Dynamic Structural Models and the High Inflation Period in Brazil: Modelling the Monetary System

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Abstract

In this paper we develop a linear, structural, dynamic, econometric model for the high inflation period in Brazil. The main goal is to obtain a parsimonious model that accounts for a complex dynamic present in the monetary system during the period describing the relationships among output, inflation rate, interest rates and real money. We start the analysis after the Cruzado plan cast in 1986 following a progressive strategy in deriving the econometric model. The results show that we can identify a long run money demand equation and the model describes parsimoniously and in detail the relationship among the variables despite all the instability present in the second half of the 1980's in Brazil with special attention to the role played by nominal wage inflation in determining the dynamics observed in price inflation.

Keywords: VAR, Cointegration, Money Demand, Simultaneous Equation Models

JEL Classification: E31, E41, C32

Resumo

O presente artigo desenvolve um modelo estrutural linear dinâmico para o período de alta inflação no Brasil. O objetivo é obter um modelo estrutural parcimonioso que leve em consideração a complexidade da dinâmica presente no sistema monetário no período em questão descrevendo as relações entre o estoque de moeda real, o produto, taxas de juros e a taxa de inflação. Para tal, utiliza-se o plano Cruzado como ponto inicial da amostra, seguindo uma estratégia progressiva ao especificar o modelo econométrico. Os resultados mostram que podemos identificar uma curva de demanda por moeda no longo prazo para o período, sendo que o modelo proposto descreve de maneira parcimoniosa o fenômeno no período detalhando as relações entre as variáveis modeladas a despeito da instabilidade presente na segunda metade da década de 80 no Brasil com atenção especial para o papel desempenhado pela inflação de salários nominais enquanto determinante da dinâmica observada na inflação de preços.

1. Introduction

The empirical analysis of the money demand in Brazil has received some interest in the past given the relative sophistication observed in the monetary system. If we consider Cagan's classical definition of hyperinflation, Brazil experienced a very short lived hyperinflation from December 1989 to March 1990, a period which is not even close to the shortest hyperinflations in Europe studied by Cagan (Austria, Greece and Poland) which lasted for 17 months. Nevertheless if we use the definition of high inflation as in Fisher (2002) namely those periods where the annual rates crosses 100% and only ends when it stays below 100% for more than one year, then the high inflation period lasted 15 years and 2 months (between April 1980 and May 1995) and the accumulated inflation rate for the period is 20,759,903,275,651% as noticed in Franco (2004).

Remarkably the literature on this subject does not seem to present a detailed analysis of such phenomenon, which is comparatively even rarer than hyperinflations. One strand in the literature has dedicated attention to develop empirical models for the money demand or monetary sector in Brazil rather than testing Cagan's model adequacy and includes Cardoso (1983), Gerlach and de Simone (1985), Calomiris and Domowitz (1989), Fadil and Mac Donald (1992).

A second approach in the empirical literature has been testing Cagan's model adequacy in describing the demand for money in Brazil, as in Phylatkis and Taylor (1993), Engsted (1993b), Rossi (1994), or a variant of the theoretical formulation, as in Feliz and Welch (1997) and Tourinho (1996).

Whilst most of the evidence in this vein has been favourable to the Cagan model, its relative simplicity in describing the money demand as a function of expected inflation does not allow a more detailed analysis of the long-run relationships present in the sector. The Cagan model was originally proposed to describe short periods of very high or explosive inflation rates, whereas the history in Brazil has shown a different phenomenon, namely high inflation over long periods, which turns its use questionable in empirically modelling the money demand. Furthermore, Phylatkis and Taylor (1993) and Engsted (1993b) concentrated their attention on a period when actual inflation rates were moderate, with both samples ending in 1986. Such restriction is also present in other papers that were devoted to studying the Brazilian case, as Juselius (2002), Durevall (1998), Feliz and Welch (1997), all of them imposing 1986 as the ceiling point in the sample length.

The findings in Juselius (2002), of a stable liquidity ratio and a long-run relationship where prices grow less than proportionally to the expansion of M3, only reinforces the argument that simply testing the adequacy of the Cagan model to the money demand in Brazil and arguing that it adequately describes the data,

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is a procedure that leaves out subtle economic relationships which could only be explored in a deeper econometric analysis.

An alternative methodology is adopted in this paper based on Hendry and Richard (1982), Cati et al. (1991), Hendry and Doornik (1994), Hendry and Mizon (1993) and Hendry (1995), being described in detail in Mizon (1995). The core of the analysis is to explore the assumption that valuable information in econometric modelling can come from different sources as economic theory, economic history of the period studied, as well as how data is defined and measured. A progressive strategy in the sense that we do not assume the knowledge of the complete economic structure that links the economic variables, or more specifically, that the theoretical model coincides with the DGP is followed in selecting the final model. We search for a dynamic structural model (SEM), starting from a general congruent and linear vector auto regression (VAR) which constitutes a basis for inference.

The main objective of this paper is precisely to explore the high inflation period in Brazil by constructing a small econometric model for the monetary sector. We explore the generalized indexing present, from prices to wages, assuming that it was the main force behind the inflationary spiral that started in 1986. A theoretical model based on Novaes (1991, 1993) addressing the indexation from prices to wages is tested on the data through imposing the restrictions implied by the theoretical model on the econometric model.

The paper presents a contribution to the literature, not only because it explores in more detail the period that cover the stabilization plans in Brazil, but also because it uses a methodological approach that allows subtle economic relationships to be drawn from the data. We concentrate our attention on the period that runs from March 1986 until July 1994 when the Real plan imposed the new stable currency. The relative lack of attention in the literature to the period posterior to the Cruzado plan cast in 1986 motivated our choice. The paper is divided in the following manner: In Section 2 we present the methodology and a brief discussion of its implications in the empirical modelling. Section 3 presents the statistical model used in the analysis and results obtained. Section 4 discusses the role of nominal wage inflation and Section 5 concludes. All results were generated using either Pc-Give or Ox.

2. Statistical Model

In this section, we consider the relevant aspects of the statistical model and the modelling strategy. The aim is to show in more technical details how we derive a linear, dynamic, structural econometric model (SEM) for the Brazilian data, starting from a congruent econometric model that is considered as a basis for inference.

The departure point of our analysis is a vector of stochastic variables which have a joint density function given by: $D_z(\mathbf{Z}_T^1/\mathbf{Z_0}, \Lambda_t^1, \theta)$, where the density for the vector $\mathbf{Z_t}$ comprising M variables is conditional to a set of initial values, $\mathbf{Z_0}$, and with θ

representing the parameters of interest. The joint density can be rewritten as a set of sequentially conditional densities through a sequential factorization given by:

$$D_z\left(\mathbf{Z_T^1}/\mathbf{Z_0},\theta\right)=\Pi_{t=1}^{T}D_z\left(\mathbf{z_t}/\mathbf{Z}_{t-1},\theta\right) \tag{1}$$

Such a conditioning process assumes that the joint density is a statistical representation of the economy and indeed that the vector stochastic process $\mathbf{Z_t}$ represents the Data Generation Process (DGP).[1] It is worthwhile to notice that by mapping the economic mechanism, namely all the agent actions of each single agent in a span of time, into a joint density of a vector of stochastic variables, comprises a considerable reduction since we are assuming that (1) is a statistical representation of all actions in the economy. Notice that despite this, the DGP is indeed still unmanageably large.

The DGP has this property at this stage because the economy is represented as a system where, ultimately, all the economic variables represented in the stochastic vector $\mathbf{Z_t}$ are endogenous and determined by interactions among each other. Such representation needs further reductions since usually the sample size of the macroeconomic time series does not allow the estimation of these large econometric models that would represent the DGP. Furthermore, the DGP is assumed to be unknown since the observed variables which constitute the macroeconomic time series available for modelling are, in general, aggregations of the original set of variables in $\mathbf{Z_t}$ across time and individuals. This aggregation implies that some level of marginalization is inevitable in the sense that we do not have their correct representation, and such marginalization does not necessarily mean that we can assume that any marginalization is an adequate representation of the DGP. Indeed, it is at this point that the approach followed here has its strength. Any final model resulting from the analysis here is subject to test and will provide evidence of the adequacy of the reduction from the DGP that it assumes to be representing. Considering that the final model is statistically congruent, theoretical hypotheses are subject to testing and not simply subject to empirical validation.

We assume therefore that the set of relevant variables $\mathbf{Y_1^T}$ with $p<m$ is derived by marginalization from Equation (1) in such way that we have a partition of $\mathbf{Z}_T^1$ into $\mathbf{Z_T^1}=(\mathbf{W}_T^1,\mathbf{Y}_T^1)$ and further marginalization of (1) into:

$$D_z\left(\mathbf{Z}_T^1/\mathbf{Z}_0,\theta\right)=D_{w/y}\left(\mathbf{W}_T^1/\mathbf{Y}_T^1,\mathbf{W}_0,t_t\right)D_y\left(\mathbf{Y}_T^1/\mathbf{Y}_0,\Lambda_t^1,\phi\right)$$

This marginalization therefore, defines the Haavelmo distribution as:[2]

[1] Assumes that $\mathbf{Z_t}$ is a $n\times 1$ vector. The vector stochastic process $\mathbf{Z_t}$ is formed then by $\mathbf{Z}(t)\equiv[Z_t(t),Z_2(t),\ldots,Z_n(t)]$ where each one of the $Z_n(t)$ is a stochastic process, namely, for a given probability space define the function $Z(.,.):S\times T\rightarrow\Re$. The ordered sequence of random variables $\{Z(.,t),t\in T\}$ is then called a stochastic random process. In this case for each $t\in T$ we have a different random variable and for each $s\in S$ we have a different realization of the process. In this sense the marginalization imposed by (2) and the restrictions imposed in the econometric model that is assumed to represent (2) either in terms of time-heterogeneity (stationarity) or in terms of the memory of the process ($VAR(p)$) are placed ultimately for tractability.

[2] Such indiscriminate use of the term Haavelmo distribution deserves a more careful definition. According to Cati et al. (1991), a well specified statistical model within which, it is possible to test

$$D_Y\left(\mathbf{Y_T^1}/\mathbf{Y_0}, \Lambda_t^1, \phi\right) = \Pi_{t=1}^{T} D_y\left(\mathbf{y_t}/\mathbf{Y}_{t-1}^{t-s}, \mathbf{D_t}, \lambda\right) \tag{2}$$

where λ represent a set of deterministic variables.[3] The sequence of densities in Equation (2) are conditioned on $\mathbf{D_t}$, representing the use of contemporaneous or lagged information only in deterministic terms, on $\mathbf{Y}_{t-1}^{t-s}$ representing a maximum lag length imposed for tractability, and finally, on $\lambda = f(\phi)$ meaning that the original set of parameters of interest ϕ might contain many transient parameters.

