Inertial Inflation, Indexation and Price Stickiness:

Evidence from Brazil

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Abstract

This paper evaluates the inertial inflation hypothesis for Brazil during the 1970s and the first half of the 1980s. According to this hypothesis, (wage) indexation created a feedback mechanism such that one-time supply shocks were fully transmitted into permanent increases in inflation. First a simple theoretical model is used to show that the hypothesis is based on the assumption of perfect price flexibility. When price stickiness is introduced, indexation does not produce inertial inflation. Then, to investigate the impact of indexation on inflation, the degree of inertia (persistence) is compared between two periods, one with widespread indexation (1969-1985) and an earlier one without indexation (1945-1963). Unit root tests and the variance ratio test are used. The variance ratio test is also applied to inflation in the U.S. for the period (1969-1985) and France for (1983-1993), a period when there was no wage indexation. Finally, vector-autoregressive representations are estimated for the period 1972-1985. They differ from earlier work in that price stickiness is allowed for. The empirical results do not support the inertial inflation hypothesis; inertia does not seem to have been unusually high during the period of indexation, and impulse response analysis indicates that inflation shocks had only short-run effects on the level of inflation.

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1. Introduction

Several economists have argued that inflation in Brazil during the 1970s and the first half of the 1980s was mainly determined by its own history. According to this hypothesis, extensive use of indexation had created a feedback mechanism so strong that current supply shocks, such as the oil-price hikes, automatically carried over into permanent increases in the level of inflation. Thus, inflation was inertial. Moreover, monetary and fiscal policies were claimed to be ineffective because the inertia rendered inflation unresponsive to demand. Important contributions to the inertial inflation hypothesis have been made by, among others, Modiano (1983 and 1985), Arida and Lara-Resende (1985), and Lopes (1986).[1]

Novaes (1993) reformulated Taylor's (1979) model with staggered contracts to fit the Brazilian experience. By assuming complete backward behavior in the wage rule, the model generates an inflation rate with a random-walk component, implying that shocks have permanent effects. However, in contrast to the predictions of the inertial inflation approach, the inflationary process also contains a moving average term that has a dampening effect on shocks. This reduces the degree of inertia because only a part of a shock is transmitted into a permanent change in the level of inflation.

Novaes also evaluated the degree of inflationary inertia in Brazil for the period 1970-85. In accordance with several other studies, she found that the rate of inflation had a unit root, that is, a random-walk component. However, by estimating ARIMA and vector-autoregressive models, she showed that Brazilian inflation could be decomposed into a permanent and a temporary part, and that a 10% shock raised inflation by about 3.5% in the long run. She concluded that an inflation shock had a permanent effect on inflation, but that inertia was smaller than implied by the inertial inflation hypothesis.

This paper further evaluates the relevance of the inertial inflation hypothesis for Brazil. First, I argue that it is (wage) indexation, combined with the assumption of perfect price flexibility, which creates inertia in theoretical models. When price stickiness is assumed, there is no inertial inflation and supply shocks only raise inflation temporarily. To support this argument a simple model is simulated with both sticky and flexible prices.

Second, to find empirical evidence on the role of indexation in the inflation process, a comparison is first made between the degree of inertia during the period 1945-63, which is before indexation became common, and 1969-85, when indexation was widespread.[2] According to the inertial inflation hypothesis there should be much more inertia in the latter period. After this I follow Novaes (1993) and compare Brazil with low-inflation countries without indexation, the U.S. and France. To carry out the analysis, unit root tests and the variance-ratio test proposed by Cochrane (1988) are used.

A disadvantage with univariate analysis is that potentially important information that is readily available is ignored. For instance, if the reduced form for inflation is not a univariate autoregressive process, as it is in the typical inertial inflation model, then univariate analysis can be misleading. Thus, finally I estimate vector-autoregressive representations

(VARs) and calculate impulse responses. To allow for stickiness, variables in both log-levels and in rates of change are included.

The results of the empirical analysis do not corroborate the inertial inflation hypothesis. Inertia was only marginally higher during the period of indexation than during the one without indexation, about the same as in the U.S., and lower than in France. Moreover, the VAR analysis showed that the effect of a shock to inflation only lasts a few months.

The paper is organized as follows: In Section 2 a theoretical model is used to illustrate how the degree of inflation inertia is affected by wage indexation when assuming price flexibility or price stickiness. In Section 3 the univariate analysis is implemented, and the VAR models are estimated in Section 4. Section 5 concludes the paper.

