STRUCTURAL CHANGE IN THE BRAZILIAN DEMAND FOR IMPORTS:
A REGIME SWITCHING APPROACH

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Abstract
The aim of the present paper is to apply a Markov Switching model to check the characteristics of the Brazilian demand for imports equation based on annual data from 1947 to 2002 and on quarterly data from 1978:I to 2002:II. The results show that this model satisfactorily describes the structural and conjunctural characteristics of Brazilian foreign trade in the last decades. The long-term analysis, based on annual data, allowed for the identification of cyclic periods of trade closure and openness that coincide with the historical events of Brazilian economy. The conjunctural analysis, based on quarterly data, indicates different elasticities for a regime with rise and fall in imports.

Key words*:* Markov Switching, Error Correction Mechanism, Demand for Imports, Trade Elasticities, Co-integration.
JEL Codes: F14; F41;C51; C52.

1. Introduction

The Brazilian economic history of foreign trade policies is packed with measures aimed at adjusting the external sector or stabilizing the economy. To accomplish that, several governments used the exchange rate controls as a key issue. However, other normative instruments were also important, such as the restriction on the import of certain products, the imposition of quotas and the grant of permits, compulsory deposits, or non-tariff barriers.[3]

In the second half of the 1940s, Brazil was faced up with a dire, global-scale postwar scenario that exerted a profound impact on its balance of payments. Due to the difficulty in boosting exports, which were centered around coffee, Brazil reached the late 1940s with a large trade balance deficit, compelling the government to curb imports. Notably, such control contributed towards import substitution, which gave rise to an industry of durable consumer goods by the end of president Dutra's government.

President Vargas' second government was also characterized by foreign trade policy difficulties, since the control over imports was maintained in order to balance external accounts. At the end of 1953, the control over imports was price-based, with the implementation of multiple exchange rates, as proposed by instruction no. 70 of the Superintendence of Money and Credit (Sumoc), which introduced remarkable changes to Brazil's exchange rate system (see Abreu, 1990).

In the late 1950s and early 1960s, the government significantly adjusted exchange rates by unifying the various exchange rates in force for exports and imports at that time and also modernized the tariff system, replacing specific tariffs with *ad valorem* ones. Moreover, the government's plan (Plano de Metas) was ambitious in seeking imports substitution. While this

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Special thanks to Amanda Pimenta (CNPq), Gustavo Russomanno (CNPq) and Júlia C. Klein (FAPERGS) for their research assistantship
[3] See Portugal (1994), Pinheiro *et alli* (1995) and Velloso (1998).

development plan was in force and in the following years, a great amount of foreign investments were directly made in Brazil.

From the mid-60s onwards, several import liberalization measures were taken. After that, the foreign trade policy, whose focus was on import substitution, started to encourage exports either through the exchange rate or through fiscal subsidies or incentives (see Portugal, 1994).

The first oil price shock in the early 1970s caused Brazil to remarkably adjust the external sector. The years 1973 and 1974 were characterized by a change in the behavior of foreign trade variables, when import restrictions were back again and the government adopted an inward-looking policy. The heavy investments made during President Geisel's government resulted in the increase of foreign debt which, to a greater extent, belonged to the public sector. At that time, a new import substitution process took place.

The second oil price crisis in the late 1970s deteriorated the Brazilian terms of trade and changed the economic scenario on a worldwide basis. The Brazilian foreign debt situation worsened with the increase in interest rates in industrialized countries and, consequently, Brazil had some difficulties in financing its current account deficit.[4]

The subsequent debt crisis forced the Brazilian government to seek financing for its external accounts in trade surpluses, mainly obtained through import restrictions. The maturation of investments within the Second Development Plan (II PND) helped Brazil achieve this goal by boosting exports, especially of industrial raw material and of capital goods. Throughout the second half of the 1980s Brazil experienced great macroeconomic instability, and the three economic plans implemented[5] could not hold back the high rates of inflation.

In the early 1990s, the adjustment of the external sector through the exchange rate controls was not working out any longer, and the Brazilian foreign trade policies went through deep changes (see Portugal, 1994 and Azevedo *et alli* 1998). The international tendency towards the formation of trade blocs and the intention to increase Brazil's participation in the international trade compelled the government to implement a more consistent trade liberalization, with the aim of lifting the restriction on the import of some products, including non-tariff barriers and the reduction of import tariffs. This process was later consolidated with currency stabilization in the Real Plan.[6]

Nonetheless, this liberalization policy combined with exchange rate appreciation and the increase in aggregate demand were too aggressive for Brazil's foreign trade structure at that time. Due to international financial crises that thwarted the financing of external accounts, to the small increase in exports *vis-à-vis quantum* expansion, and to the value of imports, which caused successive trade deficits, the government was obliged to change its exchange rate policy. Once again, the exchange rate was used to adjust the external sector, and the switch to a floating exchange rate system at the beginning of 1999 wound up establishing a new pattern for Brazilian imports, which lasted until the end of 2002.

Notably, the external sector has always played a central role in the Brazilian economic policy, and has therefore encouraged several studies on the estimations of foreign trade equations in the last few years. In case of demand for imports equations, we have the studies conducted by Portugal (1992 and 1993c) and Resende (1997a, 1997b and 2000), who use an Error Correction model for total imports, of capital goods, intermediate goods, by Azevedo *et alli* (1998) for total imports, Castro *et alli* (1998) for total imports and same main types of imported products and Carvalho *et alli* (1999) for monthly imports according to category of use.

All of these studies utilize dummy variables to correct problems with structural breaks in the series under analysis; however, the nonlinearity of the data has not been properly dealt with.

[4] By the end of 1979, the government devalued the exchange rate by 30%, and in February 1983, a new 30% devaluation occurred.
[5] This was Cruzado I Plan in February 1986, Cruzado II Plan in June 1987 and the Summer I Plan in January 1989.
[6] For discussion about the effects of the Brazilian protection on industrial productivity, see Júnior *et alli* (1999)

The study conducted by Silva, Portugal, and Cechin (2003), who applies the nonlinear models of artificial neural networks in order to estimate demand elasticities for total imports, intermediate goods and electrical material for the period between 1978:I and 1999:IV, is an exception.

Quite recently, nonlinear models have been used in academic research to characterize the existence of structural breaks or of various regimes that might arise and rule the behavior of macroeconomic data. The Brazilian time series have these characteristics on account of the several shocks that have shaken the economy in the last 20 years.

With regard to Brazil's foreign trade variables, it seems that there are different patterns or regimes, especially concerning imports, which at times show high growth rates and now and then have abrupt falls (see Portugal, 1994). In the former case, this expansion is related to the increase in purchasing power that resulted from various heterodox economic plans, or to the trade liberalization process established in 1988. The decrease in imports may be due to the economic slowdown, to the restrictive policies on imports or to the excessive depreciation of the currency. These changes to the behavior of foreign trade variables, as described in Goldstein *et alli* (1985), may be both gradual or abrupt.[7]

Thus, several factors encourage the estimation of a demand for imports equation for Brazil that contemplates the regime switching characteristic. First of all, we have the hypothesis that the impact coefficients of demand for imports are not constant over time. This characteristic is associated with the change to Brazil's productive structure characterized by investment and maturation cycles, and with the implementation of different economic plans that modified family income and consumption patterns. The impact of abrupt changes in the real exchange rate and the intensification of trade liberalization on foreign trade variables, which depend on the economic behavior of other countries, is also of note.

Another factor is concerned with the great dependence of Brazilian economy on the external account balance, which also reveals the necessity to determine foreign trade elasticities that include different regimes.[8] Finally, with greater trade integration through the reduction of tariff or non-tariff barriers, it is important that short- and long-run elasticities be determined in different behaviors of the domestic economic activity, of price and of the cyclic income flow. This allows measuring future profits and losses with greater economic openness.

As in other empirical studies,[9] we will assume an imperfect substitution model where there is a slight difference between domestic and foreign prices and products. In addition, we consider the small country hypothesis with totally price-elastic import supply, absence of monetary illusion in which the real product and the real exchange rate, prices, tariffs and subsidies are grouped into one single variable and, finally, weak exogeneity for the explanatory variables of the equation of demand.[10]

The aim of the present paper is to check whether the different foreign trade policies implemented in Brazil in the last years, as well as various external shocks, produced structural changes in Brazilian imports. The Markov regime switching model is the appropriate econometric tool to identify this characteristic.

Based on the analysis of annual data it is possible to identify the cyclic behavior of the demand for imports and associate the dates of each regime with the different foreign trade policies implemented at the time. These policies, in their turn, are associated with periods of moderate or consistent trade liberalization, and also with economic closure, such as the import

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[7] To estimate the elasticity of demand for time-varying imports, see Portugal (1993c).

[8] As postulated in Resende (2000), the restrictions on imports can be major or minor in certain periods, depending on the amount of foreign currency available in the economy.

[9] See Portugal (1992), Azevedo *et alli* (1998), Carvalho *et alli* (1999) and Carvalho *et alli* (2000).

[10] Such hypothesis was already confirmed in Castro *et alli* (1998).

substitution process.[11] We may therefore say that this section analyzes the behavior of imports in a structural fashion.

On the other hand, based on quarterly data, we perform a conjunctural analysis of the evolution of imports, seeking to associate the different regimes with periods of rise and fall in imported amounts that are compatible with these structural changes the Brazilian economy has gone through in the last years.

In addition to this introduction, the present article contains another three sections. Section 2 discusses the methodology of regime switching and its application to error correction models, more specifically, to the formulation of an equation of demand for imports. Section 3 and 4 present the statistical results for the annual and quarterly data, respectively. Section 5 concludes. Tables, tests and other estimated results not shown in the text are presented in the Appendix.

2. Methodology

When a time series is amenable to a structural break, which might occur in the variable coefficient, in the intercept and in the variance, the parameters of the static model vary over time. This way, the hypothesis of stationarity and normality is violated and nonlinearity results. In the last few years, the interest in the nonlinear modeling of economic time series, especially regime shift models, has gained momentum.

The regime shift in the time series may be characterized endogenously, by the own model, or exogenously, with an intervention from dummy variables, which implies a priori knowledge of the moment of regime shift. Among the several classifications presented in the literature, we will use Markov switching.[12]

A Markov process is a classic stochastic process in which random variable X_t is particularly time-dependent. This process will be discrete or continuous depending on the states (s_t) in which the variable is. In the former case, we have $s = (1,2,3,...)$ and in the latter, $s = (-\infty, \infty)$. If a Markov process has a finite or numerable number of states, then it is called a Markov chain.

The special characteristic of these models is the hypothesis that the realization of unobserved state $s_t \in \{1,...,k\}$ is determined by a Markov stochastic process in discrete state and discrete time, defined by transition probabilities.

The probability that X_{t+1} will be in state *j*, considering that X_t was in state *i* – called one-step transition probability– is then represented by 2.1:

$$P_{ij}^{t,t+1} = \Pr\{X_{t+1} = j / X_t = i\} \qquad (2.1)$$

As we can observe, transition probability $P_{ij}^{t,t+1}$ is not only a function of the state, but also a function of the transition time. However, if $P_{ij}^{t,t+1}$ is independent of time, the Markov process has a stationary transition probability, and $P_{ij}^{t,t+1} = P_{ij}$.

Since *k* states might exist, the transition probabilities between these states may be represented by a matrix of transition probability $P = [p_{ij}] \in M(kxk)$, exactly as in 2.2:

$$P = \begin{bmatrix} p_{11} & p_{21} & .. & p_{k1} \\ p_{12} & p_{22} & .. & p_{k2} \\ \vdots & \vdots & \vdots & \vdots \\ p_{1k} & p_{2k} & .. & p_{kk} \end{bmatrix} \qquad (2.2)$$

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[11] For an analysis of the foreign trade policies between 1947 and 1988, see Portugal (1994).

