Dynamic Macroeconometric Modelling:

Evidence on the Brazilian Monetary System

Wilson Luiz Rotatori[1]

Federal University of São João del Rei (UFSJ), Brazil

Jan M. Podivinsky

University of Southampton, United Kingdom

April 2007

Abstract

We explore in an empirical exercise the relative tension in the macroeconometric literature regarding the role played by theoretical models in modelling macroeconomic time series. We criticize the current VAR approach in giving distinctive behavioural interpretations to stochastic disturbances. Instead we propose following an alternative approach based on the 'LSE methodology' in modelling the Brazilian monetary system for the period that precedes the Cruzado plan of 1986. This period has been discussed in the literature mostly through testing the adequacy of the Cagan model to the data. We show that the 'LSE approach' allows a much more interpretable analysis based on two theoretical hypotheses, namely the Administered Exchange Rate Policy and a model for nominal wage inflation. The results show that any negative shock to the economy such as an increase in the oil price would magnify the inflation rate and represent in the short run a change in its level. The economic recovery observed from 1984 onwards seems to have represented this negative role with its associated perverse connotation. The impact on industrial activity originated in the nominal wage inflation, probably because of the accommodating monetary policy implemented in the period. In the short run inflation dynamics were linked to past changes in the industrial activity which then produced the acceleration in the inflation rate.

[1] The first author would like to thank the financial support from CAPES through a doctoral research grant. This paper is based on the first author's PhD thesis.

1) Introduction

Macroeconometric modelling has a long tradition in applied work which dates back to the Cowles Commission's large scale exercises in developing econometric models through the use of simultaneous equations models. In the early 1980's Sims criticized the use of zero restrictions on lags of specific variables included in the model to attain identification. Such restrictions were usually placed without correct attention to the underlying economic structure in such a way that some variables in these models were considered exogenous to the parameters of the model by the use of these restrictions as a result of the identification problem. Furthermore simultaneous equations models typically suffered from several problems including misspecification and predictive failure. Since then much applied work has concentrated on using Vector Auto-regression (VAR) models to model macroeconomic time series exactly because such models bring all variables to be modelled to the status of endogenous variables. In principle if we disregard the fast growth in the number of parameters to be estimated in large scale VARs its use could solve the identification problem present in the use of simultaneous equations models. VARs also represent a well specified statistical basis for testing exogeneity hypothesis on the variables of interest.

Canova (1995) presents the salient aspects of the economic of VAR models and discusses the role played by economic theory in specifying VAR models. In particular he emphasizes that unrestricted VAR models are natural representations of macroeconomic time series which emerge under mild conditions imposed on these time series. Further, solutions for many stochastic dynamic-equilibrium models come in the form of a restricted VAR and the specification of VAR models can be obtained by tying down its parameterization to the underlying theory, or in other words imposing and testing cross-equation restrictions to recover the deep parameters of agents' preferences This is also known in the literature as the hallmark of rational expectations models.

Canova (1995) uses the role played by the theoretical restrictions imposed on VAR models to establish a clear distinction between VAR econometricians and rational expectations ones. The former according to Canova would use a smaller number of restrictions imposed on the VAR than a rational expectation econometrician would use to acquire identification of structural parameters. In this case economic theory is confirmed by the evidence and is endowed with a priori veracity. Such practice in carrying out empirical modelling of macroeconomic time series has been criticized in the literature by Gilbert (1986) and Mizon (1995). In particular, according to the latter, coherence of an econometric model to only economic theory does not guarantee necessarily that such modelling strategy would result in a good econometric model.

Alternatively the London School of Economics (LSE) methodology assumes that valuable information in econometric modelling can come from different sources as economic theory, economic history of the period studied, and 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) is followed in selecting the final model. The concept of structural model is, as a consequence, differs from that associated with the rational expectations program. A model to the rational expectations program is considered structural when its parameters have economic interpretations in terms of preferences and technologies. It is also different from the concept used by the so called VAR econometrician which defines an econometric model as structural "if it is possible to give distinct behavioural interpretations to the stochastic disturbances of the model" (Canova, 1995, p. 67). Rather, to the LSE methodology 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, namely an econometric model which is to be contrasted with a reduced form system that presents parameters which are invariant (constant across interventions) and constant (time independent). Such structure may also include agents' decision rules but there is no *a priori* assumption of these decision rules being derived from inter-temporal optimization. This conceptual difference is essential in the sense that it avoids the selected econometric model selected being used for corroborating the underlying economic theory only, by being also evaluated for congruence to additional sources of information rather than solely to theory[2].

We explore this relative tension regarding the role of *a priori* economic theory in macroeconometric modelling in an empirical attempt to model the monetary system

[2] The congruence to these additional sources of information beyond *a priori* theory can usefully be divided into three categories according to Mizon (1995): observed sample information, the properties of the measured system and rival models.

in Brazil during the period that precedes the introduction of the Cruzado plan in February 1986. The reason for dedicating attention to this period is that it has been explored by different authors who followed different approaches in modelling the Brazilian monetary system. Cardoso (1983), Gerlach and Simone (1985), Calomiris and Domowitz (1989), Fadil and MacDonald (1992) developed empirical models for the money demand or the monetary sector in Brazil but did not test the adequacy of the Cagan (1956) model, whereas Phylatktis and Taylor (1993), Engsted (1993a) and Feliz and Welch (1997) all share the same underlying theoretical model, namely Cagan's money demand specification under hyperinflation. The evidence in the literature regarding Cagan's model has been favourable despite the fact that the period that precedes the Cruzado plan cannot be characterized as a hyperinflationary one. On the contrary, it is exactly in the following period from 1986 onwards that inflation followed an ascending path common to hyperinflations. In particular the exact rational expectations Cagan model is usually rejected in the Brazilian case. This result has support in the fact that rejection of the rational expectations restrictions is present in several hyperinflation episodes (Germany, Austria, Russia, Greece, Yugoslavia and China)[3].

Faced with the problem of rejections the usual escape route in the literature has been to develop measures of the adequacy of the Cagan's model to the data in an attempt to simply corroborate a theoretical model through successive empirical analysis. Such a pattern is also present in the Brazilian case. In particular Rossi (1994) tests the adequacy of Cagan's specification, performing the test using several sub-samples and different prices indexes and seeking confirmation through sequentially running the testing regressions. We propose alternatively to develop a structural model following the LSE methodology while also exploring the role of nominal wage inflation and the Administered Exchange Rate (AER) policy followed by the Brazilian central government.

From November 1980 onwards Brazil's central government adopted a policy for the monetary correction where it would follow the consumer price index, and the exchange rate devaluations would cover the difference between the internal and external inflation rates. Such policy can be thought of as a return to the policy of devaluations in the nominal exchange rate in line with the difference between the internal and external

[3] See Engsted (2002) for a survey about the subject.

inflation rates, as initiated in 1968 and interrupted in 1979/1980 when a new scheme for the monetary correction – including the exchange rate – based on pre-fixed values had been implemented by the new government[4]. Interestingly Durevall (1998) using an identical theoretical model, but looking for the effects of the Purchasing Power Parity on price dynamics in Brazil in the long run find it holds well until 1979 when it then breaks down.

The hypothesis that wage inflation influences inflation dynamics comes directly from the theoretical models developed to account for inflation inertia as in Novaes (1991). Such models are derived from the seminal paper by Taylor (1979) and include a wage equation where the wages follow periodically the past inflation rates. The model is completed with two further equations: one determining the aggregate demand and another defining a monetary rule. Whilst the hypothesis of pure inertial inflation has been rejected in most of the empirical developments, some degree of inertia has been assumed to be present in the inflation process in Brazil as pointed out in empirical studies by Barbosa and McNelis (1990) and by Novaes (1991, 1993). Barbosa and McNelis found evidence in favor of wage inflation being linked to increasing inertia in the Brazilian inflationary process, whereas Novaes (1991) found that the degree of inertia in the inflationary process was far less than that implied by the theoretical model, a conclusion similar to that obtained by Durevall (1998) when modelling persistence in inflation. However the findings in Durevall (1998) with respect to the role played by wage inflation in particular, show that there is no evidence of this variable being significant in his econometric model. The opposite results obtained by Barbosa and McNelis and Durevall leave open the hypothesis that indeed wage policy was not a strong inflation-propagating mechanism as argued by Durevall (1998) quoting Baer (1989) and merits attention in developing a structural econometric model for the period.

Using these two theoretical underpinnings we not only demonstrate the role played by *a priori* theory within the LSE methodology but also show how congruence to additional sources of information plays an essential role in deriving structural econometric models (SEM) contrasting to the approach followed in the rational expectations program. This paper also attempts to present a more detailed empirical

[4] Interestingly Blejer and Liederman (1981, p.138) make a similar assumption in implementing their empirical model, namely: " With a view to the empirical implementation of the model for Brazil, we postulate here that the policy objective is to avoid long run changes in the real exchange rate and that the nominal rate is therefore altered to maintain purchasing power parity."

analysis of the dynamics observed in this period in Brazil, which seems to be richer than the results obtained by simply testing the restrictions derived from solving the Cagan model under the hypothesis of rational expectations and constructing auxiliary measures of the adequacy of the model to the data.

In section 2 we discuss the theoretical model for the AER policy that gives support in deriving the structural model following the LSE methodology. Section 3 presents the empirical results in deriving a baseline SEM. In section 4 we present the results concerning the SEM including the theoretical assumption presented in section 2 establishing a comparative analysis and discussing how to compare the distinct models. Since the main objective of the paper is to present a discussion of the LSE methodology based on an empirical exercise we omit the technical details concerning how we can derive a SEM. A more detailed analysis of the argument can be found in Hendry and Richard (1982), Hendry and Doornik (1994), Mizon (1995), Clements and Mizon (1991), Hendry and Mizon (1993), Hendry (1995), Bontemps and Mizon (2003) and Johansen (1994).

2) Theoretical Model

2.1) The Administered Exchange Rate Policy (AER)

The AER hypothesis establishes that from November 1980 onwards the central government adopted a policy for the monetary correction where it would follow the consumer price index, and the exchange rate devaluations would cover the difference between the internal and external inflation rates. Considering the real exchange rate definition we would have:

$$E = \frac{\kappa P^*}{P} \quad (1)$$ where κ is a proportionality factor.

In equation (1) if κ is a constant we can take logs on both sides and differentiate to obtain:

$$\ln e = \ln \kappa + \ln P - \ln P^{*} \quad (2)$$

$$\frac{de/dt}{e} = \frac{d\kappa/dt}{\kappa} + \frac{dP/dt}{P} - \frac{dP^{*}/dt}{P^{*}}$$

$$\Delta\%\, e = \Delta\%\, p - \Delta\%\, p^{*} \quad (3)$$

where lower cases represent the natural logarithm of the variable

Equation (3) states that the exchange rate devaluation equals the excess of the internal inflation to the external one and is particularly useful in investigating the Brazilian inflation in the first period. One of the goals of the AER analysis proposed here is exactly to identify a long run equation like (3) in the first period and test its relevance in the SEM developed in section 4.

As it stands equation (3) basically states that the deliberate exchange rate policy implementing devaluations in the nominal exchange rate should follow the rule represented by the equation which is theoretically identical to the (weak form of the) Purchasing Power Parity (PPP) hypothesis.

In general the evidence in the literature on the validity of the PPP hypothesis is weak in tests spanning long periods. It might be precisely because of considering that the hypothesis would hold in the Brazilian economy without paying attention to the specific implications of the economic measures implemented in the period that Durevall (1998) argues that the PPP hypothesis is no longer valid after 1979. So in this sense the test implemented here is specifically intended to account for using a priori information from the institutional background and to test its implications for the price formation in the econometric model being in accordance to the LSE methodology.

