital goods (and raw materials and intermediary goods) in total imports, it is intuitive to foresee a strong association between the gross capital formation and total imports. In addition, the domestic supply of

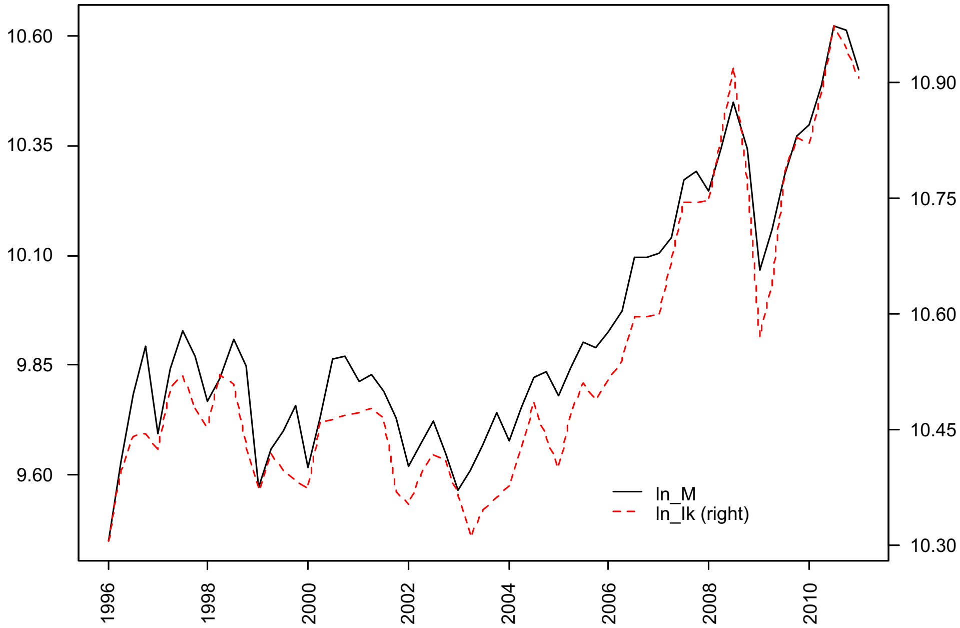

Fig. 1. Total imports and gross capital formation (1996:1–2011:1).
Source: IBGE (quarterly national accounts).

this type of good is limited and, therefore, imports increase (decrease) every time investment rises (declines). Fig. 1 shows the strong association between these variables using chained quarterly national account data (reference 1995, series in log).

The previous discussion led us to try two empirical models:

$$M^D = f\left(\overset{+}{Y}, \overset{-}{\theta}\right), \tag{3}$$

$$M = f\left(\overset{+}{I_K}, \overset{+}{C_F}, \overset{-}{\theta}\right), \tag{4}$$

where Y is domestic income; θ is the exchange rate adjusted for import tax (a measure of the relative price of imports); I_K is the gross capital formation; and C_H is the household consumption expenditure.

Fig. 2 presents the time series of GDP and household consumption expenditure. These data is also measured as chained values in 1995 price (R$ million, in log). We seasonally adjust all quarterly national account series using X12 multiplicative method before proceeding the estimations, except in the state-space models with stochastic seasonality (see Section 3).

The exchange rate time series was computed using nominal exchange rate R$/US$ (average between purchases and sales) from Brazilian Central Bank (BCB) and the wholesale price indexes from International Monetary Fund (IMF) for Brazil and the US. The average import tax rate was obtained as the ratio of total import duties collected (available at BCB website) and total nominal imports from quarterly national accounts. Fig. 3 shows the evolution of the adjusted real exchange rate.

Finally, we present the results of three distincts unit root tests (DF-GLS, Zivot-Andrews and Lee-Strazicich) in Appendix I. They indicate that all series have a stochastic trend or are very persistent, justifying the empirical exercises based on the cointegration approach.

4. Methods

We estimate equations for the total imports using the cointegration approach. The empirical model comprises the long-run equilibrium and the error correction dynamics. It is worth mentioning that the parametric stability of the

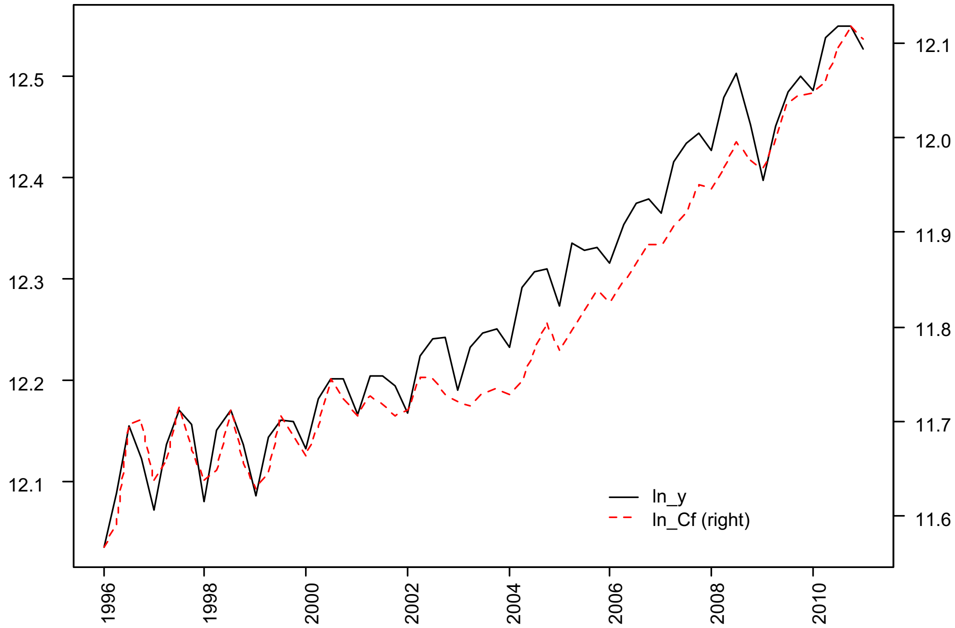

Fig. 2. GDP and household consumption expenditure (1996:1–2011:1)
Source: IBGE (quarterly national accounts).