The Haavelmo distribution plays an important role in the sense that it comprises a set of a priori information, or knowledge about the economic events and empirical observations that define the relevant variables in Equation (2). It is also important to notice that its formulation comprises the assumption that it is possible to learn the parameters of interest, μ say, from ϕ alone in the sense that $\mu = f(\phi)$ only.

Furthermore Equation (2) entails a sequential factorization of $\mathbf{Y_1^T}$ in such a way that its right hand side generates a mean innovation process given by:

$$\varepsilon_t = \mathbf{y_t} - E\left[\mathbf{y_t}/\mathbf{Y}_{t-1}^{t-s}\right] \tag{3}$$

Nevertheless a similar sequential factorization of the DGP considering the same lag truncation in (2) generates:

$$\begin{aligned} D_z(\mathbf{Z}_{t-s}^{t-1}/\mathbf{Z}_0, \theta) &= D_{w/y}(\mathbf{W}_{t-s}^{t-1}/\mathbf{Y}_{t-s}^{t-1}, \mathbf{Z_0}, \mathbf{W}_{t-s}^{t-1}, \mathbf{Y}_{t-s}^{t-1}, \theta_{a_t}) \\ &\quad D_y(\mathbf{Y}_{t-s}^{t-1}/\mathbf{W}_{t-s}^{t-1}, \mathbf{Z_0}, \theta_{b_t}) \\ \eta_t &= \mathbf{z_t} - E\left[\mathbf{z_t}/\mathbf{Z}_{t-1}^{t-s}\right] \end{aligned} \tag{4}$$

Therefore, given the partition of the vector $\mathbf{Z_t}$ into $\mathbf{Z}_t^1 = (\mathbf{W}_T^1, \mathbf{Y}_T^1)$ the condition to reducing (1) to (2) without losing information is that: $E[\varepsilon_t/\mathbf{Z}_{t-s}^{t-1}] = E[\varepsilon_t/\mathbf{Y}_{t-1}^{t-s}, \mathbf{W}_{t-1}^{t-s}] = 0$ however we only know that $E[\varepsilon_t/\mathbf{Y}_{t-1}^{t-s}] = 0$. Such conditions imply that θ_a is irrelevant in the estimation and that $D_y(\mathbf{Y}_{t-s}^{t-1}/\mathbf{W}_{t-s}^{t-1}, \mathbf{Y}_{t-s}^{t-1}, \mathbf{Z_0}, \theta_{b_t})$ can be written as $D_y(\mathbf{Y}_{t-s}^{t-1}/\mathbf{Y}_{t-s}^{t-1}, \mathbf{Z_0}, \theta_{b_t})$, or in other words, that the variables in $\mathbf{W_t}$ as per Granger do not cause the variables in $\mathbf{Y_t}$, the former being irrelevant in the analysis of the latter.

According to Hendry (1995) it is at the stage of marginalization that most empirical researches eliminate variables which are potentially relevant.[4]

In contrast to the approach taken in Phylatkis and Taylor (1993), Engsted (1993a) and Rossi (1994), we do follow a progressive strategy, or the LSE

competing structural hypotheses, defines a Haavelmo distribution. (Hendry 1995, p. 406) also defines the distribution as given by specifying the variables of interest, their status, their degree of integration, data transformations in the history of the process and the sample period.

[3] This set of deterministic variables contains seasonal dummies, step dummies, constant, etc and will be exactly defined for each model used.

[4] Interestingly this discussion is also present in the approach called *semistructural* VAR, where according to Canova (1995), the modellers are interested in identifying the behavioural shocks and in predicting the effect of a particular shock on the endogenous variables of the system. Within this approach according to the author there are situations when the observation of current and past values of endogenous variables are not sufficient to achieve the identification of behavioural shocks. Such problem is related to the fact that agents when undertaking their decisions have an information set that is larger than the one available to the econometrician. So potentially there has been some marginalization that throws away relevant variables.

methodology, when deriving any structure [5] from the Haavelmo's distribution in (4.2). The progressive strategy in the present context comprises the specification of a general linear dynamic model from which we impose and test restrictions to derive the econometric model, in doing so we expect to relate the empirical model to the actual mechanism that is generating the data, rather than to theory only, which means that theory and data form the DGP. The character of progressive comes from the fact that when imposing and testing restrictions we are checking continuously if the proposed model (more restricted) can predict the parameters of the more general model.

In particular we consider that the econometric model of interest imposes a set of restrictions on the statistical system represented by (2) which are delineated from the economic theory. The econometric model can therefore be denoted by:

$$\mathbf{f}_y(\mathbf{Y}_T^1/\mathbf{Y}_0;\xi) = \Pi_{t=1}^T \mathbf{f}_y(\mathbf{y}_t/\mathbf{Y}_{t-1}^{t-s},\xi) \tag{5}$$

where $\mathbf{f}(\bullet)$ represents the postulated sequential joint densities. In general since $\mathbf{f}(\bullet)$ is an econometric model it comprises a practical problem how to choose among the many different econometric models can be postulated to represent the joint densities in $\mathbf{f}(\bullet)$.

In our case, we firstly specify a general dynamic model that represents (2) as the Vector Autoregressive Vector (VAR) which assumes the following representation:

$\mathbf{A}(L)\mathbf{y_t} = \varpi\mathbf{D_t} + \vartheta_t$ (6) where $\vartheta_t \sim IN(\mathbf{0},\sum), \mathbf{A}(L)$ is the matrix polynomial in the lag operator such that:

$$\mathbf{A}(L) = \sum_{j=0}^{k} \mathbf{A_j} L^j = \mathbf{I}_p + \mathbf{A}^*(L)L \tag{6}$$

We have in (6) a k-th order system (VAR) because (6) can be rewritten as the following:

$$\begin{aligned} \mathbf{I}_p - \mathbf{A}^*(L)L\mathbf{y_t} &= \varpi\mathbf{D_t} + \vartheta_t \\ \mathbf{y_t} - \mathbf{A}(L)^*\mathbf{y}_{t-1} &= \varpi\mathbf{D_t} + \vartheta_t \\ \mathbf{y_t} &= \mathbf{A}(L)^*\mathbf{y}_{t-1} + \varpi\mathbf{D_t} + \vartheta_t \end{aligned} \tag{7}$$

Finally $\mathbf{D_t}$ is a vector that contains deterministic components as constant, trend, centered seasonal dummies etc. As it stands, the system can be classified as

[5] The concept of structure is defined in (Hendry and Doornik 1994, p. 9) as: "... an entity (structural model) which is to be contrasted with a system having derived parameters (reduced form) and even being a synonym for the population parameter". Structure is also defined later in the same page as: "the set of basic invariant attributes of the economic mechanism". Despite the presence of these two definitions they seem to lead to the same concept which is an econometric model as described in the first definition that presents parameters which are invariant (constant across interventions) and constant (time independent). They may also include agent's decision rules but no assumption of these decision rules being derived from inter-temporal optimization. Within the VAR literature according to (Canova 1995, p. 67): "a model is termed as structural if it is possible to give distinct behavioural interpretations to the stochastic disturbances of the model." Our intention in giving these definitions is to avoid any confusion with the term structural for the model derived in the next chapter and satisfy at least partially a plea for linguistic stability raised by Sims (1991).

complete and closed. Complete in the sense that the number of equations is equal to the number of variables and closed in the sense that all N variables are modelled despite the marginalization in terms of the deterministic variables.

VAR models have been widely applied in empirical econometrics mostly because they can be seen as the empirical counterpart of theoretical models that assumes rational agents in a framework of inter-temporal optimization. Ignoring deterministic factors, the relationships to be modelled are of the type:

$$E\left[\mathbf{A}(L)\zeta_t|I_t\right] = 0 \tag{8}$$

In Equation (8) $\zeta_{n\times t}$ represents a vector of theoretical variables of interest, I_t is the information set and $E[\bullet|I_t]$ is the conditional expectations operator.

Considering that $\mathbf{y_t}$ is the vector of observable variables which adequately describes the variables in ζ_t, only finite lags are involved in (8) and the future expectations do not affect the outcome, the empirical counterpart of (8) is: $E[\mathbf{y}_t|\mathbf{y}_{t-1}, \mathbf{y}_{t-2}, \ldots, \mathbf{y}_{t-k}] = \sum_{i=1}^{k} \mathbf{A_i y_{t-i}}$ which is exactly the VAR described in (7) if we take conditional expectations, and $\vartheta_t = \mathbf{y_t} - E[\mathbf{y_t}|\mathbf{y_{t-1}} \ldots \mathbf{y_{t-k}}]$ is an innovation process to the available information.

According to (Hendry 1995, p. 312): "Although economic theory may offer a useful initial framework, theory is too abstract to be definitive and should not constitute a strait-jacket to empirical research: theory models should not simply be imposed on the data." In this sense we do not pursue a strict analysis of the econometric model by imposing the "strait-jacket" in the model derived in the next section. We do not assume therefore any *a priori* relationship between the variables, in hope that we can gain economic intuition from the data with respect to the demand for money and the monetary system as a whole. Nevertheless, despite we follow a data driven analysis, we do not assume that economic theory has no role in the process, rather it guides the analysis throughout, given that the long run relationships are all identified on the grounds of the theoretical models for money demand. We explore further the role played by economic theory, proposing a model to account for indexation from wages to price and testing the theoretical model on the data through imposing restrictions on the econometric model. Because the theoretical model implies that nominal wage inflation is essential in driving inflation dynamics and originally was not used in the first analysis, we estimate a new system and test for nominal wage inflation exogeneity. Such strategy avoids the assumption of exogeneity based on *a priori* restrictions.

The core of the argument is to follow a progressive strategy, in the sense that knowledge of the economic structures that underlie the economy is not necessary prior to the development of the analysis, however given that the structure exists, it is possible to determine it following this progressive strategy.