3) Rising Inflation and the LSE Methodology:

We start the analysis with a brief description of the variables. The sample data are monthly seasonally unadjusted for the period 1980 (1) to 1986(2). The variables are defined as follows: M1 is the paper money held by the public plus demand deposits and measured in millions of Reais[5] CPI is the consumer price index, IP is the industrial production index both as defined in Juselius (2002), BE is the bill of exchange monthly interest rate to the payee and finally CDB is the monthly interest rate paid on the certificate of deposits. Figure 1 contains full sample time plots of the modelled

[5] The Real is currently the currency in circulation in Brazil.

variables: real money, industrial production, the bill of exchange monthly interest rate to the payee, the monthly interest rate paid in the certificate of deposits, and consumer price inflation respectively ($m1-cpi, ip, be, cdb, \Delta cpi$). Low cases indicate the log of the variable.

The choice of the two interest rates follows the decision of modelling M1 and the data availability criteria, which span the whole period under consideration. The CDB is a primary rate for the fixed income market in Brazil constituting a medium to short run investment depending on the period and legislation in course. This interest rate is also used to calculate nowadays the Reference Rate (TR) that can be considered a type of Brazilian Prime Rate. Since the TR is only available from 1991 onwards, we opted for using the CDB rate itself. In figure 1 we clearly see an upward trend following the growth in inflation.

The BE stands for an alternative to fixed income investments. It indeed represented a potentially different investment for agents' holdings of M1 than the fixed income operations represented here by the CDB interest rate to the extent that they represent bonds that were linked to goods transactions and the real sector of the economy. Its pattern in figure is similar to the CDB interest rate, though less volatile.

The industrial production index was used in the attempt to capture the impacts of M1 in the economic activity since ultimately M1 is the most liquid asset and even in the periods of high rates of inflation was never replaced by a foreign currency. Indeed the ability to carry on transactions using M1 was given by the existence of interest bearing accounts. Its pattern in figure 1 is distinctively seasonal showing clearly the impacts of the recession period in 1983 and the recovery observed from 1984 onwards.

The real money plot in figure 1 shows a steady decrease from the beginning of the sample showing the effects of rising inflation rates on money hold. The peaks observed at the end of each year represent the mandatory 13th salary paid to all employed work force.

Finally inflation presents a clear upward trend reaching a peak of approximately 14.5% in January 1986 just before the Cruzado plan initiated at the end of February 1986. The economy recovery observed from 1984 onwards seems to have produced a perverse impact on this variable with inflation accelerating during 1985 and becoming more volatile.

Figure 1 1980(1) – 1986(2) Sample Time Plots

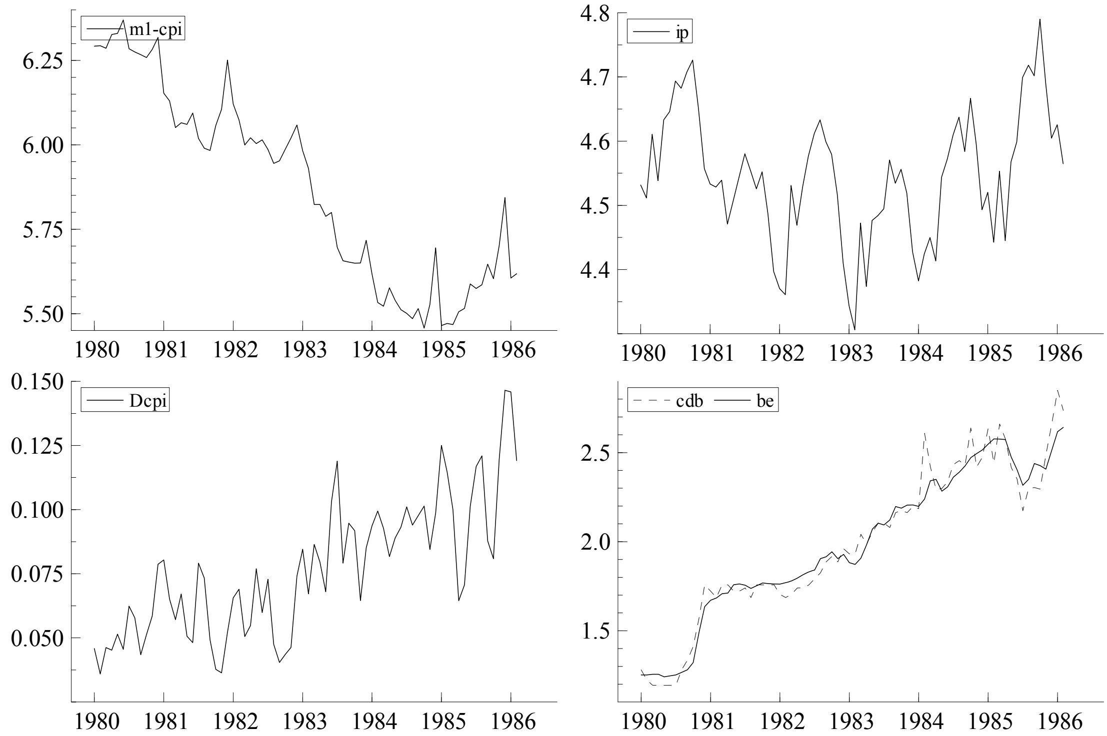


Table 1 Augmented Dickey Fuller Test (1980/1 – 1986/2)

Variable	*Lag*	*t-value*	*Critical Value (5% / 1%*)*	*Constant*	*Trend/Seasonal*
$m1-cpi$	0	-2.512	-3.44 / -4.02	Yes	yes / no
ip	0	-1.463	-3.44 / -4.02	Yes	yes / yes
cdb	2	-2.866	-3.44 / -4.02	Yes	yes / no
be	0	-1.963	-3.44 / -4.02	Yes	yes / no
Δcpi	3	-4.295	-3.44 / -4.02	Yes	yes / no

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

Table 2 Phillips Perron Test (1980/1 – 1986/2)

Variable	*Lag*	*Z statistic*	*Critical Value (5% / 1%)‡*	*Constant*	*Trend*
$m1-cpi$	1	-10.126	-20.7 / -27.4	yes	Yes
ip	2	-30.666	-20.7 / -27.4	yes	Yes
cdb	1	-17.063	-20.7 / -27.4	yes	Yes
be	1	-10.837	-20.7 / -27.4	yes	Yes
Δcpi	4	-14.331	-20.7 / -27.4	yes	Yes

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

Tables 1 and 2 present the Augmented Dickey Fuller (ADF) and the Phillips Perron (PP) tests results respectively. The null of unit root is only rejected for inflation using the ADF test and for the industrial production index using the PP test.

As a general conclusion about the order of integration of the variables the evidence found using the ADF and PP tests indicates that they are non-stationary and in particular well represented as I(1)[6] and in this case it is essential that we model them using a system so we are able to accommodate more than one cointegrating vector.

For the analysis we estimate initially a VAR (2) in the following variables: $m1-cpi, ip, cdb, be \text{ and } \Delta cpi$. 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 VAR includes also an unrestricted dummy which assumed value one for 1984 (2) and zero otherwise accounting for the exceptional growth in the first quarter of the year and a new waiver authorized by the IMF which basically signalled a new attitude of mind of the IMF towards the Brazilian external constraints in place after the Mexican external debt crisis of 1982. Table 3 presents the diagnostic statistics for the system.

Table 3 Diagnostic Tests VAR (1980/1 – 1986/2)

Test\Equation	$m1-cpi$ *(p-value)*	ip *(p-value)*	cdb *(p-value)*	be *(p-value)*	Δcpi *(p-value)*	*System* *(p-value)*
AR 1-5	1.46	0.46	1.10	0.44	1.49	1.02
	(0.21)	(0.80)	(0.37)	(0.82)	(0.21)	(0.45)
Normality	2.99	1.23	1,55	3.08	2.74	14.76
	(0.22)	(0.54)	(0.46)	(0.21)	(0.25)	(0.14)
ARCH	0.38	0.45	2.19	0.73	1.64	
	(0.85)	(0.81)	(0.07)	(0.60)	(0.17)	—
Hetero	0.45	0.59	0.89	0.69	0.51	0.47
	(0.97)	(0.89)	(0.60)	(0.81)	(0.94)	(1.00)

The diagnostic tests assessed are the LM test for autocorrelation of order p in the residuals (AR 1-p), the normality test is the test proposed by Doornik and Hansen (1994), the ARCH test is the test proposed in Engle (1982) and finally the

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

Heteroscedasticity test is based on White (1980). The system tests are the counterparts of the individual equation tests, as described in Doornik and Hendry (1997).

The residuals in the system show no sign of autocorrelation and heteroscedasticity, no sign of non-normality and ARCH effects in both levels, namely equation and system, so we consider the system a congruent representation of the information available in the data and proceed with the analysis of cointegration as described in Johansen (1995).

In table 4 we present the modulus of the five largest eigenvalues of the companion matrix. Table 5 presents the results of the trace test for testing the hypothesis of at most *r* cointegrating vectors, where the asymptotic p-values are based on the hypothesis of an unrestricted constant and restricted trend as described in Doornik (1998). In the present case since we include an unrestricted impulse dummy the results should be interpreted with caution because its presence is likely to modify the asymptotic values. Therefore, inference should be guided not only by the statistics presented, but also by other information as provided in the eigenvalues of the companion matrix and the graphic analysis of the cointegrating vectors.

From table 4 there are two eigenvalues which are close to unity whereas the result of the trace test indicates that we cannot reject the null of $r \le 2$. Both results indicate that we have then two cointegrating vectors. We therefore proceed with the assumption that $r = 2$ and impose this restriction on the cointegrated VAR in equation (4). The system was re-estimated and the five largest eigenvalues of the companion matrix are reassessed again as shown on table 5 $(r = 2)$, since the number of unit roots is equal to 3 $(N - r = 5 - 2 = 3)$ we rule out the hypothesis of the data being I(2).

Table 6 show the two cointegrating vectors after imposing and testing over-identifying restrictions using the LR test[7]. The first cointegrating vector has a difficult

---

[7] Following the discussion in Pesaran and Shin (2002) we impose a total of 5 restrictions on the long run components of the cointegrating vectors only, corresponding to the following:

$\beta' = \begin{bmatrix} 0 & 1 & 1 & 0 & * & 0 \\ * & 0 & -1 & 1 & 1 & * \end{bmatrix}$ which is tested against the data through the LR test.

The testing result is: $\chi^2(5) = 5.236\,[0.3877]$. We can conclude therefore that there is no evidence against that on the data. Since $K > r^2$, where K is the number of restrictions then the cointegrating vector is over-identified and we can propose a particular interpretation to the vector. The final form presented in table 6 corresponds to testing further restrictions on the dynamics of the system and zero restrictions on the components of the vectors. In particular the trend was not significant in the estimated vectors as well as the real money variable. The reader should notice nevertheless that zero restrictions imposed on interpretation,[8] but might indicate that the output measured by the industrial production index is negatively related to the real interest rate measured by the CDB interest rate[9]. With the maturity of investments in this type of certificate varying between 6 and 3 months in the period, it seems likely that they would have a significant role in the long run relationship linked to the seasonal pattern in the industrial production index. Further, they represent an alternative to consumption and M1 holding what would then reflect in the industrial output.

Table 4 Five Largest Eigenvalues Companion Matrix

Eigenvalues r unrestricted	*Eigenvalues* $r = 2$
0.8976	1.0000
0.8976	1.0000
0.6678	1.0000
0.6206	0.5231
0.6206	0.5231

Table 5 Cointegration Statistics VAR (1980/1 – 1986/2)

r	*0*	*1*	*2*	*3*	*4*	*5*
Trace Test	131.00	70.128	39.795	15.259	4.131	
p-value	0.000	0.012	0.099	0.560	0.723	
Eigenvalue[10]		0.561	0.336	0.282	0.139	0.054

It is worthwhile to notice from figure 2 where the plots of both vectors are presented, that for the period from 1983 until the end of 1984 the deviations of the long run mean becomes smaller than it was in the other periods. It reflects a period when the open market interest rates were fixed by the Central Bank in the beginning of the daily operations and a stricter monetary policy was followed as consequence of an agreement with the IMF.