[12] For a discussion about the other formulations, see Tsay (1989), Granger *et alli* (1993), Dijk (1999), Teräsvirta *et alli* (1992) and Tsay (1998).

where: $\sum_{j=1}^{k} p_{ij} = 1$ for $i=1,2,....,k$ and even if, $p_{ij} \geq 0$ for $i,j=1,2,......,k$. The vector of Markov transition probability is given by $P = (P_{11},...P_{kk})'$, (k^2x1).

Finally, based on the values of transition probabilities obtained in 2.2, it is possible to calculate the length of each regime, that is, the persistency from $\frac{1}{1-p_{ii}}$. The length of each regime may differ, but with the hypothesis that the matrix of transition probability is fixed, the length of regimes will be constant in time, that is, the expected conditional length does not vary with the cycle. Filardo *et allii* (1998) extend the model proposed by Filardo (1994) by using a Bayesian methodology and consider that the evolution of the unobserved state depends on the available information in the time series y_t.[13]

In the estimate of 2.2, we presume that the variable is known, but not the states. As illustration, consider vector $y_t = (y_{1t},....,y_{nt})'$ $\{y \in \Re^n\}$ of observations with $t = 1,....,T$ and $s_t \in \{1,...,k\}$ the different unobserved states in which the variable might be. An occasional discrete shift is believed to occur in the level, variance, intercept, or in the autoregressive dynamics of y_t.

By assuming a distribution function for variable *y*, then $y_t \sim N(\mu_1, \sigma_1^2)$ if the process is in regime 1, $y_t \sim N(\mu_2, \sigma_2^2)$ if the process is in regime 2, and so on and so forth, until regime *k* with $y_t \sim N(\mu_k, \sigma_k^2)$. The parameter vector of the model is; $\theta = (\mu_1, \mu_2, ..., \mu_k, \sigma_1^2, \sigma_2^2, ..., \sigma_k^2)'$ and the density function of y_t is given by:

$$f(y_t / s_t = j; \theta) = \frac{1}{\sqrt{2\pi\sigma_j}} e^{\left\{\frac{-(y_t-\mu_j)^2}{2\sigma_j^2}\right\}} \quad j=1,2,....,k \qquad (2.3)$$

The unconditional density of y_t, based on the sum of all *k* states that may possibly occur in *t*, is described by 2.4,

$$f(y_t; \theta) = \sum_{j=1}^{k} P(y_t, s_t = j; \theta) \qquad (2.4)$$

This equation represents the sum of distributions produced by a density that depends on $P(s_t = j; \theta)$. As there are *T* observations, the log-likelihood of 2.4 is given by:

$$L = \sum_{t=1}^{T} \log f(y_t; \theta) \qquad (2.5)$$

The objective is to maximize the likelihood function of the observed data $(y_T, y_{T-1}, ...y_1 : p, \theta)$ where $p = (p_{11},....p_{kk})'$ and θ as defined above, based on the choice of population parameters (p, θ), that is, transition probabilities, mean, and variance.

This maximization is achieved by an iterative process, in which it is necessary to have the initial values of the parameter vector θ for the model so that it can be implemented. We should underscore that convergence takes place when the variation between θ in iteration $m+1$ and θ in iteration m is less than any specified value ε,, or when the first-order condition for maximum likelihood is met within a tolerance interval. The estimate of maximum likelihood is then given

[13] The authors showed that the time-varying regime switching model, from the use of Bayesian priors, produce recession probability estimates that are more in line with the dates of economic cycles found by NBER for the American economy than those produced by the model with fixed transition probability, which uses initial information about the parameters.

by $\hat{\theta}$, and therefore, it is possible to make inferences about the regimes associated with each observation y_t in time.

Although maximum likelihood estimation is a method that has optimal asymptotic properties, we do not have a theoretical solution to the likelihood equation in some applications.

In this case, it is necessary to use some numerical optimization technique applied to the likelihood in order to obtain the parameters for the model. One of the alternatives proposed by Hamilton (1990) to the use of Newton-Raphson or David-Fletcher-Powell methods is the EM (Expectation-Maximization) algorithm, which was initially introduced by Dempster, Laird and Rubin (1977).[14]

The EM algorithm is an iterative technique for the estimation of maximum likelihood designed for a general class of models where the observed time series depends on some stochastic variable that is not observed.

The application of the EM algorithm in econometrics occurs through the likelihood function. Each iteration of this algorithm consists of two steps, expectation (E) and maximization (M). Initially, the unknown parameters of the model, means, and variances for the different states, vector θ, Markov transition probabilities $p = (p_{11},....p_{kk})'$ and the probability of initial state ρ_0 are chosen. After that, also considering the vector of observations y_t, the smoothed probabilities are estimated.

The next step consists of maximization. Here, the parameter vector (mean and variance) is derived from the first-order condition of the maximum likelihood estimation. To find its value, the smoothed probabilities obtained in the previous step have to be replaced.

Thus, the mean of each state $\mu^1{}_k$ is obtained, and it may be used to obtain the covariance-variance matrix $\Omega^1{}_k$, transition probabilities, and the probability of state ρ_1. Later, ρ_1 is used to obtain $\mu^2{}_k$ and $\Omega^2{}_k$ and so on and so forth until $\mu^k{}_k$ and $\Omega^k{}_k$ is obtained.

This way, each EM iteration involves one step of the filter and one of the smoothing, followed by an updating of the first-order condition and of the estimated parameters, which guarantees the increase in the value of the likelihood function. For further details about the EM algorithm and its use in the maximum likelihood estimation, see Hamilton (1990) and Ruud (1991).

Therefore, if we know θ, then the probability that the process is in some regime s_t based on the information available up to t, $p(s_t / y_1,...y_t;\theta)$, is called filtered probability. On the other hand, if all the information is used to determine s_t, $p(s_t / y_1,...y_T;\theta)$, then we have a smoothed probability.

The univariate regime shift models can be extended to the multivariate case, where the aim is to check the existence of some similar behavior of unobservable components throughout time. In the models with MS-VAR regime, we assume that regime s_t is generated by an ergodic Markov chain and with homogeneous discrete states defined by transition probabilities 2.2.

MS-VAR models are considered a generalization of finite order autoregression for the vector of time series $y_t = (y_{1t},...,y_{kt})'$ of order k, with $t=1,...,T$, given by:

$$y_t = A_0 + A_1(y_{t-1}) + ... + A_p(y_{t-p}) + \varepsilon_t \qquad (2.6)$$

assuming that $\varepsilon_t \sim NID(0,\Sigma)$.

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[14] One of the problems that might arise in this maximization process is related to the fact that as in $f(y_t;\theta)$ there is a distribution sum, local, instead of global, log likelihood maximums may be found in many applications. Thus, once the quality of initial estimates may strongly influence the final result, it is advisable that the maximization be carried out for different initial values for parameter vector θ.

The idea behind the class of models with multivariate regime shift is that the parameters of process 2.6 (intercept, coefficient and variance) depend on a regime variable that is not observed:

$$y_t = A_0(s_t) + A_1(s_t)(y_{t-1}) + ... + A_p(s_t)(y_{t-p}) + \varepsilon_t \qquad (2.7)$$

where $\varepsilon_t \sim NID(0, \Sigma_{s_t})$.

It is important to emphasize that 2.7 does not wholly describe the data generating process. However, we still lack the formulation of a regime generating process s_t that may also be given by 2.2. In this case, the intercept is not a simple parameter, and it is actually generated by a stochastic process.[15] The mechanism of dynamic propagation of impulses in the MS-VAR model for the system consists of linear autoregression, which characterizes the transmission of shocks and the regime shift that represents common shocks.

The basic finite VAR model is given by a change in intercept, that is, MSI – Markov switching intercepts:

$$y_t = A_0(s_t) + A_1 y_{t-1} + ... + A_p y_{t-p} + \varepsilon_t \qquad (2.8)$$

that is equivalent to $y_t = A_0 + A_1 y_{t-1} + ... + A_p y_{t-p} + \varepsilon_t$ in a structural VAR model without regime shift. By subtracting y_{t-1} on each side, we obtain the model in the form of vector error correction:

$$\Delta y_t = A_0(s_t) + \pi y_{t-1} + \sum_{i=1}^{p} A_i \Delta y_{t-i} + \varepsilon_t \qquad (2.9)$$

This model is called MSCI(*k*,*r*)-VAR(p), autoregressive vector of order *p* with regime shift with *k* states and cointegration rank *r*. Aside from this formulation, another way to implement the regime shift is in the parameter of the mean $\mu(S_t)$ such that:

$$\Delta y_t - \mu(s_t) = \pi(\beta' y_{t-1} - \delta) + A_1(\Delta y_{t-1} - \mu(s_t)) + ... + A_p(\Delta y_{t-p} - \mu(s_t)) + \varepsilon_t \qquad (2.10)$$

where $\beta' y_{t-1} - \delta$ determines the correction of the long-term equilibrium. The shift may occur only in the long-term equilibrium $\delta(s_t)$, in the long-term equilibrium and in the joint mean, in coefficients A_i, or in the variance of residuals Σ_{s_t}.

Just as described in Krolzig (1997b), the estimation process of an MSCI(*k*,*r*)-VAR(*p*) model consists of two stages. First, we have a maximum likelihood estimation, the Johansen Procedure, to determine the number of cointegration vectors *r*. After that, the EM algorithm is used to obtain the model parameters, also by means of maximum likelihood estimation.[16]

The function of demand for imports used herein follows the theoretical framework specified in Portugal (1992), Ferreira (1994), Resende (1997b) and Azevedo *et alli* (1998), with the small country hypothesis, but with the following linear format:

$$\ln q_t = v + \beta \ln Y_t + \delta \ln U_t + \phi \ln e_t + \varepsilon_t \qquad (2.11)$$

where q_t is the imported quantum, v is a constant, ϕ is the price elasticity of the demand for imports, e_t is the real exchange rate given by $e_n \frac{p^*}{p}(1+T)$, where e_n is the nominal exchange rate, p^* represents the external prices, p is the domestic price index and T is the tariff, β is the secular income elasticity and Y_t is the measure of the level of activity, δ is the cyclic income elasticity

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[15] For detailed specification of this model's properties, see Krolzig (1996).

[16] Even when the VAR model is of infinite order, its finite order representation still preserves the asymptotic results for cointegration relationships. For detailed specification, see Saikkonen (1992) and Lütkepohl and Saikkonen (1995). Under the hypothesis that the data generating process is an MSCI(*M*,*r*)-VAR(*p*) process, Johansen structure may be used for cointegration, as in Krolzig (1996 and 1997b). Hansen and Seo (2002) propose a test for the presence of threshold effect in a VEC model.

with U_t equal to the use of installed capacity and ε is a white noise, where we expect that $\phi<0$, $\beta>0$ and $\delta>0$.

The variable of state s_t, which characterizes the regime switch, may be inserted both in the intercept or in the elasticities and in the variance of ε. This way, the most general error correction model, with regime switch in all coefficients, follows equation 2.12.

$$\Delta \ln q_t = v_0(s_t) + \sum_{i=0}^{p} \beta_i(s_t)\Delta \ln Y_{t-i} + \sum_{i=0}^{p} \delta_i(s_t)\Delta \ln U_{t-i} + \sum_{i=0}^{p} \phi_i(s_t)\Delta \ln e_{t-i} + \alpha vec_{t-1} + \varepsilon_t \qquad (2.12)$$

where $\varepsilon_t \sim NID(0,\Sigma(s_t))$ and α is the error correction vector coefficient. This model is known as *MSIAH(k)-VEC(p)* that is, an error correction vector model with p lags, and with regime switch in the intercept, in autoregressive components and in the variance of the residual, where k is the number of regimes. As with previous studies, we assume the hypothesis of price exogeneity and of cyclic secular income.