Fig. 3. Adjusted real exchange rate (1996:1–2011:1). Adjusted real exchange rate is given by $\theta = (1+\tau)\cdot\varepsilon\cdot P_{US}/P_{BR}$, where τ is the average import tax rate, ε is the nominal exchange rate and P is the wholesale price index (*US* and *BR* means United States and Brazil, respectively).
Source: BCB, IMF and IBGE.

imports equations might be violated in some macroeconomic contexts such as in periods of changing trade balance composition or exchange rate regime shifts, among other reasons. That being said, it should be noted that we cannot determine a priori what is the best way to address such possible nonlinearities. In the present context, this is in itself an empirical matter. Hence, to address such uncertainty, we use three distinct econometric techniques. For easy of exposition, the following presentation is based on Eq. (3) (the imports demand equation), but the same approaches were employed for the alternative model of Eq. (4).

4.1. Cointegration and structural break

A variable x_t is said to be integrated of order d ($x_t \sim I(d)$) if it becomes stationary after taking differences d times. When exist a linear combination among $I(d)$ variables that is $I(d-b)$ for $b>0$, they are said to be cointegrated of order d, b; i.e. if $x_t = (x_{1t}, x_{2t}, \ldots, x_{nt})$' and $x_{1t}, \ldots, x_{nt} \sim \mathrm{I}(d)$ and $\beta x_t = \beta_1 x_{1t} + \beta_2 x_{2t} + \ldots + \beta_n x_{nt} \sim I(d-b)$, then $x_t \sim \mathrm{CI}(d, b)$ (Granger, 1981).

The logarithms of imports, GDP and real exchange rate have stochastic trends, i.e., they are $I(1)$. If there is a common trend among them, the vector composed by these variables will be cointegrated or $C(1, 1)$.[2]

To help the interpretation of the results we weigh the cointegration vector by the variable ln$M$. Then, the long-run equation with a constant as the only deterministic term is given by:

$$\ln M_t = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 \ln \theta_t + \varepsilon_t, \tag{5}$$

where $\varepsilon_t \sim \mathrm{I}(0)$ represents deviations from the equilibrium relationship; α_0 is the level of the function; α_1 is the long-run income elasticity; and α_2 is the elasticity with regard to the real exchange rate.

Any cointegration vector can be stated in an error correction form (Granger representation theorem). This representation displays both the long-run relation and the short-run dynamics. But to work in a single-equation setup, that is, with error correction models (ECM's) rather than the full system of equations, we need to assume that the right-hand side variables in Eq. (5) are weakly exogenous. In practice, this means that, in the face of a disturbance in the long-run relation, the equilibrium must be restored through adjustments in total imports only. This assumption allows us to make inference using the single equation approach. The ECM model is as follows:

$$\begin{aligned}\Delta \ln M_t = \delta_0 + \sum_{i=1}^{p-1} \delta_{1i} \Delta \ln M_{t-i} + \sum_{j=1}^{p-1} \delta_{2i} \Delta \ln Y_{t-j} \\ + \sum_{k=1}^{p-1} \delta_{3k} \Delta \ln M_{t-k} - \lambda_M (\ln M_{t-1} - \alpha_0 - \alpha_1 \ln Y_{t-1} - \alpha_2 \ln \theta_{t-1}) + u_t,\end{aligned} \tag{6}$$

where δ's are the impact coefficients (short-run elasticities), $\lambda_M > 0$ measures the speed of adjustment of ln$M$ toward the equilibrium, and $u_t \sim$ i.i.d. $(0, \sigma_u{}^2)$ are the residuals.

Engle and Granger (1987) (EG) proposed a procedure in two stages to estimate equations (5) and (6). First, the cointegration vector is estimated by ordinary least squares (OLS).[3] Then the residual series is tested for unit root using the ADF test (with adjusted critical values). If they are stationary, we conclude that the variables are cointegrated, and the residuals can be used in the second stage as a *proxy* for the disequilibrium term of the ECM.

Gregory and Hansen (1996) (GH) developed a way to deal with unknown structural breaks in the cointegration vector. The following *dummy* variable is used in the empirical models:

$$\varphi_{ty} = \begin{cases} 0, & \text{if} \quad t \le \gamma T \\ 1, & \text{if} \quad t > \gamma T \end{cases} \tag{7}$$

where $T_B = \gamma T$ is the period when the break occurs. For the purpose of this paper, the structural breaks can be of two types.

[2] With n variables, it is possible to have up to n-1 cointegration vectors. Hence, it is possible in this case to test and provide maximum likelihood estimates for two vectors following Johansen (1988) and Johansen and Juselius (1990). However, given the studies reviewed in section 2, we can argue that there is no theoretical reason to proceed in that direction using the Brazilian data. Given our strict empirical motivation, we also considered applying the Johansen methodology, but the results presented in Appendix II indicate that the hypotheses needed to apply it are violated by our dataset.

[3] Regarding endogeneity issues, it should be noted that if the variables are cointegrated they are simultaneously determined by definition. This will not invalidate the results as the bias disappears asymptotically. In finite samples we can observe simultaneity bias but the convergence rate is greater than in regressions with stationary variables. See Stock (1987) for an exposition of the super consistency of the OLS estimator with cointegrated variables.

First, a change in the level of the long-run relationship:

$$\ln M_t = \mu_1 + \mu_2\varphi_{ty} + \alpha_1 \ln Y_t + \alpha_2 + \ln\theta_t + \varepsilon_t, \tag{8}$$

where the intercept of the above equation changes from μ_1 to $\mu_1 + \mu_2$ after T_B (GH also consider a similar specification with the inclusion of a deterministic trend).

Second, the full break model:

$$\ln M_t = \mu_1 + \mu_2\varphi_{ty} + \eta_1 \ln Y_t + \eta_2\varphi_{ty} \ln Y_t + \psi_1 \ln\theta_t + \psi_2\varphi_{ty} + \ln\theta_t + \varepsilon_t, \tag{9}$$

where, besides the intercept, the elasticities change from η_1 and ψ_1 to $\eta_1 + \eta_2$ and $\psi_1 + \psi_2$, respectively.

The test follows the EG procedure with the difference that it is necessary to take into account the possibility of a structural break. The procedure consists of choosing the value of γ for which the test statistic is minimized, as the chance of rejecting the null hypothesis of cointegration is maximized in this case. In the same way as for the unit root tests with breaks, the critical values of the test statistic (simulated by the authors) take into account the selection algorithm. If the null hypothesis is rejected, it is possible to estimate the cointegration vector by OLS imposing the break in the period indicated by the test. As in the EG procedure, the residuals obtained in the first step can be used as a *proxy* for the disequilibrium term in the ECM model.

4.2. Markov-switching models

Markov-switching models (MS) can be seen as a generalization of structural break models in which there are regime shifts among a given number of states instead of an once-and-for-all change (Hamilton, 1989). This flexibility in dealing with nonlinearities is definitely an advantage over the structural break models. The downside is that it requires more complex estimation techniques as will be shown below.