2. Inertial inflation, indexation and price stickiness

In this section I first present a simple model to illustrate how indexation creates inertial inflation when prices are assumed to be perfectly flexible. Then I introduce price stickiness to show that this removes the inertial response of inflation to supply shocks.

A number of different models have been used to demonstrate how indexation creates inertial inflation in the Brazilian context. Nevertheless, the central features can be

summarized by the following model, which is a slightly altered version of the one analyzed by Novaes: [3]

$$\Delta w_t = \Delta p_{t-1} + b_1 \Delta y_t \qquad b_1 < 1 \tag{1}$$

$$\Delta p_t = \Delta w_t + v_{1t} \tag{2}$$

$$\Delta y_t = \Delta m_t - \Delta p_t + v_{2t} \tag{3}$$

$$\Delta m_t = \Delta p_{t-1}, \tag{4}$$

where all variables are in logarithms and Δ is the backwards difference operator. Eq. (1) shows that wage inflation, Δw_t, is determined by lagged inflation, Δp_{t-1}, and excess demand in the goods market, Δy_t. Lagged inflation enters with a coefficient of unity because compulsory wage settlements are assumed to have created 100% backward-looking indexation.[4] Eq. (2) shows that inflation is determined by a constant markup over wage growth plus supply shocks, represented by v_{1t}. Since Eq. (2) implies that the price level adjusts instantaneously to changes in the wage level, prices are assumed to be flexible in this model. Excess aggregate demand is described in Eq. (3) as a function of the rate of change of the real money stock, $\Delta(m\text{-}p)$ and random demand shocks, v_{2t}. Finally, in Eq. (4) the monetary policy rule is shown; by setting Δm_t equal to lagged inflation it is assumed that past inflation is fully accommodated.

This model aims at describing the dynamics of inflation in an economy that has adapted to high, and persistent, levels of inflation. Two important properties are of interest here. First, current shocks raise the level of inflation permanently, due to the fact that wage inflation and money-supply growth are functions of past inflation; there is no nominal anchor that ties inflation down. This is easily seen in the reduced form for the inflation rate,

$$\Delta p_t = \Delta p_{t-1} + v_{3t} \tag{5}$$

where $v_{3t} = (v_{1t} + b_1 v_{2t})/(1 + b_1)$. Thus, in the inflation process, today's inflation is determined by last period's inflation and an error term. Assuming that v_{3t} is white noise, the degree of inertia is unity, since all of past inflation is transmitted to current inflation. This implies that inflation follows a random walk.

Second, demand restraint, such as the abolishment of the monetary rule and the adoption of tight monetary policy, has a relatively strong transitory impact on output, as evident from Eq. (3), but weak short- and medium-term effects on inflation because of the feedback mechanisms. This is apparent in the reduced form for inflation obtained with exogenous money supply,

$$\Delta p_t = \frac{1}{(1 + b_1)} \Delta p_{t-1} + \frac{b_1}{(1 + b_1)} \Delta m_t + v_{3t} \tag{6}$$

where $1/(1+b_{1\,t}) > b_1/(1 + b_1)$ since b_1 must be less than one for the dynamics of the model to make sense. Various studies on the Phillips curve in Brazil have provided support for this view by showing that the impact of excess demand on inflation, measured by b_1 in this model, was either zero, or too low to be empirically relevant (see Lara-Resende and Lopes

1981, Lopes 1982, Modiano 1983; 1985, and Marshall and Morande 1989).

The assumption of perfect price flexibility made in this model is questionable, however. For instance, Bresser Pereira and Nakano (1987, pp. 27-60) argued that prices in Brazil are sluggish because the Brazilian economy is dominated by large corporations, and the government sets several prices. Moreover, there exists an extensive literature on price stickiness in general, and a number of reasons can be given to support the idea, such as menu costs, time delays in the spread in information, adaptive expectations, etc, (see Blanchard and Fisher 1989, ch. 8, and Ball et al. 1988).

One way to include price stickiness in the model is to replace Eq. (2) by

$$p = c + w \tag{7}$$

$$\Delta p_t = b_2(w - p + c)_{t-1} + v_{4t}, \tag{8}$$

where Eq. (7) shows that in equilibrium the price level is equal to a constant markup, c, plus the wage level, i.e., the unit cost of labor. Eq. (8) depicts how deviations from equilibrium are corrected each period at a speed given by b_2, and that there are supply shocks affecting this process.

