Overshooting agricultural prices and the importance of economic structure: evidence from Brazil

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Summary: 1. Introduction; 2. Overshooting agricultural prices: theory and evidence; 3. Exchange rates and monetary policy in Brazil; 4. Methodological aspects; 5. Data and results; 6. Conclusions.

This paper uses the case of Brazil to demonstrate that findings of overshooting in agricultural prices in response to monetary shocks depend on the type of time series formulation selected and on the inclusion of the exchange rate in the model. Alternative choices in terms of the degree of recursiveness or simultaneity of the models made relatively little difference to the results.

Este trabalho utiliza o caso brasileiro para demonstrar que o fenômeno de *overshooting* de preços agrícolas em resposta a choques monetários depende do tipo da formulação de séries temporais selecionadas e da inclusão ou exclusão da taxa de câmbio no modelo. Observou-se que a utilização de diferentes combinações em termos do esquema recursivo ou das relações simultâneas dos modelos faria pouca diferença para os resultados.

1. Introduction

Much recent work has focused on the effects of monetary policy on the agricultural sector, and particularly on the question of the overshooting of agricultural prices in response to a monetary shock. The empirical question of whether relative prices do in fact overshoot their long run equilibrium has been intertwined with an ongoing debate over the proper methodological approach to be used in evaluating the question. In general, the debate has involved those who rely on structural econometric models and those who eschew what they regard as *ad hoc* prior restrictions required in building structural models in favor of reduced form models emphasizing the time series characteristics of a relatively small number of related variables.

This paper analyzes the case of Brazil using several alternative methodologies to demonstrate the sensitivity of the results to the modelling approach used. In particular, it is shown that while it is essencial to take into account the long run properties of the relevant time series, it is equally important to consider the structural properties of the economy under analysis when estimating reduced form vector autoregression (VAR) models or vector error correction (VEC) models. Failure to do so can result in spurious findings of overshooting responses due to misspecification errors.

These findings lend support to the idea that the abstractions necessary in any modelling effort must rely both on theory and on the actual structural characteristics of the economy. Several authors have emphasized the basically arbitrary nature of the variance decomposition methods underlying VAR and VEC models due to the need to specify *a priori* the caus-

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al ordering of the variables included to obtain an identified model capable of estimation. Orden & Fackler (1989) have emphasized the arbitrariness of the assumptions underlying particular orderings, and the need to consider a wide range of possible assumptions in terms of the degree of recursiveness in any particular model. Particular orderings and selections of variables have implications for the form which the relevant supply and demand functions can take. For example, Orden and Fackler show that assuming money is exogenous (ordered first) in a three variable system is tantamount to assuming either a perfectly inelastic money supply or a perfectly elastic money demand (with respect to the interest rate). McTaggart (1988) shows that alternative assumptions regarding the parameters of these supply and demand functions are important in determining the variability of relative agricultural prices.

While the question of the proper ordering can be addressed in small models by testing all possible orderings of variables, it leaves open the question of which variables to include in the first place. In evaluating the question of the response of relative agricultural prices to monetary shocks, it has been common for time series analysts to consider only three variables: money supply, agricultural prices, and non-agricultural prices. (See for example, Robertson & Orden, 1990, and Bessler, 1984.) This relative parsimony is to some extent a result of data constraints and also of the difficulty in interpreting large scale VAR or VEC models. In the case of Brazil, there is strong evidence from structural models that the exchange rate is an important omitted variable in such specifications. This empirical evidence has firm theoretical foundations; there is ample support for the importance of an independent effect for the exchange rate in a model including money and relative prices.

The next section reviews the literature on overshooting and monetary impacts on agricultural prices, together with studies focusing on Brazil *per se*. Next, methodological aspects are discussed. This is followed by an evaluation of empirical results for the Brazilian case, using various time series models. These results are then reconsidered on the basis of an expanded model allowing for exchange rate effects. Last, the results are summarized and conclusions presented.

2. Overshooting agricultural prices: theory and evidence

Overshooting was first emphasized as an important phenomenon by Dornbusch (1976) in the context of exchange rates. The basic phenomenon illustrates what is known in physics as the "Le Chatelier" principle: if a system is composed of two or more parts which adjust to shocks at different speeds, then the effects of a shock are initially more pronounced on the more flexible part of the system even if in the long run all parts are affected proportionately. This may (but will not necessarily) result in an overshooting of the long-run equilibrium for the more flexible part of the system. In an economy with a fast-adjusting flex-price agricultural sector and a slow-adjusting "fix-price" non-agricultural sector, a positive monetary shock, for example, can result in agricultural prices rising above their long run equilibrium in the short run.

So, an economy in which money is neutral in the long run may still exhibit non-neutralities over shorter periods. Methodologies to evaluate the existence of such phenomena must therefore allow for non-neutralities in the short-run while imposing monetary neutrality in the long run. It is precisely this problem which VEC models are designed to address as fully discussed in the methodological aspects section. If money is indeed neutral in the long run, then we should find cointegration between series of money, agricultural prices, and non-ag-

ricultural prices. We can use this cointegration relationship, if it exists, to impose restrictions on reduced form models allowing consistent and efficient estimation of parameters.

As noted by Burnquist (1992) and Robertson & Orden (1990), a VAR model in levels provides consistent estimates but is less efficient than a VEC model which includes the error correction terms which represent the relative flexibility of the adjustment mechanisms in each price series. A VAR model in differenced data is misspecified because it omits the information regarding the level of each variable which allows the imposition of long run neutrality shown by a finding of cointegration.

While previous analysts have emphasized the error inherent in ignoring the information content of the levels of the variables of interest, many have neglected potential problems arising from the method of selecting variables to be included in the model, often focusing instead on designating exo- or endogeneity whichever variables were chosen. Just as ignoring the information in levels can produce spurious results, so too can omitting important causal variables.

Theory admits of a multitude of potentially important candidates for inclusion; ultimately the decision, if it is to be based on anything other than simple speculation, must rest on the structure of the economy being considered and the resulting characteristics of the data it generates. Examination of the structure of the mechanism generating the data will not only provide a factual basis for choosing which variables to include, but will often provide a basis for choosing when and where to use recursive *vs.* structural orderings for variables in reduced form models.

Accordingly, there are two potential problems with an evaluation of the existence of overshooting agricultural prices in Brazil (or elsewhere). First, the model should be based on the structural characteristics of the case at hand. Failure to include important variables may result in a spurious determination of causality or it may be that a causal variable is relatively unimportant even if it is significant. This is of particular importance for those analysts interested in providing relevant advice to policy makers. It makes little sense to get excited about a significant result for one causal variable if in fact most of the variation in the object of interest is the result of something entirely different.