The main distinction between regressions and MS models is the twofold stochastic feature of the latter. A good exposition of this class of models is found in Krolzig (1998). Along with the conditional data generation process, it is necessary to define the unobservable state process. A gaussian distribution is assumed for the first, meanwhile the second is assumed to follow a markov chain.

An autoregressive MS model of order p, cointegration rank r and M regimes is used in the following exposition. Using Krolzig's (1997, Chap. 13) terminology we can represent this model as MSCI(M,r)-VAR(p) or MS(M)-VECM(p-1).

The general representation is given by:

$$\begin{aligned}\Delta \ln M_t = {} & \delta_0(s_t) + \sum\nolimits_{i=1}^{p-1} \delta_{1i}(s_t)\Delta \ln M_{t-i} + \sum\nolimits_{j=1}^{p-1} \delta_{2j}(s_t)\Delta \ln Y_{t-j} \\ & + \sum\nolimits_{k=1}^{p-1} \delta_{3k}(s_t)\delta \ln \theta_{t-k} - \lambda_M(s_t)(\ln M_{t-1} - \alpha_0(s_t) - \alpha_1(s_t)\ln Y_{t-1} - \alpha_2(s_t)\ln\theta_{t-1}) + u_t,\end{aligned} \tag{10}$$

where $u_t \sim n.i.d.\ (0, \sigma_u^2(s_t))$ and $s_t \in \{1, \ldots, M\}$.

This unrestricted specification is a MSIAH(M)-VECM(p-1), where I, A and H mean that the intercept, slopes and covariance matrix are dependent of state s_t. Assuming a time-discrete homogeneous Markov chain for the unobservable variable s_t, we have:

$$\begin{aligned} & Pr(s_t = j|s_{t-1} = i, s_{t-2} = k, \ldots) = Pr(s_t = j|s_{t-1} = i) = p_{ij}, \\ & \sum_{j=1}^{M} p_{ij} = 1 \quad \forall \quad i, j \in \{1, \ldots, M\} \end{aligned} \tag{11}$$

It should be noted that the homogeneity assumption is necessary in order to obtain time invariant p_{ij}. Working with M regimes leads to $M \times M$ probabilities whose P transition matrix is given by:

$$P = \begin{pmatrix} p_{11} & p_{12} & \cdots & p_{1M} \\ p_{21} & p_{22} & \cdots & p_{2M} \\ \vdots & \vdots & \ddots & \vdots \\ p_{M1} & p_{M2} & \cdots & p_{MM} \end{pmatrix} \quad (12)$$

It is also assumed that the Markov chain is ergodic and irreducible. The first assumption assures that there is a stationary distribution for the regimes. The second one implies a strictly positive unconditional probability vector (Hamilton, 1994, Chap. 22).

The likelihood functions of these models are recursive, as the optimum inference in t depends on the optimum inference $t-1$, and traditional maximum-likelihood estimation is not possible. However, maximum-likelihood estimation can be performed using the EM algorithm together with the BHLK (Baum–Lindgren–Hamilton–Kim) filter. For more details, see Krolzig (1997, Chap. 5 and 6).

Regarding the MS-VECM, we can estimate the error correction model analogously to the MS regressions with stationary variables. Nonetheless, the properties of the estimator using integrated series are unknown. Besides reporting MS estimates of the long-run static relationship, we also estimate the MS-VECM model using the EG or GH regression residuals as *proxy* for the disequilibrium term.

Model selection in this approach is not trivial, however. As an example, tests to determine the number of regimes have nonstandard distributions. The general rule is to consider a small number of states in order to preserve degrees of freedom. In our estimations we used only two regimes and decided which parameters would be allowed to change using information criteria. We also performed out-of-sample forecasts to avoid overparameterization.

4.3. Stochastic parameters

Stochastic parameter regression models can be estimated using a state-space representation and the Kalman filter. Since Harvey's (1989) seminal contribution, this approach has been used to model univariate stationary time series, but its application to cointegrated multivariate cases has been increasing recently.

In the general specification we have:

$$\ln M_t = \alpha_{0,t} + \alpha_{1,t} \ln Y_t + \alpha_{2,t} \ln \theta_t + \chi_{1,t} + \varepsilon_t, \quad (13)$$

where $\varepsilon_t \sim n.i.d.\ (0, \sigma_\varepsilon^2)$ is the irregular component.

The seasonal component $\chi_{1,t}$ can be modeled as deterministic or stochastic and its representation for quarterly time series is given by:

$$\begin{aligned} \chi_{1,t+1} &= -\chi_{1,t} - \chi_{2,t} - \chi_{3,t} + \omega_t, \\ \chi_{2,t+1} &= \chi_{1,t}, \\ \chi_{3,t+1} &= \chi_{2,t}, \end{aligned} \quad (14)$$

where the disturbances $\omega_t \sim n.i.d.\ (0, \sigma_\omega^2)$ make the seasonal component stochastic.

We also allow the other parameters (α_0, α_1 and α_2) to change over time. Then, the level and slopes follow a random-walk representation:

$$\begin{aligned} \alpha_{0,t+1} &= \alpha_{0,t} + \xi_t, \\ \alpha_{1,t+1} &= \alpha_{1,t} + \zeta_t, \\ \alpha_{2,t+1} &= \alpha_{2,t} + \varsigma_t, \end{aligned} \quad (15)$$

where $\xi_t \sim n.i.d.\ (0, \sigma_\xi^2)$, $\zeta_t \sim n.i.d.\ (0, \sigma_\zeta^2)$ and $\varsigma_t \sim n.i.d.\ (0, \sigma_\varsigma^2)$.

The parameters above can be estimated recursively using the Kalman Filter and the maximum-likelihood estimator. A short presentation of the procedures is done by Commandeur and Koopman (2007, Chap. 8), while the more detailed discussion can be found in Hamilton (1994, Chap. 13) and Harvey (2006).

It is important to highlight that the observation (ε_t), seasonal (ω_t), level (ξ_t), and slopes (ζ_t and ς_t) disturbances must be serially and mutually independent in order to obtain valid estimates. Thus, besides implementing the normality and homoscedastic tests, we checked the independence of disturbances for all estimated models given by (13). When the maximum likelihood estimate of the variance of any component could not be distinguished from zero, we tried more parsimonious specifications with deterministic components. We also used information criteria and out-of-sample performance tests for model selection.