A straightforward way to see how price stickiness affects the dynamics of the model is to simulate the response of inflation to supply shocks. To do this I set $b_1 = b_2 = 0.4$ and imposed a supply shock with the value of unity during the third period.[5] In Fig. 1 the impact of a supply shock in the model with flexible prices is shown for the cases with and without indexation. When there is indexation, the shock raises inflation permanently as predicted by the inertial inflation hypothesis, while the effect of the shock only lasts for a couple of periods when there is no indexation. Fig. 2 depicts the outcome for the model with price stickiness. Here supply shocks only have short-run effects on inflation in both cases, and there is no indication that indexation increases the persistence of the shocks. The reason for this result is that supply-shock induced growth in current inflation leads to a decline in real wages, and this reduces inflation the following period.

Nevertheless, the introduction of indexation alters the response of inflation to demand shocks in the model with price stickiness. This is shown in Fig. 3, where the responses to a one-time increase in demand are depicted for our two cases. When there is indexation, a large part of the initial increase in inflation is continuously transmitted into future periods. This result is dependent on the assumption of instantaneous adjustment in the log-levels of the variables in Eqs. (3) and (4), which might not be all that realistic; introducing stickiness in aggregate demand would make demand shocks have temporary effects on inflation as well.

3. Univariate analysis

One implication of the inertial inflation hypothesis model described above is that the inflation rate must be a random walk process, or more generally, must have a unit root. Several authors have tested samples from the period 1964-85 and found that the inflationary process in Brazil at that time did have a unit root, which corroborated the inertial inflation model (see Cardoso 1983, Novaes 1993, and Rossi 1993). However, if the presence of a unit root in the model were due to indexation, one would expect inflation during periods without indexation to be stationary. The indexation of the Brazilian economy started in 1964, and by 1968 indexation had became a general feature of the economy (see Fishlow 1974, and Simonsen 1986). Before the mid 1960s there was no wage indexation, and the exchange rate was used to stabilize inflation; the only use of indexation was to calculate an excess-profits tax on corporate returns to capital (Baer 1989, p. 277, and Simonsen and Cysne 1989, p. 437). Thus, to begin I want to test whether inflation prior to 1964 had a unit root, and then compare the test results with those for the period of indexation.

To carry out the unit-root tests two different methods were applied to monthly inflation data for the periods 1945:1-1963:12 and 1969:1-1985:12.[6] First the augmented Dickey-Fuller (ADF) test was used: As reported in Table 1, the null hypothesis of a unit root was not rejected in either of the samples. Then the KPSS test proposed by Kwiatkowski et al. (1992) was applied. This test is based on the fact that a series can be decomposed into a deterministic trend, a random walk, and a stationary error. The null hypothesis is trend stationarity, which implies that the variance of the error of the random walk-component is zero. To allow for error autocorrelation, the test was carried out for values of the lag truncation parameter, l, from 0 to 36. Table 1 only reports the results of the test with l=36 since the test values decreases as l grows. The null of stationarity is rejected in both cases.

Hence, there appears to be a unit root in both series. Nevertheless, even though inflation has a unit root in the period without indexation, one would think that shocks to inflation would have had more persistent effects when indexation was in place. To investigate this issue I used the variance-ratio test suggested by Cochrane (1988).[7]

To see how this test works, let us assume that the variable of interest follows a first-difference stationary linear process. We can then write its moving-average representation as

$$\Delta\pi_t = \tau + A(L)\varepsilon_t = \tau + \sum_{j=0}^{\infty} \psi_j \varepsilon_{t-j} \quad \textbf{(9)}$$

where π_t is the variable to be analyzed, ε_t is a white noise sequence, L is the lag operator, and τ is a constant. The permanent effect of a change in ε_t on π_t is determined by $\Sigma^{\infty}_{j=0}\psi = A(1)$, which is a measure of the degree of persistence (inertia) in the series. When π_t contains a unit root $A(1)$ has a finite value, and when π_t is stationary $A(1)$ is equal to zero; in the special case when π_t is a pure random walk with drift then $\psi_0 = 1$, $\psi_j = 0$ for j = 1, 2.., and A(1) = 1.

The variance ratio test uses the fact that when a series is a random walk the variance of its k-differences grows linearly with k, i.e., $\text{var}(\pi_t - \pi_{t-k}) = k\sigma^2$, where $\text{var}(\pi_t - \pi_{t-1}) = \sigma^2$. Thus, a plot of $(1/k)\text{var}(\pi_t - \pi_{t-k})/\text{var}(\pi_t - \pi_{t-1})$ against k shows a value of unity when the series is a pure random walk. If the process is a combination of a permanent and a stationary series, the plot approaches a constant between one and zero as k grows, showing the relative importance of

the random walk component. Finally, if the inflation rate is stationary, the plot approaches zero.