Second, it is important to evaluate the robustness of the results with respect to alternative methodologies; VAR in differences or levels *vs.* VEC. Here it is important to note that even though a VAR in differences may be misspecified, it is entirely possible that this makes no practical difference. That is, one may still obtain the same parameter estimates if the misspecification bias is small. This implies that each of the models should be evaluated not only on their ability to explain the variance in the system (i.e. reduce the mean squared error), but also on whether the estimates obtained are materially different.

Previous analyses of overshooting agricultural prices have for the most part confined themselves to examining three variable systems with money and two prices via various methods. In the Brazilian case, Bessler (1984) estimates a three variable VAR in levels for Brazil while Robertson & Orden (1990) estimate a three variable VEC for New Zealand. It is clear that more complex formulations are possible since there are clearly more potentially independent sources of variation in the financial/monetary sector than simply money supply and the two relevant prices. Indeed, even the most convinced monetarist would admit that, in an open economy, *foreign* monetary shock as well as domestic ones can be important, motivating examination of the exchange rate as an important variable to be considered.

It is this fact, plus numerous empirical studies (see e.g. Taylor & Lysy, 1980, and the citations therein, or Krueger, Schiff & Valdés, 1991), that attests to the independent impor-

tance of exchange rate as an important omitted variable in the commonly used three variable model. Burnquist (1992) constructed structural models of Brazilian economy and was in accord with previous authors in finding the exchange rate to be far more important in determining agricultural sector prices than is money supply. The importance of considering a broad range of explanatory variables in the Brazilian case is demonstrated in Burnquist's analysis, where the exchange rate and level of production accounted for more than 70 percent of the total variation in food prices. This finding, based on the forecast error variance decomposition for a quarterly general macro model including nine variables in the agricultural system, is robust to various reformulations of the structural relationships. (Details on the full model are available from the authors on request or can be found in Burnquist, 1992.)

The theoretical foundations of these empirical studies include economic structure allowing differential effects of exchange rates on different sectors of the economy. A simple traded/non-traded goods formulation will generate this result analytically, while different sectoral import requirements on the input side can do so as well. See, for example, Bruno (1976) for an exposition of the standard Australian model including a monetary sector and making the distinction between traded and non-traded goods. Kyle (1992) extends this analysis to include different imported input requirements.

In terms of Orden and Fackler's analysis of implicit assumptions resulting from exclusion of important variables, omitting the exchange rate is tantamount to assuming that agricultural prices in this model are perfectly inelastic with respect to the exchange rate or that agricultural output is perfectly elastic. Neither assumption is warranted on the basis of the facts presented in numerous empirical analyses of the Brazilian economy such as those cited above.

The next section examines exchange rates and monetary policy in Brazil over the period of interest in order to gain some insight into the structure of the relevant portions of the economy. The following section discusses methodological aspects. Next, the sections present results obtained from a structural model and compares them with those from three and four variable VEC models (with and without the exchange rate) to evaluate the impact these different approaches have on the results obtained. The VEC results are also compared with VAR models in levels and in first differences to evaluate the effects of different assumptions regarding the time series properties of the variables included.

3. Exchange rates and monetary policy in Brazil

During the 70's and early 80's, Brazil maintained a "crawling peg" system under which the exchange rate was fixed in nominal terms and devalued on frequent but unannounced basis. The frequency of devaluations during this period was from three to 14 days; after 1984 this was increased to virtually a daily basis. The objective of this system was to maintain the real value of the cruzeiro, and thereby stabilize export receipts and speculation.

However, the rate of devaluation did not in fact keep pace with the difference between domestic and international deflation, with the result that the cruzeiro became increasingly overvalued after the mid-70's. Domestic pressure resulted from two factors: devaluations were seen as adding further to inflationary pressures and the local currency cost of servicing the mounting foreign debt of Brazilian firms and state entities mounted with each decrease in the value of the currency (see Zini, 1988, and Baer, 1983).

In terms of modeling the exchange rate, the evolution of government policy justifies treating it as an exogenous factor. Though the stated intent linked this policy with inflation, the evolution of the nominal exchange rate did not in fact behave in this manner over the period of interest in this study.

The money supply in Brazil increased steadily throughout the period under examination. After 1973, the government budget was perenially in deficit, fueled in large part by large government payrolls and subsidy programs, particularly that for agriculture. Given the limitations of domestic tax capacity, and the constraints on borrowing all that was required to balance the budget, frequent recourse to money emission resulted (see Pereira & Nakano, 1987).

The 1981/82 period marked a sharp change in government policy as money supply growth was restricted to below the rate of inflation. The resulting high interest rates led to an explosion in external debt, and were not ultimately successful in restraining inflation. After the onset of the international debt crisis in 1983, recourse to foreign borrowing was effectively cut off and the government was forced to fund itself through a combination of internal debt and money emission. The resulting inflationary spiral was eventually brought to a temporary end in 1986 with the advent of the Cruzado Plan, which combined a new currency with a price freeze.

The upward ratcheting of Brazil's rate of inflation over the period is among the more extreme on record. Fueled by massive positive monetary shocks, its progression was the result of numerous government policy initiatives combined with chronic overspending. Unlinked as this spending was to actual economic conditions, there is ample justification for regarding money supply growth as an exogenous factor in the models developed below, just as there is a basis for considering the various controlled exchange rate regimes as exogenous also. Consequently, various formulations were tested to determine whether the results were sensitive to alternative assumptions regarding the exo-endogeneity of these variables.

4. Methodological aspects

The methodological approach used to estimate the models in this work is a Bernanke-type vector autoregression model. Vector autoregressions (VARs) have been proposed as an alternative to large-scale macroeconometric models by Sims (1980). Sims' proposed alternative was to fit a VAR, where all variables were endogenous, in order to avoid "arbitrary identifying restrictions". In a VAR, each variable is regressed on all variables lagged, which are considered to be exogenous. A set of orthogonal shocks is usually expressed through a diagonal covariance matrix, which can be considered "equivalent" to imposing a set of identifying restrictions that assume a recursive structure for the economy (Orden & Fackler, 1989). The problem is then to justify the recursiveness of the economic system and to determine the correct ordering of the variables, as discussed above, as well as to overcome arbitrary interpretation of the results, given that different forms can be "easily rearranged" for the variance covariance decomposition matrix. Mount (1988) has argued that the determination of the adequate ordering of the variables in the recursive system is basically arbitrary, since there is no mechanism in the VAR approach which allows the determination of the correct ordering based on data information.