Considering that the presence of unit roots in macroeconomic time series is frequent and that in the presence of integrated variables, non-optimal inference might result, the system could be re-parameterized to account for non-stationary behaviour in the data. From Equation 7 we have:

$$\triangle\mathbf{y_t} = \sum_{i=1}^{k-1} \Pi_i \triangle\mathbf{y_{t-i}} + \Pi\mathbf{y_{t-k}} + \varpi\mathbf{D_t} + \vartheta_t \tag{9}$$

where:

$$\Pi_i = -\left(\mathbf{I}_p + \sum_{j=1}^{i} \mathbf{A_j}\right) \text{ and } \Pi = -\left(\mathbf{I}_p + \sum_{j=1}^{k} \mathbf{A_j}\right) = -\mathbf{A} \tag{10}$$

is the matrix of long-run responses.

Notice therefore that we are considering the system as Equation (9), to which the derived SEM is contrasted. The advantage in using (9) is that we can investigate the presence of cointegration between the variables in $\mathbf{y_t}$ by testing the rank (r) of Π, following Johansen (1994).

Apart from non-stationary like trends or level shifts presented in $\mathbf{D_t}$, if Π has full rank, then all variables in $\mathbf{y_t}$ are $I(0)$ stationary, if Π has rank, $0 < r < p$, then there exist linear combinations of variables in $\mathbf{y_t}$ which are stationary; finally, if rank of Π is zero, then all variables in $\mathbf{y_t}$ are $I(1)$ and $\triangle\mathbf{y_t}$ is $I(0)$.

It is worthwhile to note that Equation (9) represents a $I(0)$ parameterization of the system in (7) and it is essential that the system presents a congruent representation of the available information so it can be considered as coherent statistical basis for further assessments.

The class of SEM that we consider here has the form $\Theta\mathbf{f_t} = \mathbf{u_t}$ (10), where Θ is a $n \times N^*$ matrix, $N^* = pk \times r, n$ is the number of restrictions and $\mathbf{f_t}$ is the companion form of (7) given by:

$$\mathbf{f_t} = \Gamma\mathbf{f_{t-1}} + \omega_t \tag{11}$$

[6] where:

$$\mathbf{f_t} = \begin{pmatrix} \triangle\mathbf{y_t} \\ \vdots \\ \triangle\mathbf{y_{t-k+1}} \\ \beta\mathbf{y_{t-k}} \end{pmatrix}_{(pk\times r)\times 1} \quad \text{and } \Gamma = \begin{pmatrix} \Pi_1 & \Pi_2 & \cdots & \Pi_{p-1} & -\alpha\beta' & -\alpha \\ I & 0 & \cdots & 0 & 0 & 0 \\ \cdots & \cdots & \cdots & \cdots & \cdots & \cdots \\ 0 & 0 & \cdots & & \beta' & I \end{pmatrix}$$

$$\text{and } \omega_t \sim IN(0, \Omega)$$

In contrasting the different SEM models, we use the concept of encompassing formalized in Mizon and Richard (1986). Consider two rival SEMs of the form $H_1 : \Theta_1\mathbf{f_t} = \mathbf{u_{1t}}$ and $H_2 : \Theta_2\mathbf{f_t} = \mathbf{u_{2t}}$ which are over identified relative to the congruent statistical system (11). Let τ_2 denote the vector of parameters in Θ_2 and let τ_p be what H_1 predicts τ_2 to be if H_1 were the DGP.

[6] Notice that Equation (5) is an alternative re-parameterization of Equation (4) presented here only to facilitate the notation.

Then H_1 encompasses H_2 if and only if $\tau_2 - \tau_p = 0$. From this condition it is possible to derive that the VAR, (H_1), say, encompasses H_2 once it is congruent, since according to Bontemps and Mizon (2003), a general model (VAR) being congruent is a sufficient condition to encompass all simplifications derived from itself, so H_1 predicts what τ_2 is to be.

In the context of the general to specific modelling strategy, the question of interest is whether the SEM encompasses the system, since, if it does, a simpler model nested within the general model (system) is accounting for the characteristics of a more general model. The VAR then provides the framework within which we access the properties of the SEM.

It is worthwhile to notice that according to Hendry and Mizon (1993) we can circumvent the problem of establishing encompassing theorems about the SEM which are usually difficult because of exogeneity assumptions about rivals SEM may differ, by assuring that the system under analysis is closed and that:

$$E[\mathbf{u_t f'_{t-1}}] = 0 \tag{12}$$

Furthermore the condition in (12) implies that:

$$E[\Theta \mathbf{f_t f'_{t-1}}] = \Theta\Gamma E[\mathbf{f_{t-1} f'_{t-1}}] = 0 \tag{13}$$

Since $\Theta\mathbf{f_t} = \Theta\Gamma\mathbf{f_{t-1}} + \Theta\omega_t = \mathbf{u_t}$. But indeed for condition (13) to be valid, then following condition must be true:

$$\Theta\Gamma = 0 \tag{14}$$

But Equation (14) is indeed the condition that assures the absence of dynamic misspecification in the VAR in the $I(0)$ space ($\Theta\omega_t = \mathbf{u_t}$).

Consequently, if condition (14) holds, so the SEM is congruent, the SEM is a valid reduction of the VAR and parsimoniously encompasses the VAR. Therefore, parsimoniously encompassing the VAR and being congruent is a sufficient condition for the SEM to encompass rival models, (Hendry and Mizon 1993). It is worthwhile to note that condition (14) can be tested given that it coincides with the known condition for the validity of over-identifying restrictions.

Still, nevertheless to be investigated, is the fact that many models could satisfy the encompassing property derived from (14). However, as noticed in Hendry and Mizon (1993) policy regime changes will induce changes in the parameters of the different SEM (Γ_i), destroying the observational equivalence and mutual encompassing which can only be assessed in a constant parameter world, therefore, only the representation that corresponds to the actual structure of behaviours will remain constant. Such assertion is very interesting in our case where a sequence of policy regime changes took effect, so if we can derive a constant SEM for the period it will be unique given the information set. Therefore, any encompassing analysis carried out should account for this restriction which implies in our case that the two SEMs proposed are not directly comparable, in the sense that the information sets differ by including nominal wage inflation in the second model.

3. Empirical Results

The sample data are monthly/seasonally unadjusted for the period 1986 (2) to 1991(12). The period between 1992(1) and 1994(7) is used for out-of-sample forecasts and assessment of the model congruency to the sample data. $m1$ is the log of M1, defined as paper money held by the public plus demand deposits, *cpi* is the log of the consumer price index, *ip* is the log of the industrial production index both as defined in Juselius (2002), [7] *be* is the log of the bill of exchange interest rate to the payee and finally, *cdb* is the log of the interest rate paid in the certificate of deposits. We justify the use of these two interest rates based on previous analysis of Cardoso (1983), and Gerlach and de Simone (1985), whose findings indicated that the bill of exchange interest rate was relevant in modelling the money demand in Brazil. Given these definitions, we construct the real money series as: $m1 - cpi$, the inflation rate as: $\triangle cpi$, where $\triangle$ stands for the first difference operator. Figure 1 contains full sample time plots of the modelled variables: $m1 - cpi, ip, be, cdb, \triangle cpi$.

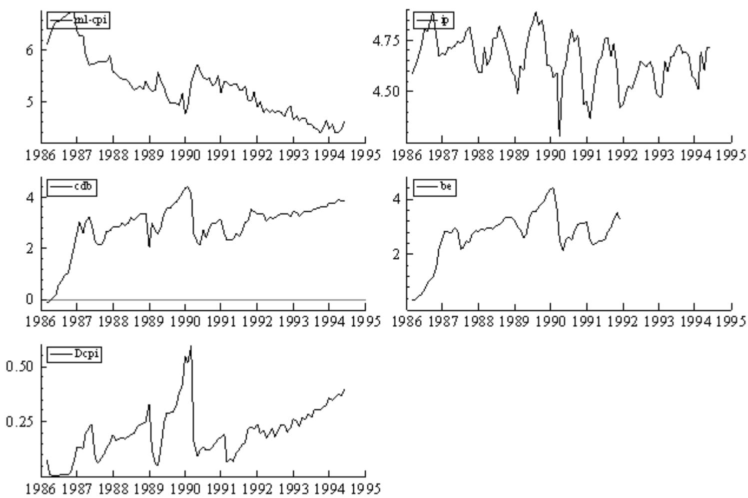

Fig. 1. Full sample time plots

In Tables 1 and 2, we present the Augmented Dickey Fuller [8] and Phillips Perron unit root tests, respectively, for the sample. The presence of stabilization plans comprises a further challenge to the analysis since the series presents a

[7] The author would like to thank Katarina Juselius for providing the dataset used in Juselius (2002). The source for the *CDB* and *BE* is the Institute of Applied Economic Research (IPEA) at: `www.ipeadata.gov.br`.

[8] In defining the lag length for the ADF we follow Maddala and Kim (1998) and use the general to specific rule starting with a lag length of 6 and test the significance of the last coefficient reducing the lag iteratively until a significant statistic is encountered. According to the authors such rule is preferable to other criteria as AIC or BIC.

sequence of structural changes in their levels. According to Cati et al. (1999) the presence of structural changes should bias the conventional unit root tests towards a non-rejection of the null of unit root. When analysing inflation in Brazil, Cati et alii found exactly the opposite, namely a bias towards a rejection of the null hypothesis, what led the authors to develop an alternative unit root test that accounts for this bias. Nevertheless, in our sample there is no evidence against the null hypothesis of unit root – the alleged bias found in Cati et alii – that justifies the use of the corrected tests despite the sequence of stabilization plans as Tables 1 and 2 show. As a general conclusion, the evidence found in the data indicates that they are non-stationary except for the industrial production index. [9]

Table 1
Augmented Dickey Fuller test (1986/4-1991/12)

Variable	Lag	t-value	Critical value (5%/1%*)	Constant	Trend/Seasonal
$m1-cpi$	0	-1.981	-3.44/-4.02	yes	yes/no
ip	0	-4.388	-3.44/-4.02	yes	yes/yes
cdb	0	-2.656	-3.44/-4.02	yes	yes/no
be	0	-2.271	-3.44/-4.02	yes	yes/no
$\triangle cpi$	0	-2.752	-3.44/-4.02	yes	yes/no

*The asymptotic critical values are as tabulated in Maddala and Kim (1998) for a sample size of 100 observations.