The test is implemented by first estimating the sample autocorrelations, ρ_j, then the estimate of the ratio of the variance, V^k, is obtained as a weighted sum of the autocorrelations

$$\hat{V}^k = 1 + 2\sum_{j=1}^{k-1}\left(\frac{k-j}{k}\hat{\rho}_j\right). \tag{10}$$

To be able to relate this test of persistence to the results of the ARIMA models estimated by Novaes and to the impulse responses presented in the next section, I also report $A(1)$, the sum of the parameters of the moving average representation of the first difference of the series tested. $A(1)$ is related to the estimate of V^k in the following way:

$$\hat{A}(1)^k = \sqrt{\frac{\hat{V}^k}{(1-\hat{\rho}_1^2)}} \tag{11}$$

where $\hat{\rho}_1^2$ is a conservative estimate of the fraction of the variance that is predictable by using all the lagged values of $\Delta\pi_t$ (see Campbell and Mankiw 1987).

The estimates of the persistence for Brazil, as well for the U.S. and France, along with asymptotic standard errors are presented in Table 2. The maximum number of differences reported is 60, since there are only small changes in the estimates at greater differences. For the indexation period 1969-85, the results basically confirm Novaes' results: Temporary shocks account for a notable share of the variance, i.e., the variance of the random-walk component only accounts for about 8% of the month-to-month variance; and the permanent effect of a unit shock is 30%. For the pre-indexation period, 1945-63, the estimates of the random-walk component and the permanent effect of a unit shock are 4% and 21%, respectively. Hence, the differences in the values of the estimates for the two periods are negligible, in particular if you consider that the two measures of persistence for the 1969-85 period would have been 100% had inflation behaved as described by Eq. (5). Moreover, if the standard deviations of the estimates are taken into account, it is not possible to tell whether the estimates are higher for the period with indexation than for the one without.

In Novaes (1993) a comparison was made between Brazil and Germany to find support for the hypothesis of inertial inflation. The expectation was that the degree of persistence would be very low in Germany; this was found to be the case. However, Germany is the low-inflation country par excellence, and may not be the best one to use for a comparison. Hence, the variance-ratio test was applied to data from the U.S. and France. In France wages were indexed to inflation until the beginning of the 1980s (OECD 1984); the sample thus starts 1983:1, when the indexation had been dismantled, and runs through 1993:12. For the U.S., the sample is 1969:1 - 1985:12. The results of the test are also presented in Table 2. Somewhat surprisingly, at 40 differences and above, the point estimates of persistence are slightly higher for France than for the period with indexation in Brazil. For the U.S., the estimates are a bit higher than for Brazil from 30 to 45 differences, then they drop below. At $k = 60$, A(1) is 34% for France, 26% for the U.S., and 30% for Brazil.

The results of the univariate analyses thus indicate that the unit root observed in the

Brazilian inflation rate is likely to have been due to other causes than indexation, and they cast doubt on the importance of indexation for creating inflationary inertia; in fact, the implementation of indexation does not seem to have increased the degree of persistence much, if at all. Furthermore, the degree of persistence in France, without indexation, is higher than in Brazil, with indexation, and there is only a small difference between the estimates of the U.S. and Brazil.

4. VAR analysis

A considerable amount of potentially relevant information is ignored in univariate time series analysis. It is possible, for example, that the unit root observed is the result of marginalizing with respect to other variables that have unit roots, as could happen if the reduced form for inflation in Eq. (5) does not correctly describe the inflationary process. When that is the case, finding a high degree of persistence using univariate analysis says little about how shocks are actually transmitted in the economy. In this section I therefore extend the analysis to a multivariate setting.

Novaes estimated a VAR model and found that the permanent part of a shock to inflation was about 40%. Since she based the analysis on a theoretical model with flexible prices, the VAR was estimated in first differences. However, this specification ignored any information contained in the log-levels of the variables, which is important for the inflationary process when prices are sticky. In fact, several studies have found long-run relationships between the log of the price level in Brazil and other nominal macroeconomic variables (for example Rossi 1993, and Durevall 1998). The existence of this long-run relation implies that the variables have stochastic trends that move together over time, i.e., they are cointegrated (see Engle and Granger 1987, and Banerjee et al. 1993 on cointegration). When some variables are cointegrated, one can take a linear combination of them and form a stationary variable, which shows the deviations from the long-run relationship. The question is then whether these deviations lead to changes in the inflation rate, as implied by Eq. (7), or in the other variables, during the adjustment process back to the long-run path. If inflation is affected, a theoretical model assuming flexible prices is likely to be misspecified, and a VAR analysis based on it might give an inadequate measure of the degree of inflationary inertia.

