

Optimum pricing policy, government induced shocks and the dispersion of relative prices in Brazil

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This paper examines the optimum pricing policy under an inflationary environment and discusses the resulting dispersion of relative prices in Brazil during the last decade. The paper estimates through the variability of relative prices, first, the temporary effectiveness of major government induced macroeconomic shocks, such as the maxidevaluation of 1983, the three successive price freezes and the demand shock produced by the Collor Plan; second, the anticipation of these government actions by agents; third, the causality between inflation and dispersion of relative prices. The results indicate that the past income policies in Brazil have not performed adequately the role of coordinator of relative prices. Anticipation of the second and third price freezes is shown to have occurred. Only instantaneous causality between the mean and the dispersion of the inflationary process is observable in the sample considered. It is argued that as a result of the lack of a practical feasible device to cluster relative price decisions other than constraining the choice of agents, there should be no control of relative prices apart from the one that results from controlling the causes of the underlying inflation.

1. Introduction; 2. Some empirical remarks; 3. A model of optimum pricing policy under inflation; 4. Nash improving coordinating devices and relative prices; 5. Intervention analysis; 6. Causality; 7. Concluding remarks and policy implications.

1. Introduction

Recent income policies in Brazil have attempted to cluster relative price decisions based on the need to coordinate individual behavior. This position is quite generally accepted among government economists in Brazil and is best justified in Simonsen (1988) where income policies and their compatibility to rational expectations models are exposed to conclude: "The

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Government should play the role of the Walrasian auctioneer, speeding up the location of Nash equilibria” and “The central problem for the government is not to discover Nash strategies, but to coordinate the simultaneous playing”.¹ This coordination, to be effective, requires not only *an appropriate clustering of the timing of decisions regarding the setting of relative prices* but also *implicitly assumes that the dispersion of these prices is not caused, with or without statistical observable delay, by the change in the inflation rate*.² This paper performs a theoretical discussion and an empirical assessment of these two requirements.

The literature on optimal pricing policies under uncertainty provides a model where relative prices’ dispersion increases with the average rate of inflation, even in the absence of variability in inflation,³ Sheshinski and Weiss (1983), Benabou (1986). The certainty equivalence result established that firms under stochastic inflation choose an optimal pricing policy as if they faced a certain and fixed rate of inflation higher than the actual expected rate, the difference being a risk premium. This is equivalent to say that an increase in the variance of expected inflation or in other words, an increase in the uncertainty about future inflation leads firms to choose a pricing policy with larger amplitude in real price. This result, the dependence of individual price policies on the magnitude of inflation is tightly tied to the idea of a correlation between the relative price changes variability and the inflation rate. The second part of this essay analyzes this latter relationship for Brazil for the decade of the 1980’s.⁴ The estimation of causality is performed on section 6.

The empirical work in the paper differs from the literature that explains the unconditional variability of the overall inflation rate rather than the dispersion of relative prices. Most of the models following Lucas (1973) predict a link between the inflation rate and the variability of the overall inflation rate as opposed to the variability of relative prices. The basic assumption behind this theoretical relationship is the one of asymmetric information, taking into account the fact that decisions of a specific agent do depend on the decisions of all the other agents. Otherwise we should not encounter any dependence of relative price changes on absolute price changes. This theoretical relation is the

¹ Simonsen (1988), p. 501-32.

² Otherwise, the dispersion of relative prices could not be considered a *control* variable. The distinction between acts, states and consequences are causal distinctions and a policymaker should not ignore them.

³ The overall variability is defined for the average inflation rate overtime. The dispersion is defined for relative prices across industries at each point in time. For a description of the formulas used, please refer to the appendix.

⁴ The previous case study, Silva (1982) for the 1970’s in Brazil, focused on the effect of concentration ratios of some industries on the above relationship.

result of an informational problem. It is not an implication of a model of decentralized pricing with adjustment costs as considered in this paper. There has been some empirical evidence for the significant relationship between the mean and variance of the inflationary process. One of the earliest articles on the subject was by Vining (1976). By the early 1980's a couple of papers were published by Cukierman (1983, 1982), Taylor (1981), Fischer (1981) and Hercowitz (1981).