Table 2
Phillips Perron Test (1986/4-1991/12)

Variable	Lag	Z statistic	Critical value (5%/1%‡)	Constant	Trend
$m1-cpi$	1	-9.232	20.7/27.4	yes	yes
ip	3	-61296.52	20.7/27.4	yes	yes
cdb	1	-11.038	20.7/27.4	yes	yes
be	1	-9.989	20.7/27.4	yes	yes
$\triangle cpi$	1	-18.948	20.7/27.4	yes	yes

‡The asymptotic critical values are as tabulated in Maddala and Kim (1998) for a sample size of 100 observations.

We estimate initially a VAR (2) with the following variables $m1-cpi, ip, cdb, be$ and $\triangle cpi$. The VAR also included centered seasonal dummies, an unrestricted constant and a restricted trend, so we avoid the unlikely presence of a quadratic trend in the levels. The model further includes the following unrestricted dummies that were significant in the test statistics:

[9] We do not discard the possibility of the data being well described as $I(2)$ as well, however this hypothesis is investigated using a system rather than a univariate analysis.

$D3 = \begin{cases} 1 \;\; t = 1989(1) \textbf{ to } 1989(4) \\ 0 \qquad\quad \textbf{otherwise} \end{cases}$ corresponding to period when the Summer plan actually took place;

$dfm(3) = \begin{cases} 1 \quad 1989(1) \\ 0 \; \textbf{otherwise} \end{cases}$ corresponding to first month of which the Summer plan actually took place;

$dfma(4) = \begin{cases} 1990(6) \\ \quad 0 \qquad \textbf{otherwise} \end{cases}$ corresponding to the first month after the end of the Collor plan (fourth plan);

$dfm(4) = \begin{cases} 1990(3) \\ \quad 0 \qquad \textbf{otherwise} \end{cases}$ corresponding to the first month of which the Collor plan actually took place. [10]

Table 3
Diagnostic tests VAR 1986(4)-1991(12)

Test/Equation	$m1 - cpi$ (p-value)	ip (p-value)	cdb (p-value)	be (p-value)	$\triangle cpi$ (p-value)	System (p-value)
AR 1-5	1.15	2.91*	0.53	1.09	5.53**	1.18
	(0.35)	(0.02)	(0.75)	(0.38)	(0.00)	(0.23)
Normality	0.56	2.91	2.91	6.75	6.48*	13.79
	(0.75)	(0.23)	(0.23)	(0.03)	(0.03)	(0.18)
ARCH	0.25	0.47	0.10	0.47	0.72	
	(0.93)	(0.79)	(0.99)	(0.80)	(0.60)	—
Hetero	0.52	0.34	0.23	0.21	0.66	0.25
	(0.93)	(0.99)	(0.99)	(0.99)	(0.83)	(1.00)

*indicates rejection at 5% level and **at 1% level.

[10] The precise dates observed for the plans are the same as in Cati et al. (1999) paper. The figures obtained here were derived from initially setting the following dummies: $D(i), dfm(i)$ and $dfma(i)$ for each one of the plans (Cruzado, Bresser, Summer, Collor and Collor I, where $i = 1, 2, 3, 4$ respectively) where:

$D(i) = \begin{cases} 1 \text{ for those months when the plan took place} \\ 0 \text{ otherwise} \end{cases},$

$dfm(i) = \begin{cases} 1 \text{ for the first month of the plan beginning date} \\ 0 \text{ otherwise} \end{cases}$ and

$dfma(i) = \begin{cases} 1 \text{ for the first month after the plan ending date} \\ 0 \text{ otherwise} \end{cases}.$

In Table 3 we present the diagnostic statistics for the system. The individual diagnostic tests for the system show the presence of autocorrelation and non normality in the residuals for the $\triangle cpi$ equation and autocorrelation in the residuals of the equation for the industrial production index. In contrast, at the system level the tests suggest that there is no departure from the null hypothesis of no autocorrelation, normality and homoskedasticity. Because Pc-Give performs the individual tests using the system residuals, as they were the residuals for each one of the individual equations, the interpretation of the individual tests is compromised. At best, according to Doornik and Hendry (1997), the individual tests are usually valid only when the remaining equations are problem free, so we decided to carry out the analysis based on the results of the system test.

3.1. *Cointegration Analysis*

We based our cointegration inference on the trace test statistics and on the analysis of the eigenvalues of the unrestricted companion matrix. The trace test statistics are presented in Table 5. Since we are dealing with a relatively small sample, we decided to include the small sample correction as proposed in Johansen (2002),[11] also known in the literature as Bartlett corrections. Both the corrected and the trace test statistics led to the same result, namely, not rejecting the null of $r = 3$, indicating therefore, that we have three cointegrating vectors. In contrast, the eigenvalues of the companion matrix (unrestricted) presented in Table 4 admit an opposite interpretation, because the third eigenvalue of the companion matrix seems to be a bit far from one which would suggest that the system has only two unit roots and possibly only two cointegrating vectors.

Table 4
Five largest eigenvalues companion matrix

Eigenvalues r unrestricted	Eigenvalues $r = 3$
0.8898	1.0000
0.8898	1.0000
0.6519	0.7154
0.6519	0.5299
0.4425	0.5299

Considering that:

a) the trace test results and the eigenvalues of the companion matrix analysis are non conclusive;
b) the period presents several structural breaks imposed by the sequence of stabilization plans and that these breaks could modify the Johansen's test

[11] The trace test statistics and the corresponding small sample corrections were generated using CATS in RATS, version 2 since the PC-GIVE does not support such corrections.

Table 5
Cointegration statistics VAR (1986/4-1991/12)

R	0	1	2	3	4	5
Trace test	193.87	114.23	63.09	23.67	3.54	
p-value	0.000	0.000	0.000	0.091	0.80	
Corrected	172.64	103.57	55.09	20.72	3.29	
p-value	0.000	0.000	0.002	0.194	0.832	
Eigenvalue		0.693	0.534	0.445	0.259	0.052

statistics, we decided to further investigate the hypothesis of $r = 3$ accounting for the presence of structural breaks, by testing for cointegration in the presence of structural breaks, following Johansen, Mosconi and Nielsen (2000), henceforth JMN (2000).

We based our model on the JMN (2000) H_1 model which allows for different linear trends in each sub-sample under analysis, or in other words, broken linear trends. [12] Unfortunately, exact distributions of the test are not known and asymptotic distribution approximations are needed. Quantiles of the distribution are calculated by simulation in JMN (2000), and approximated by the first two moments of the Γ-distribution. Following JMN (2000), let $v_j = \frac{t_j}{T}$ denote the break points as a percentage of the full sample and q denote the number of sample periods. For the case of $q = 3$ we can construct three relative sample lengths, $v_1 - 0, v_2 - v_1$ and $1 - v_2$, let a be the smallest and b the second smallest of these relative sample lengths, the mean and variance of the Γ-distribution, which approximates the unknown distribution quantiles, can be computed by the following:

$$\text{mean} \cong \exp[f_{mean}(p - r, a, b, \infty)] - (3 - q)(p - r)$$

where:

$$\text{variance} \cong \exp[f_{variance}(p - r, a, b, \infty)] - 2(3 - q)(p - r)$$

[12] We omit technical details of how JMN(2000) models are defined and concentrate our attention on the test implementation itself, considering that this paper is ultimately an empirical exercise. Model H_1 is defined by JMN (2000) as Equation (2.2) in their paper. In practice we estimated Equation (2.6) of JMN's paper over the whole sample, namely 1986(2)-1991(12), and tested for cointegration using the trace statistics. The test statistics are in last row of Table 6.

$$
\begin{aligned}
f_{mean} = {} & 3.06 + 0.456(p-r) + 1.47a + 0.993b - 0.0269(p-r)^2 \\
& -0.0363(p-r)a - 0.0195(p-r)b - 4.21a^2 - 2.35b^2 \\
& +0.000840(p-r)^3 + 6.01a^3 - 1.33a^2b + 2.04b^3 - 2.05(p-r)^{-1} \\
& -0.304a(p-r)^{-1} + 1.06b(p-r)^{-1} + 9.35a^2(p-r)^{-1} \\
& +3.82ab(p-r)^{-1} + 2.12b^2(p-r)^{-1} - 22.8a^3(p-r)^{-1} \\
& -7.15ab^2(p-r)^{-1} - 4.95b^3(p-r)^{-1} + 0.681(p-r)^{-2} \\
& -0.828b(p-r)^{-2} - 5.43a^2(p-r)^{-2} + 13.1a^3(p-r)^{-2} \\
& +1.50b^3(p-r)^{-2}
\end{aligned}
$$

$$
\begin{aligned}
f_{variance} = {} & 3.97 + 0.314(p-r) + 1.79a + 0.256b - 0.00898(p-r)^2 \\
& -0.0688(p-r)a - 4.08a^2 + 4.75a^3 - 0.587b^3 - 2.47(p-r)^{-1} \\
& +1.62a(p-r)^{-1} + 3.13b(p-r)^{-1} - 4.52a^2(p-r)^{-1} \\
& -1.21ab(p-r)^{-1} - 5.87b^2(p-r)^{-1} \\
& +4.89b^3(p-r)^{-1} + 0.874(p-r)^{-2} - 0.865b(p-r)^{-2}
\end{aligned}
$$

Unfortunately reported results in JMN (2000) are only for three sample periods, corresponding to two breaks, whereas we have five in our case because of the sequence of stabilization plans, namely, Bresser, Summer, Collor I and Collor II. Considering this restriction we opted to carry out the test taking into account only the Summer and Collor I plans. This choice is justified on the grounds that the Summer plan represented the last plan before we actually had undergone explosive inflation rates whereas the Collor I plan was the landmark of the new monetary policy regime implemented by President Collor under a hyperinflation. Considering Cati et al. (1999) definitions, we have the last observation of the first period as February 1989 and of the second period as March 1990. Such dates imply that $v_1 = 0.5142$ and $v_2 = 0.7$. Consequently $a = 0.1858$ and $b = 0.30$.

Test results are presented in Table 6 where the quantiles are calculated using the GAMMA (1999) software which computes the quantiles and p-values for a Gamma distribution using numerical methods. JMN (2000) suggests a sequential procedure, testing the hypothesis: $H_1(0), H_1(1), \ldots, H_1(p-1)$ against the unrestricted model $H_1(p)$. If $H_1(r)$ is the first hypothesis to be accepted, then the cointegrating rank is estimated by r. Following this test strategy the figures in Table 6 indicate that we have $r = 3$ as the first hypothesis to be accepted. This result supports our previous findings regarding the trace test in Table 5 and consequently we carry out our analysis under the hypothesis that the system has three cointegrating vectors.