I estimated three different VAR models using the same variables as Noveas, and a proxy for foreign prices.[8] In the first model the endogenous variables are inflation and the real variables, m-p, $e+p^*$-p, and w-p. In the second the endogenous variables are Δp^*, Δm, Δp, Δe, and Δw, and the log-levels of the real variables are included as equilibrium-correction terms lagged one period. The final model has only variables in rates of change. All models include seasonal dummies, and dummies for the devaluations in December 1979 and February/March 1983. The VARs contain three lags as in Novaes; increasing the number to, say, 12 does not alter the long-run effects of shocks. Since the wage series is only available from 1972:1, the estimation period is 1972:4 - 1985:12. The specification of the contemporaneous relations between the variables was obtained with the standard Cholesky decomposition, which produces a recursive model. Inflation was put first in the VAR with

real variables, and third in the other two models. Changing the order of the contemporaneous relations does not influence the results.[9]

Fig. 4 shows the reaction of inflation to a shock in the inflation equation in the three models, together with the standard errors multiplied by two. As evident, when attention is paid to the long-run information, as in the first two models, there is only a temporary effect of a shock on the rate of inflation; the increase in inflation is significantly different from zero only for about two months. These results can be compared to that for the VAR in first differences, where about 40% of the initial shock remains after 24 months. Hence, by adding log-levels to the VAR model we have obtained results that contradict the predictions of the hypothesis of inertial inflation.

6. Conclusions

The purpose of this paper has been to evaluate the relevance of the inertial inflation hypothesis for Brazil. In the first part, a simple theoretical model was used to show that the hypothesis is based on the assumption that prices are flexible. When this assumption was replaced by price stickiness, indexation did not create inertia and supply shocks only raised inflation temporarily. In the second part, the empirical relevance of the inertial inflation hypothesis was tested.

One implication of the inertial inflation hypothesis is that the rate of inflation has a unit root because of indexation and monetary accommodation. This feature was evaluated by comparing the degree of persistence in the inflation rate in the period 1969-85, when there was widespread indexation, with the period 1945-63, when there was no wage indexation. A unit root was found in the inflation rate during both periods, and the degree of persistence was quite low and roughly the same. Although the power of these tests might not be high, they indicate that indexation is not likely to have given rise to the unit root during the period 1969-1985. Furthermore, by comparing the degree of persistence in inflation for this period with the same one in the U.S., and with a period in France when there was no wage indexation, more evidence was found against the hypothesis of inertial inflation.

Finally, to analyze how shocks to inflation are transmitted in a multivariate framework, VAR models were estimated and the impulse responses calculated. In contrast to Novaes (1993), the information contained in the log-levels of the variables was used in the analysis. The inclusion of the log-levels of the variables was motivated by the possibility of price stickiness in the goods market. The analysis showed that shocks to inflation raise inflation for about two months, when the information in the log-levels of the variables enters, while about 40% of the shock is permanent in the VAR formulated in first differences only. Thus, by conditioning inflation on the log-levels of the variables we are able to remove the long-run impact of shocks to inflation present in the univariate analysis and in the VAR analysis in first differences only.

The empirical results are not favorable to the inertial inflation hypothesis. Instead they corroborate the hypothesis that prices are sticky, and consequently, that indexation did not create inertial inflation. An interesting question is thus why inflation in the 1970s and the first half of the 1980s exhibited a random-walk like behavior. This issue has not been addressed directly here, but a reasonable answer is that a combination of several large demand and supply shocks, and accommodating monetary policy, was the main source for this feature of the inflationary process.