5. Empirical results

In this section we present the results for the estimations of both the canonical and alternative models. In the canonical specification, total imports depend on GDP and real exchange rate. The alternative specification in turn explores the strong correlations among total imports, gross capital formation and household consumption expenditures. In order to make the exposition easier, we report the results according to estimation technique. At the end of the section we carried out the out-of-sample evaluation.

5.1. OLS estimates of cointegration and short-run adjustment

According to the cointegration tests, we cannot reject the hypothesis that there is a long-run relationship among total imports, exchange rate and GDP (canonical specification) and an even stronger relationship among total imports, exchange rate, household consumption expenditures and gross capital formation (alternative specification). The estimated long-run relationships for the alternative specification show remarkable performance in projections, providing additional evidence against spurious regression problems. This result is robust to the use of different techniques.[4]

In what follows, for the EG specifications, C denotes the presence of a constant, and C/T stands for constant and trend, whereas, for the GH specifications, C denotes structural break in the intercept, C/T indicates the presence of a deterministic trend but structural break only in the intercept, whereas FB generalizes with a trend break. The null hypothesis of non-cointegration was not rejected only for the EG-C specification of the canonical model. In all other cases we found evidence of cointegration. Table 4 summarizes the results of the cointegration tests and presents the estimated coefficients for selected models.[5]

For the canonical model, the GH-FB specification reveal a structural break in the long-run relationship in the second quarter of 2002, with a rise in the level of the function and in the long-run income elasticity. The income elasticity increased from 2.2 to 2.9 after 2002. It should be noted that those values are not much different from the ones encountered in the previous literature (Table 1). The elasticity with regard to the real exchange rate is in turn close to zero after 2002, making imports less sensitive to changes in this variable.

In the alternative model, even though the elasticity with regard the household consumption expenditures is smaller than the elasticity of gross capital formation, the former variable is important to explain the quarterly dynamics of imports once its exclusion leads to evidence of non-cointegration. This fact may be explained by the significant weight of the household consumption expenditures in GDP.

We report the results of the GH-C and GH-FB models. Although the GH-C test indicates a structural break in the third quarter of 2007, the estimated coefficients are not significantly different from the ones estimated without the break, being in the order of −0.12 (exchange rate), 0.23 (household consumption expenditures) and 1.20 (gross capital formation). When we allow for breaks in both the intercept and the elasticities (GH-FB specification), there is evidence

[4] For the alternative model, the static long-run vector performed very well in the out-of-sample assessment. Meanwhile, the full specification including the short-run dynamics did not predict well. This may be due to overfitting. But, on the other hand, the performance of the error correction representation of the canonical model was among the best. Hence, one possible additional explanation lies in the different speeds of adjustments. The low speed of adjustment in the canonical model renders the short-run dynamics important for predicting total imports, whereas in the alternative specification this coefficient is relatively high and the dynamic part of the model turns out to be irrelevant for improving forecasts. This point will be further explored ahead.

[5] The results not reported are available under request.

Table 4
Cointegration tests and OLS estimates of long-run elasticities.

	Canonical				Alternative			
	EG-C		GH-FB		GH-C		GH-FB	
	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat
C	−1.00	−0.61	−16.72	−3.38	−7.60	−8.45	−13.82	−6.34
DU(2002:2)	–	–	−9.22	−1.39	–	–	–	–
DU(2002:3)	–	–	–	–	–	–	10.02	3.73
DU(2007:3)	–	–	–	–	−0.05	−2.30	–	–
lnθ	−0.53	−7.51	−0.52	−5.47	−0.12	−3.61	−0.10	−1.83
DU(2002:2)*lnθ	–	–	0.56	3.41	–	–	−0.16	−1.98
ln$Y$	0.89	6.69	2.19	5.36	–	–	–	–
DU(2002:3)*ln$Y$	–	–	0.72	1.33	–	–	–	–
lnIk	–	–	–	–	1.28	15.69	1.54	10.04
DU(2002:3)*lnIk	–	–	–	–	–	–	−0.49	−2.70
lnCf	–	–	–	–	0.35	4.51	0.64	2.84
DU(2002:3)*lnCf	–	–	–	–	–	–	−0.42	−1.46
Cointegration: test statistic		−2.93		−5.98***		−6.06***		−6.96***
Durbin–Watson		0.44		1.35		1.45		1.86

Source: Authors' elaboration.
*Cointegrate at 10%.
**Cointegrate at 5%.
*** Cointegrate at 1%.

that the effect of the real exchange rate might have actually increased after the second quarter of 2002, changing from −0.10 to −0.26. The values of the other elasticities for the periods after the break are in line with the ones obtained for the other specifications.

In Table 5 we present the results for the ECM's that include the short-run dynamics using the residuals from the estimated cointegrated vector. We follow a general-to-specific procedure, starting with four lags of each variable. The most striking difference between the estimates from the canonical and alternative models is related to the estimated coefficients for the speeds of adjustment. The speed of adjustment in the canonical model is estimated to be 0.41, which is nearly 50% of those obtained using the alternative specification (0.83 and 0.72). Therefore, given a deviation from the long-run relationship, the return to equilibrium via the error correction mechanism occurs much faster in the alternative case.

We also note that the influence of the exchange rate seems to be only transitory in the canonical model, as its long-run multiplier is close to zero. Moreover, the immediate impact of income is more relevant when compared to the exchange rate, the same happening in the alternative specification. This conclusion can be drawn from the fact that the short-run coefficients of the household consumption expenditure and the gross capital formation are higher than the exchange rate one. Also, there are more significant lags of those variables in the ECM's.

5.2. *Estimates of the Markov-switching models*

The static Markov-switching regressions, which represent the long-run relationships, were estimated allowing for changes in the intercept (MSI), in the intercept and variance (MSIH), in the intercept and the stochastic regressors (MSIA) and in all terms (MSIAH). The same procedure was adopted for the error correction representation when we estimate MS-VECMs to assess possible non-linearities in the short-run relationships. We present the results for models estimated using only 2 regimes, since the results for models with 3 regimes were hard to interpret and the inclusion of a third regime does not improve prediction performance. Table 6 shows the main results of the static MS regressions, including point estimates of the canonical and alternative models as well as the transition probabilities and the regime dates. It is worth noting that it was always possible to reject the null hypothesis of linearity. Moreover, the selected specifications indicate changes in the parameters in 2002. However, unlike the Gregory-Hansen specifications, at some point in time the systems return to the previous state of nature.

Table 5
Short-run dynamics, ECM's.