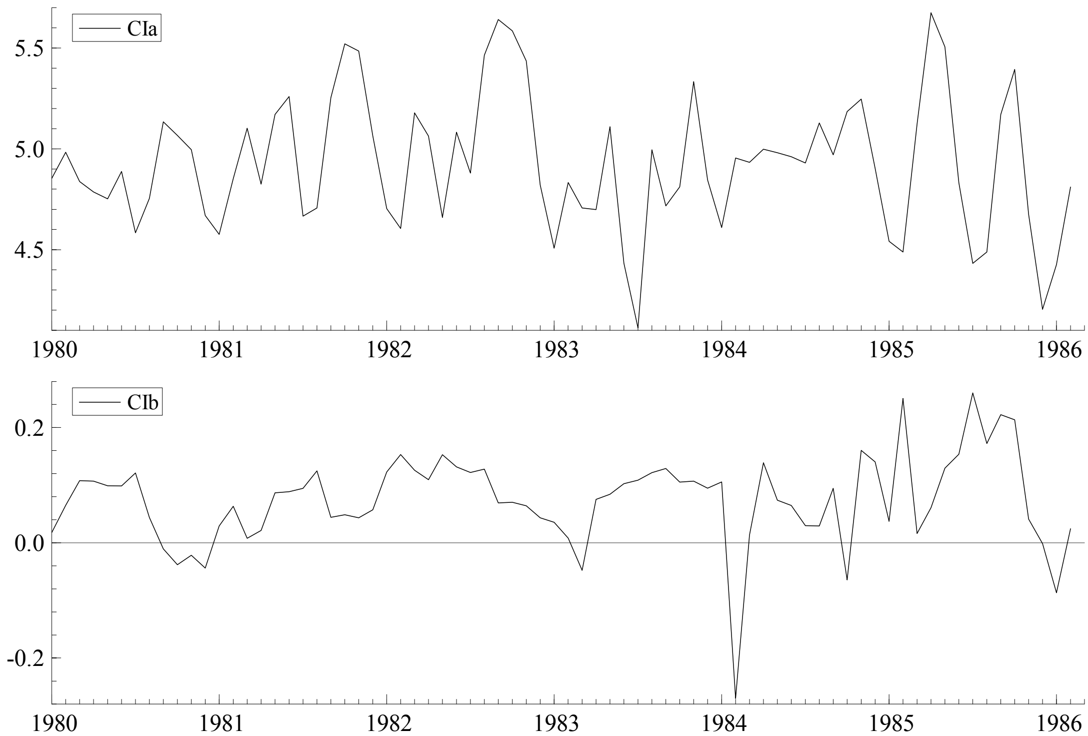
components in both vectors cannot be used as (over) identifying restrictions as pointed out in Boswijk and Doornik (2004) and Greenslade *et al l*(2002).

[10] Notice that the eigenvalues here are those referring to solving the eigenvalue problem generated by implementing the reduced rank regression as proposed in Johansen (1995).

Table 6 Cointegrating Vectors and Adjustment Coefficients VAR 1980(1) – 1986(2)

Cointegrating Vectors	$\hat{\alpha}_i$ (se)	$i=1$	$i=2$
CIa: $ip + cdb - 20.908\Delta cpi_t$	$m1 - cpi$	0	0
CIb: $\Delta cpi_t - cdb_t + be_t$	ip	-0.0492 (0.015)	0
LR test of Restrictions	cdb	0	0.557 (0.127)
Equilibria and Feedback: $\chi^2(13) = 19.285 \quad [0.1145]$	be	-0.030 (0.015)	0
	Δcpi	0.039 (0.006)	0

Figure 2 Cointegrating Vectors VAR 1980(1)-1986(2)



In general, the first cointegrating vector appears to contradict the argument that during the period that preceded the Cruzado Plan launching the economy would not react to classical measures such as control over consumption through higher interest rates which had been implemented in the beginning of the 1980's and allegedly had been innocuous. Indeed it seems that the passive monetary policy followed by the Central Bank would lead to an increase in consumption and consequently in the output establishing therefore the link between inflation and output.

This seems particularly true towards the end of the period when according to Andima (1997) shortage of liquidity in the financial system led the Central Bank to inject large sums of money in the system to guarantee the solvency. This policy nevertheless imposed on the Central Bank the cost of loosing its control over the monetary aggregates starting a long period where the main goal of the Brazilian Central Bank was to avoid large portfolio losses for the banks[11].

The second vector indicates that the spread between the interest rates on the bill of exchanges and the certificates of deposits cointegrates with inflation, proxying a Fisher relationship between the nominal spread and inflation. The vector presented in figure 2 has mean close to zero and deviations of the nominal spread from inflation are not sustainable in the long run consistent therefore with the hypothesis of rational expectations.

The adjustment coefficients in table 6 show that real money is weakly exogenous for the parameters of both cointegrating vectors. The persistence of this result in the simultaneous equations analysis where the growth in real money appeared to be driven only by seasonal dummies reinforces the previous results and suggests that the model proposed is describing the path of inflation and its links with the real sector of the economy where money can be considered strongly exogenous. The rate of growth in the industrial production not surprisingly reacts to the first vector but does not to the second whilst *cdb* reacts to second vector only. The effect of the first cointegrating vector on the rate of growth of *be* has a possible explanation on the fact that the bills of exchange are related to goods transactions and in this case deviations from the long run equilibrium would affect the operations with this type of bond. Finally, the rate of growth in inflation reacts to the first cointegration vector only, showing that eventually disequilibria in the long run have effects on inflation.

The following vector equilibrium correction model was then estimated: $\Delta\mathbf{y_t} = \Delta\mathbf{y_{t-1}} + \alpha\left(\beta\mathbf{y_{t-1}}\right) + \varpi\mathbf{D_t} + \upsilon_t$, where $\beta\mathbf{y_{t-1}}$[12] comprises the two cointegrating vectors. The misspecification tests presented signals of non-normality of the residuals which led us to re-estimate the system excluding $\Delta m1 - cpi_{t-1}$ and Δcdb_{t-1} from $\Delta\mathbf{y_{t-1}}$ since both variables were not significant in the whole system according to the F-test.

[11] Indeed this has increased from US$ 610 millions in December of 1983 to US$ 1.6 billions in the first two months of 1984.

[12] It is worthwhile to notice 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 lay on the cointegration space.

Testing the reduction led to the results presented in table 7 representing then the Parsimonious VAR against which the SEMs are tested.

The reduced system appears to be congruent to the information available with only a small sign of non-normality of residuals in the Δbe equation where the test is rejected at the 5% significance level but not at 1%. Otherwise the system has no sign of serial correlation and heteroscedasticity in the residuals. The reduction constituted a valid simplification in the system[13], where the number of parameters was reduced from 100 to 90 and constitutes therefore the basis from which the SEM is tested.
Although presenting no sign of misspecification there is evidence of parameter instability with the equation for the interest in the certificates of deposits presenting signals of parameter instability comprising evidence that a change in the monetary sector was in place during the 1980's.

Interestingly parameter instability is detected before actually a change in the variables levels had taken place as a consequence of the Cruzado Plan, which usually is the easiest detectable form of parameter non-constancy[14].

The SEM[15] derived imposes a total of 15 over-identified restrictions which were not rejected based on the results of the LR test ($\chi^2\left(15\right) = 22.276$). The diagnostic tests are shown on table 8 whereas the final SEM is presented in table 9.

The SEM presents only a slight sign of autocorrelation in the residuals of the Δcdb equation but the vector autocorrelation test did not rejected the null hypothesis of no autocorrelation. In figure 3 the time plots of the breakpoint Chow test are presented[16] (for each one of the equations and for the SEM as a whole) and one step residuals with ± 2 standard deviations. This plot shows that the model has no sign of break points and individually there is signal of parameter instability in the equation for Δcdb only. The

[13] The reduction tests led to the following results:
SYS(29) --> SYS(30): F(10,100)= 0.73429 [0.6906], where SYS(29) stands for (5.11) and SYS(30) is (5.11) with Δcdb and $\Delta m1 - cpi$ removed.

[14] See to this respect Hendry (2000) where the author argues that using the type of test for parameter instability we use here shifts in the unconditional expectations such as those observed in our case after the Cruzado Plan (see figure 5.1) are the most easily detectable.

[15] All SEMs presented in this paper were estimated by Full Information Maximum Likelihood (FIML)

[16] This test is based on the following statistics: $\frac{\left(RSS_T - RSS_{t-1}\right)(t-k-1)}{RSS_{t-1}(T-t+1)}$ which under the null of constant parameters has a F distribution as $F\left(T-t+1, t-k-1\right)$ for $t = M, \ldots, T$. The sequence of forecasts run from $T - M + 1$ to 1 where M here is 1982(10).

plot of the one step residuals does not indicate also signals of outliers so we consider that the SEM is a congruent representation of the data and parsimoniously encompasses the VAR.

Table 7 Diagnostic Tests Reduced VEqCM 1980/1 – 1986/2

Test\Equation	$\Delta m1-cpi$ *(p-value)*	Δip *(p-value)*	Δcdb *(p-value)*	Δbe *(p-value)*	$\Delta\Delta cpi$ *(p-value)*	*System* *(p-value)*
AR 1-5	1.45	0.60	1.41	0.62	1.73	0.88
	(0.22)	(0.70)	(0.24)	(0.68)	(0.14)	(0.76)
Normality	0.38	3.78	0.38	8.46	3.20	24.77*
	(0.83)	(0.15)	(0.83)	(0.015)*	(0.20)	(0.016)
ARCH	0.78	0.41	1.56	0.42	1.46	
	(0.57)	(0.84)	(0.19)	(0.83)	(0.22)	—
Hetero	1.06	0.54	0.58	0.58	0.53	0.63
	(0.41)	(0.85)	(0.83)	(0.82)	(0.86)	(0.99)

*indicates rejection at 5% level

Table 8 Diagnostic Tests SEM 1980/1 – 1986/2

Test\Equation	$\Delta m1-cpi$ *(p-value)*	Δip *(p-value)*	Δcdb *(p-value)*	Δbe *(p-value)*	$\Delta\Delta cpi$ *(p-value)*	*System* *(p-value)*
AR 1-5	1.80	1.66	4.50**	1.36	2.05	0.86
	(0.12)	(0.16)	(0.001)	(0.25)	(0.08)	(0.79)
Normality	0.95	4.59	1.37	0.01	1.33	11.45
	(0.62)	(0.10)	(0.50)	(0.99)	(0.51)	(0.32)
ARCH	0.84	0.44	1.05	0.67	0.99	
	(0.52)	(0.81)	(0.39)	(0.63)	(0.42)	—
Hetero	0.97	0.47	2.12	1.32	0.39	0.81
	(0.48)	(0.90)	(0.03)*	(0.24)	(0.95)	(0.93)

The equation for $\Delta m1-cpi$ shows the growth of real money being driven basically by the centered seasonal dummies with Δbe_{t-1} appearing to be marginally significant, nevertheless excluding it from the equation led to a non-congruent representation of the data. Such result does not match exactly the weakly exogeneity observed in the cointegration analysis in the sense that we cannot condition on

$\Delta m1 - cpi$, but Δbe_{t-1} being only marginally significant in some extent shows that the model is likely to be capturing the long run relationship among industrial output, represented by the industrial production index, interest rates and inflation.

Interestingly the results obtained by Cardoso (1983) who investigated the money demand in Brazil using quarterly data spanning the period between 1966 and 1979 showed that indeed the bill of exchange monthly interest rate was relevant in the estimated money demand equation whereas inflation was not[17]. Under this assumption, given the increasing use of indexation which gave the impression to the agents that they would be able to cope with inflation without facing the cost, it could have been that they were only adjusting their cash holds according to inflation to keep real money balances constants through financial operations.