More recently, some authors have used the features of the VAR models without imposing recursiveness (Bernanke, 1986; Barros, 1992; Burnquist, 1992, and Sims, 1986).

These were pioneered by Bernanke, and were popularized as the Bernanke VAR approach. The basic difference between the Bernanke method and a conventional VAR approach is that non-recursive structural assumptions can be considered, which allows economic interpretations of the results.

The controversies about VAR models have gained force as small changes in their specification have yielded conflicting answers (Robertson & Orden, 1990). In the literature, it has been suggested that these conflicting results may be due in part to behavioral differences among economies (Bessler, 1984; Devadoss & Meyers, 1987). Others have argued that it can be due to failure to adequately evaluate the time series characteristics of the data, or take into account the implications for model specification and hypothesis testing of non-stationary due to unit roots common among aggregate macroeconomic variables (Robertson & Orden, 1990).

Until recently, even though the levels of various economic time series are known to be non-stationary, not much consideration has been given to the phenomenon in most econometric studies. Instead, it has simply been assumed that the data processes are stationary (Hendry, 1986). This same author points out that to overcome the non-stationarity problem, some researchers have suggested differencing the data to remove random walk and trendlike components, while others have argued that this loses valuable "long-run" information. Advances in understanding time series data which are non-stationary because they contain unit roots have been related to the concept of cointegration and to error correction mechanisms.

The concept underlying cointegrated series can be viewed, on the one hand, as a resolution of the debates about the treatment of non-stationarity of time series. On the other hand, cointegrated series have an error correction mechanism representation which can be evaluated as a way to impose testable restrictions on the parameters of VAR models. The concept of cointegration refers to economic variables with long-run equilibrium relationships which are manifested by stationarity of linear combinations of the variables in a multivariate model. Non-stationarity series with independent unit roots have no tendency to move together in the long-run. Hence, independent unit roots are inconsistent with the existence of long-run equilibrium among individual time-series. However, some variables can have common unit roots. When this is the case, the linear combination of these variables will contain fewer unit roots than the individual series. Cointegration is the term used for such a relationship. If instead of evaluating for cointegration, the time series are differenced until they are all stationary, overdifferencing can result and information is lost (Robertson & Orden, 1990).

In an empirical framework, economic theory typically provides information about variables that move together and should not diverge too much from each other at least in the long run. These variables are expected to have a determined linear relationship in the long run. Therefore, empirical tests for cointegration allow econometricians to estimate and test for the relationships among variables postulated by economic theory.

An important implication of the cointegration theorem presented by Engle and Granger (1987) is that, if a long-run relationship has been established between a pair or set of variables, there always exists a dynamic vector error correction model (VEC) of the relationship and conversely, that VECs generate cointegrated series. The vector error correction model encompasses a VAR model in differences since it includes variables which have a long-run relationship. In essence, it represents a reparameterization of a VAR which takes into account long-run relationships among the variables. If these long-run relationships exist and are not taken into account, the model is misspecified (Robertson & Orden, 1990).

Derivation of the error correction model involves three steps. The first consists on examining the stationarity properties of the individual series. This is done by using statistical methods for models with unit roots. There are several methods that have been developed for this purpose. The most popular have been the Dickey-Fuller and the Augmented Dickey-Fuller approaches, which are based on a least squares estimator of a long AR regression. The presence of unit roots is determined under the null hypothesis that the series is integrated of order one. The test is based on distributions derived for the AR coefficients and regression t-statistics in models with pure random walks driven by i.i.d. innovations and simple autoregressive integrated (ARI) processes (Fuller, 1976; and Dickey & Fuller, 1979, 1981). The test for cointegration is based on the maintained hypothesis that the time series are integrated, or say, contain at least one common unit root. Cointegration is defined as a linear combination, which has a residual that is integrated of order zero of two or more series which are integrated of order one. Engle and Granger (1987) investigated those tests which consider for the null hypothesis that a pair of time series which are integrated were non-cointegrated against the alternative that they were cointegrated. In this context, the null hypothesis is that the linear combination of the series is non-stationary, while the alternative is stationarity due to common unit roots. The residuals from these regressions provide estimates of the equilibrium errors, and tests for cointegration among the series of residuals (Robertson & Orden, 1990). For first-order systems, the Durbin-Watson test and a Dickey-Fuller test were found to be the best. For higher-order systems, the Augmented Dickey-Fuller test (ADF), which is a generalization of the Dickey-Fuller test, was recommended by these authors. Finally, the residual from a cointegrated regression is entered into the model in order to obtain the error correction model (VEC).

Therefore, the models estimated are of the general form:

$$A_o\, y_t = \sum_{s=1}^{p} A_s\; y_{t-s} - \lambda\, Z_{t-1} + U_t \tag{1}$$

where:

y_t = vector of random variables;

A_o = matrix of contemporaneous coefficients;

A_s = matrix of coefficients for lagged relations;

Z_t = cointegration term, taken from the cointegrating equations $\mu' y_t$;

U_t = vector of 1 step ahead forecast errors;

$\mu = N \times r$ matrix of parameters where N is the number of variables and $r \leq N$ is the number of linearly independent cointegrating equations.

It should be noted that structural assumptions regarding exogeneity can be introduced by allowing non-zero off-diagonal elements for A_o. Interrelations among lagged values of each series are represented by non-zero elements for A_s. The error correction terms Z_t affect the process through and, following Engle and Granger, are estimated as the errors from the cointegrating regressions. For a "standard" VAR with no error correction structure, $\mu = 0$.

5. Data and results

This section presents first the sources of time-series data used for estimation of the econometric models. The results are presented in three sub-sections. First, the results of the integration test are presented. Next, the results of a three variable model including money, industrial prices and agricultural prices are presented. This formulation is identical to that used in previous studies of Brazil such as Bessler's, as well as analyses of other countries. In the last sub-section the results of a four variable model including the exchange rate are presented.

Data

The agricultural food price index used in this study is a weighted average of nominal prices at the wholesale level (given in cruzeiros/kg). The weights are the volume of production in each year. Until the 70's, the price data used to compose the index are those received by agricultural producers in São Paulo, as published by the Instituto de Economia Agrícola do Estado de São Paulo. Average prices for the whole country were compiled beginning in 1966 by the Fundação Getulio Vargas, but are only used for the period after 1970 in the present study. The observation that the tendency of average food prices for Brazil are compatible with those of São Paulo state allowed the composition of a price series which is a combination of these two different sources.