	Canonical		Alternative			
	ECM GH-FB		ECM GH-C		ECM GH-FB	
	Coef.	*p*-value	Coef.	*p*-value	Coef.	*p*-value
C	−0.02	0.01	0.00	0.97	0.00	0.69
ε_{t-1}	−0.41	0.00	−0.72	0.00	−0.83	0.00
$\Delta\ln M_{t-2}$	0.19	0.03	0.29	0.01	0.16	0.01
$\Delta\ln M_{t-3}$	–	–	0.22	0.01	0.16	0.01
$\Delta\ln\theta_t$	−0.14	0.05	−0.16	0.01	−0.17	0.00
$\Delta\ln Y_t$	2.25	0.00	–	–	–	–
$\Delta\ln Y_{t-1}$	1.54	0.00	–	–	–	–
$\Delta\ln Ik_t$	–	–	0.87	0.00	0.83	0.00
$\Delta\ln Ik_{t-2}$	–	–	−0.33	0.04	–	–
$\Delta\ln Cf_t$	–	–	0.64	0.07	0.76	0.01
$\Delta\ln Cf_{t-1}$	–	–	0.54	0.08	0.49	0.06
$\Delta\ln Cf_{t-4}$	–	–	−0.71	0.01	−0.67	0.01
Adjusted R^2	0.71	–	0.82	–	0.85	–
Breusch–Godfrey	1.04	0.36	0.29	0.75	0.59	0.56
White	0.42	0.98	0.31	0.97	0.20	1.00
ARCH	0.55	0.58	0.16	0.85	0.43	0.65
Ramsey RESET	1.85	0.18	0.38	0.54	2.59	0.11
Jarque–Bera	1.16	0.56	3.19	0.20	0.57	0.75

Source: Authors' elaboration.

Table 6
Long-run relationships estimated by Markov-Switching.

	Canonial		Alternative			
	MSIAH (2)		MSI (2)		MSIA (2)	
	Regime 0	Regime 1	Regime 0	Regime 1	Regime 0	Regime 1
C	−5.56***	−2.50*	−7.22***	−7.284***	−3.17***	−9.61***
$\ln\theta$	−0.31***	−0.43***	−0.11***	–	−0.26***	0.01
$\ln Y$	1.25***	1.02***	–	–	–	–
$\ln Ik$	–	–	1.31***	–	1.05***	1.35***
$\ln Cf$	–	–	0.28***	–	0.17**	0.45**
Standard Deviation	0.02***	0.06***	0.02***	–	0.02***	–
p(0|0)	0.87***	–	0.93***	–	0.86***	–
p(0|1)	0.02	–	0.35**	–	0.22	–
Dates (regime 0)	1996(1)–1996(1); 2002(4)–2005(4)		1996(2)–2002(3); 2003(2)–2007(4); 2008(4)–2009(2)		1996(4)–2000(2); 2002(1)–2005(3); 2007(4)–2008(3); 2009(3)–2010(1)	
LR linearity test	42.73***		10.37*		18.58***	
Non-normality	0.31		39.35		0.20	
Autocorrelation: Portmanteau	28.89***		17.27		11.96	
ARCH F	7.09**		0.004		0.04	

Source: Authors' elaboration.
* Reject H0 at 10%.
** Reject H0 at 5%.
*** Reject H0 at 1%.

The best representation for the canonical specification seems to be the MSIAH. The estimation suggests that there is a persistent regime (regime 1) that occurred from the second quarter of 1996 to the third quarter of 2002 with average duration of 21.5 quarters. In 2002, there is a transition to a state with a lower level, variance and exchange-rate elasticity, but with a higher income elasticity (regime 0). However, three years later the previous regime returns and goes from the first quarter of 2006 onwards. The similarity with the GH-FB is restricted to the parametric change in 2002 and the increase in the income elasticity. The difference in the point estimates is quite large, being the OLS one more consistent with the literature (Table 1). There is also some indication of autocorrelation and conditional heteroskedasticity. Moreover, some properties of these models are unknown for integrated variables.

For the alternative model, the MSI and MSIA specification were chosen. In this case we do not observe any problems in the regression diagnostic tests, and the long-run elasticities are similar to the ones estimated in the GH cointegration vectors. In the MSI specification, that allows changes only in the intercept, the persistent regime (regime 0) has a slightly higher level than in the other regime. The average duration of this regime is 16 quarters and the economy would have switched to regime 1 only in three periods (from 2002 to 2003, between the first and third quarters of 2008 and from the third quarter of 2009 onwards). It is clear that there are strong similarities between the MSI specification and the GH-C estimates. The main difference refers to the dates in which changes in regimes are observed since the GH specification is restricted to an once-and-for-all change to "regime 1" in the third quarter of 2007. The MSIA estimates are in turn similar to the results of the GH-FB regressions with structural breaks in the third quarter of 2002. The elasticities estimates are again very similar and the main difference is in the transition dates. The estimates indicate regime 0 has also occurred from the start of 2002 through the third quarter of 2005, then from the end of 2007 to the third quarter of 2008, and finally from the third quarter of 2009 onwards.

We also estimated MS specifications for the short-run relationships using the estimated residuals from the OLS long-run relationships. Even with only one lag of each variable, allowing regime switches brings about important improvements in the degree of adjustment of the error correction representation. Considering up to four lags, we noticed that the best predictions occur in specifications with two and one lags in the canonical and alternative model, respectively.

Table 7 presents the results for the selected specifications: the MSIAH for the canonical model and the MSIA for the alternative one. The characterization of the regimes is similar in both models. In regime 1 the equilibrium growth rate (i.e. the constant of the error correction representation) is positive and the adjustment to short-run deviations is very fast. In regime 0, the constant is negative and the deviations are corrected slowly. Moreover, in the MSIAH of the canonical model the variance is higher under regime 0 and the short-run multipliers are similar to the estimates of the linear ECM's. Finally, we should note that the MSIA specification shows an absolute coefficient for the disequilibrium term higher than one in regime 1, indicating that there is no error correction mechanism.[6] This result, however, might be related to the fact that there is a very fast transition between regimes implying that the system does not stay in the unstable state for a long period.

5.3. Stochastic parameters regressions estimated using the Kalman filter

In this section we report the results of the state-space models estimated via the Kalman filter. As discussed in Section 4.3, this methodology allows the parameters to change at any point in time. We started estimating the canonical and alternative models allowing all parameters to be stochastic. When we could not distinguish the estimated variance from zero, we fixed that parameter. All models that do not converge during the estimations were discarded. Table 8 shows the results of the selected models, presenting their point estimates in the final state when the parameter is stochastic.