Table 9 SEM 1980/1 – 1986/2

$$\underset{SE}{\Delta m1 - cpi} = -\underset{0.005}{0.004} - \underset{0.099}{0.147}\,\Delta be_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.041$$

$$\underset{SE}{\Delta ip} = \underset{0.004}{0.004} - \underset{0.108}{0.269}\,\Delta ip_{t-1} - \underset{0.248}{0.439}\,\Delta\Delta cpi_{t-1} + \underset{0.164}{0.488}\,CIb_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.031$$

$$\underset{SE}{\Delta cdb} = \underset{0.001}{0.024} - \underset{0.530}{0.802}\,\Delta\Delta cpi_{t-1} + \underset{0.406}{1.760}\,CIb_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.080$$

$$\underset{SE}{\Delta be} = \underset{0.108}{0.586} + \underset{0.094}{0.341}\,\Delta ip_{t-1} + \underset{0.074}{0.270}\,\Delta be_{t-1} - \underset{0.014}{0.075}\,CIa_{t-1} - \underset{0.199}{0.572}\,CIb_{t-1}$$
$$+\hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.032$$

$$\underset{SE}{\Delta\Delta cpi} = -\underset{0.051}{0.323} + \underset{0.041}{0.081}\,\Delta ip_{t-1} + \underset{0.121}{0.279}\,\Delta\Delta cpi_{t-1} + \underset{0.007}{0.042}\,CIa_{t-1} + \underset{0.077}{0.167}\,CIb_{t-1} + \hat{\varpi}\mathbf{D_t}$$
$$\hat{\sigma} = 0.013$$

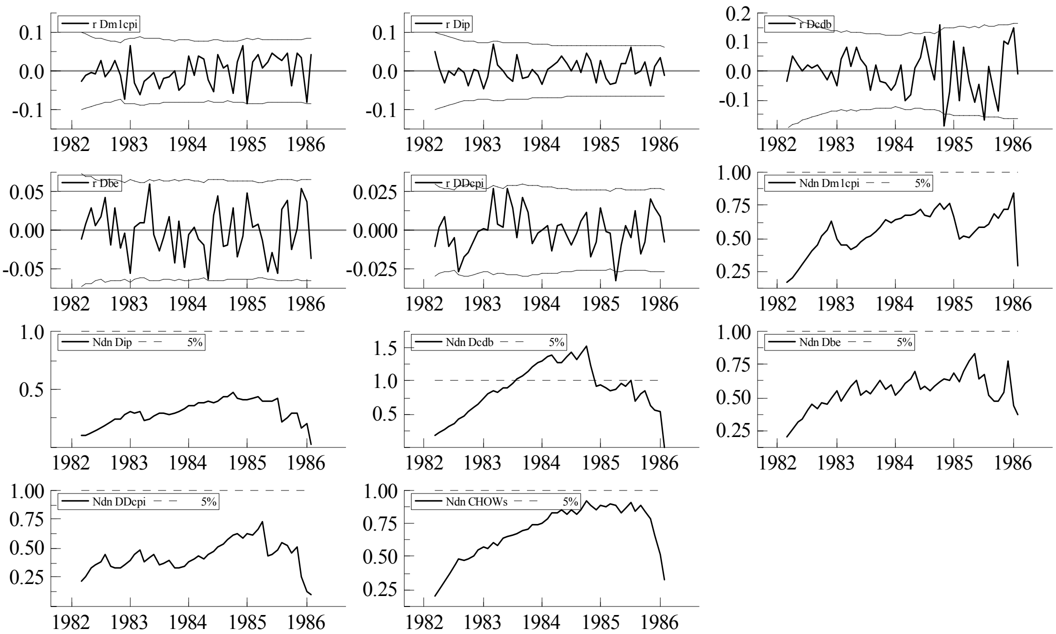
$\mathbf{D_t}$ stands for a vector with centered seasonals a pulse dummy for 1984/2 and in the equation for Δbe also a pulse dummy for 1980/11 which represented the end of pre-fixed monetary correction policy. This measure implied that the monetary correction would follow the consumer price index and the exchange rate devaluations would cover the difference between the internal and external inflation rates.

In the Δip equation the rate of growth in the industrial production index is a negative function of its lagged value most likely being influenced by the 1982-1983 recession and the rate of growth in inflation in a such way that increases in the rate of growth in inflation would lead to reductions in Δip. The relation shows that the level of indexation in the economy did not allow increases in inflation to have a positive impact in the industrial sector and consequently on the real sector of the economy. If we

[17] Cardoso's estimations were severely criticized by Gerlach and Nadal de Simone (1985) given the econometric techniques used by Cardoso. This discussion given the elapse of time has lost its appeal and will not be reproduced here but the main fact here is that accounting for the problems of simultaneity and non-stationarity not discussed in the previous two works a similar result has been found.

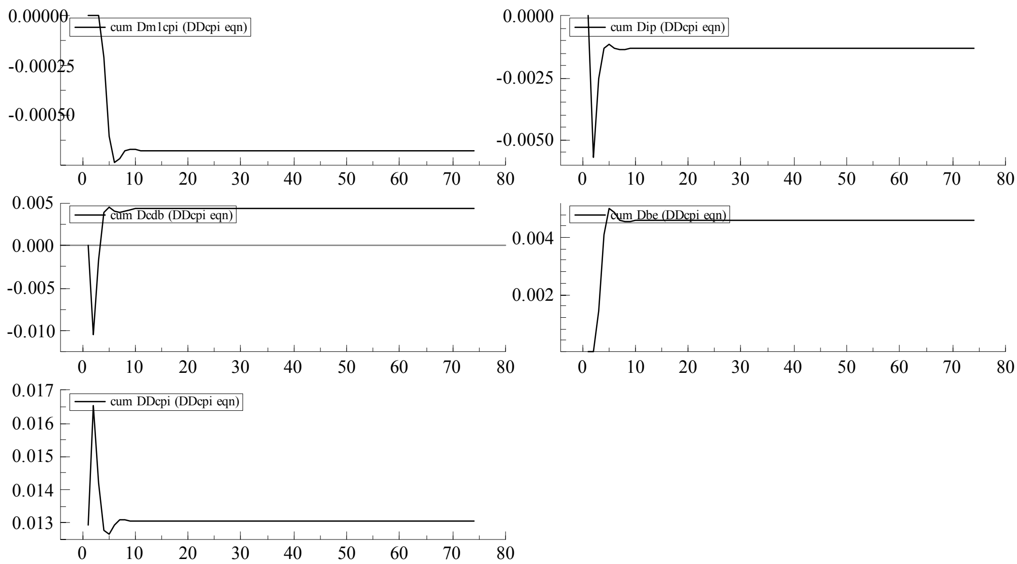
consider the cumulating impulse response functions to a one standard deviation shock in figure 4 the long run effect is negligible and in the context of the Brazilian economy indicates that the indexation has removed any impact of inflation on the economy except perhaps in its level. Such conclusion contrasts with the findings for the first cointegrating vector and exposes the risk of interpreting cointegrating vectors without considering the short run properties of the system as discussed in Lütkepohl (1994).

Figure 3 One Step Residuals and Chow Breakpoint Tests (SEM 1980/1 – 1986/2)



When we consider the equation for $\Delta\Delta cpi$ this variable is positively influenced by Δip_{t-1} and $\Delta\Delta cpi_{t-1}$. Such pattern explains why inflation rate gained momentum and why inflation assumed a pattern of persistence in a period where the Government followed a loose monetary policy by simply accommodating the inflation pressures. Indeed when we consider the impulse response functions to a one standard deviation shock to $\Delta\Delta cpi$ in figure 4 the effect is basically concentrated on the $\Delta\Delta cpi$ equation with an initial high level of impact on the rate of growth of inflation. In the long run the shock shows some degree of inertia in the inflation rate in the $\Delta\Delta cpi$ equation. This conclusion follows because the shock leads to increasing *rates of growth in inflation* for five periods approximately and considering that we are modeling the rate of growth in the inflation rate and the accumulated shocks have the magnitude of one standard deviation they represent accelerating inflation rates.

Figure 4 Cumulative One Standard Deviation Impulse Response Functions
Shock To (From) SEM 1980/1 – 1986/2



Juselius (2002, p. 30) traces the following scenario to explain the inflation growth in Brazil: *"The expansion of money stock need to finance recession and devaluations in the first case increased inflationary expectations in the black market, which then gradually spread to the whole domestic economy"*. She derives this conclusion after using the exchange rate in a trend liquidity ratio, namely given by $\beta x_t = m_{t-1} - s^b_{t-1} - y^r_{t-1} - 0.005t$ where *m* is the log of M3 stock, s^b is the log of the black market exchange rate and y^r is the log of the industrial production index. In her comments she stresses that the inflationary expectations were strongly affected by the expansion of money stock and that explains why the exchange rate has been used instead of the consumer price index. Nevertheless M3 in Brazil includes public securities, which after the currency devaluation in 1983 started to be issued indexed to the US dollar explaining therefore the role of the exchange rate in the trend liquidity ratio as a parameterization to the expected inflation. The expansion of money observed by Juselius is indeed the growth in non-monetary (but indexed) assets observed following the increases in inflation and after expectations of a currency devaluation materialized in 1983. They were representing an increasing demand for indexed money in an environment of soaring rates of inflation.

A more plausible scenario given the results found here is that the accommodating monetary policy (M1) followed in the period after the 1984 economic upturn and represented here by the growth in the industrial production index in a highly indexed economy allowed the increase in the inflation rates as shown by the equation for $\Delta\Delta cpi$. Given the model's dynamics presented in figure 4 any negative shock to the economy as an increase in the oil prices would magnify the inflation rate and represent in the short run a change in its level as we can infer from the blip in figure 4 for the $\Delta\Delta cpi$ equation. The economy recovery observed from 1984 onwards seems to have represented this negative role with its associated perverse connotation. The balance of payments adjustment and the constant surplus in the trade balance were followed by a monetary policy that only accommodated the demand pressures generating then a spiral on the inflation rates[18].

Such pattern in the dynamic properties of the estimated model is clearly in accordance with a highly indexed economy and with the hypothesis that inflation had an inertial component as defended by Resende and Lopes (1981 and 1982), Lopes (1983) and Modiano (1983). However we do not derive the policy implication that inflation's path was following an independent process. Rather the SEM is showing that we cannot ignore the links between the real variables and inflation in the period. The results of the empirical modeling throw light on the debate about the level of inertia present in the Brazilian inflation for the period. According to Novaes (1993) the notion that inflation in Brazil was largely dominated by its past values and by inertia being further insensitive to demand policies had became widespread. Nevertheless she found using a univariate approach that indeed the persistence of a one percent shock to the inflation rate was only 35% after 30 months. The author then argues to a reassessment of the importance of the inertial component in the Brazilian inflation in line therefore to the conclusions found here despite the different approaches implemented.

As a final remark the equations for the interest rates seems to describe their conceptual difference with Δcdb being influenced by $\Delta\Delta cpi_{t-1}$ and the second

[18] The impact of the 1984.2 dummy in both equations, Δip and $\Delta\Delta cpi$ despite being marginal is positive. $\underset{(0.034)}{0.071}$ and $\underset{(0.014)}{0.016}$ respectively.

cointegration vector only, whereas Δbe is influenced by Δip its own lag appearing to be linked to operations in the real sector.

4) The Role of Nominal Wage Inflation and the AER Policy

The empirical model's dynamic properties presented in last section describe clearly the pattern of relative persistence in inflation and the long run cointegrating. The theoretical hypotheses presented in section 2 were not addressed in the analysis and one implication is that the encompassing analysis was carried by concentrating on finding empirical reductions of the VAR and did not consider empirical versions of these theoretical models on aiming to specify parsimonious encompassing models, namely the SEMs compared to the VAR termed as the system. Therefore each SEM (the empirical reduction of the VAR) was in turn derived by eliminating non-significant variables from the system following a general to specific modelling strategy. From the previous section analysis we can conclude using the results of the over-identifying test statistics that the SEM derived in last section consisted of a congruent model that parsimoniously encompassed the VAR.