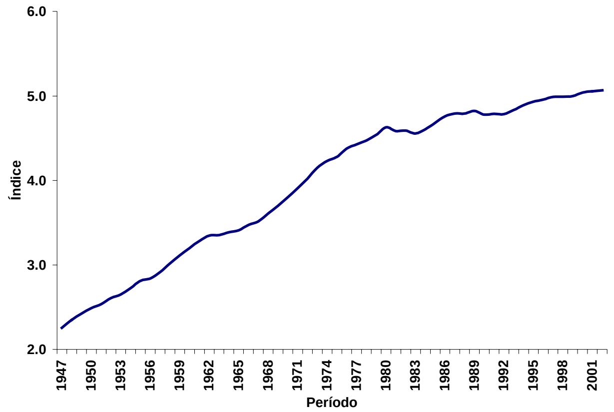
3. Statistical results with annual data

In this section, the tests and linear and nonlinear models are made with annual data, starting in 1947 and ending in 2002, with a total of 56 observations. To determine the Error Correction Mechanism for the Brazilian demand of imports, we use the imported quantum as dependent variable and the GDP, real exchange rate and installed capacity utilization as independent variables.

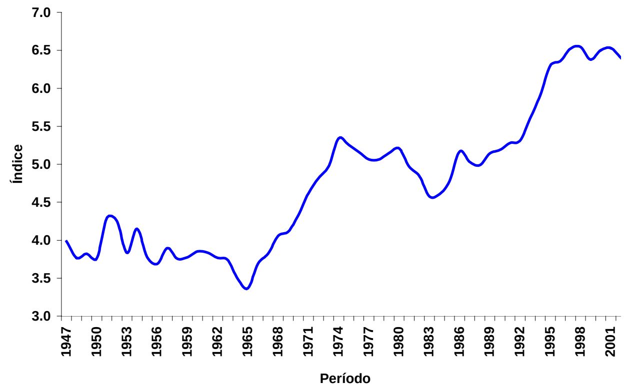
The GDP is used to determine the level of domestic economic activity, from which an income elasticity with positive sign is expected. For the real exchange rate, we initially considered two formulations, one that does not include the incidence of an import tariff, and another one that contains this variable,[17] in which price elasticity is expected to have a negative sign. Finally, the installed capacity utilization is seen as a cyclic component of the income with positive-sign elasticity. As there have been no available data for this variable since 1947, it was calculated in three different ways according to Portugal (1993b).[18]

Graphs 3.1 through 3.4 show the behavior of each of these series.

Graph 3.1 - GDP



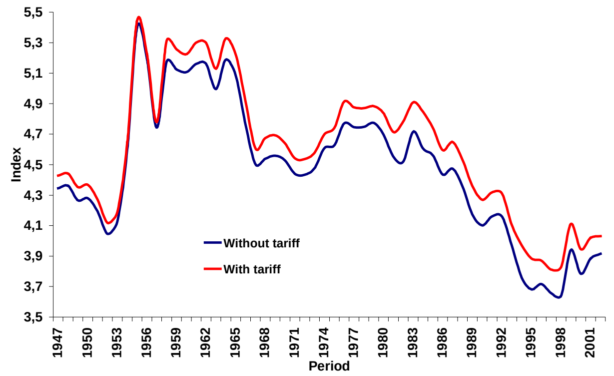
Graph 3.2 – Imported *Quantum*



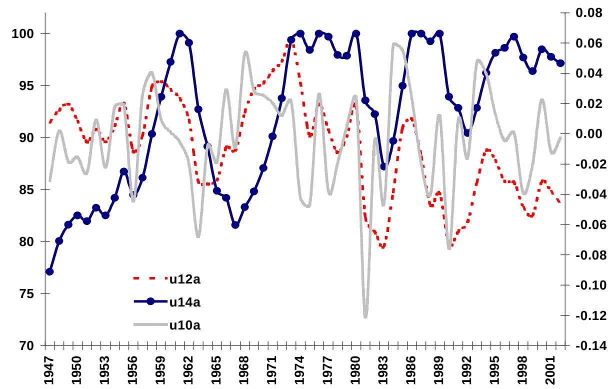
[17] In this section *Qm* is the imported *quantum*, *GDP* is the gross domestic product, *er* is the real exchange rate without tariff, *et* with tariff and the installed capacity utilization is given by variables *u10a*, *u12a* and *u14a*..

[18] The first estimate is obtained from the regression of the GDP logarithm with a trend where the residual is seen as the output gap. After that, one calculates the series of capacity utilization denominated u14a. The second estimate is made using the model of unobserved components, such as the trend, the cycle, and the irregular component, and the resultant series is denominated u10a. Finally, the third estimate of capacity utilization is made using a two-period moving average, and the resultant series is denominated u12a.

Graph 3.3- Real Exchange Rate



Graph 3.4- Installed Capacity Utilization



Only the graphical inspection indicates that the GDP and imported *quantum* should not be stationary. The results of the ADF test, performed in the level and in differences, are shown in table A.1 (see Appendix), where we note that the variables GDP, *Qm*, *er* and *et* are clearly nonstationary in the level but stationary in the first difference.[19] The three forms derived from the installed capacity utilization are stationary in the level.[20]

We should not forget that the Brazilian economy suffered several macroeconomic shocks in the 1980s and 1990s, and that the data in this section include a longer time period, of which the 1970s stand out as the probable moment of structural break, due to the first international oil price shock and to the industrialization process Brazil went through. Therefore, Perron's (1997) is the most appropriate test to filter these shocks. Tables A.2 and A.3 (see Appendix) present the results of this test.

According to the test results, the imported *quantum* series has two probable dates of structural break: 1950 and 1974. For the GDP, the date was 1974 and, for the exchange rate, both series (with or without tariffs), it was 1952. All of these series appear to be stationary in first difference even in the presence of these structural breaks. Castro *et alli* (1998) used Perron's test (1997) and found out that the imported *quantum*, GDP and the real exchange rate were *I(1)* even in the presence of structural break, and included a dummy variable in 1974 and another one in 1995 in the error correction model.

TABLE 3.1. – CORRELATION MATRIX

	U10a	*U12a*	*U13a*	*U14a*
U10a	1.00			
U12a	0.39	1.00		
U13a	0.12	0.84	1.00	
U14a	-0.06	0.23	0.53	1.00

In the case of capacity utilization different results exist, where series u10a and u14a are stationary and have break dates of 1962 and 1972, respectively. The series u12a shows a structural break in 1972, but has signs of nonstationarity in the level. Another way to test these series is by comparing them with the quarterly data of the conjunctural analysis collected by Getúlio Vargas Foundation (FGV) since 1970 for the industrial capacity utilization, hereinafter referred to as u13a. This is accomplished with the correlation matrix between this series and the other three formulations estimated herein, whose results are shown in table 3.1. Just as described

[19] Castro *et alli* (1998) also report that *Qm*, *GDP* and the real exchange rate are *I(1)* in the level and stationary from a difference.
[20] All the tests were performed on the three *Icu* transformed variables.

by Portugal (1993b), the series u12a was also highly correlated with the conjunctural analysis series, being used in the subsequent estimations.

To select the number of lags of the VAR model we used Akaike and Schwarz criteria and chose the order *p=0*. The causality test shows bi-causality for all variables, with exception of GDP and u10a.

3.1. Linear Model

According to the ordinary least squares method, the results are very close when we consider both the exchange rate with or without tariff. We opted for the model with the tariff since it contains more economic information. As shown in table 3.2 below, the sign of the long-run estimated income and price elasticity is just as expected, and the values are close to those found by Portugal (1993c), [21] where price elasticity is –0.648 and the income elasticity is 0.675.

TABLE 3.2. – LONG-RUN ELASTICITIES
OLS ESTIMATION

Variable	*Coefficient*	*Std Error*
GDP	0.692	0.052
et	-0.943	0.106
Constant	6.353	0.609

Johansen (1988) test was performed with one lag; the results of this test are shown in table A.7 (see Appendix), where the only cointegration vector, normalized for *Qm*, is given by:[22]

$$\beta = (1 \quad -0{,}940 \quad 1{,}695 \quad -8{,}83c)$$

The short-run dynamics in the linear model is determined by the error correction mechanism. The results, outlined in table 3.3 below, show that both the constant and the elasticity of installed capacity utilization are not significant. One should emphasize that models with one and no lag were estimated considering both exchange rates, but the results were not satisfactory as far as the coefficients are concerned.[23]

TABLE 3.3. – ESTIMATION RESULTS
LINEAR VEC(0) MODEL

VARIÁBLE	*COEFFICIENT*	*STD DEVIATION*
Constant	-0.047	(0.041)
ΔGDP	1.706	(0.676)
ΔU12a	0.339	(0.797)
Δet	-0.426	(0.109)
vec_{t-1}	-0.228	(0.073)

Note: The residual of OLS was obtained with variables *Qm, GDP, et.*

The income elasticity estimated here is much lower than that observed by Portugal (1993c) and Castro *et alli* (1998); price elasticity, however, shows a closer value, with the same sign, as shown in table 3.4.

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[21] The results of the model that considers variable u10a result in negative elasticity for capacity utilization. Although the elasticity for u14a is positive, it is much greater than the one found in the quarterly estimates. The group of tables A.6 (in the appendix) show the results for another six different OLS formulations. Portugal (1993c) also considered the capacity utilization variable for long-term estimates, and found the value of 2.307.

[22] That is, $Qm = 8.83 + 0.94GDP - 1.695et$. Castro *et alli* (1998) found 2 cointegration vectors.

[23] The results can be seen in tables A.8 to A.12 in the Appendix.

TABLE 3.4. – ELASTICITIES ESTIMATED IN OTHER STUDIES

	Constant	*GDP*	*Real Exchange*	*Vec$_{t-1}$*
Portugal (1993c)		2.149	-0.332	-0.337
Castro et alli (1998)	0.95	2.030	-0.450	-0.150

The error correction coefficient found here (-0.228) is significant, and its value is greater than that estimated by Castro *et alli* (1998) but smaller than that found in Portugal (1993c). This result shows that some of the short-run disequilibrium is corrected in each period.

3.2. An MS-VEC model

Nonlinear formulation includes the selection of a model that best adapts to the regime switch in the data. For this reason, general models are initially estimated, allowing for regime switches in the mean, intercept, and variance, and with two and three regimes, for the selection of the specific form and of the appropriate number of regimes, based on Akaike, Hanna-Quinn and Schwarz criteria, as well as on the likelihood ratio test.

According to the results in table A.15 (see Appendix), the H-Q and SC comparison criteria are used to select the model with two regimes and zero lag, a result also observed in the likelihood ratio test with $\chi^2_{(10)}$= [0.0063]. Due to the inaccuracy of the general model estimations,[24] several restrictive formulations were tested, with three and two regimes and with one and zero lag. The results are shown in tables A.16 and A.17 (see Appendix).

The comparison between several functional forms with two regimes and one lag and the general model is impaired as it was not possible to estimate the MSIAH(2)-VEC(1) model. In the relationship between zero-lag formulations, the selected model is given by MSIH(2)-VEC(0). For the three-regime model, the comparison criteria select the MSIH(3)-VEC(1) model among all formulations with one lag, and the MSI(3)-VEC(0) model among zero-lag formulations.

In the two-regime model, the intercept coefficient of regime 2 and the elasticity of installed capacity utilization were insignificant, as shown in table A.19 (see Appendix), while in the three-regime model, only the coefficient related to the intercept in the third regime is insignificant. The mean square error was also calculated, and its value is close to that of the models with two and three regimes. The LR test, carried out to verify the restriction of the model from three to two regimes, rejects the hypothesis H_0 with $\chi^2_{(4)}$= [0.1967]. Therefore, the most appropriate nonlinear formulation of the VEC model seems to be MSI(3)-VEC(0), described in equation 3.1, whose elasticities are shown in table 3.5.

$$\Delta qm_t = v(s_t) + B_1 \Delta GDP_t + C_1 \Delta ICU_t + D_1 \Delta e_t + \alpha vec_{t-1} + \varepsilon_t \qquad (3.1)$$

with $s_t=1,2,3$ and $\varepsilon_t \sim NID(0,\Sigma)$.

The income elasticity has a high value compared to previously made linear estimations for annual data, but its sign is as expected; the value of price elasticity, however, is close to that observed in linear models. On the other hand, the negative sign of the coefficient of installed capacity utilization shows an unexpected association. The three values of the constants for each regime reveal the difference of level among imports in each one of them. Regime 1 describes the moments of adjustment of the external sector, with a more temporary characteristic than other regimes, herein called moderate liberalization. Regime 2 represents periods of economic closure in Brazil, and regime 3 stands for more consistent moments of trade liberalization. Table 3.6.