In the canonical specification, the estimated variance of the elasticities and slope were zero (or close to) and these parameters were then fixed in the estimations. Even though the slope was not significant, if we drop this term the models show misspecification problems once the null hypothesis of residual independence is rejected using the Box–Ljung statistic. The automatic identification of outliers and/or structural changes showed a positive break in the fourth quarter of 1996 and a negative one in the first quarter of 2009. The estimated elasticities were similar to the ones obtained when using OLS. It is important to note that even allowing the income elasticity to vary (its estimated variance was equal to zero) the point estimate in the final state was 2.83, very close to the OLS estimation for the period after 2002.

[6] We gratefully acknowledge an anonymous referee for bringing this point to our attention.

Table 7
Short-run relationships estimated by Markov-Switching, MS-VECM's.

	Canonical GH-FB/MSIAH		Alternative GH-C/MSIA	
	Regime 0	Regime 1	Regime 0	Regime 1
C	−0.04***	0.03***	−0.02***	0.03***
ε_{t-1}	−0.10	−0.84***	−0.39***	−1.35***
$\Delta \ln M_{t-1}$	0.07	−0.25**	−0.27**	0.37***
$\Delta \ln M_{t-2}$	0.30***	0.27***	–	–
$\Delta \ln \theta_t$	−0.20***	0.06	−0.16***	−0.02
$\Delta \ln \theta_{t-1}$	−0.31***	0.01	−0.11***	0.50***
$\Delta \ln \theta_{t-2}$	0.09	−0.27***	–	–
$\Delta \ln Y_t$	1.61	0.67	–	–
$\Delta \ln Y_{t-1}$	1.29	0.73	–	–
$\Delta \ln Y_{t-2}$	0.18	−0.06	–	–
$\Delta \ln Ik_t$	–	–	1.21***	1.37***
$\Delta \ln Ik_{t-1}$	–	–	−0.33**	−0.74***
$\Delta \ln Ik_{t-2}$	–	–	–	–
$\Delta \ln Cf_t$	–	–	0.78***	0.14
$\Delta \ln Cf_{t-1}$	–	–	1.66***	−0.54**
$\Delta \ln Cf_{t-2}$	–	–	–	–
Standard Deviation	0.02***	–	0.01***	0.01***
p(0|0)	0.58***	–	–	0.56***
p(0|1)	0.77***	–	–	050***
LR linearity test	40.28***	–	37.23***	–
Non-normality	14.12*	–	4.40	–
Autocorrelation: Portmanteau	7.35	–	8.83	–
ARCH F	0.54	–	0.02	–

Source: Authors' elaboration.
* Reject H0 at 10%.
** Reject H0 at 5%.
*** Reject H0 at 1%.

However, when this parameter was fixed from the beginning, its point estimate (2.24) was similar to the OLS before the break. The exchange-rate elasticity, in its turn, was again less important to explain the quarterly behavior of imports, what also happened for the alternative specification.

In the alternative case, all models estimated with stochastic elasticities did not converge. Then both models presented in Table 8 were obtained fixing those parameters. The only difference between the two models is a slope term. The coefficients of household consumption were 0.75 and 0.84, and that of the gross capital formation was 0.95 in both cases. Thus, the estimates for the elasticities using the Kalman filter were much closer to one another when compared to those obtained using other methods.

5.4. Out-of-sample prediction

Before proceeding, we highlight some of the preliminary conclusions: (i) there is strong evidence of nonlinearities in the long-run relationship; (ii) while in the canonical model the instability has affected both the level and the elasticities of the functions, in the alternative model it is not clear whether or not the same has happened; (iii) the correction of short-run disequilibria seems to occur much faster in the alternative model when compared to the canonical one; and (iv) when we allow for regime switches in the error correction model, there is a clear improvement in the in-sample fit.

As it is always possible to improve the in-sample fit of a model by increasing the number of variables, lags or non-linearities, we needed better criteria to choose among different specifications. To avoid overparameterization, we assessed the models in their out-of-sample performance, which also helps to avoid choosing spurious regressions since

Table 8
Long-run relationships estimated by Kalman filter.

	Canonical		Alternative	
	EE-1 Stochastic level, slope and elasticities; fixed seasonality	EE-2 Stochastic level; fixed slope and elasticities; no seasonality	EE-3 Stochastic level, slope and seasonality; fixed elasticities	EE-4 Stochastic level and seasonality; fixed elasticities; no slope
Level	−25.27	−17.82	−10.18	−9.09
Slope	−0.01	−0.01	0.00	–
Outlier (date)	–	–	2008:4	–
Level break (date)	1996:4	1996:4; 2009:1	–	–
lnθ	−0.20	−0.20	−0.17	−0.17
ln$Y$	2.83	2.24	–	–
lnIk	–	–	0.95	0.95
lnCf	–	–	0.84	0.75
Seasonality	20.3***	–	2.35	2.15
Autocorrelation: Box–Ljung	3.84	9.06	6.51	6.47
Heteroscedasticity	0.77	1.35	1.27	1.12
Non-normality	0.16	3.77	1.94	1.48

Source: Authors' elaboration.
*Reject H0 at 10%.
**Reject H0 at 5%.
*** Reject H0 at 1%.

we know they have poor predicting performance.[7] Additionally, we can get a better sense of how helpful these models can be for constructing conditional scenarios in predicting exercises.

Predictions were made for the horizon of one year from the second quarter of 2010 to the first quarter of 2011. All predictions are on-step-ahead predictions, i.e., the observed values update the equation at every new prediction. The models were compared according to well-known metrics in the time series literature. The Mean Average Percentage Error (MAPE) equally penalizes positive and negative errors. The Mean Square Error (MSE), as well as its square root (RMSE), gives more weight to bigger errors. We also report the usual MSE decomposition into bias, variance and covariance. The bias (variance) proportion shows by how much the mean (variance) prediction diverges from the series mean (variance) in the prediction horizon. The U-Theil coefficient ranges between 1 and 0, where the lower its value the higher the model predictive power.

Table 9 shows the out-of-sample statistics for the best specifications for the long-run vectors of the canonical and alternative models. We note that both the MAPE and the RMSE are much smaller in the alternative case, which also have smaller proportion of non-random errors as the covariance proportion is much higher. The difference in performance between the GH-C and MSI in the alternative model is subtle because, even though the MAPE is slightly higher in the GH-C, this regression shows a better performance by the MSE criteria. Moreover, the proportion of random prediction errors is higher in this specification, reinforcing that its good prediction performance is not due to some statistical artifact. Therefore, for the long-run relations, the predictive power of the alternative model is better.