Although this can be considered a central property for any econometric model an interesting question in this framework is if the results obtained in section 3 are robust to extensions in the information set comprised in relevant theoretical hypotheses not included in the analysis previously. The sample size in the present case prevents an extensive analysis through including more endogenous variables in the system. In particular I focus on the AER policy followed by the Brazilian central government implemented in the first half of the 1980's and nominal wage inflation impact on price dynamics.

Given the restricted sample size the strategy that we follow is to perform the extended analysis of the econometric models developed in section 3 through testing for the presence of omitted variables in these models in the case of the AER. The constraint imposed by the sample size is therefore the reason for not including the relevant variables implied by the AER hypothesis in a VAR framework comprising the estimation of a full system to re-assess the model derived in section 3 in a comparative analysis through the use of the over-identifying restrictions test as an encompassing test as suggested in Hendry and Mizon (1993).

Alternatively, for the nominal wage hypothesis we carry out a test of weak exogeneity of this variable by the parameters of interest and assess the significance of this variable in the SEM. In this case if the variable is weakly exogenous there is no loss of information in estimating the marginal system only instead of carrying on the full system analysis[19]. The weak exogeneity test in a VAR framework and its implications for the present analysis is discussed in section 1.

The strategy pursued of testing if the SEM is robust to extensions in the information set raises the following subtle question beyond the exogeneity assumption: can the test for over-identifying restrictions be used on the same grounds as it was in section 3? This question is relevant exactly because in the presence of exogenous variables the SEM is no longer a closed model and the results shown in section 3 are valid only for closed systems. Put in another way using the test for over-identifying restrictions in an encompassing analysis is in principle only valid if the VAR in a closed form works as a catalyst model for testing the reductions. This question in particular is also addressed in section 4.1.

4.1) Encompassing the VAR in Open Systems

The encompassing analysis pursued in section 3 used the VAR as a catalyst for testing the different SEMs. The testing procedure works by specifying a congruent VAR and comparing the different SEMs using the over-identified restrictions test. In Hendry and Mizon (1993) it is shown that a sufficient condition for the SEM to encompass its competitors in the same class is that it parsimoniously encompasses the VAR and be congruent. Such condition is valid on the assumption that the VAR is closed, or more specifically, that all variables are modelled. In particular we have for a closed system:

$$\Delta \mathbf{y_t} = \Pi \mathbf{y_{t-1}} + \sum_{i=1}^{p-1} \Pi_i \Delta \mathbf{y_{t-i}} + \kappa + \upsilon_{\mathbf{t}} \quad (4)$$

where: $\Pi = \alpha\beta'$, α and β are $n \times r$ matrices, Ω is the covariance matrix of $\upsilon_{\mathbf{t}}$ and finally $0<r=rank(\Pi)<n$ so there is cointegration (assuming that the variables in $\mathbf{y_t}$ are I(1) which represents the case of interest in our analysis).

[19] It should be noticed that the AER hypothesis include in the estimation the variable *DDcpi* , so testing the hypothesis of weak exogeneity to the system in chapter 5 is not a feasible option.

It is worth noticing that testing the theoretical hypothesis represented by the nominal wage inflation poses an extra difficulty in our case regarding the encompassing test based on 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, or more specifically that it 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. In section 3 we test the hypothesis that the SEM encompasses the VAR considering that the VAR is the unrestricted model on which restrictions are imposed by the SEM. Alternatively, in the present section we test first the hypothesis that nominal wage inflation is weakly exogenous to the system (the VAR not the SEM). Doing that we can test the weak exogeneity of the relevant variable to the VAR and at the same time, construct an open VAR that corresponds to the unrestricted model.

Johansen (1994) shows that by partitioning the vector $\mathbf{y_t}$ in (4) such that we have $\mathbf{y_t} = (\mathbf{s_t}, \mathbf{v_t})$ and conditioning $\mathbf{s_t}$ on $\mathbf{v_t}$ there is a very special case where estimating the conditional model only is efficient so that there is no loss of information. This case arises when $\alpha_v = 0$ in (6)

$$\Delta\mathbf{s_t} = \lambda\Delta\mathbf{v_t} + \left(\alpha_s - \lambda\alpha_v\right)\beta'\mathbf{y_{t-1}} + \sum_{i=1}^{p-1}\left(\Pi_{si} - \lambda\Pi_{vi}\right)\Delta\mathbf{y_{t-i}} + \kappa_s - \lambda\kappa_v + \upsilon_{\mathbf{st}} - \lambda\upsilon_{vt} \quad (5)$$

where $\lambda = \Omega_{sv}\Omega_{ss}^{-1}$

$$\Delta\mathbf{v_t} = \alpha_v\beta'\mathbf{y_{t-1}} + \sum_{i=1}^{p-1}\Pi_{vi}\Delta\mathbf{y_{t-i}} + \kappa_v + \upsilon_{\mathbf{vt}} \quad (6)$$

Testing the hypothesis that $\alpha_v = 0$ constitutes therefore the test of weak exogeneity of $\mathbf{v_t}$ for β whereas testing the additional hypothesis that the coefficients of $\Delta\mathbf{s_t}$ are zero in equation (6) constitutes the test of strong exogeneity of $\mathbf{v_t}$ for β [20].

Johansen (1994) suggests an alternative weak exogeneity test for the case when there are many variables or the sample size is short. He argues that an alternative test involves evaluating (5) by the reduced rank procedure determining the cointegrating

[20] In the present case we focus on testing weak exogeneity since we are interested in statistical inference.

vectors, and then testing if their coefficients are different from zero using a F-test in the marginal model given in (7), or more specifically that $\varphi = 0$ in (7).

$$\Delta\mathbf{v_t} = \varphi\hat{\beta}\mathbf{y_{t-1}} + \sum_{i=1}^{p-1}\Pi_{vi}\Delta\mathbf{y_{t-i}} + \kappa_v + \upsilon_{vt} \quad (7)$$

We follow this alternative test when testing if nominal wage inflation is weakly exogenous instead of re-estimating the VECM's including the nominal wage inflation and carrying out the reduced rank analysis that equations (5) and (6) would suggest. The justification is the reduced sample size in the first and second periods mainly, where the VAR's in section 3 were estimated with 5 equations each whereas the test proposed in equations (5) and (6) would imply re-estimating the VAR's with 6 equations.[21]

Hendry and Mizon (1993) suggest a two-step encompassing test in the open systems case, namely, firstly test each closed model against the VAR using the over-identifying restriction test, and then as a second step test the validity of the weak exogeneity reduction of each closed model. Such encompassing analysis is nevertheless unfeasible in the present case because we am not formulating a closed model including Δnw in $\mathbf{y_t}$ in equation (4), where nw is a measure for nominal wage inflation. If nominal wage inflation is weakly exogenous for the parameters of the cointegrating vectors we can rewrite (4) as (5) and without loss of information concentrate our analysis on (5) which will act as the catalyst VAR relative to which we can test restrictions 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, as conditioning implies. In this case we preserve the specification order pursued in section 3 where the SEM arises as a valid reduction of the system rather than the reduced form being derived from the SEM. If this were the case the reduced form would have been identified on incredible

[21] The other theoretical hypothesis of interest in the present chapter, namely the AER hypothesis, does not allow a test of weak exogeneity because the Brazilian consumer price index enters in both cointegrating vectors, the AER VAR analysis and in the system analysis carried on in section 3 so we can not decompose equation 9 as equations 10 and 11. Further alternatively testing $\varphi = 0$ in equation (7) does not ensure that $\mathbf{v_t}$ is weakly exogenous for β. Consequently, in the cases where the model incorporates the cointegrating vector derived from the AER VAR analysis we test encompassing through a forecast encompassing test.

restrictions which are in other words assumptions on the exogeneity of the variables of interest.

The approach followed here consists in running the extended VAR and test if nominal wage inflation is weakly exogenous. If this variable is not exogenous it is advisable to restart modelling by including the variable in the system. On the contrary if the variable is weakly exogenous we define an unrestricted model which is an extended system to that used in section 3 and assess the congruence of the system relative to the data. This analysis is nevertheless restricted in such way we cannot compare among different econometric models. Given this sort of restriction that prevents the use of the over-identifying restriction test as an encompassing test, we carry on the encompassing analysis based on the forecast encompassing tests.

4.2) Forecasting Encompassing Tests: Encompassing the SEM

Forecast encompassing tests have a widespread use in applied econometrics given their simplicity. The basic hypothesis states that model 1 forecast encompasses model 2 if model 2 forecasts provide no further useful information for the prediction made by model 1. Let ϕ_1 and ϕ_2 be two forecasts from two different SEM's of interest for the variable of interest $y_t \in \mathbf{y_t}$ in (4) and let e_1 and e_2 be the corresponding forecast errors. According to Newbold and Harvey (2002) a natural test for model 1 forecast encompassing model 2 is a test for $\theta = 0$ in (8).

$$e_{1t} = \theta\left(e_{1t} - e_{2t}\right) + v_t \quad (8)$$

Chong and Hendry (1986) propose a slightly different test given by (9) where the rationale is to test if the information provided by the forecast in model 2 is of no value to the forecast error in model 1. Alternatively, tests based on (10) have also been used in applied work such as testing the hypothesis $(\alpha, \beta_1, \beta_2) = (0,1,0)$ in Diebold and Lopez (1996) and $\beta_2 = 0$ in Fair and Shiller (1989).

$$e_{1t} = \alpha\phi_{2t} + v_t \quad (9)$$
$$y_t = \alpha + \beta_1\phi_{1t} + \beta_2\phi_{2t} + v_t \quad (10)$$

We do use the three different ways of testing forecast encompassing, namely Hendry and Chong (HC), Diebold and Lopes (DL) and Fair and Shiller (FS) when comparing the SEM's. A particular comment of interest here is that if the variables involved in (9) and (10) are integrated it is likely that the time series represented by ϕ_{it} will be integrated. Nevertheless this is not of particular concern here since all models were mapped into an I(0) representation before the forecasts had been generated.

4.3) The Rising Inflation Period: 1980-1986(2)

For the analysis in this sub-sample following equation (6) we estimate initially a VAR (2) in the following variables: $m1-cpi, ip, cdb, be\ and\ \Delta cpi$ plus Δnw which represents the first difference of the nominal wage index calculated to the São Paulo manufacturing industry by the FIESP[22]. The VAR includes also a restricted trend and an unrestricted dummy which assumed value one for 1984 (2) and zero otherwise accounting for the exceptional growth in the first quarter of the year and a new waiver authorized by the IMF. Table 10 presents the diagnostic statistics for the system.

The residuals in the system have no sign of autocorrelation and heteroscedasticity, no sign of non-normality and ARCH effects in both levels, namely equation and system, so we consider the system a congruent representation of the information available in the data and proceed with the analysis of cointegration as described in Johansen (1995).

In table 11 we present the modulus of the five largest eigenvalues of the companion matrix and table 12 presents the results of the trace test for testing the hypothesis of $r \leq k$. Despite the clear rejection of the hypothesis of the number of cointegrating vectors being less than two in table 12, the moduli of the eingenvalues in table 11 suggest that we have only one unit root with the remaining values being far lower than the first largest modulus which led us to assume that we have just one cointegrating vector in the system. The over-identifying restrictions were not rejected and led to the estimated vector in equation (11).

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

Table 10 Diagnostic Tests VAR Exogeneity Test (1980/1 – 1986/2)

Test\Equation	$m1-cpi$ *(p-value)*	ip *(p-value)*	cdb *(p-value)*	be *(p-value)*	Δcpi *(p-value)*	Δnw	*System* *(p-value)*
AR 1-5	1.17	0.47	1.12	0.41	1.22	0.40	1.36
	(0.33)	(0.79)	(0.36)	(0.83)	(0.31)	(0.84)	(0.0524)
Normality	1.42	1.48	2.09	2.93	1.58	0.42	10.10
	(0.49)	(0.47)	(0.35)	(0.23)	(0.45)	(0.80)	(0.60)
ARCH	0.39	0.42	1.64	0.65	1.25	0.97	—
	(0.85)	(0.82)	(0.17)	(0.65)	(0.30)	(0.44)	
Hetero	0.46	0.52	0.69	0.50	0.27	0.69	0.20
	(0.96)	(0.94)	(0.81)	(0.95)	(0.99)	(0.81)	(1.00)

$$\frac{m_1}{cpi} - ip - cdb + \left(41.64\Delta cpi - 6.56\Delta nw\right) \quad \chi\left(5\right) = 3.52 \qquad \left(11\right)$$

Table 11 Five Largest Eigenvalues Companion Matrix

Eigenvalues r *unrestricted*
0.9351
0.7484
0.7484
0.6369
0.5965

Table 13 presents the results of estimating equation (6) including the over-identified cointegrating vector in (11). Since the F-test 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.