[24] MSIAH(3)-VEC(2) and MSIAH(2)-VEC(1) models did not converge.

shows the different dates for each of these regimes, while the graphs (further below) show the smoothed probabilities.

TABLE 3.5. - RESULTS OF ML ESTIMATION FOR MSI(3)-VEC(0)

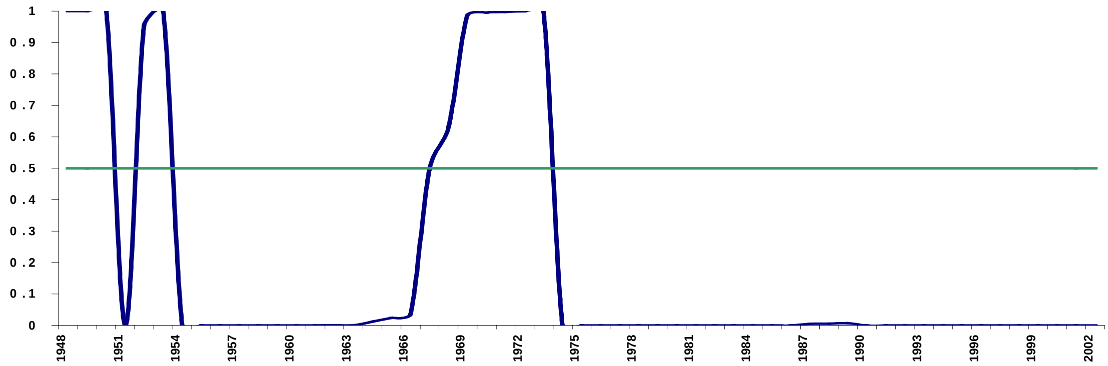
VARIABLE	*Coefficient*	*Std Deviation*
$Constant_1$	-0.418	0.043
$Constant_2$	-0.250	0.026
$Constant_3$	0.023	0.022
ΔGDP	4.704	0.418
ΔU12a	-1.623	0.411
Δet	-0.522	0.054
vec_{t-1}	-0.335	0.044

TABLE 3.6.- DATES OF EACH REGIME

REGIME 1	*REGIME 2*	*REGIME 3*
Moderate liberalization	*Economic closure*	*Consistent liberalization*
1948-1950	1955-1967	1951
1952-1953	1975-1989	1954
1968-1973		1974
		1990-2002

It is important to note the narrow relationship between the determination of these regimes and the moments of closure and openness of Brazilian economy to international trade. Despite the control over imports and exchange rate in the second half of the 1940s, we may say that the Brazilian economy was under a moderate liberalization regime (regime 1). This situation lasted until the mid-50s when a system of multiple exchange rates was adopted, restricting imports heavily and encouraging exports.

Graph 3.5. – Probability of Regime 1 – Moderate liberalization



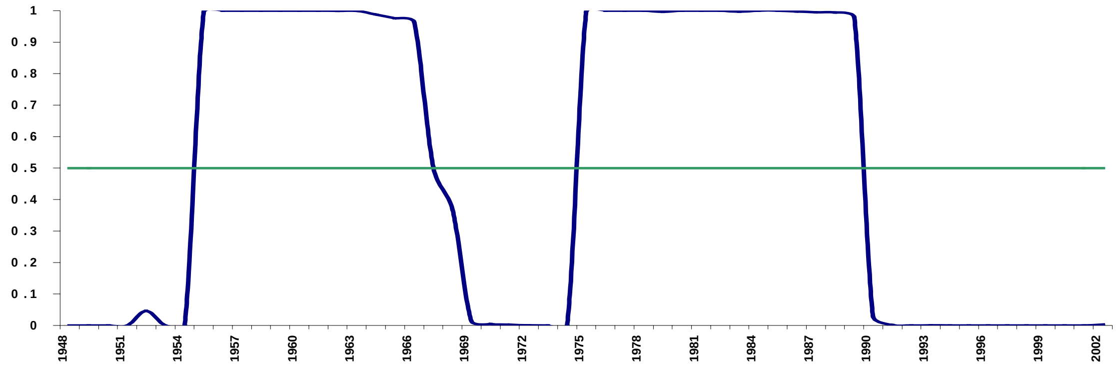
The Plano de Metas, implemented in the late 1950s, aimed at stimulating import substitution, especially of capital goods. The indication, by the regime switching model, that the Brazilian economy was in regime 2 in 1955 to 1967 coincides with this period of economic history. In 1961, a process of conversion to a single and free exchange rate was implemented and a 100% devaluation of the Brazilian currency occurred. Between 1968 and 1973, the Markov model signals that the Brazilian economy was under a regime of moderate liberalization (regime

1). Again, this result coincides with the foreign trade policy in force at that time, when import substitution was replaced with export incentives, see Portugal (1994).

The first oil price shock meant a great change in relative prices for the Brazilian economy and, the rapid deterioration of external accounts forced the government to promote a macroeconomic adjustment. Therefore, it is understandable that the years 1973 and 1974 represent a rupture in economic variables involved here and, as observed, the estimated model indicates the 1975-1989 period as a time of economic closure (regime 2).

In fact, in 1974, president Geisel's government, within the II PND, started an import substitution process with the aim of adjusting the external sector, restricting imports, increasing tariffs and implementing the obligatoriness of "down-payment" deposit, a policy that contributed to the increase of exports and reduction of imports (see Velloso, 1998).[25]

Graph 3.6. - Probability Regime 2 – Economic closure



The late 1980s and early 1990s is marked by trade liberalization of the Brazilian economy with reduction of nominal tariffs and the end of non-tariff barriers. This process was consolidated with the Real Plan, and the regime switching model can clearly associate this period with a more consistent trade liberalization (regime three). However, the Markov model also relates the years 1951, 1954 and 1974 to this regime. The sudden shift from regime 1 to regime 3 in these years may be associated with the behavior of dependent variables in the error correction mechanism equation.

At the beginning of 1951, the foreign trade policy adopted by president Vargas' second government consisted of a fixed and overvalued exchange rate, with restrictions on imports. However, in that year, import permits were largely granted due to the fear that the Korean War could spread into other countries and result in a new supply tightness in the domestic market, as occurred during World War II.

Therefore, the assumption that the Brazilian economy was under a consistent liberalization regime in 1951 is directly related to the increase in the imported quantum of that year, and should not be seen as a definitive trade liberalization. As a matter of fact, the deterioration of the balance of payments at the end of 1951 pressed the government on reestablishing strict control over the imported quantum.

During the year of 1954, the control over imports was price-based, with multiple exchange rates implemented through instruction no. 70 of the Superintendence of Money and Credit (Sumoc). This instruction also lifted the quantitative control over imports. To some extent, this policy explains the more consistent trade liberalization in 1954. Nevertheless, the economic

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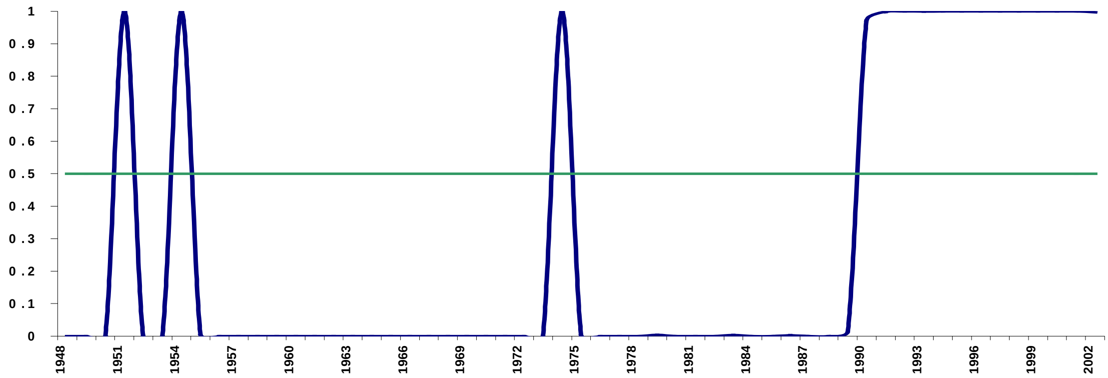
[25] Portugal (1993c) found that the imports-income elasticity was unstable. The author associated these movements with import substitution effects of the Plano de Metas (1955-1965) and II PND (1974-1982) and with the trade liberalization period (1966-1974).

and political developments in the subsequent years showed that Brazil was not ready yet for the consolidation of a broader trade liberalization.

Portugal (1993c) found out that 1974 was the only year that represented a structural break for the elasticity of installed capacity utilization, and associated this movement to the excessive economic activity in that year. Given the fact that the oscillation interval of this variable is limited, there was no other possibility than its fall in the following moment. Thus, when economy is "at full throttle", this variable loses importance, turning the relationship between installed capacity utilization and the imported amount into a nonlinear one. Therefore, the fact that the Markov model associates that year with the regime of consistent liberalization and, in subsequent years, with the regime of economic closure, may be related to the behavior of installed capacity utilization in those years.

Graph 3.7. - Probability Regime 3 – Consistent liberalization



The MSI(3)-VEC(0) formulation was also tried out, with a dummy variable with value 1 in 1951, 1954 and 1974 and zero in the other years. However, the results were not satisfactory due to nonconvergence. Furthermore, the Markov model without the use of dummy variables is more elegant and informative, representing different characteristics of the Brazilian foreign trade policy.

The estimated transition matrix between regimes shows greater persistence for regime 2 (economic closure), with average duration of 13 years. Actually, they are only two long periods of reduction in the level of economic liberalization. Also interesting is the nearly null probability of transition from a regime of moderate liberalization to one of closure, which is given by P_{12}. In other words, once the process of trade liberalization is achieved, it is less likely that it will be reversed. Given the conjunctural necessity, it is a bit more likely to go from economic openness to a process of adjustment of the external sector, regime 1, a probability characterized by P_{13} and whose value is 0.253.

$$P = \begin{bmatrix} 0.747 & 0.000 & 0.253 \\ 0.039 & 0.925 & 0.036 \\ 0.068 & 0.137 & 0.795 \end{bmatrix}$$

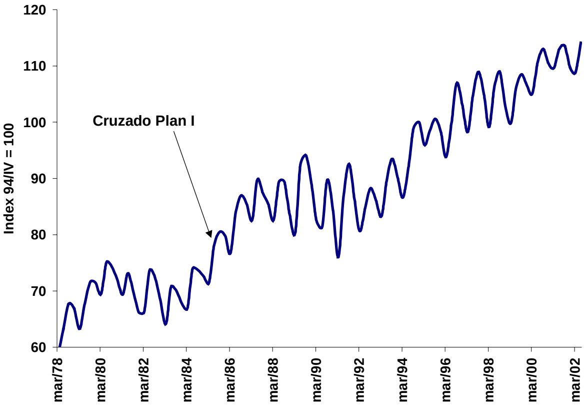
	Ergotic Probability	*Length of Regimes*	*Observations*
Regime 1	0.163	3.95	11.1
Regime 2	0.539	13.29	27.9
Regime 3	0.297	4.89	16.0

It should be underscored that the process of liberalization of Brazilian economy is still recent and that its length estimated by the model used herein is of approximately 5 years. In addition, this low value is associated with the short periods in which it was possible to characterize regime 3, as in 1951, 1954 and 1974. However, as the results show, there has seemed to be a consolidation of Brazilian economic liberalization in the last 12 years.