The wholesale food price index was calculated on a quarterly basis using the individual data series available for São Paulo state. The weight used to calculate the index was the annual value of production of each crop.

The prices for the industrial sector correspond to the wholesale price index for the Brazilian manufacturing industry compiled by the Fundação Getulio Vargas and published in *Conjuntura Econômica.*

The source of the data for the monetary system is the *Estatísticas históricas do Brasil — séries econômicas, demográficas e sociais/1550-1988 — IBGE.* The series for M1 was used for this study, and includes money in circulation plus cash deposits in commercial banks and Banco do Brasil.

Information on the exchange rate was taken from various issues of the *Boletim do Banco do Brasil.* The rate *vis à vis* the US dollar was used. This series is identical to that published in *Conjuntura Econômica.*

Integration test results

This section analyzes the question of whether the series are stationary, employing integration tests. Next, tests for cointegration are performed based on the maintained hypothesis that all the series are integrated of order one.

Table 1 reports the Augmented Dickey-Fuller (ADF) test statistics for integration relative to the three series listed. The ADF test was carried out after fitting various lags to the data until these were sufficient to ensure that white noise residuals were obtained. The appropriate number of lags chosen to be tested were based on previous similar studies (Robertson & Orden, 1990; Palaskas & Varangis, 1989). The conclusions about the number of lags to ensure white noise residuals have been determined by applying portmanteau Q-tests of

the RATS program.[1] The portmanteau results are presented along with the ADF test for integration.

The results of the test for stationarity, summarized in table 1, suggest that the hypothesis of a unit root cannot be rejected at a 90 percent significance level for all the three series when the Augmented Dickey-Fuller (ADF) statistical test for integration is employed.

Table 1
Integration tests for the long-run (ADF)

Series	Unit root test	Nº of lags in the ADF test	Portmanteau	
			Q-test	P
M1	1.17	4	15.33	0.22
IP	2.44	4	12.12	0.44
FP	2.22	4	6.19	0.96
ER	0.61	4	5.36	0.94
Critical Value (ADF)[1]				
90%	-2.63			

[1] Critical Value based on Fuller for a regression including an intercept and without a trend variable (1976: 371).

A structural model of food prices

The importance of considering a broad range of explanatory variables in the Brazilian case is demonstrated in Burnquist's analysis, where the exchange rate and level of production accounted for more than 70 percent of the total variation in food prices. This finding, based on the forecast error variance decomposition for a quarterly general macro model including nine variables in the agricultural system, is robust to various reformulations of the structural relationships. The agricultural block of the model reported in Burnquist (1992) is reproduced here. (The contemporaneous coefficient matrix showing the structure used is presented in the appendix. Details on the full model are available from the authors on request or can be found in Burnquist, 1992.) As shown in figure 1, the model tracks food prices extremely well, while the variance decomposition shown in table 2 demonstrates that the exchange rate appears to be far more important than the money supply in explaining the variance.

Figure 2 presents impulse response functions showing the response of food prices to shocks in money and the exchange rate. It is apparent that there is quite pronounced overshooting in response to the exchange rate while there is little evidence of the phenomenon in response to monetary shocks.

[1] From Doan Associates. Regression Analysis for Time Series, Version 3.11.

Three variable VAR and VEC models

A first step in constructing a model of money supply, agricultural prices, and industrial prices is an examination of the statistical properties of the variables to determine whether they are stationary and if there are any cointegration relationships between them. All three series were shown to have a unit root based on an Augmented Dickey-Fuller test of the series over the 1970-85 period. Table 3 presents the results of tests of cointegrating regressions among the three variables. It can be seen that the results support the hypothesis that there are long-run cointegrating relationships between money and each of the price variables.

Based on these results, a three variable VEC model was estimated, including four lags of each variable. The choice of four lags makes this study comparable to previous work (e.g. Robertson & Orden, 1990) and is confirmed as appropriate by portmanteau Q-tests on the residuals of the integrating regressions. Table 4 shows the contemporary coefficient estimates for the model, demonstrating the recursive ordering adopted (comparable to previous studies such as Bessler's).

Table 5 shows the forecast error variance decomposition of the estimated VEC model, where it can be seen that food prices are dependent mainly on own sources of variation, while this is even more pronounced in the case of money supply. Industrial prices depend both on own sources of variation as well as food prices. It is striking that money does not account for a significant portion of the variance in either of the two price series.

Figure 3A shows impulse responses in terms of percent changes for each of the two price series in response to a one standard deviation shock in money supply. Overshooting of agricultural prices is evident, but is not of a great magnitude at approximately 1 percent of the initial price level. Figures 3B and 3C show impulse responses for models estimated in differences (which is expected to yield biased results due to omission of information on levels), generating results in which overshooting is barely apparent.

A four variable model including the exchange rate

Based on the numerous considerations cited above, the three variable model was expanded to include the exchange rate. The exchange rate was found to have a unit root, and the cointegration tests reported in table 6 show that it is cointegrated with both industrial and agricultural prices. In this, as well as in the analysis that follows, the exchange rate is assumed to be an exogenous variable, an assumption in keeping with both prior analyses (see Barros and Burnquist) and with the history of exchange rate policy in Brazil, where the rate has in fact been set exogenously by the monetary authorities throughout the period under consideration. Also in accord with these analyses, food prices are assumed to affect industrial prices but not vice versa. These relationships are shown in table 7, where contemporaneous coefficient estimates for the model are presented.

Table 8, showing the forecast error decomposition for the model, demonstrates that exchange rates play a far more important role than does money in determining variations in food prices. After 24 months the exchange rate accounts for about 58 percent of this variation, while money accounts for a little more than 3 percent. In addition, there does seem to be a relatively important result when considering the effect of omitting the exchange rate from a model intended to evaluate the impacts of money supply, since the extent of bias in

estimating the coefficient for an included variable depends on its correlation with the omitted variable.

The impulse response functions obtained from this model are presented in figure 4. Panel (A) compares the response of food prices to shock in the money supply and the exchange rate respectively. No overshooting in response to monetary shock is apparent, while the plot for exchange rate shocks does in fact show signs of overshooting. In addition, the size of the exchange rate effect is far larger in absolute terms than that from money. This result is to be expected, given the relative unimportance of money in the forecast error decomposition presented above.