However, we should also assess the short-run adjustment dynamics. Table 10 presents the out-of-sample statistics for those specifications. In the alternative case, the linear ECM associated with the GH-C cointegration vector (of the one with the best out-of-sample assessment) is the worst in terms of prediction. Possible explanations for such a result lie in the high speed of adjustment estimated and the presence of nonlinearities in the short-run dynamics. This interpretation also finds support in the fact that the MS-VECM of this long-run vector shows a small speed of adjustment in one of the regimes and a better prediction performance (compared to the ECM GH-C and to the GH-C long-run vector). We

[7] As highlighted by an anonymous referee, it is important to mention that a bad performance in the out-of-sample analysis might indicate other sources of misspecification and not only the spurious regression one. Anyway, the out-of-sample analysis can be used as an assessment tool also in those cases.

Table 9
Out-of-sample statistics for the selected long-run vectors.

	Actual values		Canonical model				Alternative model			
	M_{sa}	M	GH-FB		EE-2		GH-C		GH-CMSI	
			Predicted value	Error%	Predicted value	Error%	Predicted value	Error%	Predicted value	Error%
2010q2	37,155.0	35,853.2	35,948.5	−3.25	36300.31	1.25	36858.46	−0.80	36877.58	−0.75
2010q3	38,603.4	41,125.7	36149.32	−6.36	37,314.19	−9.27	37,677.08	−2.40	38,053.09	−1.43
2010q4	39,366.4	40,730.4	37,473.47	−4.81	37,527.92	−7.86	38,879.65	−1.24	39,435.9	0.18
2011q1	39,742.6	37,133.3	39,144.96	−1.50	35,411.36	−4.64	40,676.02	2.35	41,442.17	4.28
MAPE				3.98		5.75		1.70		1.66
RMSE				1689.58		2643.33		716.63		904.58
Bias				0.83		0.61		0.07		0.07
Variance				0.03		0.29		0.39		0.61
Covariance				0.14		0.10		0.54		0.32
U-Theil				0.02		0.04		0.01		0.01

Source: Authors' elaboration.

Table 10
Out-of-sample statistics for the selected error correction representation.

	Actual values	Canonical model				Alternative model					
	M_{sa}	ECM GH-FB		MS-VECM GH-FB (MSIAH)		ECM GH-C		MS-VECM GH-C (MSIA)		ECM GH-FB	
		Predicted value	Error%	Predicted value	Error%	Predicted value	Error%	Predicted value	Error%	Predicted value	Error%
2010q2	37,155.0	37514.09	0.97	37097.3	−0.16	36,710.62	−1.20	36,927.06	−0.61	36,345.43	−2.18
2010q3	38,603.4	38,192.96	−1.06	38,570.78	−0.08	38,846.91	0.63	37,713.74	−2.30	37,859.46	−1.93
2010q4	39,366.4	38,726.03	−1.63	39,525.88	0.40	40,684.93	3.35	38,655.66	−1.81	39,428.81	0.16
2011q1	39,742.6	39,806.84	0.16	40,899.21	2.91	41,710.18	4.95	40,142.65	1.01	39,916.84	0.44
MAPE			0.95		0.89		2.53		1.43		1.18
RMSE			421.82		584.70		1211.03		614.15		557.46
Bias			0.14		0.27		0.41		0.34		0.35
Variance			0.13		0.46		0.57		0.11		0.55
Covariance			0.73		0.27		0.03		0.55		0.10
U-Theil			0.01		0.01		0.02		0.01		0.01

Source: Authors' elaboration.

also note that the best prediction of the alternative model is the one obtained by the ECM estimated using the long-run vector that includes a change in the level and elasticities in the third quarter of 2002.

The canonical model in turn makes clear the relevance of the short-run dynamics in predictions when there is a slow adjustment process. In this case, the MS-VECM specification has the smaller MAPE. When we consider the RMSE and the proportion of random prediction errors, the ECM GH-FB has the best performance. Therefore, one should not discard the error correction representations of the canonical model.

6. Concluding remarks

In this paper we provide empirical estimates of the quarterly dynamics of aggregate imports in Brazil using national accounts data. The results show the major role played by the domestic income in determining its dynamic as well as point out its strong association with the gross capital formation of the economy. Furthermore, the econometric techniques employed indicate the existence of parametric instability in the estimated total import equations, although it is hard to characterize precisely the regimes given the multiplicity of different results.

Regarding the out-of-sample exercises, we highlight the good performances of the long-run vectors of the alternative models and the error-correction models of the canonical ones. Differences in the speed of adjustment toward equilibrium might be one of the reasons for such result. While it takes more than two years for a shock to return to the long-run equilibrium in the canonical model, our estimates imply convergence within a year on average in the alternative specification. Therefore, the short-run dynamics seems to be relevant only for the demand equation, as in the alternative case a similar performance can be obtained only among the long-run vectors and their error-correction representations are overparameterized.

The results of the alternative specification should be interpreted with caution given that this specification is not based in a theoretical model. In that respect, our results should be interpreted as providing empirical evidence on the importance of such relationship for the Brazilian economy in the sample period. Besides this, future research should also study import equations for the end-use categories, since the aggregate does not properly account for the distinct dynamics in each group. Inasmuch as the composition of demand matters for total imports, this seems to be a promising line of research for macroeconomists interested in the sectoral dynamics of the Brazilian economy.

Appendix I.

We performed three unit root tests before pursuing the cointegration analysis. It is worth remembering that, except in the exchange rate case, all series were seasonally adjusted. The results are reported in Table I.1 and suggest that all five time series are $I(1)$.

The DF-GLS test, developed by Elliott et al. (1996), was used because it is more powerful, compared to the traditional ADF, in the presence of deterministic terms in the data generating process. According to the results, it is not possible to reject the unit root hypothesis. However, this test is biased in the presence of structural breaks. We employed two other tests to address this possibility.

Zivot and Andrews (1992) test allows for breaks in both the level and slope of the trend function. The period of the break is endogenously chosen in such a way that the probability of rejecting the null hypothesis is maximized. This procedure is then incorporated when tabulating the test statistics. In general, the results corroborate the idea that there is a stochastic trend in the series. For the domestic income, however, the null is rejected at 5% significance level. This led us to consider a third test in order to check the results.

It should be noted that the rejection of the null in the Zivot and Andrews procedure does not warrant the conclusion that the series is stationary as there is an ambiguity in the alternative hypothesis. As the break is specified only under the later, the rejection of the former can be related to the fact that the series has a stochastic trend and a regime change.