Imposing the rank condition on the model and re-estimating generates the eigenvalues of the companion matrix presented in the second column in Table 4. The eigenvalues show that the number of unit roots is equal to two with the third eigenvalue being far from the unit, so we rule out the hypothesis of the data being $I(2)$.

Table 6
Cointegration test in the presence of structural breaks

R	0	1	2	3	4
$Mean$	112.85	82.17	55.63	33.17	14.57
$Variance$	179.35	132.99	92.54	57.30	27.38
95% quantile	135.74	102.01	72.34	46.52	24.11
Trace test	230.47	129.55	77.80	41.78	14.59

Table 7 shows the cointegrating vectors after imposing and testing over-identifying restrictions using the LR test. The first cointegrating vector shows the interest rate on the bill of exchange cointegrated with inflation, admitting the interpretation of a Fisher relationship between the nominal interest rate and inflation with a small but significant trend. A possibly explanation for the rate of interest of the certificate of deposit absence is that on several occasions the Brazilian Central Bank changed the scope of the operations with CDBs and BEs. [13]

Nevertheless, the operations with CDBs seem to have been more affected than those with the bill of exchange. As observed in ANDIMA (1997), in 1989 the market share for the CDB underwent a significant reduction since this type of investment could not follow the high, real interest rates, offered by the overnight operations after the change in the monetary correction index in November 1987, and possibly was no longer reflecting the real inflation rates.

The second cointegrating vector represents a long-run demand for real money being positively influenced by increases in the economic activity represented by the industrial production and negatively affected by increases in inflation. Therefore, both coefficients have the correct signal.

Despite all the instability of the period, the long-run relationship appears to be remarkably stable, as can be seen in Figure 2 where we depict the time plot of the three cointegrating vectors.

We consider the identification of this vector that assumes the interpretation of a money demand equation, a remarkable result in the sense that during the 1980s a debate in literature took place with respect to the money demand in Brazil, and authors as Rossi (1989), argued that the money demand was unstable and consequently estimations were not reliable. For the period posterior to the Cruzado Plan, the attempts in the literature were very restricted to a Cagan specification given the high levels of inflation. Nevertheless, using a different approach it was

[13] Indeed several interventions took place during this period and all were valid for both CDB and BE. The first one in January 1986 fixing the minimum period for investment at 90 days at market determined rates or market rates plus monetary correction. Then in February 1986 the period was reduced to 60 days but only for those investments with ex-ante interest rates. In December 1986 the period was extended to 90 days again but the investment could have the same nominal yield as the Brazilian Central Bank Bills, plus negotiable interest rates. These bills were negotiated in the open market and should give a closer nominal correction to the rates of CDB and BE. In November 1987 nevertheless, the CDB and BE rates were again linked to official indexation rates. Finally in May 1989 the minimum period of investment was reduced to 30 days.

possible to identify a long-run money demand equation leading to a much richer analysis than that allowed by the Cagan model, in the sense that the real money demand equation is linked to the level of activity in the economy and not only to inflation or the rate of growth in M1, as in Engsted (1993a). Further, the dynamic properties of the SEM investigated below allow a much more detailed analysis of the monetary system in Brazil. Such results point out that the approach adopted here is more adequate than the Cagan model in empirically describing the phenomenon observed in Brazil.

Table 7
Cointegrating vectors and adjustment coefficients VAR 1986(4) – 1991(12)

Cointegrating vectors	$\hat{\alpha}_i$	$i = 1$	$i = 2$	$i = 3$
	(se)			
CIa: $-0.197be + \triangle cpi_t + 0.001t$	$m1 - cpi$	3.937	-0.179	0.550
		(0.737)	(0.057)	(0.108)
CIb: $m1 - cpi - 5.47ip_t + 4.799\triangle cpi_t$	ip	0	0	0
CIc: $cdb - 4.508\triangle cpi - 1.930ip_t - 0.012t$				
LR test of restrictions	cdb	-7.035	0	-1.321
		(0.781)		(0.133)
Equilibria and Feedback: $\chi^2(9) = 5.205[0.8160]$	be	0	-0.376	-0.221
			(0.522)	(0.049)
Equilibria only: $\chi^2(2) = 0.005[0.997]$	$\triangle cpi$	0	-0.110	0
			(0.014)	

Finally, the third vector has a difficult interpretation. Theoretically, we expect that increases in the real interest rate would lead to a reduction in the economic activity reflected in the industrial production index, but indeed, with this, cointegrating vector increases in the real interest rate would lead to an increase in the economic activity.

Nevertheless, we need to consider here that the Brazilian Central Bank followed a very loose monetary policy during the years of 1986, 1987, and 1990 with negative real interest rates whereas during 1988, real interest rates were close to zero. Such loose monetary policy possibly influenced the long-run relationship expressed by the vector.

The adjustment coefficients show that the output is weakly exogenous for the parameters of the cointegrating vectors. The real money equation reacts to the three vectors but only error corrects to the second one which exactly represents the long run money demand. An interesting result appears in the equation for $\triangle cpi$, where it error corrects only to the money demand.

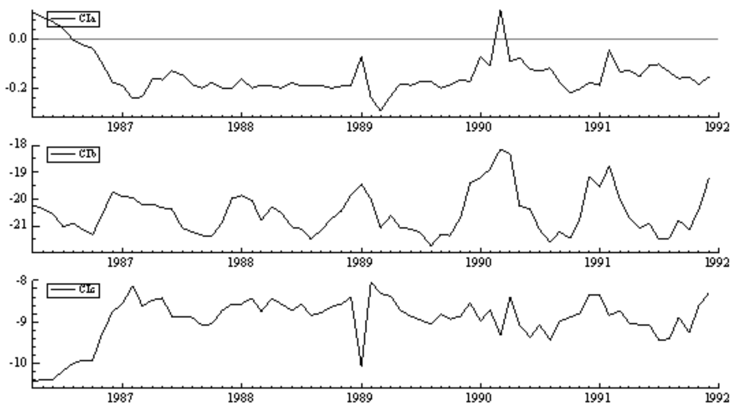

Fig. 2. Cointegrating vectors VAR 1986/4 – 1991/12

Table 8
Cointegrating vectors VAR 1986/4 – 1991/12

Test/Equation	$\triangle m1 - cpi$	$\triangle ip$	$\triangle cdb$	$\triangle be$	$\triangle\triangle cpi$	System
	(p-value)	(p-value)	(p-value)	(p-value)	(p-value)	(p-value)
AR 1-5	0.37	1.44	0.38	0.56	2.46	1.23
	(0.86)	(0.23)	(0.86)	(0.72)	(0.048)*	(0.14)
Normality	1.98	1.98	3.83	7.96	4.24	11.89
	(0.37)	(0.37)	(0.15)	(0.018)*	(0.12)	(0.23)
ARCH	0.40	0.24	0.54	0.36	0.84	
	(0.85)	(0.93)	(0.74)	(0.87)	(0.53)	—
Hetero	0.49	1.17	0.24	0.35	2.24*	0.53
	(0.92)	(0.35)	(0.99)	(0.98)	(0.03)	(1.00)

*indicates rejection at 5% level and **at 1% level.

In the equation for the bill of exchange, surprisingly, there is no reaction to the first vector; whereas, in the equation for the interest rates in the certificate of deposits it reacts to the first and third vectors, the latter reinforcing the rule of the real interest rate based on *cdb* proxied by this vector. A result that is difficult to interpret is the reaction to the first vector since this vector links *be* and $\triangle cpi$ only. In the sequence we estimate a VEC including the three cointegrating vectors. [14]

The misspecification tests presented signals of autocorrelation in the residuals of the equation for $\triangle\triangle cpi$ which led us to reestimate the system excluding

[14] It is worthwhile to note that in both cases the trend was not significant, so the variables present a long-run growth given by the fact that the constant is not restricted to lie on the cointegration space.

$\triangle ip_{t-1}$ from $\triangle \mathbf{y_{t-1}}$ since this variable was not significant in the whole system according to the F-test. Testing the reduction led to the results presented in Table 8. The diagnostic tests show that there are signals of autocorrelation and heteroskedasticity in the equation for $\triangle\triangle cpi$ and non-normality in the equation for $\triangle be$. In contrast, at the system level, the VEqCM appears to be congruent with the information available with normal residuals – no signal of autocorrelation and heteroskedasticity.

Such exclusion constituted a valid simplification in the system, [15] where the number of parameters was reduced from 120 to 115 and therefore constitutes the basis from which the SEM is tested.

3.2. *Econometric model*

The SEM derived imposes a total of 19 restrictions which were not rejected based on the results of the LR test ($\chi^2(19) = 8.90$) and a total reduction of 19 parameters. The diagnostic tests are shown in Table 9 whereas the final SEM is presented in Table 10.

Misspecification test statistics shown in Table 9 indicate that the SEM has no signal of autocorrelation, non-normality and heteroskedasticity at the system level. In contrast, test results at the individual equations level show signals of misspecification in the equation for $\triangle\triangle cpi$, which are restricted to the presence of heteroskedasticity and autocorrelation, with this last test being rejected at a 5% confidence level.

In Figure 3 we present the test for parameter instability based on the Chow break point test and one step residuals. Remarkably, there is no signal of instability in the parameters over the period under consideration. The *ex ante* dynamic forecasts from January 1992 to July 1994 are shown in Figure 4 plus $\pm 2se$. The SEM closely tracks the growth in real money and industrial production index but not so close the rate of growth in the interest rate for certificate of deposit and in inflation, the latter most likely because it is the second difference of the price index.

We qualify this result as remarkable given that between 1986 and 1989 the economy underwent three stabilization plans, one short-lived hyperinflation plus the financial embargo enforced in the Collor Plan. Considering the instability in the period when the president stepped down in the middle of a political crisis, and the upcoming of a new stabilization plan in 1994, the SEM is considered congruent with the information available and parsimoniously encompasses the VAR.

The SEM short-run dynamic shows that the equation for $\triangle m1 - cpi$ is affected positively by changes in the bill of exchange interest rate, a signal that has a difficult interpretation. A possible interpretation lies in the fact that this interest rate is catching the effects of growth in the economy in the presence of low levels of M1 holdings in such way that agents would need more money in their demand accounts

[15] The reduction test led to the following result: SYS(29) $\rightarrow$ SYS(32): $F(5, 41) = 1.7221[0.1512]$.