Panel (B) shows the same results based on a model in levels with no error correction term. Here the results are even clearer; if there is overshooting in agricultural prices, it is a result of exchange rates, not money. Panel (C), presenting the results from a model estimated in first differences only, again confirms that exchange rates can be an important source of overshooting. However, here the results for money can be interpreted as showing some signs of overshooting, though not nearly as pronounced as that for exchange rates.

Figure 1
Wholesale food price — actual *versus* stimulated

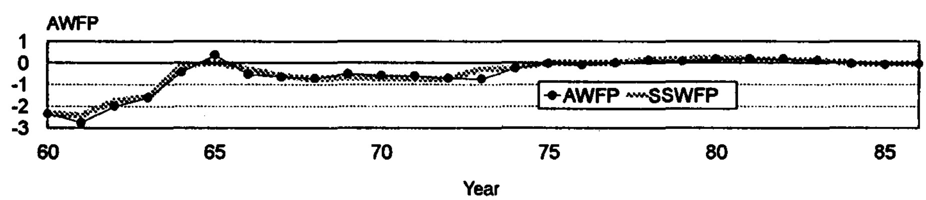
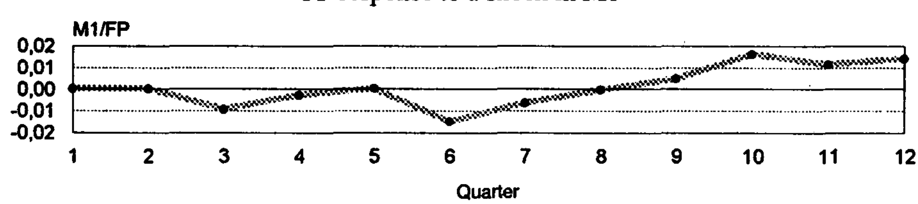


Figure 2
Response of food prices to money and exchange rate shocks — structural model

FP response to a shock in M1



FP response to a shock in ER

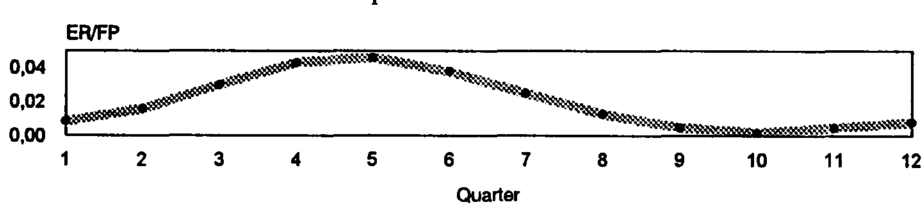
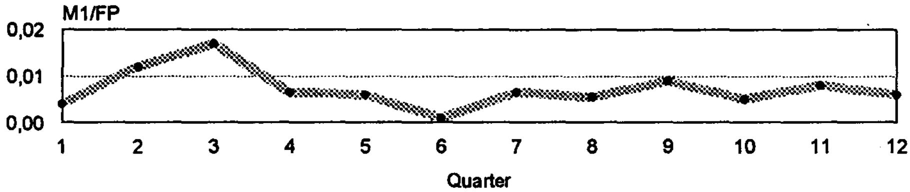


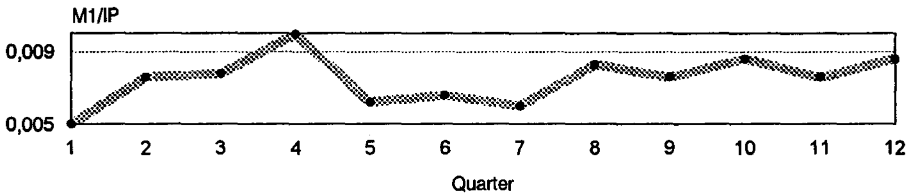
Figure 3
Response of food prices and industrial prices to money supply shocks — three variable mode

FP response to a shock in M1

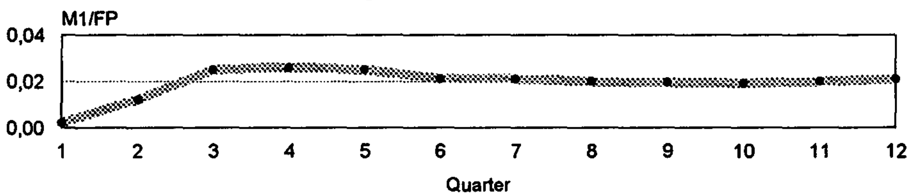


(A)

IP response to a shock in M1

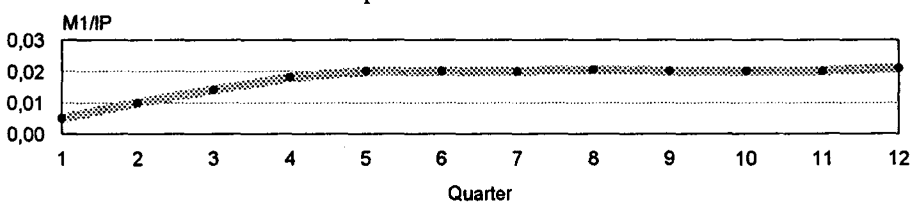


FP response to a shock in M1

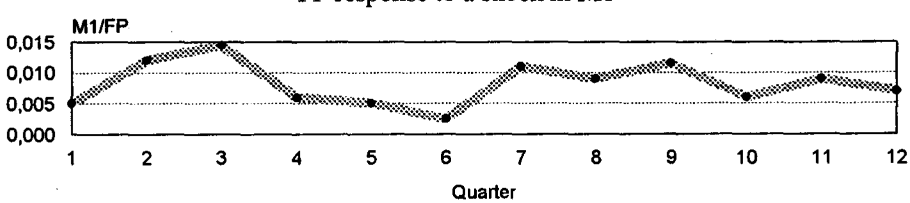


(B)

IP response to a shock in M1



FP response to a shock in M1



(C)