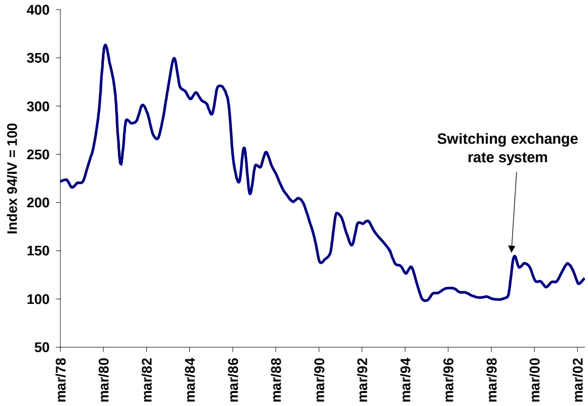
4. Statistical results with quarterly data

As in the previous section, we used the imported *quantum* index as dependent variable, and the installed capacity utilization, real exchange rate and GDP as independent variables to determine the level of domestic economic activity. The efficiency of Industrial Production as variable for the determination of economic activity was also tested, but its results were not satisfactory. The sample period includes from the first quarter of 1978 to the second quarter of 2002, totaling 98 observations. The behavior of each variable is shown in graphs[26] 4.1 to 4.4.

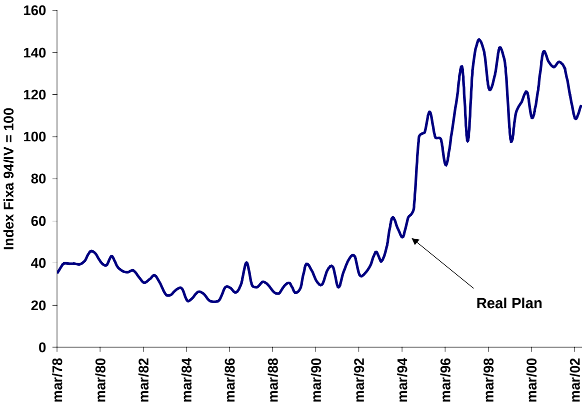
Graph 4.1. - GDP



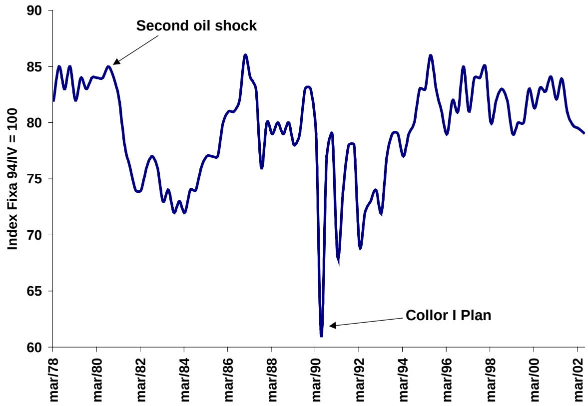
Graph 4.2. – Real Exchange Rate



Graph 4.3. - Imported *Quantum*



Graph 4.4.-Installed Capacity Utilization



The graphical inspection of data indicates that they do not seem to be stationary. Moreover, the fact that Brazilian economy has gone through several shocks suggests a structural change in the coefficients of the equation of demand for total imports. This parameter instability was already investigated in Fachada (1990), Portugal (1992) and Ferreira (1994), who observed parameter instability in 1981:IV. Castro *et alli (1998)* and Resende (1997a and 1997b) found a structural break in this equation for 1981:I justified by the complementation of investment cycles made in the Brazilian economy, especially in the capital goods industry. Azevedo *et alli* (1998)

[26] Quarterly observations were obtained through arithmetic mean of available monthly data.

found a change in the coefficients of GDP and of installed capacity utilization[27] in 1990:I and Resende (2000) spotted a break in parameters in 1990:I and 1994:IV. Silva, Portugal and Cechin (2001) performed linearity tests for the demand for imports and concluded for the rejection of the null hypothesis of linearity, in addition to identifying 1989:IV and 1994:III as the most significant structural break dates.

4.1. Linear Model

The first step is to determine the order of integration of variables; the results of the unit root tests are shown in table B.1 (see Appendix). As we may note, it is not possible to reject the hypothesis H_0 of existence of unit root when the variables are determined in the level; but in differences, they become stationary.[28] Aside from this test, the seasonal unit root test of Hylleberg *et alli* (1990) is also done; the results of this test are shown in table B.2 (see Appendix) and reveal that at frequency zero, for tests $t{:}\pi_1$, the series are nonstationary, that is, $y_t \sim I_0(1)$, as verified in the ADF test. However, for the biannual test, $t{:}\pi_2$, we may conclude that all series are stationary, that is, $y_t \sim I_{1/2}(0)$; the same occurs for the annual frequency $t{:}\pi_3$, where all series are believed to be stationary. On the other hand, the results indicate that all series, except for GDP, are $y_t \sim I_{1/4}(1)$, and when the auxiliary regression has deterministic seasonality, the results are ambiguous for $\pi_4=0$ and $\pi_3 \cap \pi_4=0$. This way, it is possible to conclude for the nonexistence of seasonal unit root.[29]

It should be emphasized that the non-rejection of the unit root hypothesis at frequency zero in the series in question may be related to possible structural breaks caused mainly by stabilization plans and by the periods of strong exchange rate devaluation that assailed Brazilian economy in the last years.[30] Thus, Perron's test (1997) seems to be highly recommended; its results are shown in table B.3 (see Appendix).

According to the estimations in levels the three methods used do not reject the unit root hypothesis at frequency zero, even in the presence of structural break. The identification of the moment of the break is also important, since it differs among the series, especially 1992:IV for the imported *quantum* and installed capacity utilization, 1989:III for GDP and industrial production and 1998:III for the real exchange rate.[31]

The selection of VAR lag is based on Akaike and Schwarz information criteria, both indicating order p=1 for a VAR approximation to the system. When we assume that the true model is subject to regime switching, any order of finite VAR is only an approximation (Krolzig, 1997b, p.314). The causality test with one lag, whose results are in table B.6 (see Appendix), shows that the imported *quantum*, *GDP* and industrial production are not related and that there is bi-causality between the exchange rate and all the other variables.

Here we will use a single-equation model according to the hypotheses described in the introduction. The cointegration regression, obtained through ordinary least squares, produced results for coefficients of price, income and capacity utilization with signs and magnitude compatible with those reported in the literature, as shown in tables 4.1. and 4.2.

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[27] These authors used both the GDP and the industrial production to measure income elasticity, finding better results with the GDP.

[28] To test the unit root in capacity utilization, the following transformation, $\log(icu/(100-icu))$ was necessary.

[29] Portugal (1992) found seasonal unit root in the GDP and installed capacity utilization series.

[30] Ferreira (1994) uses dummy variables to check the instability in the parameters of the equation of demand for imports. Resende (1997) uses a *dummy* variable in 1986-IV to filter a possible speculative demand for imports of capital goods, Resende (2000) uses dummy variables in 1986:IV, 1989:I and 1994:III to correct structural breaks in the parameters of the equation of demand and Azevedo *et alli* (1998) use them in 1990:I for the GDP and for the capacity utilization.

[31] In the test proposed by Perron (1997) the data of the possible structural change is not fixed *a priori*, and is considered unknown.

TABLE 4.1. – LONG-RUN ELASTICITIES
OLS ESTIMATION

Variable	*Coefficient*	*Std Error*
GDP	0.821	0.309
ICU	2.622	0.508
E	-0.910	0.130
Constant	-6.470	2.810

Next, we have Johansen (1988) analysis for cointegrated linear systems for VAR(1) model, whose results are shown in table B.7 (see Appendix).[32] The tests reveal the existence of only one cointegration vector, which is normalized for *Qm*, and given by:[33]

$$\beta = (1 \quad 5.833 \quad -3.514 \quad 4.150 \quad -36.11c)$$

TABLE 4.2. – ELASTICITY OF DEMAND FOR TOTAL IMPORTS ACCORDING TO OTHER STUDIES (QUARTERLY)

	Income Elasticity	*Price Elasticity*	*Installed Capacity**	*Trend*	*Period*
Zini Jr .(1988)	3.28	-0.460	3.310	-	1970 – 1986
Fachada (1990)	1.186	-0.376	1.563	-	1976/II-1988/IV
Fachada (1990)	1.186	-0.376	1.563	-0.0095	1976/II-1988/IV
Portugal (1992)	0.344	-0.910	3.865	-	1976/I-1988/IV
Ferreira (1994)	-0.212	-1.323	2.210	-	1981/IV-1989/IV
Azevedo et alli (1998)	2.106	-0.576	2.541	-	1980/I-1994/IV
Resende (1997b)	-0.89	-0.007	0.75	-	1974/II-1988/IV
Resende (2000)	3.310	-1.39	-	-	1978/I-1998/IV
Silva *et alli*(2001)**	1.277	-1.175	0.290	-	1994/III-1999/IV

Note: * In Resende (2000) this variable was not statistically significant. **The other income, price and installed capacity elasticities found by the author are: -0.006, -0.275 and 0.05 for the 1978/I-1989/III period and 0.179, -0.905 and 0.04 for the 1984/IV and 1994/II period.

TABLE 4.3. – ESTIMATION RESULTS
LINEAR VEC(1) MODEL

VARIABLE	*COEFFICIENT*	*STD DEVIATION*
Constant	0.0106	(0.009)
Δqm_{t-1}	-0.3286	(0.103)
ΔGDP	1.0241	(0.241)
ΔGDP_{t-1}	-0.7184	(0.216)
ΔIcu	0.9767	(0.235)
ΔIcu_{t-1}	0.8964	(0.230)
Δe	-0.227	(0.102)
Δe_{t-1}	-0.2069	(0.105)
vec_{t-1}	-0.1240	(0.034)

32 The results reject the cointegration test in equation $\Pi y_{t-1} + Bx_t = \alpha(\beta y_{t-1} + \nu)$, that is, they consider the presence of the intercept term.

33 That is, $Qm = 36.11 - 5.833GDP + 3.514Icu - 4.150e$

For the determination of the short-run dynamics, an error correction mechanism is estimated and these results are shown in table 4.3 below. Except for the constant, all coefficients were significant. The only result that differed from expected was that of lagged income elasticity, with negative sign. In case of price elasticity, we note that less than half of the adjustment occurs in the first two quarters.[34] On the other hand, there is a strong impact of the level of capacity utilization.

The error correction coefficient is –0.124, less than that found in previous studies, as shown in table 4.4, but with the expected sign. This means that only a small part of the short-run disequilibrium is corrected in each period.

TABLE 4.4. – ERROR CORRECTION VECTOR COEFFICIENT OBSERVED IN OTHER STUDIES

	Portugal (1992)	*Azevedo et alli (1998)*	*Resende (1997b)*	*Resende (2000)*
Vce_{t-1}	-0.182	-0.459	0.593	-0.735

Although linear estimations prevail in the empirical modeling of time series in the Brazilian literature, there is no reason to presume that the true dynamic data structure is linear. The hypothesis of linearity implies that the model's variables do not vary with size and that the sign of shocks and parameters do not vary over time.

Therefore, to test the nonlinearity of the series used here, we implemented Hansen's linearity test (1999), whose results arte shown in table B.8 (see Appendix).[35] We may note that when the test is performed between one-regime models, SETAR(1), and two-regime models, SETAR(2), the null hypothesis is rejected, that is, the hypothesis of nonlinearity is accepted for all series, which is similar to what occurs when we test one regime against another three. However, the test between 2 regimes and 3 regimes indicates that the null hypothesis is accepted. Thus, we may conclude that the data used here have a nonlinear structure and may be represented by two regimes.

4.2. An MS-VEC model

The cointegration results obtained in the previous section are now used to determine the error correction mechanism in a regime switching model. The procedure used to select the best nonlinear data formulation consisted in going from a general model to a specific one. This way, formulations with two and three regimes are estimated, considering simultaneous changes in the intercept, mean and variance. Akaike, Schwarz and Hanna-Quinn information criteria and the likelihood ratio test, as mentioned in the previous section, are used to select the number of regimes.[36]

Table B.9 (see Appendix) shows the information criteria for the general model with one or two lags and three regimes, where we note that they select the model with one lag, just as in the linear model, but are inaccurate in the selection of the number of regimes. The likelihood ratio test was carried out between three and two regimes, whose result was $\chi^2_{(14)}$=46 [0.000] that is, the restriction is accepted. This way, the general model has one lag and two regimes where regime 1 is characterized by a decrease in imports, while regime 2 shows an increase in imports.