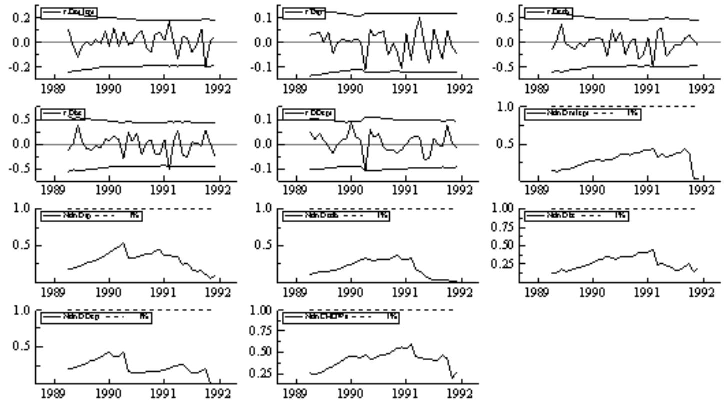

Fig. 3. One step residuals and breakpoint Chow test SEM 1986/4 – 1991/12

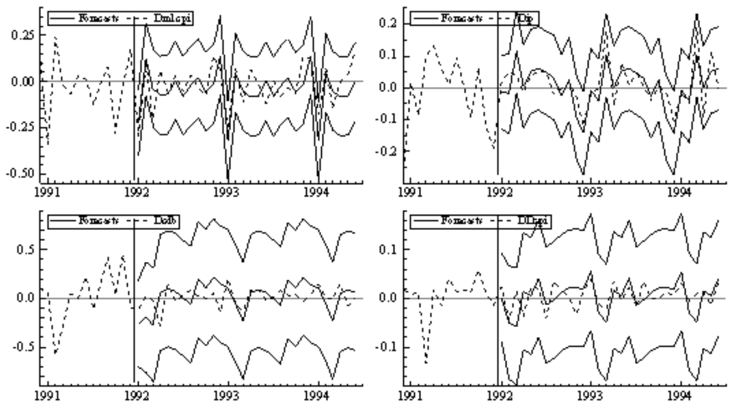

Fig. 4. Out of sample forecasts SEM 1986/4 – 1991/12

to apply in a different asset which would be otherwise protected from inflation in interest bearing accounts.

The sign for the rate of growth in inflation is, as expected, negative. The sign for the rate of growth in the interest rates and for the certificate of deposits seems to reflect simply the substitution between M1 and a fixed income investment which had its maturity period reduced in relation to the previous period, in line with inflation rates growth. Overall, the impact of shocks in $\triangle m1 - cpi$ appears to be restricted to $\triangle m1 - cpi$, only as we can infer from the accumulated impulse response functions in Figure 5.

The equation for $\triangle ip$ shows a negative relationship with the growth rate in inflation and is certainly expressing the nominal impacts of inflation only. It is also represented in the impulse response functions, where the dynamic properties of the

Table 9
Diagnostic tests SEM 1986/4 – 1991/12

Test/Equation	$\triangle m1 - cpi$	$\triangle ip$	$\triangle cdb$	$\triangle be$	$\triangle\triangle cpi$	System
	(p-value)	(p-value)	(p-value)	(p-value)	(p-value)	(p-value)
AR 1-5	0.89	2.07	1.00	1.45	2.63*	1.28
	(0.49)	(0.08)	(0.42)	(0.22)	(0.037)*	(0.09)
Normality	2.63	1.98	3.64	4.63	4.18	11.37
	(0.26)	(0.37)	(0.16)	(0.09)*	(0.12)	(0.32)
ARCH	0.28	0.09	0.76	0.47	0.86	
	(0.91)	(0.99)	(0.57)	(0.79)	(0.51)	—
Hetero	0.54	1.48	0.42	0.53	2.62**	0.61
	(0.91)	(0.16)	(0.96)	(0.91)	(0.0089)	(0.99)

*indicates rejection at 5% level and **at 1% level.

estimated model indicate a negative fast response and posterior stabilization to a shock in $\triangle\triangle cpi$, with the impact reaching approximately half of the standard deviation for the equation to $\triangle ip$. The findings of the cointegrating VAR were not reproduced with the equation for $\triangle ip$ since it reacts to the first cointegrating vector, and not surprisingly, to the third one.

The equation for $\triangle cdb$ basically depicts this variable as negatively related to the rate of growth in real money, reinforcing the interpretation that investing in fixed income would represent an alternative to holding money. The negative signal in the interest rates for the bill of exchange coefficient represents the substituting effect between the two interest rates and finally, the negative signal in $\triangle ip$ has no clear interpretation.

Table 10
SEM 1986/4 – 1991/12
[16]

$$\triangle m1_{(SE)} cpi = 0.94_{0.57} - 0.22_{0.06}\triangle cdb_{t-1} + 0.29_{0.08}\triangle be_{t-1} - 0.63_{0.15}\triangle\triangle cpi_{t-1}$$
$$+3.97_{0.70}CIa_{t-1} - 0.21_{0.05}CIb_{t-1} + 0.52_{0.10}CIc_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.090$$
$$\triangle ip_{(SE)} = -0.69_{0.29} - 0.59_{0.09}\triangle\triangle cpi_{t-1} - 0.49_{0.18}CIa_{t-1} - 0.07_{0.03}CIc_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.056$$
$$\triangle cdb_{(SE)} = -11.50_{0.83} - 0.43_{0.19}\triangle m1 - cpi_{t-1} - 0.13_{0.04}\triangle be_{t-1} - 6.62_{0.47}CIa_{t-1}$$
$$-1.19_{0.09}CIc_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.22$$
$$\triangle be_{(SE)} = -8.32_{0.87} - 0.19_{0.16}\triangle m1 - cpi_{t-1} - 0.34_{0.03}CIb_{t-1} - 0.14_{0.05}CIc_{t-1}$$
$$+\hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.21$$
$$\triangle\triangle cpi_{(SE)} = -2.37_{0.26} - 0.017_{0.010}\triangle cdb_{t-1} - 0.29_{0.07}\triangle\triangle cpi_{t-1} - 0.50_{0.21}CIa_{t-1}$$
$$-0.09_{0.01}CIb_{t-1} - 0.06_{0.03}CIc_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.045$$

[16] The reader should notice that the vector $\mathbf{D_t}$ comprises a different set of variables for each equation. We use the same notation only to save space. Indeed for all equations it comprises the centered seasonal dummies but for the first equation it includes also all dummies. For the second equation it comprises

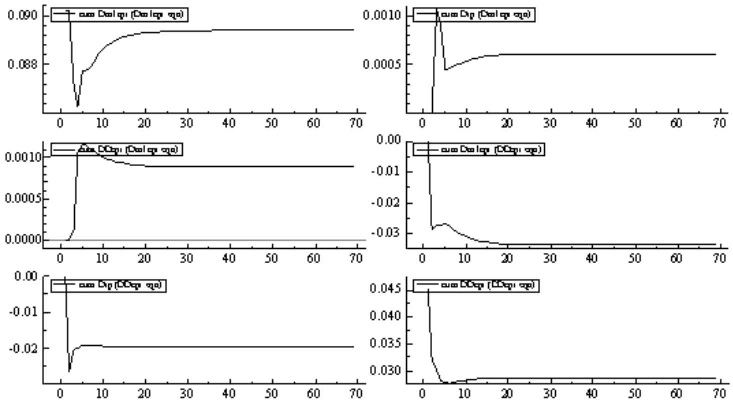

Fig. 5. Accumulated one standard deviation impulse responses functions shock to (from) SEM 1986/4 – 1991/12

The equation for $\triangle be$ has a difficult interpretation since it is not clear why the coefficient for $\triangle m1 - cpi$ should have a negative sign; nevertheless, this variable is only marginally significant, implying therefore that $\triangle be$ is basically driven by the second and third cointegrating vectors and the deterministic variables.

Finally, for the rate of growth in inflation, the negative sign in the $\triangle\triangle cpi_{t-1}$ coefficient depicts the presence of memory in the process but possibly reflecting the sequence of attempts for bringing down inflation. Nevertheless, when we consider the impulse response function, the picture is reversed and shocks on $\triangle\triangle cpi$ present a positive impact which shows the contrasting short-run impacts of the stabilization plans and a long-run effect of increasing inflation rates, given their successive failures. The coefficient for $\triangle cdb_{t-1}$, despite having a difficult interpretation is only marginally significant, which led us to conclude that the equation for $\triangle\triangle cpi$ is basically driven by its past values and the cointegrating vectors. This result seems to be describing the presence of inertia in inflation despite all the attempts to break down this component with the stabilization plans. The cumulating impulse response functions show an initial impact, stabilizing after less than ten periods, in values close to the standard deviation of the $\triangle\triangle cpi$ equation.

Overall, the SEM main strengths are the correct characterization of the growth in real money as a negative function of $\triangle\triangle cpi$ and the negative sign of $\triangle cdb_{t-1}$, in this equation showing the trade off between fixed income and cash holding. The SEM is also able to disentangle the short and long-run effects in $\triangle\triangle cpi$ of a shock in $\triangle\triangle cpi$. In the short run, the SEM correctly describes a negative effect reflecting the sequence of stabilization plans, whereas, in the long run it shows increasing rates which dominated the period.

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the $dfm4$, $dfma4$ and $D3$ dummies. For the third equation it includes only $dfm3$ and $D3$. For the fourth equation the vector comprises all dummies, and finally in the last equation $dfm4$, and $D3$ only.

4. Nominal Wage Inflation

The role played by nominal wage inflation in the Brazilian high inflation is captured in a theoretical perspective by readdressing Taylor's (1979) model as proposed in Novaes (1991). The model is composed by a wage rule, a mark-up price rule for prices, a monetary rule and an aggregate demand equation. The key assumption in modifying the original model is to assume that the wage rule follows a backward adjustment, as well as the monetary rule which accommodates immediate past inflation. Following Novaes (1991), we write:

$$\triangle w_t = \triangle p_{t-1} + \gamma \triangle y_t \tag{15}$$

$$\triangle p_t = \left(\frac{\triangle w_t + \triangle w_{t-1}}{2} \right) \tag{16}$$

$$\triangle m_t = \triangle p_{t-1} + \varphi \left(\triangle p_t - \triangle p_{t-1} \right) \tag{17}$$

$$\triangle m_t = \triangle p_t + \triangle y_t + \varepsilon_t^d \tag{18}$$

In equations (14) through (16), low cases indicate variables in log so that they are expressed in their variation rates given the $\triangle = (1 - L)$ operator. $w$is the nominal wage, y is a measure of demand excess, p is the price index, m is money and ε is a random shock in the demand equation. Equation (15) represents the wage rule, (16) the price rule, (17) the monetary rule and finally, Equation (18), the aggregate demand equation.