After concluding that the nominal wage inflation is weakly exogenous, the VAR in the I(0) space represented by equation (12) is estimated The estimation results are presented in table 14 where we focus on the diagnostic tests only.

Table 12 Cointegration Statistics VAR Exogeneity Test (1980/1 – 1986/2)

R	*0*	*1*	*2*	*3*	*4*	*5*	*6*
Trace Test	185.99	123.95	74.510	37.320	14.308	3.56	
p-value	0.000	0.000	0.004	0.163	0.637	0.799	
Eigenvalue		0.567	0.487	0.395	0.267	0.135	0.046

Table 13 Nominal Wage Weak Exogeneity Test[23]

$$\Delta nw = -\underset{0.008}{0.016}\,CIa + \sum_{i=1}^{3} \hat{\Pi}_i \Delta nw_{t-i} + \sum_{i=1}^{3} \hat{\Theta}_i \Delta \mathbf{y}_{\mathbf{t-i}} + \hat{\upsilon}_t \qquad \hat{\sigma} = 0.03403$$

$$F(1,55) = 3.7668[0.057]$$

$$\Delta \mathbf{s}_{\mathbf{t}} = \lambda \sum_{i=0}^{1} \Delta v_{t-i} + \alpha_s \beta' \mathbf{y}_{\mathbf{t-1}} + \Theta \Delta \mathbf{s}_{\mathbf{t}-1} + \omega_t \quad (12)$$

$$\text{where } \mathbf{y}_{\mathbf{t}} = \frac{m_1}{cpi} - ip - cdb + (41.64\Delta cpi - 6.56\Delta nw)$$

Table 14 Diagnostic Tests Open VAR (1980/1 – 1986/2)

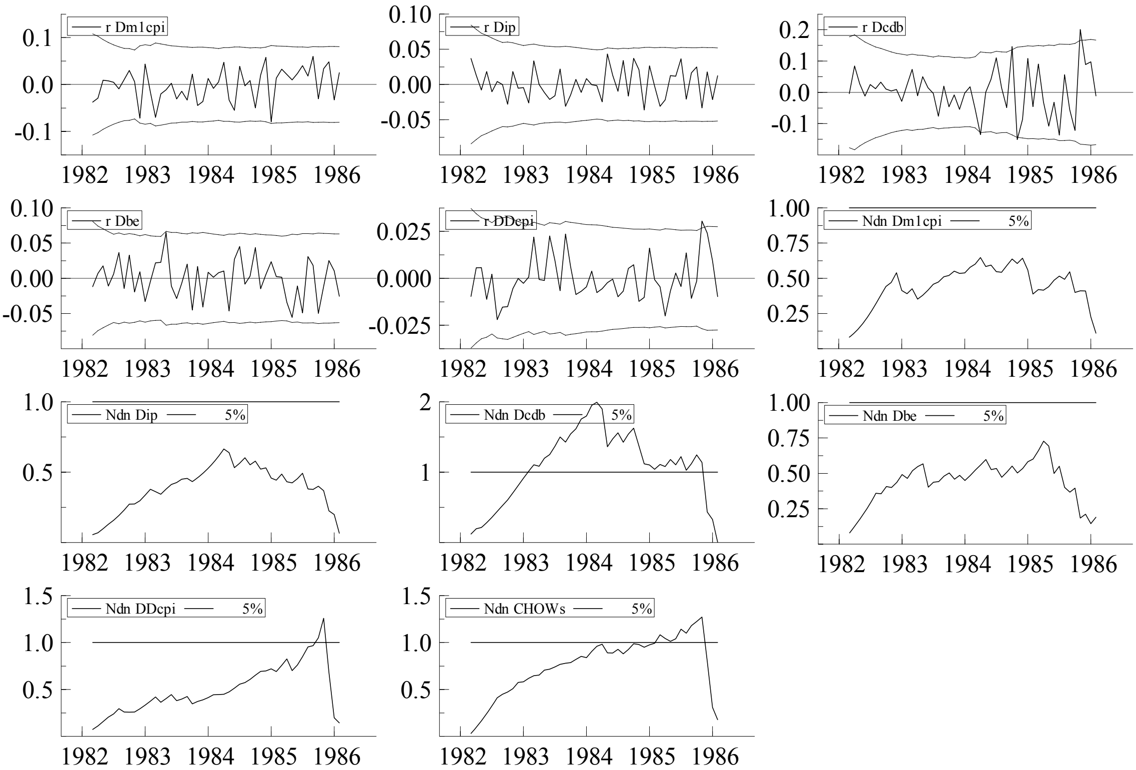
Test\Equation	$m1-cpi$ *(p-value)*	*ip* *(p-value)*	*cdb* *(p-value)*	*be* *(p-value)*	Δcpi *(p-value)*	*System* *(p-value)*
AR 1-5	1.08	1.21	1.34	0.30	0.96	0.92
	(0.37)	(0.31)	(0.26)	(0.90)	(0.45)	(0.66)
Normality	0.02	1.94	2.78	0.55	4.54	11.09
	(0.98)	(0.37)	(0.24)	(0.75)	(0.10)	(0.35)
ARCH	0.42	0.54	2.65*	0.61	0.96	
	(0.83)	(0.74)	(0.035)	(0.69)	(0.45)	
Hetero	0.55	0.69	0.73	0.92	0.17	0.55
	(0.90)	(0.78)	(0.75)	(0.55)	(0.99)	(1.00)

The long run cointegrating vector admits the interpretation of a long run money demand equation which is positively related to the output and real interest rate where inflation is measured as a weighted average between consumer and wage inflation. The positive signal on the real interest rate might be reflecting the agent’s reactions to the loose of control observed in the monetary policy in such way that increases in the interest rate would sign for a tighter monetary police more in line with fighting the

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

soaring inflation rates. Such conclusion is related to the use of M1 as the money stock of analysis and the reduction observed in the holdings of M1 as a proportion of the GDP during the first half of the 1980's.

Figure 5 One Step Residuals and Breakpoint Chow Tests Extended System (1980/1-1986/2)



The diagnostic tests show only a small sign of conditional heteroskedasticity in the residuals of the CDB equation, otherwise there is no sign that the system is not congruent with the data. The one-step residual and break point Chow tests shown in figure 5 have a similar pattern in the certificate of deposit equation to the observed in section 3 where there was sign of instability. The system also presents some sign of instability at the end of period only in the inflation equation and in the system test as a whole most likely reflecting the accelerating inflation rates that preceded the Cruzado plan cast in February 1986.

4.3.1) The AER Policy: Empirical Results

The AER hypothesis test for the first period was focused on equation (3) and included the nominal exchange rate in Reais to U.S dollar (e)[24], the consumer price index (cpi) as defined before and the U.S. wholesale price index (p^*)[25]. Initially we ran a fourth order VAR on the first difference of the variables including two dummies, the first one assuming the value one for 1983-2 and 1983-3 and zero otherwise and the second dummy assuming the value one for 1984-3 and zero otherwise. Table 15 presents the diagnostic test results.

In general there is no sign of non-congruence of the model with the data except in the equation for Δcpi where there is a sign of autocorrelation in the equation residuals with the autocorrelation test rejecting the null of no autocorrelation at a 5% significance level but not at 1% level. Such result nevertheless, is not reproduced in the system test for autocorrelation which led us to conclude that the system is a congruent representation of the information available in the data and to proceed with the analysis of cointegration as described in Johansen (1995).

In table 16, we present the modulus of the five largest eigenvalues of the companion matrix and in table 17 we present the results of the trace test for testing the hypothesis of $r \leq k$ where the asymptotic values are based on the hypothesis of an unrestricted constant and restricted trend as described in Doornik (1998).

From table 16 there is only one eigenvalue close to unity whereas the result of the trace test indicates that we cannot reject the null of $r \leq 1$. Both results indicate that we have then one cointegrating vector. We therefore proceed with the assumption that $r = 1$ and impose this restriction on the cointegrated VAR. The system was re-estimated and the five largest eigenvalues of the companion matrix are re-assessed again as shown on table 16 $(r=1)$, since the number of unit roots is equal to 2 $(N-r=3-1=2)$ we rule out the hypothesis of the data being I(2). Imposing the restriction implied by equation 3 lead

[24] This series is measured as the monthly average of daily asking price and corresponds to the official exchange rate for international trading.
[25] All series are available upon request to the first author or from the IPEADATA at www.ipeadata.gov.br

to the rejection in the LR test of over-identifying restrictions[26] which led us to re-estimate the VAR imposing the following restriction $\beta'=[1,*,-1]$.

. Table 15 Diagnostic Tests VAR (1980/1 – 1986/2)

Test\Equation	Δcpi *(p-value)*	Δp^* *(p-value)*	Δe *(p-value)*	*System* *(p-value)*
AR 1-5	2.63*	1.99	0.94	1.02
	(0.036)	(0.09)	(0.46)	(0.44)
Normality	0.46	0.11	1.54	3.78
	(0.79)	(0.95)	(0.46)	(0.71)
ARCH	0.82	0.08	0.79	
	(0.54)	(0.99)	(0.56)	
Hetero	0.39	0.68	0.40	0.47
	(0.98)	(0.81)	(0.98)	(1.00)

Table 16 Five Largest Eigenvalues Companion Matrix

Eigenvalues r unrestricted	*Eigenvalues* $r=1$
0.9971	1.0000
0.7997	1.0000
0.6948	0.7222
0.6948	0.7222
0.6874	0.6989

Table 17 Cointegration Statistics VAR (1980/1 – 1986/2)

r	*0*	*1*	*2*	*3*
Trace Test	120.04	7.33	0.008	
p-value	0.00	0.54	0.93	
Eigenvalue		0.782	0.094	0.0001

The estimated long run relationship became then $\Delta cpi - 1.9374\Delta p^* - \Delta e$, since the estimated parameter for Δp^* is close to 2 we re-estimated once more the VAR imposing

[26] The over-identifying restrictions test in this case means the LR test for identifying the cointegrating vector and not the LR test used for testing if the SEM encompass the VAR.

this value to the variable. The results are presented in table 18 from which we can infer that the restriction is not rejected leading to the identified cointegrating vector which has the signals in accordance to equation (3). We consider henceforth this estimated cointegrating vector as the empirical counterpart of (3) and use this vector in testing if it is relevant in the SEM derived in section 3.

Table 18 Cointegrating Vector and Adjustment Coefficients VAR 1980(1) – 1986(2)

Cointegrating Vector	$\hat{\alpha}_i$ (se)	$i=1$
$\Delta cpi - 2\Delta p^* - \Delta e$	Δcpi	-0.028 (0.080)
LR test of Restrictions	Δp^*	0.014 (0.023)
Equilibria and Feedback: $\chi^2(2) = 1.2163[0.5444]$	Δe	0.783 (0.063)

4.3.2) Forecast Encompassing: Empirical Results

In carrying on the forecast encompassing analysis two different models are considered. The first model (M1) is the SEM derived in section 3 and therefore without considering the role played by the nominal wage inflation and the AER hypothesis. The second model (M2) is the augmented SEM derived from testing the significance of both the nominal wage inflation and the AER hypothesis. The estimation results are presented in table 19 whereas the diagnostic tests are presented in table 20. The model includes the nominal wage inflation in the equation for Δip only. The AER cointegrating vector derived in section 4.3.1 was not significant at all in the system and consequently excluded from all equations. Such result is somewhat in line with the findings in Durevall (1998) where the author pointed out that from 1979 onwards the PPP hypothesis broke down given the fact that the AER test is identical to the PPP test in its weak version.