The question of coordination of relative prices setting and the resulting appropriate clustering of the timing of decisions to speed up the location of Nash equilibria is also addressed in the paper. It is best illustrated on theoretical grounds by the simple example provided by Aumann's Nash improving correlated equilibria exposed in section 4. The coordination adds information, signalling to each player what probable moves the other player(s) will do. As a result, the solution for the game may be one in which players are better off compared to the Nash solutions including strategies where players randomize. The section discusses the nature of the coordination implied by such a correlated equilibria.

The past experience with income policies in Brazil have been advocated on the basis of being such a coordinating device. *The paper argues that the theoretical inappropriateness of price freezes is based on the fact that it does not add information to a multi agent game but simply constrains individual decision making.* The empirical results corroborate the ineffectiveness of such coordinating devices even on a temporary basis. *Based on the dispersion of relative price changes we fail to reject the hypothesis that price clustering occurs as a result of temporary price freezes.* This is robust evidence against the essence of recent income policies in Brazil. The lack of impact on the dispersion of relative prices of such policies is an empirical argument against the somewhat settled idea that income policies are temporarily successful. *Evidence is also provided on the occurrence of anticipation of the price shocks after agents learned in the sample that such price shocks could take place with certainty.* This points to the stabilizing effect that a non reneging commitment of government not to implement new alternative versions of past price freezes may have.

The paper is organized as follows: some initial empirical remarks on the behavior of individual industries' relative prices in Brazil over the last decade are presented in section 2. Section 3 exposes a model of optimum pricing with adjustment costs and its implication to the dispersion of relative prices. Section 4 discusses the Nash improving correlated equilibria and its implications for coordination of relative prices. Section 5 performs an empirical assessment of the temporary effectiveness of the government induced shocks to cluster relative prices. Evidence on the anticipation of government actions after repeated use of controls is provided through the observation of the impact on

the dispersion of relative prices. Section 6 tests for the causality — Granger sense — of inflation on the relative price changes variability. Section 7 presents policy implications and the concluding remarks.

2. Some empirical remarks

This section provides some background information on the evolution of relative prices in Brazil during the last decade. Brazil is a propitious country to review the relationship between the variability of relative prices and the inflationary process, since changes in the variance of relative prices in a short span of time cannot be attributed to the differential gains in productivity, since the magnitude of changes is very high⁵ (see definitions in the appendix).

a) Both absolute variability and relative variability (normalized by the inflation rate) are greater in the 1990's compared with the 1980's, with a sharp increase by the end of the 1980 decade (charts 1 to 1b).

b) The relative price variability is affected by price shocks. It is sharply increased under maxidevaluations and is temporarily reduced during negative demand shocks. Indeed it is observed a rather peculiar characteristic for the variance of prices when we focus on its rate of change. Notice the spikes corresponding to the adjustment of administered prices immediately before all the observed shocks (charts 2 and 2a).

c) There is a significant positive correlation value between the variance of relative price changes and the inflation rate. The contemporaneous correlation found for the rate of change of the variables is 0.51 for the period January 1980 to February 1986. A more detailed analysis of the whitened series is provided in section 6. Chart 3 shows the comovements of the variance growth and the acceleration observed in annual rates of change of the wholesale price index.

d) The evolution of real relative prices shows that some industries repeatedly benefited from the "general" price freezes in Brazil: vegetables, fibers, oilseeds, roots and tubers, poultry and dairy products and meat industry, while others suffered the most — lime and silicates, industrial equipment, non-ferrous metal, house appliances, paper, rubber, plastics, fabrics, cereals and bakery products, oils. This pattern, where relatively competitive sector managed to keep prices well above more concentrated sectors, was reversed as a result of the demand shock imposed by the Collor Plan in March 1990.

⁵ Indeed, we make extensive use of the variance of relative prices *changes* instead of the variance of relative prices where the build up of differential gains of productivity would be cumulative and would require appropriate analysis, Silva (1982).