The main features of the model are the monetary rule that simply accommodates, in totality, the previous inflation rate but with some discretionary power in such a way that the growth in money is adjusted by the acceleration in the inflation rate in the current period. Further, given this hypothesis, the model does not assume that the agents have rational expectations. Novaes, in her paper, solves the model defining a reduced form equation for inflation as a function of its own past since her objective was test persistence in inflation. We follow though an alternative route, finding a reduced equation form for inflation as a function of w and y in solving the model. We justify that because, originally, Novaes carries out a univariate analysis of time series data, whereas we are interested in the equation for p derived from the SEM in a multivariate context. Solving the model yields then:

$$\triangle p_t = \left(1 - \frac{\varphi}{2}\right) \triangle w_t - \left(\frac{\varphi}{2}\right) \triangle w_{t-1} + \left(-\varphi\gamma - \gamma - 1\right) \triangle y_t + \vartheta_t^d \text{ where } \vartheta_t^d = -\varepsilon_t^d \tag{19}$$

Equation (19) relates inflation to the nominal wage growth rate and the excess in the aggregate demand.

4.1. *Nominal wage inflation and exogeneity*

It is worth noting that testing the theoretical hypothesis represented by the nominal wage inflation poses an extra difficulty in our case regarding the over-identifying restrictions test. In particular, including an extra variable in the SEM and testing encompassing implies an assumption about the variable's status. More specifically, that nominal wage inflation is exogenous for the parameters of interest in the system, namely the cointegrating vector's parameters (not exogenous for the parameters of interest in the SEM). This is because the system (VAR) provides the framework within which we can assess the properties of the SEM. The LR test presented in Section 3 ($\chi^2(19) = 8.90$), tests the hypothesis that the SEM encompasses the VAR considering that the VAR is the unrestricted model on which restrictions are imposed (SEM). Alternatively, in this section we first test the hypothesis that nominal wage inflation is weakly exogenous to the system (the VAR not the SEM). Testing that nominal wage inflation is weakly exogenous allows us to test the weak exogeneity of the relevant variable to the VAR and at the same time to construct an open VAR that corresponds to the unrestricted model.

We use the São Paulo manufacturing industry nominal wage index calculated by the FIESP as a proxy to the nominal wage inflation in Brazil.[17] The test follows Johansen (1994) and is based on the following regression:

$$\triangle \mathbf{v_t} = \varphi \hat{\beta} \mathbf{y_{t-1}} + \sum_{i=1}^{p-1} \Pi_{vi} \triangle \mathbf{y_{t-i}} + \kappa_v + \upsilon_{vt} \tag{20}$$

where we assume a partition of $\mathbf{y_t}$ into $\mathbf{y_t} = (\mathbf{s_t}, \mathbf{v_t})$.

In the present case, the proposed partition is given by conditioning on nominal wage in such a way that if $\varphi = 0$ then we can estimate the system (VECM) conditional to nominal wage inflation, without losses of information. This system will act as a catalyst relative to which we can test restrictions implied by the theoretical model using the over-identified restrictions test. According to Hendry and Mizon (1993), if the variable is validly weakly exogenous, the relevant underlying congruent statistical system is a VAR, even if it is open (not all variables being modelled), as conditioning implies. In this case, we preserve the specification where the SEM arises as a valid reduction of the system as in Section 3 rather than the reduced form being derived from the SEM. If this were the case, the reduced form would have been identified based on incredible restrictions which are in other words, assumptions on the exogeneity of the variables of interest.

We therefore initially estimate a VAR (2) in the following variables: $m1 - cpi, ip, cdb, be$ and $\triangle cpi$ plus $\triangle nw$. The VAR also includes a restricted trend, centered seasonal dummies and the following unrestricted dummies: $dfm2, dfm3, dfm4, dfm5, D1, D3$ and $D4$ as defined in footnote 10.

Table 11 displays diagnostic test results which show some sign of non-normality in two equations and in the system test, nevertheless the test rejects only at 5% significance level but not at 1%, which led us to conclude that the system is reasonably a congruent representation of the data generation process. Furthermore,

[17] Sao Paulo Industry Union, data available from IPEA at `www.ipeadata.gov.br`.

Table 11
Diagnostic tests VAR Exogeneity test (1986/4 – 1991/12

Test/Equation	$m1 - cpi$ (p-value)	ip (p-value)	cdb (p-value)	be (p-value)	$\triangle cpi$ (p-value)	$\triangle nw$ (p-value)	System (p-value)
AR 1-5	0.67	2.73*	0.47	0.33	2.29	1.35	0.96
	(0.64)	(0.0236)	(0.78)	(0.89)	(0.06)	(0.26)	(0.57)
Normality	1.38	1.23	11.41**	17.79**	3.24	3.93	23.05*
	(0.50)	(0.53)	(0.003)	(0.0001)	(0.19)	(0.14)	(0.027)
ARCH	0.15	0.018	0.20	0.12	0.16	0.07	
	(0.97)	(0.96)	(0.95)	(0.98)	(0.97)	(0.99)	—
Hetero	0.22	0.14	0.18	0.13	0.16	0.22	509.88
	(0.99)	(1.00)	(0.99)	(1.00)	(0.99)	(0.99)	(0.86)

the assumption of normality in the residuals is not an essential hypothesis in the cointegration analysis as proposed in Johansen (1994).

In Table 12 we present the modulus of the five largest eigenvalues of the companion matrix and in Table 13 we present the trace test statistics for testing the hypothesis of $r \leq k$. Despite the clear rejection of the hypothesis of the number of cointegrating vectors being less than three in Table 13, the moduli of the eingenvalues in Table 12 suggest that we have only two unit roots with the remaining values being far lower than the first two largest modulus, leading us to assume that we have just two cointegrating vectors in the system.[18] The over-identifying restrictions were not rejected and resulted in the cointegrating vectors in Equations (21) and (22).

Table 12
Five largest eigenvalues companion matrix

Eigenvalues r unrestricted
0.8896
0.8896
0.6713
0.6279
0.6279

[18] We also tested for cointegration in the presence of structural breaks, nevertheless, estimations using the same dates as in Section 3 led to non-conclusive results with non-positive residual's variance in the VAR where cointegration is tested, most likely related to the high instability in the data and the estimation of a system with 6 variables, which led us to ommit test results.

Table 13
Cointegration statistics VAR Exogeneity Test (1986/4 – 1991/12)

R	0	1	2	3	4	5	6
Trace test	276.23	182.02	113.90	54.114	16.043	2.4357	
p-value	0.000	0.000	0.000	0.002	0.497	0.921	
Eigenvalue		0.7447	0.6274	0.5795	0.4240	0.1789	0.0346

$$\triangle cpi - \triangle nw + 0.016be \qquad (21)$$

$$\frac{m1}{cpi} - 2ip - 2cdb + 3.20be - 2\triangle nw \qquad (22)$$

$$X^2(8) = 7.3535$$

Table 14
Nominal wage weak Exogeneity Test
[19]

$$\triangle\triangle nw = 0.210_{0.297} CIa - 0.0376 CIb_{0.0375} + \sum_{i=1}^{1} \hat{\Pi}_i \triangle nw_{t-i}$$
$$+\sum_{i=1}^{1} \hat{\Theta}_i \triangle \mathbf{y_{t-i}} + \hat{\nu}\hat{\sigma} = 0.0868$$
$$F(2,49) = 0.559[0.575]$$

Table 14 presents the results of estimating Equation (20) including the over-identified cointegrating vector in (21) and in (22). Since the F-test clearly does not reject the null, we cannot find evidence in the data against excluding the cointegrating vector and therefore it is possible to conclude that nominal wage inflation is weakly exogenous to the system in the sample. Consequently we proceed therefore with the analysis conditioning on nominal wage inflation.

After concluding that the nominal wage inflation is weakly exogenous, the VAR in the $I(0)$ space is estimated. Estimation results are presented in Table 15 where we focus on the diagnostic tests only. The VAR does not present any sign of non-normality, autocorrelation, heteroskedasticity or ARCH effects in the residuals.

The first cointegrating vector has a similar specification as the first vector in Section 3, showing a long-run relationship between the bill of exchange interest rate and inflation, but now extended to include the nominal wage effect. The second cointegrating vector admits the interpretation of a long run money demand equation which is positively related to the output and a weighted interest rate being discounted by nominal wage inflation. The signs are nevertheless difficult to interpret since the industrial production index, the nominal wage inflation and the CDB interest rate all have the same coefficient.

The second over-identified SEM (M2), derived using Equation (19) as a benchmark, is presented in Table 16. The diagnostic tests in Table 17 show the

[19] Figures below coefficients are standard deviations and inside square brackets are p-values.

Table 15
Diagnostic tests open VAR (1980/1 – 1986/2)

Test/Equation	$m1 - cpi$ (p-value)	ip (p-value)	cdb (p-value)	be (p-value)	$\triangle cpi$ (p-value)	System (p-value)
AR 1-5	0.67	2.24	1.41	0.66	1.10	1.24
	(0.64)	(0.07)	(0.24)	(0.64)	(0.37)	(0.17)
Normality	3.32	1.55	1.10	5.83	4.65	13.56
	(0.18)	(0.45)	(0.57)	(0.054)	(0.09)	(0.19)
ARCH	0.77	0.08	0.98	0.19	0.39	
	(0.57)	(0.99)	(0.44)	(0.96)	(0.85)	—
Hetero	0.29	0.33	0.39	0.45	0.28	0.25
	(0.99)	(0.98)	(0.97)	(0.95)	(0.99)	(1.00)

presence of autocorrelation in residuals in the equations for $\triangle ip$ and $\triangle\triangle cpi$. In contrast, at the system level there is no sign of non-normality or autocorrelation in the residuals, which led us to conclude that the model is a congruent representation of the DGP.