IP response to a shock in M1

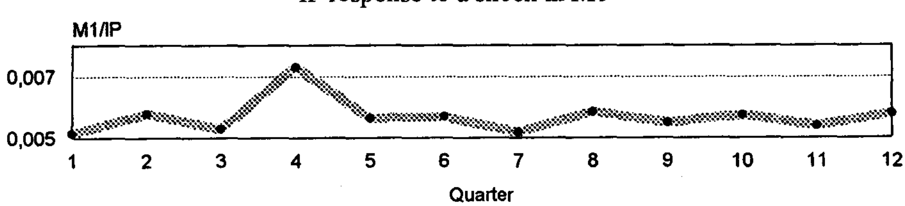
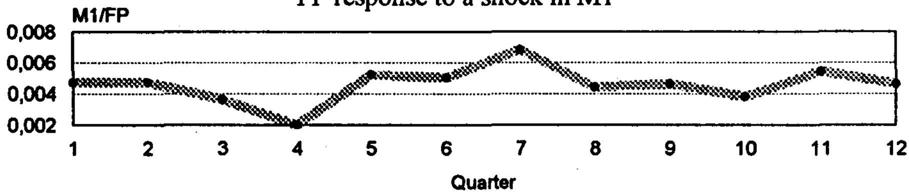


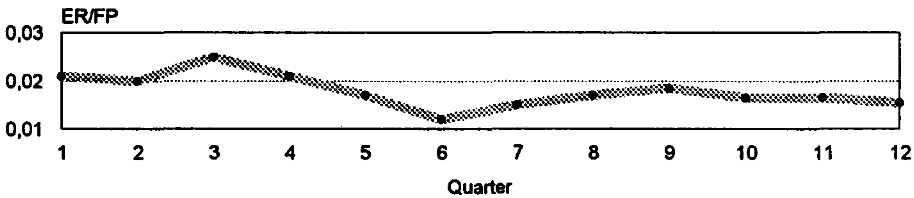
Figure 4
Response of shocks in money supply and exchange rates — four variable model

FP response to a shock in M1

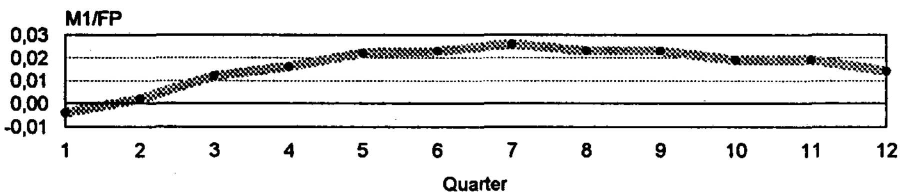


(A)

FP response to a shock in ER

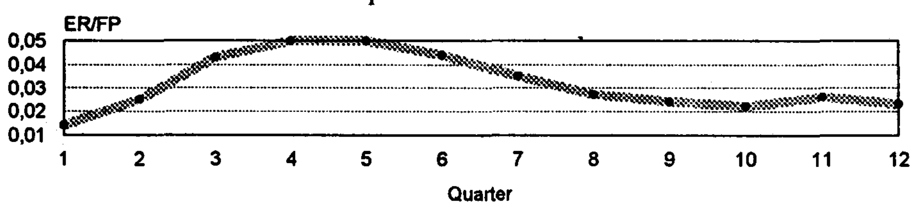


FP response to a shock in M1

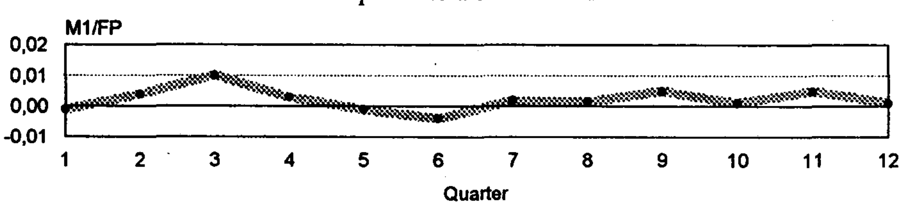


(B)

FP response to a shock in ER



FP response to a shock in M1



(C)

FP response to a shock in ER

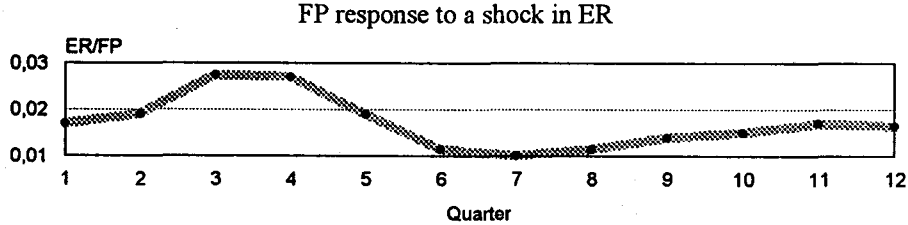


Table 2
Decomposition of variance of k-months ahead forecast errors
— structural model (%)