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Table 1
Unit Root Tests

		Augmented Dickey-Fuller test		
Period	Specification	Ho:	Test-value	5% sig. level
1945:1 - 63:12	constant and trend	unit root = 1	-2.14	-3.41
1945:1 - 63:12	constant and trend	unit root = 1, trend = 0	2.92	6.25
1945:1 - 63:12	constant	unit root = 1	-0.61	-2.86
1945:1 - 63:12	constant	unit root = 1, const. = 0	0.66	4.59
1969:1 - 85:12	constant and trend	unit root = 1	-1.76	-3.41
1969:1 - 85:12	constant and trend	unit root = 1, trend = 0	2.48	6.25
1969:1 - 85:12	constant and trend	unit root = 1	0.54	-2.86
1969:1 - 85:12	constant and trend	unit root = 1, const. = 0	2.17	4.59
		KPSS test, *l* = 36		
Period	Specification	Ho:	Test-value	5% sig. level
1945:1 - 63:12	constant and trend	root<1, trend ≠ 0	-0.73	-0.15
1969:1 - 85:12	constant and trend	root<1, trend ≠ 0	-1.31	-0.15

Notes:
1. The ADF and KPSS tests were carried out with the procedures URAUTO.SCR and KPSS.SCR in RATS.
2. In the ADF test 12 lags were used to correct for serial correlation.
3. The KPSS test was calculated for values of the lag truncation parameter, *l*, from 0 to 36. Only the test with *l* = 36 is reported since the test values decline as *l* increases.

Table 2
Nonparametric Estimates of Persistence

	Brazil 1945:1-63:12		Brazil 1969:1-85:12		France 1983:1-93:12		USA 1969:1-1985:12	
k	V^k	A(1)	V^k	A(1)	V^k	A(1)	V^k	A(1)
1	0.67 (0.07)	0.87	0.70 (0.08)	0.88	0.74 (0.11)	0.89	0.57 (0.07)	0.84
2	0.52 (0.07)	0.77	0.43 (0.07)	0.68	0.41 (0.07)	0.67	0.42 (0.06)	0.72
4	0.36 (0.06)	0.63	0.33 (0.06)	0.60	0.25 (0.06)	0.52	0.28 (0.05)	0.59
10	0.16 (0.04)	0.42	0.18 (0.05)	0.45	0.14 (0.05)	0.39	0.12 (0.03)	0.39
20	0.09 (0.03)	0.31	0.15 (0.06)	0.40	0.08 (0.04)	0.30	0.11 (0.04)	0.38
30	0.06 (0.02)	0.27	0.09 (0.04)	0.31	0.09 (0.05)	0.31	0.10 (0.05)	0.35
40	0.06 (0.03)	0.25	0.08 (0.04)	0.29	0.09 (0.05)	0.30	0.09 (0.05)	0.33
45	0.05 (0.02)	0.23	0.07 (0.04)	0.27	0.09 (0.06)	0.30	0.08 (0.06)	0.30
60	0.04 (0.02)	0.21	0.08 (0.05)	0.30	0.10 (0.08)	0.34	0.05 (0.03)	0.26

Notes:
1. The standard errors are in parenthesis.
2. The downward bias of the estimate of V^k was corrected for by multiplying it by (T-k)/T, as suggested by Campbell and Mankiw (1987).

Notes

1. See also Lara-Resende and Lopes (1981); Bresser Pereira and Nakano (1987); Cardoso (1983; 1987); Cardoso and Dornbusch (1987); Lopes (1989); and Marshall and Morande (1989).

2. The analysis does not cover the period after 1985 because a series of stabilization programs, based on de-indexation and price freezes, were implemented, and they contributed to a marked increase in inflation instability (see Kiguel and Liviatan 1991). Inflation during the latter half of the 1980s and the beginning of the 1990s is better described as hyper, or mega, than chronic inflation (Cardoso 1991).

3. This model differs from the one analyzed by Novaes (1993) in three ways. First, inflation is a function of current wage inflation, not the average of current and last period's wage inflation. This formulation simplifies the dynamics somewhat without changing the conclusions, and is in line with other models of inertial inflation (see Marshall and Morande 1989). Second, since supply shocks, interacting with inertia, are believed to have generated the accelerations in inflation, a term representing supply shocks is included in Eq (3) (see Cardoso and Dornbusch 1987). Finally, the money supply equation does not include the rate of change of inflation. This simplification has no effect on our conclusions.

4. For details about the different indexation schemes used in Brazil, see for example Macedo (1986) and Simonsen (1986).

5. The results of the simulations are not dependent on the choice of values for b_1 and b_2 as long as they are less than one. When they are set to one or higher, the dynamics of the model does not make economic sense.

6. Inflation was measured as the first difference of the log of the general price index. The source is Estatísticas Historicas do Brasil, IBGE, Rio de Janeiro, (1988).

7. The variance ratio test is a nonparametric alternative to the ARIMA model used by Novaes. See Cochrane (1988) and Campbell and Mankiw (1987) for comparisons of the two methods to measure persistence.