The model consists of the system presented above where we impose restrictions aiming to identify a short run equation such as (19). In particular, the equation of interest is $\triangle\triangle cpi$, where this variable is a function of nominal wage inflation with the same lag structure and signs as Equation (19), and the long-run cointegrating vectors presented in Equations (21) and (22). Such results show the nominal wage relevance in explaining the rate of growth in inflation rate and differently from the SEM derived in Section 3, $\triangle\triangle cpi$ lagged one period, enters in the equation with a positive sign highlighting the short-run increases in inflation dynamics that were present in the period. The positive coefficient in $\triangle\triangle cpi_{t-1}$ contrasts with the SEM in Section 3 where the same variable had a negative coefficient, a difficult result to interpret, that is now clarified by using a theoretical model.

Nevertheless, it should be noticed that Equation (19) is a reduced-form equation from the general model proposed by Novaes (1991), and as it stands, it implies a theoretical equation that imposes a causality direction from nominal wage and a measure of excess demand to prices. In contrast, the present analysis does not assume such causality with respect to the excess demand on the extent that the system used is a VAR wherein the demand variable, namely the industrial production, is modeled.

Similarly to the $\triangle\triangle cpi$ equation, nominal wage inflation enters into the equation for the two interest rates (CDB and BE) with the same lag structure. Further, the industrial production index only enters into its own equation being the money demand, the inflation dynamic, and the two interest rate equations driven basically by the nominal variables (interest rates and price and wage inflation). Such dynamic

Table 16
SEM (M2) 1986/4 – 1991/12

$$\triangle m1_{(SE)} - cpi = 0.13_{0.02} - 0.28_{0.04}\triangle cdb_{t-1} + 0.30_{0.05}\triangle be_{t-1}$$
$$-0.121_{0.126}\triangle nw_{t-1} - 0.14_{0.02}CIb_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.080$$
$$\triangle ip_{(SE)} = 0.02_{0.008} - 0.37_{0.09}\triangle ip - 0.029_{0.019}\triangle be - 0.22_{0.13}\triangle\triangle cpi_{t-1}$$
$$+0.55_{0.11}\triangle\triangle nw + \hat{\varpi}\mathbf{D_t}\sigma = \hat{0}.042$$
$$\triangle cdb_{(SE)} = 0.23_{0.03} - 0.29_{0.07}\triangle cdb_{t-1} - 0.34_{0.11}\triangle be_{t-1} + 1.65_{0.59}\triangle\triangle cpi_{t-1}$$
$$+1.72_{0.34}\triangle nw - 1.60_{0.42}\triangle nw_{t-1} - 2.80_{0.58}CIa_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.187$$
$$\triangle be_{(SE)} = -0.002_{0.03} + 1.05_{0.40}\triangle\triangle cpi_{t-1} + 2.07_{0.21}\triangle\triangle nw - 0.60_{0.30}\triangle nw_{t-1}$$
$$-1.13_{0.39}CIa_{t-1} - 0.24_{0.04}CIb_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.150$$
$$\triangle\triangle cpi_{(SE)} = 0.05_{0.009} + 0.21_{0.08}\triangle\triangle cpi_{t-1} + 0.45_{0.05}\triangle\triangle nw_t - 0.22_{0.05}\triangle\triangle nw_{t-1}$$
$$-0.55_{0.08}CIa_{t-1} + 0.03_{0.009}CIb_{t-1} + \hat{\varpi}\mathbf{D_t}\hat{\sigma} = 0.026$$

The vector $\mathbf{D_t}$ comprises a different set of variables for each equation. Indeed for all equations it comprises the centered seasonal dummies but for the first equation it includes also $dfm3, dfm4, dfm5, D3, D4$, as defined in Chapter 5 and 1989.1. For the second equation includes $dfm2, dfm3, dfm4, dfm5$ and $D4$. For the third equation includes $dfm2, dfm3, dfm4, dfm5$ and 1989.1. For the fourth equation it comprises $dfm3, dfm4, dfm5, D1, D3$ and 1989.1. Finally for the fifth equation it comprises $dfm3, dfm4, D3$ and $D4$.

properties indicate the high level of indexation present in the economy, a contrast result to the model presented in Section 3 and in accordance with the hypothesis of nominal wage playing a central role in inflation dynamics in the period fuelling the consumer price index. Such a conclusion follows, basically because the BE equation is driven by past inflation whereas the money demand and CDB equations are driven by past inflation and the interest rates lagged one period.

The impulse response functions analysis reflects this clear cut between nominal and real impacts in the long run, with one standard deviation, shocks to $\triangle\triangle cpi$ being concentrated to inflation dynamics itself as well as the two interest rates, as Figure 6 shows.

Table 17
Diagnostic tests SEM (M2) (1986/4 – 1991/12)

Test/Equation	$m1 - cpi$ (p-value)	ip (p-value)	cdb (p-value)	be (p-value)	$\triangle cpi$ (p-value)	System (p-value)
AR 1-5	1.96	3.66**	2.29	1.37	2.65*	1.35
	(0.10)	(0.009)	(0.06)	(0.25)	(0.03)	(0.06)
Normality	3.60	0.67	1.84	3.74	3.63	11.19
	(0.16)	(0.71)	(0.39)	(0.15)	(0.16)	(0.34)
ARCH	0.66	0.12	1.01	0.23	0.40	
	(0.65)	(0.98)	(0.42)	(0.94)	(0.83)	—
Hetero	0.45	0.45	0.35	0.31	0.40	0.26
	(0.96)	(0.96)	(0.99)	(0.99)	(0.98)	(1.00)

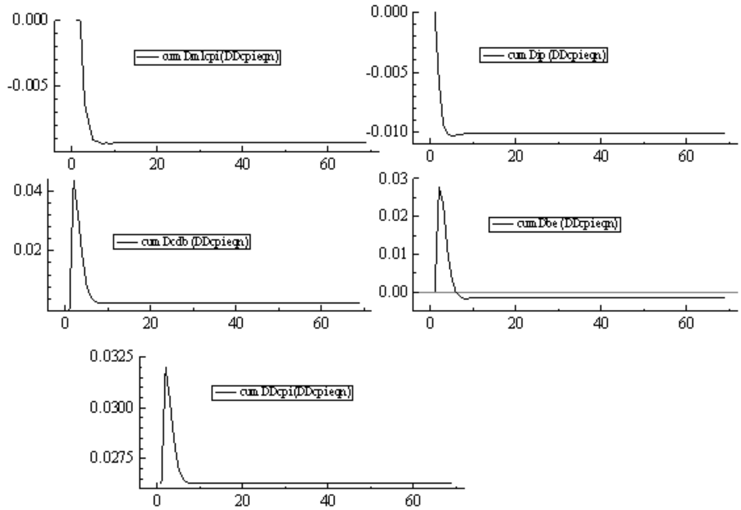

Fig. 6. Cumulative one standard deviation impulse response functions shock to (from) M2 1986/4 – 1991/12

5. Conclusion

The present paper investigates the long-run properties of two small macro econometric models in explaining the demand for money and inflation in Brazil during the period of high inflation in the 1980's. We model the monetary system following a progressive strategy as derived in Hendry and Richard (1982), Cati et al. (1991), Hendry and Doornik (1994), Hendry and Mizon (1993), Hendry (1995). This strategy contrasts to those followed in the literature for the Brazilian case where, in general, the Cagan model or variants of it have been tested. We derive two SEMs which parsimoniously encompass the respective underlying VARs. The models have relatively complex dynamics and despite all the instability represented by the short-lived hyperinflation, three stabilization plans and a political crisis that culminated with President Collor's stepping down in 1992, displayed constant parameters.

The equation for the money demand in the first model shows the rate of growth in real money reacting positively to changes in the output and negatively to changes in the inflation rate. This variable also error corrects to the long run equilibrium, represented by the second cointegrating vector that exactly admits the interpretation of a long-run money demand with real money cointegrating with inflation and output with the corrected signals.

The equation for the growth in the industrial production index shows it reacting negatively to increases in the rate of growth of inflation, an unexpected result but one which has a reasonable interpretation on the grounds that the combination of price freezing and low interest rates present in most of the plans had the effect of generating a rapid growth in the economy in the aftermath of the plan launching; however, in the sequence with the return of the inflation and the end of the consumption bubble, the economy faced a contraction, therefore explaining the apparent contradiction observed.

Finally, the equation for the rate of growth in inflation shows the presence of memory in the process with lagged $\triangle\triangle cpi$ but with a negative sign, possibly representing the several attempts to bring down inflation. The major finding however, is that the rate of growth in inflation error corrects to the second cointegrating vector, which shows that there had possibly been an equilibrium in the level of real money, output and inflation itself in the economy and that departures from this equilibrium had an impact in the rate of growth of inflation. Such result again reinforces the link between inflation and output in the period.

Whilst the first SEM structure is generally in accordance to the theoretical models of persistence in inflation, the first model dynamics in the short run indicated that inflation was negatively related to its own past, a result that might be connected to the sequence of stabilization plans but has no counterpart in theoretical models which usually proposed an inflation dynamics based on the hypothesis of widespread use of indexation, from wage to prices, resulting in an autoregressive pattern. We address this difficult, empirical, model output, proposing and testing a theoretical model on the grounds of the literature developments on this subject.

The extended SEM emphasizes the importance of nominal wage inflation in determining the short run structure and more specifically in the equation for $\triangle ip$ that is now completely determined by the short-run impacts of nominal wage and price inflation, highlighting the demand pressures exerted by wage inflation in driving the growth of industrial activity. Indeed, the short-run model dynamics clearly depict the growth in the industrial production and inflation rate being driven by past inflation (wage and prices), a result that is much more in line with the theoretical models that emphasized the role played by the backward looking mechanism represented by widespread use of indexation. Interestingly, the new SEM presents basically the same structure in the long run as that observed in the SEM derived in Section 3. Since the results from Section 3 point to the inertial pattern of inflation present in the long run, such behaviour only strengthens the conclusion that the backward-looking mechanism of indexation played a significant role in determining the inflation pattern.

More generally, the second model proposed allows that a theoretical model be tested on the data without imposing the *a priori* hypothesis that the data generation process and the theoretical model coincide and, consequently avoiding that, more subtle relationships among the relevant variables be ignored, as would be the case had we simple tested Equation (19) on the data.

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