The next step involves the selection of the functional form. Table B.10 (see Appendix) shows the information criteria for four different formulations, where all select the general model, a result

[34] In Portugal (1992) the price elasticity was –0.476 for the first quarter.

[35] Hansen's (1999) linearity test in SETAR models is a null hypothesis test of SETAR(1) against an alternative of SETAR(*m*), where m is the number of regimes and with *m>1*.

[36] Ang *et alli* (1998) use a likelihood ratio test to select between two models with different number of regimes from a an asymptotic distribution approximated by a chi-square distribution.

that is also confirmed by the likelihood ratio test. Thus, the final model selected is MSIAH(2)-VEC(1), given by equation 4.1:

$$\Delta qm_t = v(s_t) + A_1(s_t)\Delta qm_{t-1} + \sum_{i=0}^{1} B_i(s_t)\Delta GDP_{t-i} + \sum_{i=0}^{1} C_i(s_t)\Delta Icu_{t-i} + \sum_{i=0}^{1} D_i(s_t)\Delta e_{t-i} + \alpha vec_{t-1} + \varepsilon_t \tag{4.1}$$

with $s_t=1,2$ and $\varepsilon_t \sim NID(0,\Sigma(s_t))$, where the estimated parameters are shown in table 4.5 below.

TABLE 4.5. - RESULTS OF ML ESTIMATION FOR MSIAH(2)-VEC(1)

VARIABLE	*REGIME1*	*REGIME 2*
	Decrease in imports	*Increase in imports*
Constant	-0.03(0.006)	0.0088(0.015)
Δqm_{t-1}	0.084(0.093)	-0.302(0.224)
ΔGDP	1.146(0.165)	0.08(0.374)
ΔGDP_{t-1}	1.317(0.193)	1.171(0.298)
ΔIcu	-0.140(0.102)	2.616(0.668)
ΔIcu_{t-1}	-0.708(0,116)	-0.100(0.392)
Δe	0.122(0.076)	-0.424(0.128)
Δe_{t-1}	-0.219(0.077)	0.126(0.195)
vec_{t-1}	-0.0056(0.019)	-0.096(0.061)
σ	0.0257	0.0747

Note: Standard deviations are shown between parentheses.

The likelihood ratio test of the linear VEC(1) model against the MSIAH(2)-VEC(1) model ($LR_{(10)}$=50.81) strongly rejects the hypothesis of linearity, just like the comparison criteria indicate the nonlinear model. Finally, other pieces of evidence in favor of the nonlinear model include the large difference between the estimated parameters for both regimes and the fact that for the linear model the residuals are normal $\chi^2_{(2)}$=1.039[0.594], not correlated $\chi^2_{(7)}$=14.62[0.04] and homoskedastic $\chi^2_{(16)}$=14.94[0.528]. As shown in table 4.5, only some parameters are not significant, Δqm_{t-1} and vec_{t-1} in regime 1, and the constant (as in the linear model), in addition to ΔGDP, ΔIcu_{t-1} and Δe_{t-1} in regime 2.

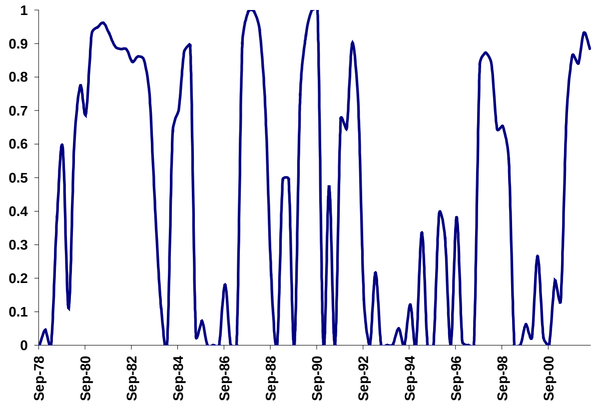
The income elasticity is positive in both regimes, as expected, and greater after one quarter, which means that when the economic activity grows, imports also increase, regardless of the regime. The elasticity of the cyclic component has negative sign in regime 1, which is different from the expected result and which was also found in the previous section, when annual data were considered. Nevertheless, in regime two, this coefficient is positive, as expected, with magnitude closer to that observed in linear models.

Price elasticity has positive sign and a small value in regime 1 and in the first quarter; however, from the second quarter onwards, the sign becomes negative, revealing that an exchange rate devaluation will reduce imports with the lag of one quarter. Interestingly enough, this impact is more significant when under a regime with an increase in imports (regime two), where its value, in the first quarter, is close to that found in linear models. This means that as imports increase, there is greater need for external adjustment in this regime, and the change in the exchange rate produces more immediate effects. Finally, we have the error correction coefficient, which is not significant in regime 1.

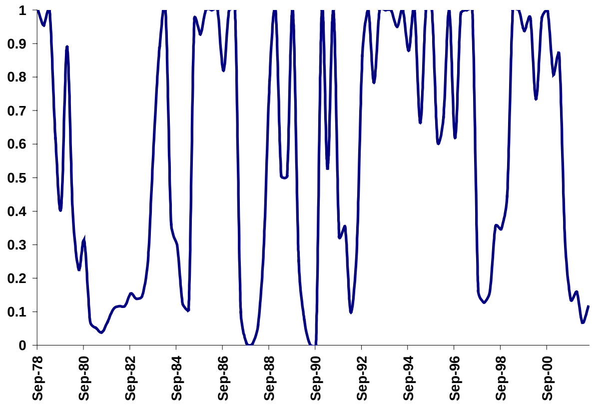
TABLE 4.6. - REGIME DATES	
REGIME 1	*REGIME 2*
Decrease in imports	*Increase in imports*
79:III – 79:III	78:III – 79:II
80:I – 83:II	79:IV – 79:IV
84:II – 85:I	83:III – 84:I
87:II – 88:II	85:II – 87:I
89:IV – 90:III	88:III – 89:III
91:III – 92:II	90:IV – 91:II
97:III – 98:IV	92:III – 97:II
01:II – 02:II	99:I – 01:I

Despite the expected sign, the speed of error correction of regime 2 is much lower than the one observed in linear estimations, but close to the one estimated in the linear model of this section. The probabilities of each regime are specified in the graph that follows, and the dates of each regime are shown in table 4.6.

Graph 4.5. - Probability of Regime 1



Graph 4.6. - Probability of Regime 2



The transition matrix shows that both regimes P_{11} and P_{22} are persistent, with the period of decrease in imports estimated to last approximately 10 months, and that of increase in imports, a little bit longer than one year.

$$P = \begin{bmatrix} 0.7097 & 0.2903 \\ 0.2209 & 0.7791 \end{bmatrix}$$

The estimated period of longer permanence in regime 1 occurred in the early 1980s soon after the second oil price shock, when international and national economic activity slowed down. The estimated period of greater increase in imports was between 1992:III and 1997:II, right after the intensification of trade liberalization, also including the initial period of the Real Plan, when the real exchange rate was strongly appreciated. Another interval of continued increase in imports (regime 2), captured by the model, goes from 99:I to 01:I, just after the exchange rate switch.

	Ergotic Probability	*Length of Regimes*	*Observations*
Regime 1	0.432	3.44	40.9
Regime 2	0.568	4.53	55.1

5. Conclusion

In the present paper, we estimated linear and nonlinear Error Correction Mechanisms for the demand of Brazilian imports using annual and quarterly data. In case of annual data, the results show that the three-regime Markov model describes the different foreign trade policies implemented in the Brazilian economy since 1947. The regime characterized as economic closure pointed to the 1955-1967 and 1975-1989 periods, which coincide with the import substitution policies of the period.

The regime designated as consistent economic openness coincides with the trade liberalization policy implemented in 1990, with reduction of taxation rates, and elimination of tariff and non-tariff barriers. This period stretches up to the end of the sample period (year 2002). The last estimated regime may be related to an intermediate level of economic openness, where we note import liberalization, but not in a consistent manner. This regime is associated with the years 1968 to 1973.

The regime switching model represents the periods of import substitution and economic openness referred to in the economic literature. Thus, we conclude that its application with annual data can be seen as representative of a structural adjustment of the foreign trade in Brazil, producing satisfactory results.

Unlike annual data, the application of the model to quarterly data aims at describing a more conjunctural behavior of the demand for imports. In this regard, the Markov model selects a two-regime formulation, herein called increase and decrease in imports. The observed results adapt well both to periods with strict control over total imports (in which imports decreased) and periods in which imports were unrestrained.

Both regimes revealed high persistence, but that of regime 2 (increase in imports) was higher, which indicates the difficulty there is when a country needs to make an external adjustment and then go quickly from regime 2 to regime 1 (decrease in imports). The estimated transition matrix shows a greater probability for us to go from a period of decreased to increased imports P_{12}, than the other way round.

Therefore, our conclusion is that the application of the Markov regime switching model to the demand for imports with quarterly data also showed satisfactory results, being able to describe the movements of conjunctural adjustment in the Brazilian foreign trade in the last two decades.

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Appendix

APPENDIX A – ANNUAL DATA ESTIMATIONS

Table A.1. Unit root test in the level and in differences

In the level					*In differences*				
	τ	τ_μ	τ_τ	I(.)		τ	τ_μ	τ_τ	I(.)
Qm	1.53	-0.29	-2.01	I(1)	*ΔQm*	-6.47[a]	-6.77[a]	-6.74[a]	I(0)
GDP	1.36	-2.27	-0.27	I(1)	*ΔGDP*	-1.75[c]	-2.73[c]	-3.62[b]	I(0)
e	-0.43	-1.55	-4.85[a]	I(1)	*Δe*	-6.60[a]	-6.54[a]	-6.56[a]	I(0)
et	-0.41	-1.80	-4.77[a]	I(1)	*Δet*	-6.79[a]	-4.00[a]	-4.25[a]	I(0)
U10a	-7.06[a]	-7.04[a]	-6.97[a]	I(0)	*ΔU10a*	-5.73[a]	-5.68[a]	-5.64[a]	I(0)
U12a	-0.94	-2.81[b]	-3.28[b]	I(0)	*ΔU12a*	-7.83[a]	-7.76[a]	-7.69[a]	I(0)
U14a	-1.68	-5.08[a]	-5.35[a]	I(0)	*ΔU14a*	-5.35[a]	-5.30[a]	-5.28[a]	I(0)

Note: τ means no constant, τ_μ means with constant and τ_τ means test with constant and trend. *I(.)* is the order of integration. Rejects at 1%, b 5% and c above 10%.