8. The variables used are the following: Money is M1; the exchange rate is cruzados per SDR; and the world price level is an index of the consumer prices of the countries included in the SDR, with the same weights as in the calculation of the SDR. The source was the IFS database of the IMF. The wage level was collected by ABDIB (Associação Brasileira de Desenvolvimento da Industria de Base) and taken from various issues of Conjuntura.

9. Inflation was placed third to make the impulse response analysis consistent with one in Novaes (1993). However, the order of the variables does not matter because we are interested in the long-run impact of a shock. Moreover, the contemporaneous correlations between the variables are low - partly because monthly data are used - so the short-run dynamics for inflation are also insensitive to the order of the variables.

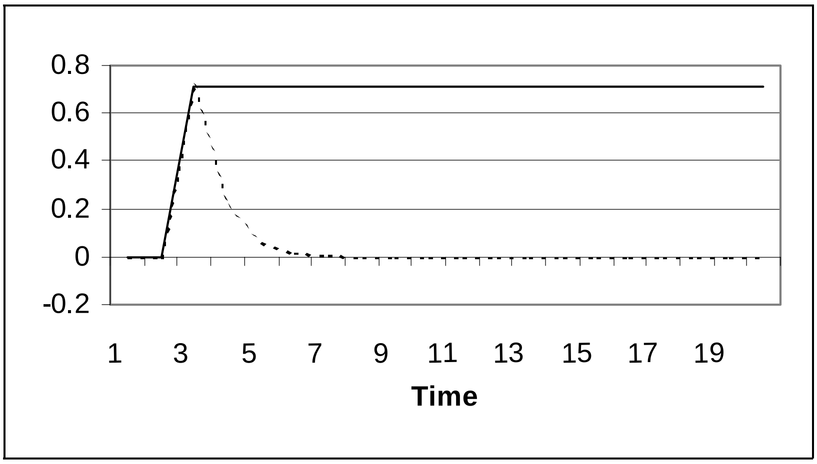

Fig.1. Response of inflation to a supply shock in period 3 in the flexible-price model with indexation (__), and without indexation **(....).**

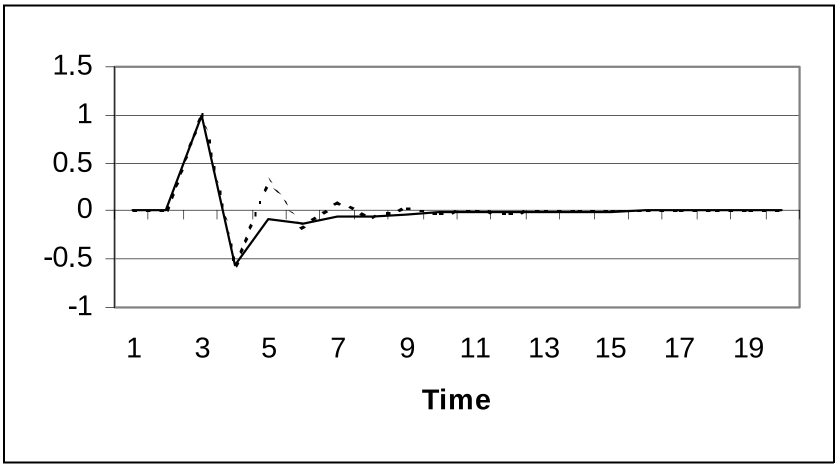

Fig. 2. Response of inflation to a supply shock in period 3 in the sticky-price model with indexation (__), and without indexation **(**......**).**

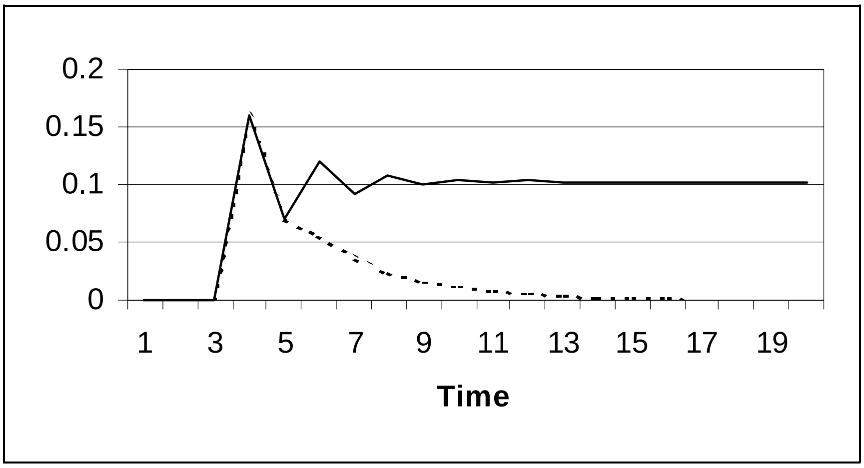

Fig. 3. Response of inflation to a demand shock in period 3 in the sticky-price model with indexation (__), and without indexation **(**......**).**