Chart 1
Variance of relative prices
Weighted and proportional to average wholesale price changes

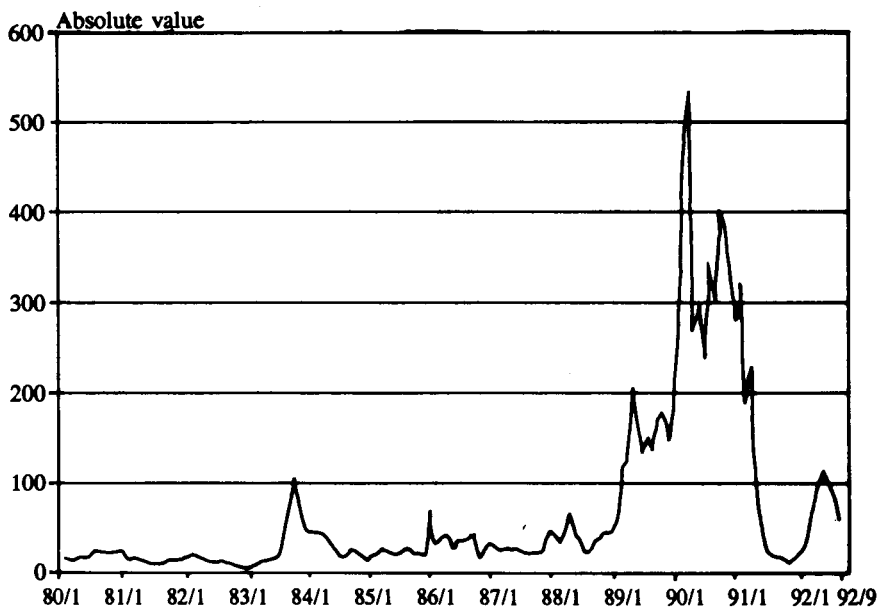


Chart 1a

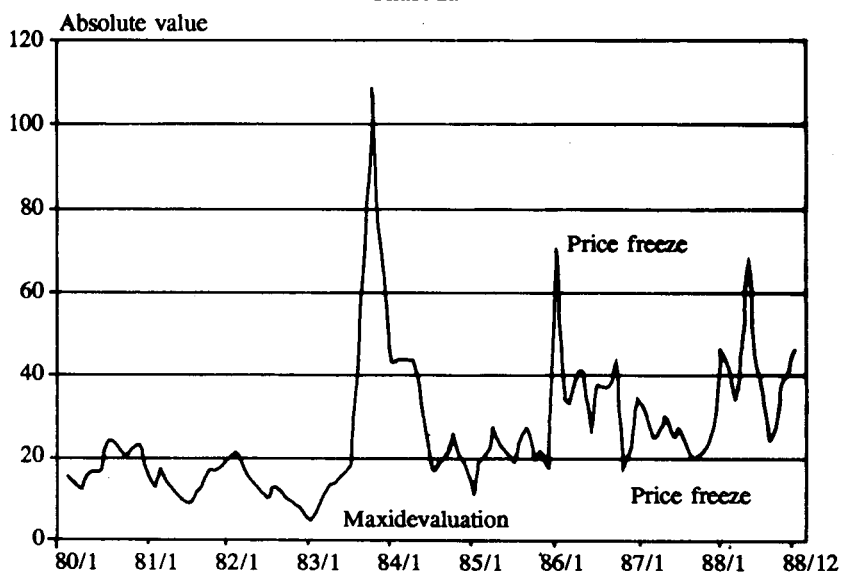


Chart 1b

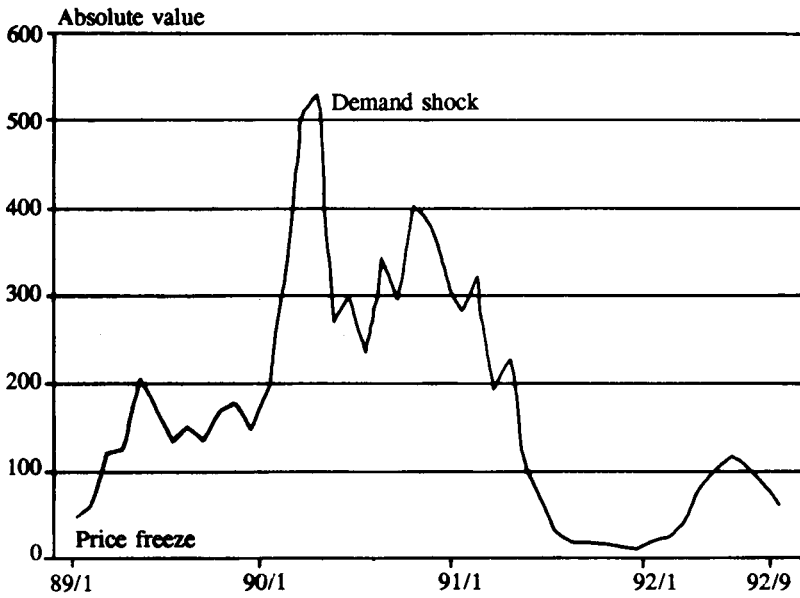