TABLE A.2. – Unit root test in the presence of structural break

Test	Model	*Qm*	*GDP*	*e*	*et*	*U10a*	*U12a*	*U14a*
UR	1	-3.47	-2.02	-5.11[b]	-4.42	-7.18[a]	-5.95[b]	-6.39[a]
$t^*_\alpha(i)$	2	-3.47	-4.36	-5.09	-4.41	-7.14[a]	-6.16[a]	-5.36[b]
	3	-3.23	-4.44	-4.07	-3.84	-4.83[b]	-3.72	-4.99[b]
STUD	1	-1.66	-0.17	-3.50	-3.40	-5.10[b]	-5.95[a]	-5.06[b]
$t^*_{\alpha,\hat{\theta}}$ and $t^*_{\alpha,\gamma}$	2	-2.88	-4.36	-1.30	-1.67	-4.85	-4.21	-4.84
	3	-2.97	-4.35	-3.94	-3.39	-4.79[b]	-3.72	-4.87[b]
STUDABS	1	-1.66	-0.17	-3.49	-3.39	-5.10[b]	-5.95[a]	-5.06[b]
$t^*_{\alpha,\lvert\hat{\theta}\rvert}$ and $t^*_{\alpha,\lvert\gamma\rvert}$	2	-2.88	-4.36	-1.30	-1.67	-4.85	-4.21	-4.84
	3	-2.97	-4.35	-3.94	-3.39	-4.79[b]	-3.72	-4.86[b]

TABLE A.3. – Structural break dates

Test	Model	*Qm*	*GDP*	*e*	*et*	*U10a*	*U12a*	*U14a*
UR	1	1950	1966	1952	1952	1962	1972	1988
$t^*_\alpha(i)$	2	1950	1974	1952	1952	1962	1972	1972
	3	1949	1981	1958	1958	1953	1972	1982
STUD	1	1974	1979	1957	1957	1973	1972	1972
$t^*_{\alpha,\hat{\theta}}$ and $t^*_{\alpha,\gamma}$	2	1955	1974	1955	1955	1979	1978	1983
	3	1963	1980	1959	1959	1955	1972	1987
STUDABS	1	1974	1979	1957	1957	1973	1972	1972
$t^*_{\alpha,\lvert\hat{\theta}\rvert}$ and $t^*_{\alpha,\lvert\gamma\rvert}$	2	1955	1974	1955	1955	1979	1978	1983
	3	1963	1980	1959	1959	1955	1972	1987

TABLE A.4. – SELECTION OF VAR LAG

VARIABLES	*ORDER*	*AKAIKE*	*SCHWARZ*
	0	227.70	227.85
Qm	1	323.40	324.14
GDP	2	345.76	347.10
e	3	344.43	346.38
U10a	4	369.16	371.74
	0	253.53	253.68
Qm	1	319.30	320.04
GDP	2	388.51	389.85
e	3	398.14	400.09
U12a	4	423.45	426.03
	0	279.37	279.52
Qm	1	322.87	323.61
GDP	2	326.82	328.16
e	3	335.50	337.45
U14a	4	360.44	363.01
	0	227.66	227.81
Qm	1	323.80	324.54
GDP	2	346.27	347.61
et	3	345.37	347.33
U10a	4	371.21	373.79
	0	253.44	253.58
Qm	1	319.78	320.52
GDP	2	388.49	389.83
et	3	398.16	400.11
U12a	4	425.24	427.81
	0	279.40	279.55
Qm	1	323.48	324.22
GDP	2	327.63	328.97
et	3	335.75	337.70
U14a	4	361.98	364.56

TABLE A.5. – CAUSALITY TEST

VAR. X	*VARIABLE Y*						
	Qm	*GDP*	*e*	*et*	*U10a*	*U12a*	*U14a*
Qm	-	0.002	1.143	0.944	2.552	0.117	0.395
		(0.965)	(0.290)	(0.335)	(0.116)	(0.733)	(0.532)
GDP	0.341	-	0.340	0.224	22.568*	7.502*	0.012
	(0.561)		(0.562)	(0.637)	(0.000)	(0.008)	(0.913)
e	1.402	0.280	-	0.016	0.515	0.551	0.309
	(0.242)	(0.598)		(0.897)	(0.476)	(0.461)	(0.581)
et	2.05	0.759	0.281	-	0.825	1.017	0.779
	(0.147)	(0.387)	(0.598)		(0.367)	(0.318)	(0.381)
U10a	0.535	3.342**	0.578	0.128	-	0.623	1.904
	(0.467)	(0.07)	(0.450)	(0.721)		(0.433)	(0.173)
U12a	2.511	0.683	0.130	0.474	10.75*	-	0.226
	(0.119)	(0.412)	(0.719)	(0.493)	(0.001)		(0.636)
U14a	0.827	3.89**	0.260	0.088	34.06*	11.15*	-
	(0.367)	(0.05)	(0.612)	(0.767)	(0.000)	(0.01)	

Note: Probability is shown between parentheses. H_0: variable x does not result in y. * Rejects at 1%, ** Rejects at 5%.

TABLE A.6. – LONG-TERM ELASTICITIES
OLS ESTIMATION

Variable	*Coefficient*	*Std Error*
GDP	0.660	0.057
e	-0.926	0.113
U10a	-0.692	1.286
Constant	6.277	0.648

Variable	*Coefficient*	*Std Error*
GDP	0.359	0.079
e	-1.125	0.104
U14a	3.966	0.835
Constant	-9.565	3.381

Variable	*Coefficient*	*Std Error*
GDP	0.692	0.052
et	-0.944	0.108
U10a	-0.618	1.233
Constant	6.354	0.614

Variable	*Coefficient*	*Std Error*
GDP	0.406	0.072
et	-1.131	0.096
U14a	3.896	0.785
Constant	-9.259	3.185

Variable	*Coefficient*	*Std Error*
GDP	0.714	0.059
e	-0.969	0.111
U12a	2.040	0.942
Constant	-2.888	4.277

Variable	*Coefficient*	*Std Error*
GDP	0.743	0.056
et	-0.978	0.105
U12a	1.865	0.902
Constant	-2.057	4.111

TABLE A.7 –JOHANSEN'S (1988) COINTEGRATION TEST

Eigenvalue	0.390	0.115	0.075
LR test	37.60	10.89	4.238
Critical value at 5%	29.68	15.41	3.76
H_0: Rank = r	r = 0	r ≤ 1	r ≤ 2

Note: LR test indicates the presence of 1 cointegration vector at 5%. Variables used, *Qm, GDP and et.*

TABLE A.8. – ESTIMATION RESULT
LINEAR VEC(1) MODEL

VARIABLE	*COEFFICIENT*	*STD DEVIATION*
Constant	-0.03	(0.04)
Δqm_{t-1}	-0.156	(0.148)
ΔGDP	-0.524	(2.331)
ΔGDP_{t-1}	2.143	(2.225)
$\Delta U12a$	1.587	(1.349)
$\Delta U12a_{t-1}$	-2.339	(2.203)
Δe	-0.450	(0.117)
Δe_{t-1}	0.076	(0.133)
vec_{t-1}	-0.233	(0.080)

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, and, u12a.*

TABLE A.9. – ESTIMATION RESULT
LINEAR VEC(1) MODEL

VARIABLE	*COEFFICIENT*	*STD DEVIATION*
Constant	-0.005	(0.042)
Δqm_{t-1}	-0.232	(0.142)
ΔGDP	0.175	(2.171)
ΔGDP_{t-1}	0.944	(2.078)
$\Delta U12a$	1.607	(1.243)
$\Delta U12a_{t-1}$	-1.079	(2.054)
Δet	-0.465	(0.109)
Δet_{t-1}	0.035	(0.125)
vec_{t-1}	-0.302	(0.077)

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, et, u12a.*

TABLE A.10. – ESTIMATION RESULT
LINEAR VEC(1) MODEL

VARIABLE	*COEFFICIENT*	*STD DEVIATION*
Constant	-0.031	(0.045)
Δqm_{t-1}	-0.189	(0.146)
ΔGDP	0.149	(2.260)
ΔGDP_{t-1}	1.437	(2.155)
$\Delta U12a$	1.325	(1.313)
$\Delta U12a_{t-1}$	-1.725	(2.127)
Δet	-0.467	(0.113)
Δet_{t-1}	0.059	(0.129)
vec_{t-1}	-0.261	(0.080)

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, et.*

TABLE A.11. – ESTIMATION RESULT
LINEAR VEC(0) MODEL

VARIABLE	*COEFFICIENT*	*STD DEVIATION*
Constant	-0.048	(0.040)
ΔGDP	1.715	(0.698)
ΔU12a	0.344	(0.821)
Δe	-0.417	(0.114)
vec_{t-1}	-0.206	(0.073)

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, and, u12a.*

TABLE A.12. – ESTIMATION RESULT
LINEAR VEC(0) MODEL

VARIABLE	*COEFFICIENT*	*STD DEVIATION*
Constant	-0.022	(0.039)
ΔGDP	1.227	(0.646)
ΔU12a	0.802	(0.731)
Δet	-0.417	(0.106)
vec_{t-1}	-0.260	(0.071)

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, et, u12a.*

TABLE A.13. – INFORMATION CRITERIA MS-VEC MODELS

MODEL	*AIC*	*H-Q*	*SC*
MSIAH(3)-VEC(1)	-3.059	-2.548	-1.734
MSIAH(3)-VEC(2)	-4.911	-4.225	-3.127
MSIAH(2)-VEC(1)	-2.492	-2.180	-1.682
MSIAH(2)-VEC(2)	-1.202	-0.773	-0.087

Here we used the variables real exchange rate without tariff and u12a

TABLE A.14. – INFORMATION CRITERIA MS-VEC MODELS

MODEL	*AIC*	*H-Q*	*SC*
MSIAH(3)-VEC(1)	-2.194	-1.682	-0.868
MSIAH(3)-VEC(2)	-3.354	-2.667	-1.569
MSIAH(2)-VEC(1)	-2.630	-2.318	-1.821
MSIAH(2)-VEC(2)	-1.134	-0.705	-0.019

Here we used the variables real exchange rate with tariff and u12a

TABLE A.15. – INFORMATION CRITERIA MS-VEC MODELS

MODEL	*AIC*	*H-Q*	*SC*
MSIAH(3)-VEC(1)	-1.832	-1.321	-0.506
MSIAH(3)-VEC(0)	-0.940	-0.602	-0.065
MSIAH(2)-VEC(2)	-1.057	-0.628	0.058
MSIAH(2)-VEC(0)	-0.858	-0.660	-0.347

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, et.*

TABLE A.16. – FUNCTIONAL FORMS MODELS

	MS(2)-VEC(1)			*MS(2)-VEC(0)*		
MODEL	*AIC*	*H-Q*	*SC*	*AIC*	*H-Q*	*SC*
MSIA	-0.916	-0.618	-0.143	-0.914	-0.730	-0.439
MSIH	-1.082	-0.884	-0.567	-1.037	-0.896	-0.673
MSI	-0.889	-0.705	-0.411	-0.877	-0.750	-0.549
MSA	-1.127	-0.871	-0.464	-0.838	-0.668	-0.400
MSAH	-1.638	-1.340	-0.865	-0.844	-0.661	-0.370
MSH	-1.106	-0.921	-0.627	-0.884	-0.757	-0.556
LINEAR	-0.721	-0.579	-0.353	-0.760	-0.675	-0.541

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, et.*

TABLE A.17 – FUNCTIONAL FORMS MODELS

	MS(3)-VEC(1)			*MS(3)-VEC(0)*		
MODEL	*AIC*	*H-Q*	*SC*	*AIC*	*H-Q*	*SC*
MSIA	-1.466	-0.983	-0.214	-0.942	-0.632	-0.139
MSIH	-1.875	-1.591	-1.138	-0.944	-0.718	-0.360
MSI	-0.911	-0.655	-0.248	-1.002	-0.804	-0.491
MSA	-1.256	-0.845	-0.188	-0.883	-0.600	-0.153
MSAH	-1.188	-0.706	0.063	-0.917	-0.605	-0.114
MSH	-1.206	-0.950	-0.543	-0.916	-0.718	-0.405
LINEAR	-0.721	-0.579	-0.353	-0.760	-0.675	-0.541

Note: The residual of OLS estimation was obtained with variables *Qm, GDP, et.*

TABLE A.18. – RESULTS OF ML ESTIMATION FOR MSIH(2)-VEC(0)

VARIABLE	*Coefficient*	*Std Deviation*
$Constant_1$	-0.162	0.023
$Constant_2$	-0.035	0.043
ΔGDP	2.767	0.336
ΔU12a	-0.183	0.395
Δet	-0.499	0.046
vec_{t-1}	-0.190	0.036
σ_1	0.052	
σ_2	0.196	

APPENDIX B – QUARTERLY DATA ESTIMATIONS

Table B.1. Unit root test in the level and in differences

	In the level				*In differences*			
	τ	τ_μ	τ_τ	I(.)		τ	τ_μ	τ_τ	I(.)
Qm	0.52	-0.73	-2.58	I(1)	*ΔQm*	-2.25^a	-2.32^c	-2.41	I(0)
GDP	1.80	-0.26	-3.97^c	I(1)	*ΔGDP*	-2.98^a	-3.49^a	-3.49^a	I(0)
Pd	0.70	-1.17	-3.24^c	I(1)	*ΔPd*	-3.07^a	-3.14^a	-3.15^a	I(0)
Icu	-0.66	-2.73	2.85	I(1)	*ΔIcu*	-5.06^a	-5.04^a	-5.02^a	I(0)
e	-0.76	-0.90	-2.69	I(1)	*Δe*	-9.21^a	-9.20^a	-9.17^a	I(0)

Note: τ means no constant, τ_μ means with constant and τ_τ means test with constant and trend. *I(.)* is the order of integration. Rejects at 1%, b at 5% and c above 10%.