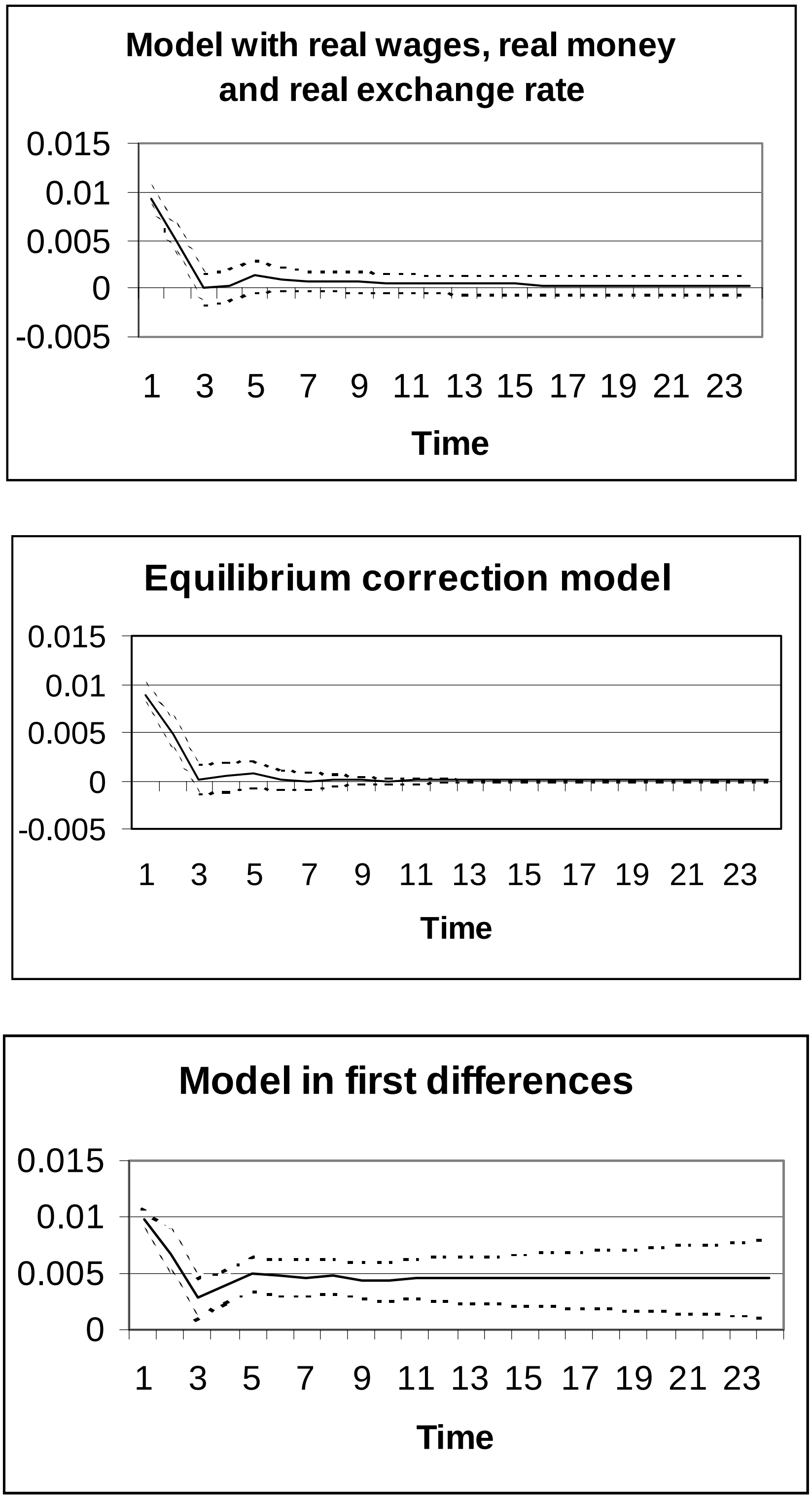

Fig 4. Impulse response functions (__) and confidence bands (- - - -) showing the response of inflation to an inflation shock over 24 months in three different VAR-models.