The sign for Δnw throws some light on the difficult interpretation of the sign for $\Delta\Delta cpi$ in section 3. Indeed this last variable presented a negative signal in the SEM derived in section 3 whilst the expected signal would be positive in such way that increasing inflation would lead to an accelerating economic activity. Now considering the nominal wage inflation it becomes clearer how the economic activity is related to inflation. The positive sign for Δnw is clearly showing that the impact on the industrial activity was originated in the nominal wage inflation likely because of the accommodating monetary policy implemented in the period. Such conclusion is also based on the fact that the coefficient for $\Delta\Delta cpi$ which is marginally significant only as in the model presented in section 3. This contrasts with the conclusion derived in section 3 that the level of indexation in the economy did not allowed increases in inflation to have a positive impact in the industrial sector, and consequently on the real sector of the economy.

Table 19 SEM (M2) 1980/1 – 1986/2

$$\underset{SE}{\Delta m1 - cpi} = -\underset{0.020}{0.040} - \underset{0.397}{0.497}\,\Delta\Delta cpi_{t-1} - \underset{0.009}{0.023}\,CIa_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.039$$

$$\underset{SE}{\Delta ip} = -\underset{0.014}{0.034} - \underset{0.030}{0.136}\,\Delta cdb_{t-1} - \underset{0.273}{0.514}\,\Delta\Delta cpi_{t-1} + \underset{0.084}{0.195}\,\Delta\Delta nw + \underset{0.077}{0.019}\,CIa_{t-1}$$

$$+\hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.025$$

$$\underset{SE}{\Delta cdb} = \underset{0.042}{0.100} - \underset{0.087}{0.218}\,\Delta cdb_{t-1} - \underset{0.841}{1.159}\,\Delta\Delta cpi_{t-1} + \underset{0.020}{0.059}\,CIa_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.082$$

$$\underset{SE}{\Delta be} = -\underset{0.016}{0.050} + \underset{0.039}{0.178}\,\Delta ip_{t-1} - \underset{0.331}{0.453}\,\Delta\Delta cpi_{t-1} - \underset{0.008}{0.031}\,CIa_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.032$$

$$\underset{SE}{\Delta\Delta cpi} = -\underset{0.007}{0.0312} + \underset{0.137}{0.216}\,\Delta\Delta cpi_{t-1} - \underset{0.003}{0.015}\,CIa_{t-1} + \hat{\varpi}\mathbf{D_t} \quad \hat{\sigma} = 0.013$$

$\mathbf{D_t}$ stands for a vector with centered seasonal dummies. For the first equation there is also a pulse dummy for 1984/2. In the second equations there is also a pulse dummy for 1981/3 accounting for the 1981 recession beyond the 1984/2 dummy. In the third equation we have only the 1984/2 dummy. In the equation for Δbe there is only a pulse dummy for 1980/11 which represented the end of pre-fixed monetary correction policy. This measure implied that the monetary correction would follow the consumer price index and the exchange rate devaluations would cover the difference between the internal and external inflation rates. Finally in the last equation we do not have pulse dummies at all.

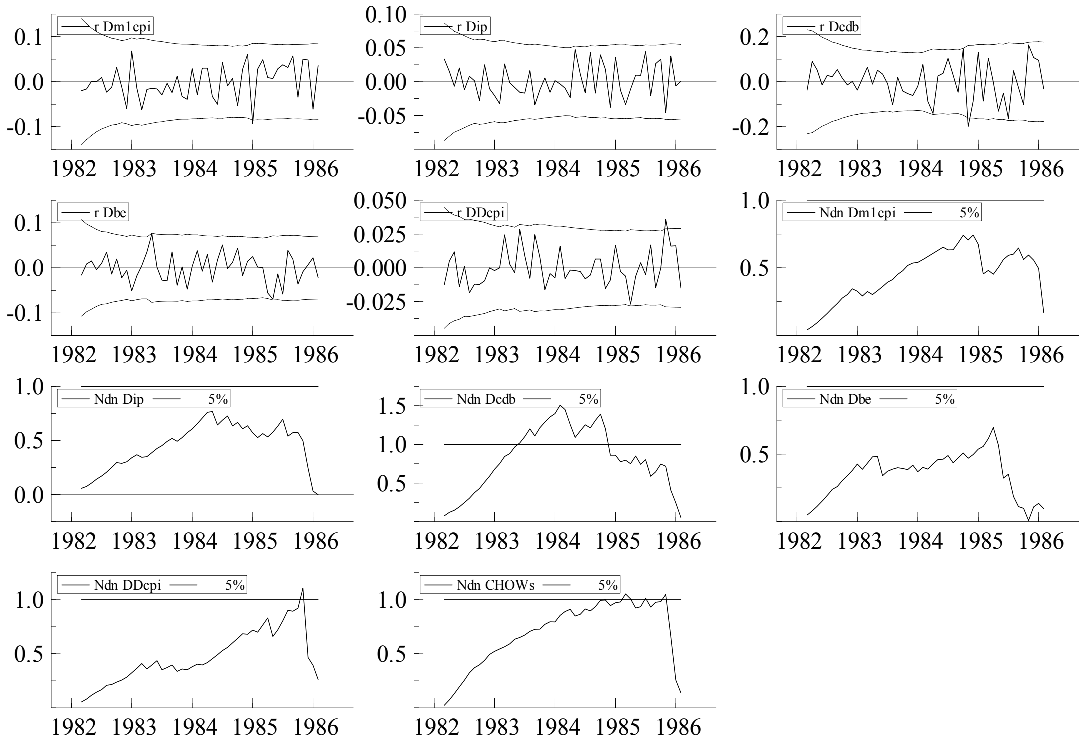
Further contrary to the conclusions presented by Durevall (1998) there is a role to play by nominal wage as a propagation mechanism for inflation in Brazil. The equation for inflation dynamics is in accordance to the hypothesis of inertial inflation being basically driven as an autoregressive process, though with a marginally significant coefficient, and the long run money demand.

The diagnostic test results show no sign of non-congruence in the residuals at the system's level and the results for the break point Chow test presented in figure 6 indicate the same structure of those presented in figure 5, namely evidence of parameter instability in equation for Δcdb and in the final sample for the system as a whole and inflation individually

Table 20 Diagnostic Tests SEM (M2) 1980/1 – 1986/2

Test\Equation	$\Delta m1-cpi$ *(p-value)*	Δip *(p-value)*	Δcdb *(p-value)*	Δbe *(p-value)*	$\Delta\Delta cpi$ *(p-value)*	*System (p-value)*
AR 1-5	2.63*	2.27	2.75*	2.26	2.78*	0.96
	(0.035)	(0.06)	(0.029)	(0.06)	(0.027)	(0.59)
Normality	0.34	3.40	0.97	0.18	3.36	10.28
	(0.84)	(0.18)	(0.61)	(0.91)	(0.18)	(0.41)
ARCH	0.71	0.15	2.07	0.55	1.28	
	(0.61)	(0.97)	(0.08)	(0.73)	(0.28)	
Hetero	1.18	0.67	1.03	1.15	0.35	0.84
	(0.31)	(0.80)	(0.44)	(0.33)	(0.98)	(0.92)

Figure 6 One Step Residuals and Chow Breakpoint Tests SEM (M2)



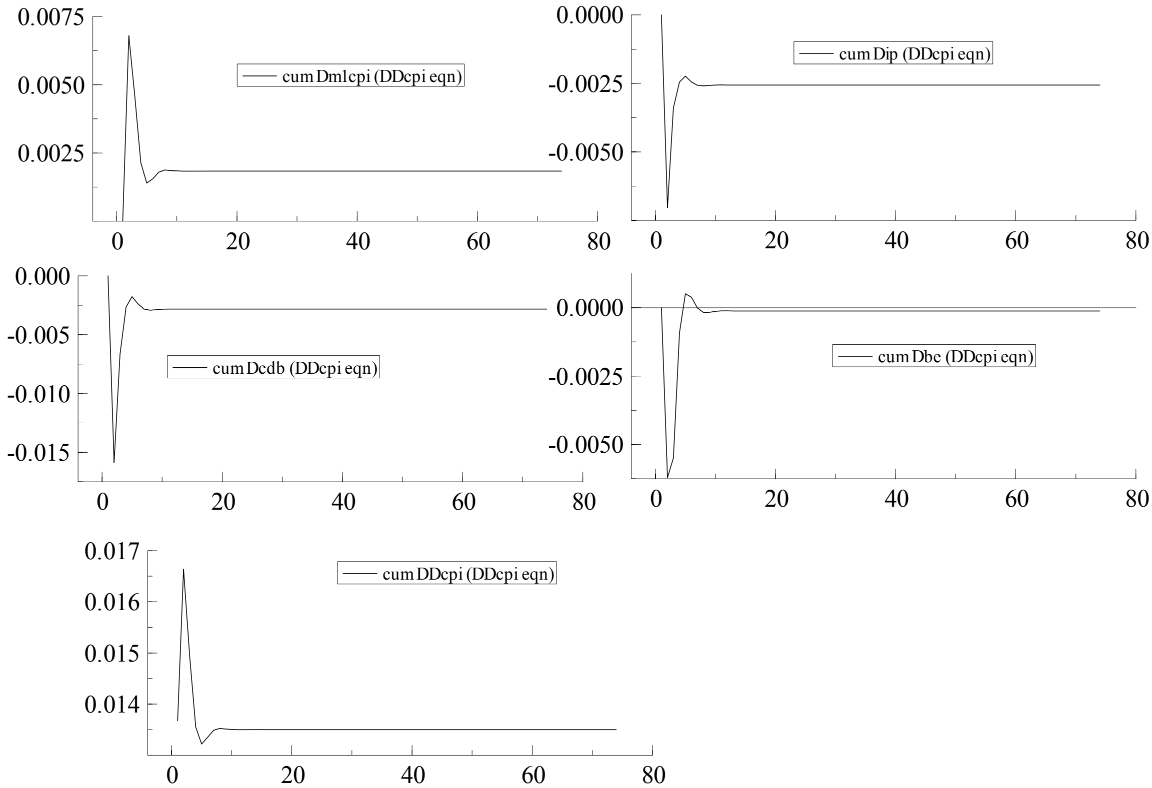
A totally different structure emerges in the impulse response functions as shown in figure 7. The effect of a one standard deviation shock on $\Delta\Delta cpi$ is basically concentrated on $\Delta\Delta cpi$ itself showing that changes in the rate of growth of inflation are incorporated in the inflation dynamics in accordance with the hypothesis of high levels of indexation with the shock on $\Delta\Delta cpi$ showing the same degree of inertia as in the

SEM presented in section 3. Such result is also in line with the findings in Barbosa and McNelis (1990) who found evidence in favor of wage inflation being linked to increasing inertia in the Brazilian inflationary process, whereas Novaes (1991) found certain degree of inertia in the inflationary process similarly to that obtained by Durevall (1998) when modelling persistence in inflation.

Table 21 Empirical Estimates Forecast Encompassing Equations[27]

$$e_{1t} = \underset{0.6742}{-0.6803}\,\phi_2 \qquad RSS = 0.0474 \quad \hat{\sigma} = 0.0454$$

Figure 7 Cumulative One Standard Deviation Impulse Response Functions Shock To (From) M2 1980/1 – 1986/02



As shown in table 21 the empirical counterpart of equation 14 in the HC test does not reject the hypothesis that $\alpha = 0$ indicating that model 2 does not add information to model one's forecast error. Unfortunately the empirical counterpart of equation 15 had no significant values for testing which led us to conclude that model 2

[27] The forecasts generated are out of sample dynamic forecasts for 24 periods (1986/3 to 1988/2). The variable under analysis is $\Delta\Delta cpi$. Values below coefficients are standard deviations and values inside square brackets are p-value. All regressions are OLS based.

cannot forecast encompass model 1. Such result is somewhat expected given that both models have very close forecasts for the period since their dynamic properties are very similar and suffer from severe forecast failure caused by the sequence of interventions that followed the Cruzado plan cast in February 1986.