Chart 2
Variance of relative prices – rate of change (weighted)

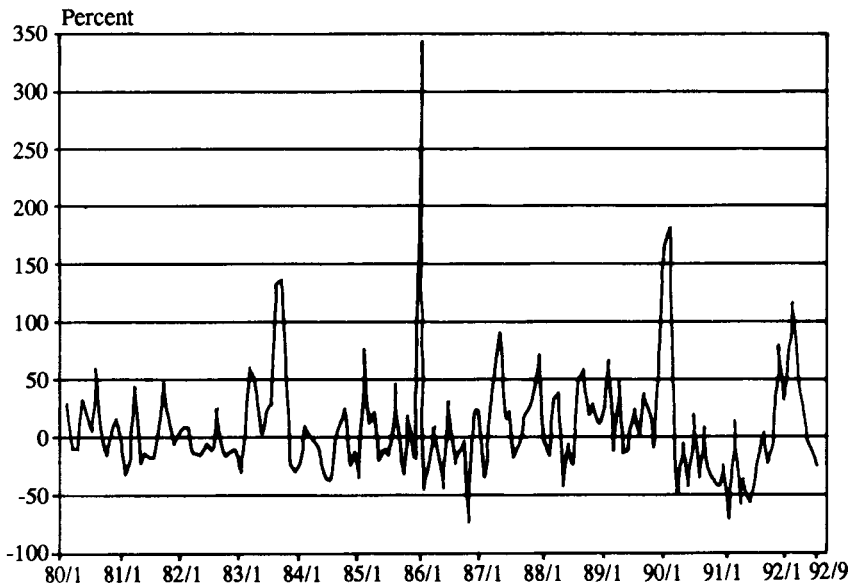


Chart 2a
Variance of relative prices – weighted and proportional to average wholesale price changes