Table I.1
Unit root tests (1996:1–2011:1).[a,b]

	DF-GLS			Zivot and Andrews				Lee and Strazicich			
	Test statistic	Critical value (5%)	*Lags*	Test statistic	Critical value (5%)	*Lags*	Break	Test statistic	Critical value (5%)	*Lags*	Break
ln$M$	−1.56	−3.16	1	−4.30	−5.08	1	2002:3	−5.36	−5.65	3	2002:2; 2008:1
ln$Y$	−1.14	−3.17	4	−5.18	−5.08	1	2002:4	−5.55	−5.65	7	2002:3; 2007:2
lnIk	−1.33	−3.16	2	−4.53	−5.08	2	2002:4	−5.66	−5.67	1	2001:2; 2004:2
lnCf	−0.65	−3.16	0	−4.52	−5.08	0	2003:1	−6.48	−5.67	6	2002:4; 2004:4
lnθ	−0.21	−1.95	0	NA	NA	NA	NA	NA	NA	NA	NA

Source: Authors' elaboration.

Notes: "NA" means "do not apply".

[a] The number of lags in the Zivot and Andrews and Lee e Strazicich tests were chosen using the AIC criteria, while for the DF-GLS the modified AIC was used.

[b] Given the properties of the time series, we chose the specification with a drift and deterministic time trend for the quarterly national accounts series and the specification with a drift for the exchange rate. The alternative hypothesis of the tests that allow for structural breaks include a constant and trend and do not apply for the real exchange rate.

Table II.1
Serial correlation in the initial VAR (alternative specification).[a]

Autocorrelation order (h)	VAR (1)					VAR (2)				
	1	2	3	4	5	1	2	3	4	5
Breusch-Godfrey[b]	1.74 (0.05)	1.13 (0.31)	1.12 (0.31)	1.67 (0.01)	1.38 (0.06)	0.41 (0.98)	0.91 (0.60)	1.60 (0.02)	1.73 (0.01)	1.43 (0.05)
Portmanteau[c]	NA	31.0 (0.01)	52.1 (0.01)	76.2 (0.01)	93.1 (0.01)	NA	NA	24.1 (0.09)	45.6 (0.06)	61.4 (0.09)

Source: Authors' elaboration.
Note: "NA" means "do not apply".
[a] Test statistics and p-values (in parentheses) take into account finite sample correction (Lütkepohl, 2005).
[b] H0: no autocorrelation in the lag h.
[c] H0: no autocorrelation up to lag h.

Table II.2
Non-normality of the initial VAR (alternative specification).[a]

	VAR (1)			VAR (2)		
	Skewness	Kurtosis	Normality	Skewness	Kurtosis	Normality
Doornik–Hansen	29.83 (0.00)	39.93 (0.00)	68.76 (0.00)	50.64 (0.00)	23.52 (0.00)	27.11 (0.00)
Lütkepohl	17.34 (0.00)	12.47 (0.00)	29.81 (0.00)	14.13 (0.08)	9.92 (0.04)	4.21 (0.38)

Source: Authors' elaboration.
[a] Test statistics and p-values (in parentheses) take into account finite sample correction (Lütkepohl, 2005).

Besides allowing for breaks in both null and alternative hypothesis, the Lee and Strazicich (2003) test also admits the possibility of two breaks in a way that is possible to test the significance of the new deterministic terms. We highlight that our sample is short, which brings about the risk of overparameterization. For the household expenditures the test rejects the null at 1%, meanwhile for the other national accounts series we can reject the stochastic trend hypothesis at 10%.

Even though some results are not conclusive, it is still valid to estimate cointegration vectors using persistent, but stationary, series. This is the case because near-integration models can still lead to spurious regressions. The main practical issue consists in the correction of the cointegration test statistic for the uncertainty in relation to the characteristic root of the data generating process of the dependent variable (Hjalmarsson and Österholm, 2007; Beechey et al., 2009). In this paper we assume that all series have a unit root since we do not know how to deal with this uncertainty in the presence of structural breaks in the series and in the long-run vector.

Appendix II.

The variables to be modeled in this study cannot be well represented as a multivariate linear process. The initial VAR will not be congruent, a requirement for the validity of the Johansen approach. We present the results of the alternative specification as they do not change substantially in the canonical case.

Table II.1 shows the serial correlation results. Initially we used a 1-lag VAR (as indicated by the Bayesian criteria), VAR (1), with an intercept. Evidence of autocorrelation is found by the Breusch–Godfrey (LM) and Portmanteau tests and the inclusion of a lag does not improve the results.

Table II.2 presents two versions of the Jarque–Bera multivariate test. The results suggest that both VAR's fail to pass the Jarque–Bera normality test.

Parameter instability tests are reported in Table II.3. The results of the Chow break-point and forecast tests indicate parametric instability, even though our small sample suggests they should be interpreted with caution.

Those results indicate that the initial VAR is not congruent. Therefore, it is not possible to apply Johansen procedures in this case.

Table II.3
Stability tests of the initial VAR (alternative specification).[a,b]

	VAR (1)		VAR (2)	
	Break-point	Forecast	Break-point	Forecast
1998	108.53 (0.06)	3.12 (0.11)	NA	NA
1999	77.17 (0.12)	1.41 (0.39)	138.66 (0.54)	1.18 (0.69)
2000	81.4 (0.05)	1.76 (0.11)	125.46 (0.10)	1.76 (0.17)
2001	98.98 (0.00)	1.99 (0.05)	123.42 (0.03)	1.65 (0.13)
2002	96.06 (0.00)	1.57 (0.11)	117.11 (0.03)	1.27 (0.29)
2003	76.89 (0.05)	1.10 (0.45)	108.05 (0.09)	1.15 (0.40)
2004	109.12 (0.00)	1.19 (0.33)	128.1 (0.03)	1.17 (0.35)
2005	104.63 (0.00)	1.27 (0.29)	172.83 (0.01)	1.20 (0.28)
2006	108.48 (0.00)	1.61 (0.09)	NA	1.36 (0.15)
2007	159.64 (0.02)	2.18 (0.02)	NA	1.79 (0.03)
2008	NA	3.27 (0.00)	NA	2.39 (0.00)
2009	NA	1.12 (0.32)	NA	1.14 (0.35)

Source: Authors' elaboration.
Note: "NA" means "do not apply".
[a] *P*-values were computed using Candelon and Lütkepohl (2001) bootstrap procedures.
[b] The tests were performed trying to find breaks in all possible points in time, but we prefer to report the values for the last quarter of each year as the results do not change.

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