k	FP	TFP	EXPR	M1	ER	Y	PP	LC	TEI	IEP
FP										
3	0.70	85.04	0.04	0.08	12.00	0.00	1.99	0.13	0.00	0.00
6	0.42	60.74	1.00	0.05	23.21	0.03	8.68	0.72	1.53	3.59
12	1.56	20.37	0.32	2.14	64.05	0.89	3.17	1.63	0.45	5.40
24	1.32	9.27	6.64	3.62	59.00	4.87	1.37	2.92	1.65	9.32
TFP										
3	58.77	0.09	3.49	7.00	3.88	0.25	15.70	10.56	0.18	0.05
6	48.84	0.72	3.02	8.80	5.11	6.01	13.84	12.92	0.62	0.10
12	28.80	0.60	2.41	18.12	10.87	5.71	10.67	11.91	8.96	1.96
24	25.19	3.99	4.26	10.38	19.81	4.29	12.46	8.45	7.66	3.49
EXPR										
3	0.00	0.00	84.24	0.00	8.46	0.00	5.94	0.00	0.00	1.35
6	0.79	1.36	52.10	10.77	5.36	2.72	19.37	0.97	0.42	6.12
12	1.97	1.11	42.59	19.41	4.50	2.41	18.99	3.12	0.93	4.94
24	4.30	5.20	20.58	20.51	6.42	3.16	8.52	16.23	11.74	3.32
M1										
3	0.00	0.00	0.00	100.00	0.00	0.00	0.00	0.00	0.00	0.00
6	0.51	0.17	0.00	64.84	6.27	4.29	0.28	21.05	1.06	1.51
12	5.55	0.37	3.20	43.04	4.37	7.90	1.14	20.35	12.72	1.34
24	1.14	9.84	8.47	23.21	4.26	7.23	1.89	10.22	24.27	6.45
ER										
3	0.00	0.00	0.00	0.00	100.00	0.00	0.00	0.00	0.00	0.00
6	0.02	0.22	0.03	0.32	91.95	0.93	0.87	1.22	1.82	2.61
12	0.24	0.49	0.72	6.51	76.05	2.58	1.41	0.55	1.53	9.90
24	0.25	6.81	1.72	9.65	54.53	1.91	0.77	5.83	9.06	9.45
Y										
3	0.00	0.00	0.00	0.00	0.00	100.00	0.00	0.00	0.00	0.00
6	5.64	0.15	1.35	3.41	0.22	65.60	7.95	8.04	7.22	0.40
12	3.81	8.08	27.28	1.67	1.15	20.81	5.38	6.88	21.98	2.95
24	1.15	16.77	22.45	22.36	1.03	6.68	2.32	1.74	18.89	6.60
PP										
3	0.00	0.00	0.00	0.00	0.00	0.00	100.00	0.00	0.00	0.00
6	3.06	9.25	3.06	0.09	2.23	1.03	64.65	0.71	8.03	7.87
12	5.26	12.56	1.99	20.17	1.79	2.53	25.58	2.94	19.43	7.74
24	14.37	6.95	2.16	12.09	4.26	1.49	12.16	15.92	21.13	7.62
LC										
3	0.00	0.00	0.00	0.00	0.00	0.00	0.00	100.00	0.00	0.00
6	0.27	5.58	0.27	12.52	4.97	0.30	3.67	69.27	1.80	1.34
12	4.08	4.67	9.93	11.84	6.84	0.41	2.08	39.04	17.72	3.38
24	1.71	10.32	1.09	27.15	9.47	1.09	2.70	16.80	14.30	5.33
TEI										
3	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	100.00	0.00
6	0.00	6.50	3.81	0.02	5.84	6.74	0.03	9.10	67.93	0.03
12	1.06	12.50	10.75	0.83	8.53	8.08	0.33	6.87	50.24	0.80
24	0.91	19.20	10.59	18.83	6.99	5.94	0.30	2.00	33.00	2.24
IEP										
3	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	100.00
6	3.62	0.24	1.09	21.57	0.30	5.01	0.37	8.82	0.65	58.32
12	4.96	2.60	10.97	15.93	3.19	2.93	5.23	23.42	5.78	24.97
24	3.17	7.72	12.96	15.20	3.46	3.39	3.96	32.88	4.49	12.76

Table 3
Tests of cointegration for bi-variate equations

t = 62 (1970:2-1985:4)	Dependent variable					
	LIP		LFP		LM1	
Regressor	LFP	LM1	LIP	LM1	LFP	LIP
	0.91	1.08	1.07	1.17	0.84	0.92
R^2	0.99	0.98	0.99	0.99	0.99	0.98
Cointegration tests						
Unrestricted model						
ADF (lag = 4)	-2.70*	-2,91*	-2.70*	-3.11**	-3.11**	-2.84**
Restricted model						
(coeff = 1)						
ADF (lag = 4)	-3.03**	-1.73	-3.40**	-1.88	-2.04	-2.06

* Reject null hypothesis at 0.10 level.
** Reject null hypothesis at 0.05 level.
*** Reject null hypothesis at 0.01 level.

Table 4
Contemporaneous relationship coefficient estimates for three variable model

Dependent variables	LM1	LFP	LIP
LM1	1.00	0.00	0.00
LFP	0.06	1.00	0.00
	(0.08)		
LIP	0.05	0.24	1.00
	(0.08)	(0.04)	

Table 5
Decomposition of k-months ahead forecast-error variance
— three variable model (%)

k	M1	FP	IP
M1			
3	100.00	0.00	0.00
6	97.32	2.11	0.56
12	94.42	3.77	1.80
24	90.03	7.47	2.49
FP			
3	0.20	99.79	0.00
6	0.26	99.39	0.34
12	1.22	89.17	9.60
24	1.92	84.37	13.70
IP			
3	1.02	24.70	74.27
6	0.79	39.08	60.12
12	1.12	36.25	62.62
24	1.42	37.05	61.52

Table 6
Tests of cointegration with exchange rate

t = 62 (1970:2-1985:4)	Dependent variable				
	LIP			LFP	
Regressor	LFP	LM1	LER	LM1	LER
	0.37	0.23	0.42	0.53	0.64
R^2	0.99	0.98	0.99	0.99	0.99
Cointegration tests					
ADF (lag = 4)		-3.18**		-6.19***	

* Reject null hypothesis at 0.10 level.
** Reject null hypothesis at 0.05 level.
*** Reject null hypothesis at 0.01 level.

Table 7
Contemporaneous relationship coefficient estimates for four variable model

EP Variables D	LM1	LFP	LIP	LER
LM1	1.00	0.00	0.00	0.00
LFP	0.07 (0.09)	1.00	0.00	0.62 (0.15)
LIP	0.18 (0.03)	-0.19 (0.05)	1.00	0.06 (0.07)
LER	0.00	0.00	0.00	1.00

Table 8
Decomposition of k-months ahead forecast-error variance
— four variable model (%)

k	M1	FP	IP	ER
M1				
3	100.00	0.00	0.00	0.00
6	86.41	0.50	11.30	1.77
12	72.56	0.66	13.31	13.46
24	65.20	2.76	12.76	19.27
FP				
3	0.72	77.10	0.00	22.17
6	1.52	63.32	0.76	34.39
12	1.48	43.16	2.20	53.14
24	3.30	34.96	5.94	57.84
IP				
3	23.86	12.77	55.14	8.23
6	20.00	12.82	41.83	25.34
12	12.77	8.23	24.29	54.70
24	9.68	8.40	18.13	63.78
ER				
3	0.00	0.00	0.00	100.00
6	3.33	3.83	2.09	90.74
12	3.44	5.27	3.64	87.65
24	4.97	5.80	6.59	82.64

6. Conclusions

The results presented in this study confirm that the exchange rate is a potentially important omitted variable in many VAR and VEC studies of the relation between money and relative agricultural prices. The Brazilian case illustrates two potential costs resulting from this omission: first, three variable models excluding the exchange rate can result in spurious findings of overshooting agricultural prices in response to monetary shocks. Second, the exchange rate is a more important determinant of variations in food prices, in terms of size of response, than is monetary policy in open economies such as Brazil's. Even if prices *did* overshoot in response to monetary shock (a result this study does not confirm), this movement is swamped by variations resulting from other variables.