Table B.2. – Seasonal unit root test

States	*Aux. Reg.*	$t{:}\ \pi_1$	$t{:}\pi_2$	$t{:}\ \pi_3$	$t{:}\ \pi_4$	$F{:}\pi_3 \cap \pi_4$
Qm	-	0.743	-3.223^a	-2.670^a	-2.040^a	5.646
	I	-0.474	-3.183^a	-2.664^a	-1.992^b	5.538
	I. SD	-0.094	-5.416^a	-5.265^a	-4.550^a	34.519
	I. Tr	-2.448	-3.174^a	-2.701^a	-1.856^b	5.367
	I. SD. Tr	-2.475	-5.311^a	-5.340^a	-4.276^a	32.752
GDP	-	1.414	-1.123	-0.804	-1.013	0.839^a
	I	-0.262	-1.118	-0.803	-1.007	0.832^a
	I. SD	-0.242	-4.516^a	-3.186^b	-3.403^a	10.831
	I. Tr	-3.974	-1.025	-0.848	-0.815	0.694^a
	I. SD. Tr	-2.596	-4.276	-3.322	-2.915	9.753
Pd	-	1.007	-8.516^a	-2.264^a	-1.518^b	3.870^b
	I	-0.880	-8.526^a	-2.294^a	-1.497	3.910
	I. SD	-1.020	-7.109^a	-4.938^a	-3.251^a	21.281
	I. Tr	-2.119	-8.630^a	-2.376^a	-1.445	4.030
	I. SD. Tr	-2.654	-7.279^a	-5.303^a	-3.134^a	23.101
Icu	-	-0.585	-2.113^b	-3.076^a	-0.252	4.764
	I	-3.380^a	-1.930	-3.244^a	-0.175	5.277
	I. SD	-3.268^a	-3.056^b	-3.965^b	-0.017	7.862
	I. Tr	-4.093^a	-1.872	-3.249^a	-0.344	5.342
	I. SD. Tr	-3.988^a	-2.983^b	-4.007^a	-0.134	8.038
e	-	-1.691	-4.177^a	-4.125^a	-6.996^a	47.074
	I	-1.219	-4.183^a	-4.196^a	-6.913^a	47.028
	I. SD	-1.194	-4.081^a	-4.148^a	-6.759^a	45.402
	I. Tr	-1.489	-4.197^a	-4.332^a	-6.663^a	46.907
	I. SD. Tr	-1.482	-4.094^a	-4.286^a	-6.516^a	45.347

Note: The deterministic term is zero (-), one intercept (I), one seasonal dummy variable (SD) and one trend (Tr). a – significance level of 1% and b of 5%. The table can be referred to in Hylleberg et alli (1990). π_1 zero frequency, π_2 biannual e π_3 annual.

TABLE B.3. – Unit root test in the presence of structural break

Test	Model	*Qm*	*GDP*	*Pd*	*Icu*	*e*
UR $t^*_\alpha(i)$	1	-3.74	-3.90	-3.92	-3.75	-4.01
	2	2.75	-3.87	-3.88	-3.73	-3.83
	3	2.43	-3.15	-2.90	-3.24	-3.36
STUD $t^*_{\alpha,\hat{\theta}}$ and $t^*_{\alpha,\gamma}$	1	-3.73	-3.90	-3.92	-3.44	-4.01
	2	-0.57	-3.52	-2.22	-2.69	-3.83
	3	-2.21	-3.12	-2.89	-3.21	-2.48
STUDABS $t^*_{\alpha,\lvert\hat{\theta}\rvert}$ and $t^*_{\alpha,\lvert\gamma\rvert}$	1	-3.74	-3.90	-3.93	-3.44	-4.01
	2	-0.58	-3.52	-2.23	-2.69	-3.83
	3	-2.21	-3.13	-2.89	-3.21	-2.48

TABLE B.4. – Structural break dates

Test	Model	*Qm*	*GDP*	*Pd*	*Icu*	*e*
UR $t^*_\alpha(i)$	1	1992:IV	1989:III	1989:III	1980:III	1998:III
	2	1982:II	1989:III	1989:III	1980:III	1998:III
	3	1979:I	1982:II	1994:III	1980:IV	1997:II
STUD $t^*_{\alpha,\hat{\theta}}$ and $t^*_{\alpha,\gamma}$	1	1992:IV	1989:III	1989:III	1992:IV	1998:III
	2	1995:IV	1983:III	1981:IV	1981:IV	1994:I
	3	1985:III	1987:III	1993:I	1982:IV	1981:III
STUDABS $t^*_{\alpha,\lvert\hat{\theta}\rvert}$ and $t^*_{\alpha,\lvert\gamma\rvert}$	1	1992:IV	1989:III	1989:III	1992:IV	1998:III
	2	1995:IV	1983:III	1981:IV	1981:IV	1994:I
	3	1985:III	1987:II	1993:I	1982:IV	1981:III

TABLE B.5. – SELECTION OF VAR LAG

VARIABLES	*ORDER*	*AKAIKE*	*SCHWARZ*
Qm	1	531.67	532.10
GDP	2	567.62	568.48
Icu	3	586.41	587.71
e	4	614.87	616.61
Qm	1	492.33	492.76
Pd	2	576.92	577.78
Icu	3	591.74	593.04
e	4	596.48	598.23

TABLE B.6. – CAUSALITY TEST

VAR. X	*VARIABLE Y*				
	Qm	*GDP*	*Pd*	*Icu*	*e*
Qm	-	28.84*	19.39*	4.522**	0.781
		(0.000)	(0.000)	(0.03)	(0.379)
GDP	47.06*	-	33.84*	15.90*	0.157
	(0.000)		(0.000)	(0.000)	(0.692)
Pd	41.63*	15.73*	-	29.94*	0.625
	(0.000)	(0.000)		(0.000)	(0.430)
Icu	1.186	12.31*	21.36*	-	0.074
	(0.278)	(0.000)	(0.000)		(0.785)
e	0.753	0.321	0.112	0.656	-
	(0.387)	(0.572)	(0.740)	(0.420)	

Note: Probability is shown between parentheses. H_0: variable x does not result in y. * Rejects at 1%, ** Rejects at 5%.

TABLE B.7. –JOHANSEN'S (1988) COINTEGRATION TEST

Eigenvalue	0.233	0.159	0.107	0.007
LR test	53.84	28.33	11.64	0.755
Critical value at 5%	47.21	29.68	15.41	3.76
H_0: Rank = r	r = 0	r ≤ 1	r ≤ 2	r ≤ 3

Note: LR test indicates the presence of only one cointegration vector.

TABLE B.8. –HANSEN'S (1999) LINEARITY TEST

	SETAR(1) vs. SETAR(2)			*SETAR(1) vs. SETAR(3)*			*SETAR(2) vs. SETAR(3)*			
Series	F_{12}	HB	HBt	F_{13}	HB	HBt	F_{23}	HB	HBt	GHBt
Qm	53.96	0.000	0.000	70.6	0.000	0.000	13.01	0.186	0.363	0.256
GDP	152.7	0.000	0.000	167.5	0.000	0.000	8.325	0.562	0.626	0.575
Icu										
e	170.1	0.000	0.000	211.6	0.000	0.000	22.17	0.209	0.208	0.174

Note: F_{12} is the test statistics given by $F_{jk} = n\left(\frac{S_j - S_k}{S_k}\right)$. HB is the test carried out with the homoskedastic bootstrap method HBt is the heteroskedastic test and GHBt is the general heteroskedastic test, see Hansen(1999).

TABLE B.9. – INFORMATION CRITERIA MS-VEC MODELS

MODEL	*AIC*	*H-Q*	*SC*
MSIAH(3)-VEC(1)	-2.437	-2.048	-1.475
MSIAH(3)-VEC(2)	-2.481	-1.960	-1.191
MSIAH(2)-VEC(1)	-2.245	-2.008	-1.658
MSIAH(2)-VEC(2)	-2.334	-2.008	-1.153

TABLE B.10. – FUNCTIONAL FORMS VEC(1) MODELS

MODEL	*AIC*	*H-Q*	*SC*
MSIA(2)	-2.14	-1.913	-1.579
MSIH(2)	-1.97	-1.819	-1.596
MSI(2)	-1.903	-1.763	-1.556
MSAH(2)	-2.202	-1.975	-1.641
MSH(2)	-1.97	-1.832	-1.625
LINEAR	-1.966	-1.858	-1.699

Appendix C – SOURCE OF DATA

ANNUAL DATA

- Laspeyres' index number for Brazilian imports except oil and wheat (*Qm*): Until 1988, we used total import except oil and wheat. After 1989, the index was stretched, considering the variations in the total amount of imports. The series until 1959 was built by deflating the values in dollars according to the U.S. wholesale price index. The series in dollars was obtained using the amount of oil and wheat in the total of imports in *cruzeiros*. For this calculation between 1947 and 1959, we used the Statistical Yearbook of Brazil and IFS/IMF. From 1960 to 1986, the data correspond to an unpublished series of Getúlio Vargas Foundation (FGV). For 1987 and 1988, the series of FGV was extended by Fachada (1990) using the same method. From 1989 to 2002, the series was extended using the variations of the *quantum* index of total imports of FUNCEX. Source: Statistical Yearbook of Brazil, IFS/IMF, data not published by FGV, Fachada (1990) and FUNCEX.
- Domestic activity indicator (*GDP*): real gross domestic product index. Source: IBGE
- Domestic price index (*WPI*): Wholesale price index, Internal availability (IPA-DI). Source: FGVDADOS.
- Import prices (*Pm*) were obtained in the same fashion and sources as the quantum index (*Qm*).
- Nominal exchange rate index (*e*): Until 1988, we use the exchange rate index Cr$/US$ for total imports. This data shows the cost-effectiveness of buying currency for the import of goods. In addition to the exchange rate, it also includes additional taxes over time, such as premiums, compulsory loans, etc. From 1989 to 2002, only the average annual nominal exchange rate. Source: From 1947 to 1953, SUMOC bulletins. From 1953 to 1957, weighted average of all categories of import tariffs presented in Simonsen (1961). From 1957 to 1960 the exchange rate cost was obtained from Von Doellinger *et alii* (1977). From 1960 to 1976 the exchange rate cost was taken from Clark and Weisskof (1967). From 1967 to 1974 the exchange rate was obtained from Central Bank bulletins. From 1975 to 1979 the exchange rate cost was calculated by Rosa *et alii* (1979). From 1980 to 2002, we used the exchange rate informed on the bulletins of the Central Bank.
- Average tariff (*t*): calculated using the ratio between import tax and total imports. Source: IBGE, Receita Federal and MDIC.
- Installed capacity utilization (*U13a*). Three other series were later created, *U10a*, *U12a* and *U14a*, according to Portugal(1993b). Source: Industrial research by Getúlio Vargas Foundation, FGVDADOS.

QUARTERLY DATA

- Domestic activity indicator (*GDP*) – Quarterly series of fixed basis index of the real gross domestic product (1994/IV=100) for the years 1978 and 1979 according to the GDP growth rates registered by NAPE. Source: FGVDADOS.
- Quarterly *quantum* number of Brazilian imports (*Qm*) fixed basis (1994/IV=100). Source: FUNCEX.
- Installed capacity utilization (*Icu*), quarterly series that measures the installed capacity utilization of the general processing industry of Brazil. Source: Industrial research of Getúlio Vargas Foundation.
- Real exchange rate index (*e*), for total imports. Source: FGVDADOS. The WPI-IAI of Getúlio Vargas Foundation and the WPI-USA of the Bureau of Labor Statistics were used.