5) Conclusions

This paper presents an analysis of the main results obtained by deriving SEMs following the LSE's methodology. The main assumption in deriving the econometric models is that valuable information in specifying them can come from numerous different sources beyond economic theory, including information from the observable sample of relevant variables and economic history in such way that we expect to relate the empirical model to the actual mechanism that is generating the data rather than to theory only which means then that theory and data form the DGP. The analysis follows assuming that modeling joint densities is essential in pursuing an empirical analysis of economic time series and that knowledge acquisitions are progressive and partial. Under this methodological framework the paper presents contributions to the literature about the Brazilian high inflation period.

In section 3 the SEM derived demonstrates the role of economic activity measured by the industrial production in the long run identified cointegrating vector and how it relates to the real interest rates. Also the short run price dynamics is influenced by the first difference of the industrial production index. Such result contrasts to the usual findings in the literature that the demand variables had no role in the accelerating inflation period as stressed in Durevall (1998) where the author using the same methodological approach, namely the LSE methodology, derived a one equation model for inflation in Brazil including the period we analyzed here and could not find any influence from demand variables in the econometric model. The results found also contrasts to those obtained by Juselius (2002) where the author concentrated on the demand for M3 and explains the accelerating inflation using a trend liquidity ratio. Whilst her model explains how inflation expectations played a significant role on the demand for M3 through the black market exchange rate it does not show any link between the economic recovery observed from 1984 onwards that had strong links with the acceleration in prices observed.

The SEM allows us to emphasize its role with the dummy 1984.2 having a positive effect on $\Delta\Delta cpi$ and on the industrial production index. Further the equation for the rate of growth in inflation shows how inflation rate gained momentum and why inflation assumed a pattern of persistence.

In section 4 we extend the analysis to discuss the role of the administered nominal exchange rate devaluations as a complementing theory explaining the price formation in the period as well as the role of wage inflation in the model following Durevall (1998). Implicitly we assume that two forces had a significant effect on the price dynamics in Brazil, namely the administered nominal exchange rate devaluations and nominal wage inflation. In particular two main questions are of interest. The first one is, to what extent did the administered exchange rate policy followed by the government during the first half of the 1980's influence the price dynamics? The second question is did the nominal wage inflation have a role to play in price dynamics during the period of high inflation? This last question has been discussed in the theoretical models that generate persistence in inflation but, in the empirical attempts, doubts were cast on the relevance of such variable in determining the price dynamics during the first half of the 1980's in Brazil.

Given the sample size we opted for including the variable in the SEM but without assuming its exogeneity to the long run parameters of the cointegrating vectors derived in chapter 5. So we carry out the weak exogeneity test of nominal wage inflation. In doing so we preserve the LSE general to specific approach, considering therefore the SEM a restricted model derived from testable hypotheses against a (congruent) general system and avoiding a priori hypothesis about the exogeneity of nominal wage inflation. The first question is addressed through the use of the AER hypothesis in deriving an equation for the administered devaluations which is estimated separately as a long run solution to the exchange rate devaluations policy equation.

In section 4 a new SEM is presented including the nominal wage inflation and the long run solution to the administered exchange rate policy. The section presents a long run solution to the exchange rate policy in accordance with the theoretical model proposed with the expected signs. The long run vector nevertheless is not significant in the SEM as a whole, a result that is somewhat in accordance to the result obtained in Durevall (1998) who shows that the PPP hypothesis which has a similar specification than the AER theoretical model proposed broke down after 1979. A long run money demand equation is identified being positively related to the output and real interest rate where inflation is measured as a weighted average between consumer and wage inflation, a result difficult to explain.

The equation for $\Delta\Delta cpi$ in the SEM with a positive sign for Δnw is clearly showing that the impact on the industrial activity was originated in the nominal wage inflation likely because of the accommodating monetary policy implemented in the period, a contrasting result to the findings in Durevall (1998) who found no role to play for nominal wage inflation in price inflation dynamics in Brazil. The equation for inflation dynamics is in accordance to the hypothesis of inertial inflation being basically driven as an autoregressive process, though with a marginally significant coefficient, and the long run money demand. The effect of a one standard deviation shock on $\Delta\Delta cpi$ is basically concentrated on $\Delta\Delta cpi$ itself showing that changes in the rate of growth of inflation are incorporated in the inflation dynamics in accordance with the hypothesis of high levels of indexation with the shock on $\Delta\Delta cpi$ showing the same degree of inertia as in the SEM presented in section 3. In the forecast encompassing exercise the model derived in section 3 does not encompass the SEM derived in section 4.

References:

ANDIMA. (1997) *Taxa de juros: Um amplo estudo sobre o mercado aberto no Brasil.* Séries Históricas, Associação Nacional das Instituições de Mercado Aberto, Rio de Janeiro.

Baer, W. (1989) *The Brazilian Economy: Growth and Development*, New York: Praeger.

Barbosa, F.H., McNelis, P. D. (1990) Indexation and inflationary inertia: Brazil 1964-1985, *The World Bank Economic Review*, 3(3): 339-357.

Blejer, M. I., Leiderman, L. (1982) A monetary approach to the crawling-peg system: Theory and evidence, *The Journal of Political Economy*, 89(1): 132-151.

Bontemps, C., Mizon, G. E. (2003) Congruence and encompassing, in Stigum, B.(ed) *Econometrics and the Philosophy of Economics*, Princeton: Princeton University Press, 354-378.

Cagan, P. (1956) The monetary dynamics of hyperinflation, in Friedman, M.(ed) *Studies in the Quantity Theory of Money,* Chicago: University of Chicago Press.

Calomiris, C. W., Domowitz, I. (1989) Asset substitution, money demand, and the inflation process in Brazil, *Journal of Money, Credit and Banking*, 21: 78-89.

Canova, F. (1995) The economics of VAR models, in Hoover, K. D. *Macroeconometrics: Developments, tensions and prospects,* Boston: Kluwer Academic Publishers.

Cardoso, E. (1983) A money demand equation for Brazil, *Journal of Development Economics*, 12: 183-193.

Clements, M.P, Mizon, G. E. (1991) Empirical analysis of macroeconomic time series, *European Economic Review,* 35: 887-917.

Diebold, F.X., Lopez, J.A. (1998) Forecast evaluation and combination, in: Maddala, G.S. and Rao, C.R. *Handbook of Statistics*. Amsterdam: North Holland, 14: 241-268

Doornik, J. A. (1998). Approximations to the asymptotic distribution of cointegration tests, *Journal of Economic Surveys*, 12(5): 573-593.

Doornik, J. A., Hansen, H. (1994) *A practical test for univariate and multivariate normality,* Discussion Paper, Nuffield College, Oxford University.

Doornik, J. A., Hendry, D. (1997) *Modelling Dynamic Systems Using PcFiml 9.0 for Windows*. London: International Thomson Business Press.

Durevall, D. (1998) The dynamics of chronic inflation in Brazil, 1968-1985, *Journal of Business & Economics Statistics*, 16(4): 423-432.

Engle, R. F. (1982) Autoregressive conditional heterocedasticity with estimates of the variance of United Kingdom inflation, *Econometrica,* 50: 987-1007.

Engsted, T. (1993) Testing for rational inflationary bubbles – the case of Argentina, Brazil and Israel. *Applied Economics*, 25(5): 667-674.

Engsted, T. (2002) Measures of fit for rational expectations models, *Journal of Economic Surveys,* 16(3): 301-355.

Fadil, F.G., MacDonald, R. (1992) *The demand for money in Brazil revisited.* Discussion Paper 92-05, Department of Economics, University of Aberdeen.

Fair, R.C., Shiller, R.J. (1989) The informational content of ex ante forecasts, *Review of Economics and Statistics,* 71:325-331

Feliz, R. A., Welch, J. H. (1997) Cointegration and tests of a classical model of inflation in Argentina, Bolivia, Brazil, Mexico and Peru, *Journal of Development Economics,* 52: 189-219.

Gerlach, S., Simone, S. F. (1985) A money demand equation for Brazil, *Journal of Development Economics,* 18: 493-501.

Greenslade, J. V., Hall, S. G. and Henry, B.S.G. (2002) On the identification of cointegrated systems in small samples: a modelling strategy with an application to UK wages and prices, *Journal of Economic Dynamics and Control,* 26 1517-1537.

Gilbert, C. L. (1986) Professor Hendry's Econometric Methodology, *Oxford Bulletin of Economics and Statistics*, 48(3): 283-307.

Hendry, D. F. (1995) *Dynamic Econometrics.* Oxford: Oxford University Press.

Hendry, D. F. (2000) On detectable and non-detectable structural change, *Structural Change and Economic Dynamics,* 11: 45-65.

Hendry, D. F., Doornik, J. A. (1994) Modelling linear dynamic econometric systems, *Scottish Journal of Political Economy,* 41:1-33.

Hendry, D. F., Mizon, G.E. (1993) Evaluating dynamic econometric models by encompassing the VAR, in Phillips, P.C.B. (ed) *Models Methods and Applications of Econometrics*. Cambridge: Blackwell.

Hendry, D. F., Richard, J.F. (1982) On the formulation of empirical models in dynamic econometrics, *Journal of Econometrics,* 20: 3-33.

Johansen, S. (1994) Testing weak exogeneity and the order of cointegration in UK money demand data in: Ericsson, N. R., Irons, J. S. (eds) *Testing Exogeneity*, New York, Oxford University Press.

Johansen, S. (1995) *Likelihood-Based Inference in Cointegrated Vector Auto-Regressive Models*, Oxford: Oxford Unversity Press.
Juselius, K. (2002) *Inflation, money growth and I(2) analysis*. Mimeo Department of Economics, University of Copenhagen.

Lopes, F.L., (1983) Inflação e nível de atividade no Brasil: Um estudo econométrico, *Pesquisa e Planejamento Econômico*, 12(3): 639-670.

Lütkepohl, H. (1994) *Introduction to Multiple Time Series Analysis*. Berlin: Springer-Verlag.

Mackinnon, J. (1991) Critical Values for Cointegration Tests, in Engle, R.F., Granger, C.W.J. *Long–Run Economic Relationships: Readings in Cointegration*, New York: Oxford University Press.

Maddala, G. S., Kim, I.-M. (1998) *Unit roots, Cointegration and Structural Change,* Cambridge: Cambridge University Press.

Mizon, G. E. (1995) Progressive modelling of macroeconomic time series: The LSE methodology, in Hoover, K. D. *Macroeconometrics: Developments, tensions and prospects,* Boston: Kluwer Academic Publishers.

Modiano, E.M. (1983) A dinâmica de salários e preços na economia brasileira: 1966/81, *Pesquisa e Planejamento Econômico,* 13(1): 39-61.

Newbold, P., Harvey, D.I.(2002) Forecast combination and encompassing, in Clements, M.P., Hendry, D. F. (eds) *A Companion to Economic Forecasting*, Oxford: Blackwell Publishers.

Novaes, A. D. (1991) Um teste da hipótese de inflação inercial no Brasil, *Pesquisa e Planejamento Econômico*, 21(2): 377-196.

Novaes, A. D. (1993) Revisiting the inertial hypothesis for Brazil, *Journal of Development Economics,* 42: 89-110.

Phylaktis, K., Taylor, M. P. (1993) Money demand, the Cagan model, and inflation tax: some Latin America evidence, *The Review of Economics and Statistics,* 75: 32-37.

Rossi, J. W. (1994). O modelo hiperinflacionario da demanda por moeda de Cagan e o caso do Brasil, *Pesquisa e Planejamento Economico,* 24(1): 73-96.

Taylor, J. B. (1979) Staggered wage setting in macro model, *American Economic Review*, 69(2): 108-113.

White, H. (1980) A heteroskedastic-consistent covariance matrix estimator and a direct test for heroskedasticity, *Econometrica*, 48: 817-838.