Indeed, if overshooting agricultural prices are of interest, this study shows that the exchange rate is a far more likely source of the phenomenon than is money supply. This is not a surprising finding in a small open economy, where the scope for independent monetary policy is far more limited than exchange rate policy. The forecast error decomposition was shown to be a very useful tool in gaining some insight into the relative importance of the variables included in the model.

Finally, it is apparent from the results presented here that in the Brazilian case (and possible others) it is at least, if not more, important to select the appropriate variables for analysis than it is to worry about methodological concerns such as whether to include error correction terms or to estimate the model in levels or in differences. This is a result of the structure of the economy at issue; no matter how the data are manipulated, the exchange rate remains a more important determinant of food prices than the money supply.

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Appendix

Contemporaneous relationship coefficient estimates
for the agricultural system — structural model

Dep. var.	FP	TFP	EXPR	M1	ER	Y	PP	LC	TEI	IEP
FP	1.00	-7.48	0.60	-7.29	6.06	-3.29	6.71	-5.75	-4.57	0
		(na)	(na)	(na)	(na)	(7.54)	(na)	(na)	(11.72)	
TFP	(na)	1.00	0	0	16.61	0	3.49	0	0	0
					(13.88)	(2.94)				
EXPR	0	0	1.00	0	4.04	0	1.26	0	0	0.33
					(1.70)		(0.68)			(0.36)
M1	0	0	0	1.00	0	0	0	0	0	0
ER	0	0	0	0	1.00	0	0	0	0	0
Y	0	0	0	0	0	1.00	0	0	0	0
PP	0	0	0	0	0	0	1.00	0	0	0
LC	0	0	0	0	0	0	0	1.00	0	0
TEI	0	0	0	0	0	0	0	0	1.00	0
IEP	0	0	0	0	0	0	0	0	0	1.00

Variable definitions

FP — Price index of agricultural food products composed as a weighted average of wholesale prices weighted by the value of production;
TFP — Total production of agricultural food products such as rice, beans, corn and potatoes;
EXPR — Total production of agricultural exportables such as soybeans, cotton, orange and sugar;
M1 — Money aggregate representing money supplied held by the public and as cash deposits in commercial banks;
ER — Exchange rate;
Y — National income;
PP — Oil price on the international market;
LC — Nominal value of market interest rates;
TEI — Total employment in the manufacturing industry — Brazil;
IEP — Price of exportable commodities in the international market.

Data Appendix

Year	FP	IP	MI	ER
1970:1	13.55	18.71	26,753.00	4.39
1970:2	13.70	19.60	29,469.00	4.52
1970:3	14.71	20.64	29,950.00	5.65
1970:4	15.60	21.31	33,638.00	4.81
1971:1	16.56	21.95	33,374.00	5.00
1971:2	17.69	23.20	37,799.00	5.19
1971:3	18.48	24.33	40,524.00	5.38
1971:4	19.28	25.05	44,514.00	5.58
1972:1	20.67	26.02	45,292.00	5.75
1972:2	21.25	26.96	49,073.00	5.89
1972:3	22.64	27.88	52,110.00	5.98
1972:4	23.51	28.84	61,550.00	6.12
1973:1	24.91	29.19	62,342.00	6.12
1973:2	25.83	30.92	73,672.00	6.08
1973:3	26.81	32.08	79,638.00	6.13
1973:4	27.34	33.75	90,490.00	6.17
1974:1	29.74	36.24	93,857.00	6.35
1974:2	34.76	40.49	100,885.00	6.59
1974:3	35.08	42.76	103,574.00	6.95
1974:4	35.97	44.99	120,788.00	7.26
1975:1	37.90	47.73	116,573.00	7.57
1975:2	38.98	50.75	133,144.00	7.91
1975:3	43.69	53.81	143,819.00	8.27
1975:4	47.66	57.54	172,433.00	7.77
1976:1	52.97	62.48	165,953.00	9.35
1976:2	61.44	67.89	192,791.00	10.41
1976:3	72.48	75.09	196,521.00	11.07
1976:4	80.65	80.74	236,506.00	11.86
1977:1	88.59	88.51	226,020.00	12.72
1977:2	104.30	97.23	260,524.00	13.70
1977:3	100.57	103.73	277,492.00	14.68
1977:4	106.60	110.50	325,243.00	15.48
1978:1	118.40	119.47	319,518.00	16.39
1978:2	138.43	129.47	360,415.00	17.41
1978:3	151.40	141.03	391,512.00	18.59
1978:4	161.43	155.57	462,655.00	19.90
1979:1	174.43	172.90	463,968.00	21.94
1979:2	202.53	192.87	538,467.00	24.55
1979:3	233.17	222.17	603,103.00	27.30
1979:4	279.70	263.27	803,113.00	33.99
1980:1	333.63	313.27	791,216.00	45.01
1980:2	401.30	374.87	987,815.00	49.83
1980:3	502.17	460.23	1,051,913.00	54.70
1980:4	653.47	555.37	1,367,017.00	61.33

(continue)

(continuation)

Year	FP	IP	MI	ER
1981:1	773.26	677.23	1,256,458.00	70.80
1981:2	924.73	809.57	1,547,094.00	83.89
1981:3	1,011.80	956.46	1,687,219.00	99.72
1981:4	1,155.60	1,122.42	2,558,479.00	118.08
1982:1	1,316.23	1,339.10	2,346,000.00	137.87
1982:2	1,601.33	1,614.37	2,859,000.00	160.18
1982:3	1,791.33	1,963.47	3,165,000.00	189.65
1982:4	2,055.13	2,286.83	4,222,000.00	230.36
1983:1	2,770.67	2,740.83	4,144,000.00	326.43
1983:2	3,976.93	3,483.30	5,200,000.00	475.89
1983:3	6,192.23	4,793.97	6,121,000.00	638.36
1983:4	9,470.73	6,523.97	8,232,000.00	867.50
1984:1	12,372.93	8,592.13	8,917,000.00	1,140.67
1984:2	17,243.57	11,176.27	11,429,000.00	1,515.42
1984:3	20,990.42	15,460.07	15,325,000.00	2,005.00
1984:4	29,665.24	21,583.37	24,853,000.00	3,732.00
1985:1	39,309.33	30,088.63	27,400,000.00	3,763.00
1985:2	51,891.46	38,121.37	38,669,000.00	5,232.00
1985:3	73,263.52	48,961.27	56,238,000.00	6,798.00
1985:4	106,254.50	68,869.63	102,413,000.00	9,009.00

FP = Food price; IP = Manufacturing industrial price index; MI = Money supply; ER = Exchange rate.