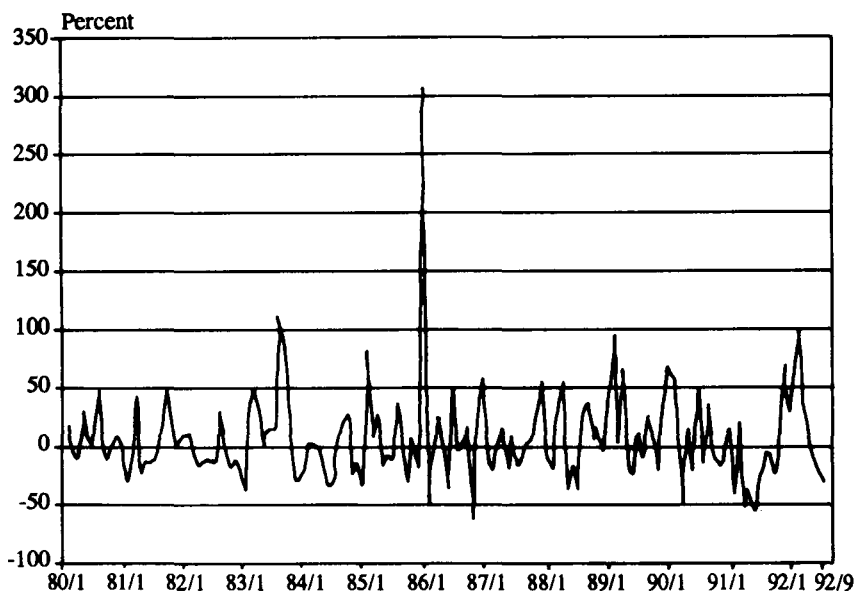
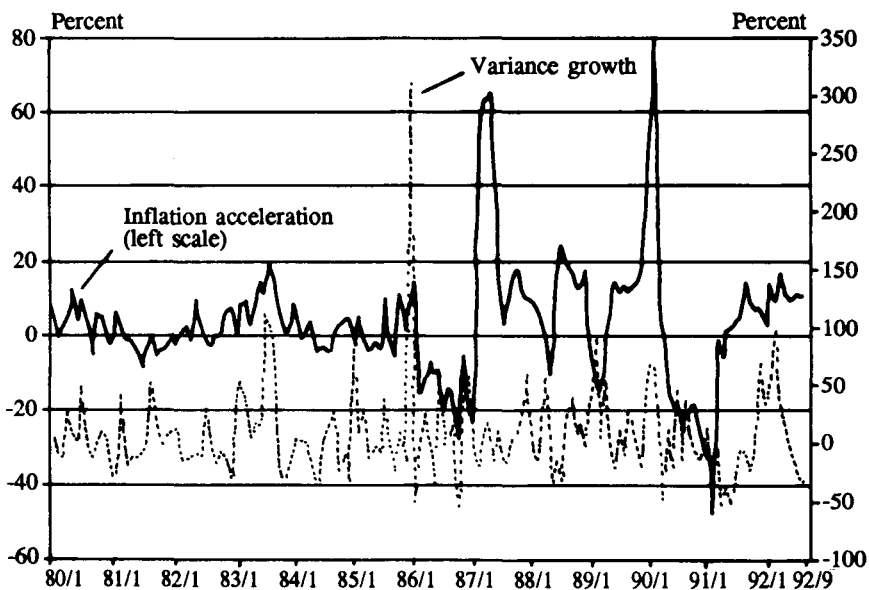


Chart 3
Variance growth and inflation acceleration



The behavior of prices in each industry during each of the three price freezes and the more recent demand shock is depicted in charts 4 to 8 in the appendix. The charts plot the cumulative real relative prices gains and losses during each of the price periods. Table 1 separates the gains at the time of the price freezes, during the Collor Plan and for the overall period Jan./80 - Sept./92. (See appendix I.)

3. A model of optimum pricing policy under inflation

This section presents a model of the behavior of a representative firm facing an exogenous inflationary process and adjustment costs. It serves to illustrate models of (s, S) pricing policies and their empirical implications for the relationship between changes in the rate of inflation and the dispersion of relative prices. The section is based on the model posed in the literature by Weiss & Sheshinski (1983). Benabou (1986) extends the model to storable goods. Blanchard (1989) examines the problem of coordination of relative prices. The main testable implication is however derived using the initial model. The basic firm is a monopolistic one that produces a non-storable product whose demand depends on its relative price to the price of rival commodities and for which price adjustments are not costless. The idea of modelling monopoly indeed seems quite satisfactory under highly inflationary environments, since the cost of searching by the consumer — proportional to the rate of money devaluation — may indeed give rise to quite an important monopoly power. It may be a matter for another paper but it must be the case that the searching will depend on the probability of a differential gain in prices, the present rate of acceleration of inflation and the nature of the product (remark that this is even more certain in the case of non-storable goods). The cost of price adjustments leads each firm to keep the nominal price constant for certain periods and to realize price adjustments discretely. The main implication for this paper is that with a constant rate of inflation, prices will increase on the average at the rate of inflation and if price adjustments are independent among firms, we should observe a variance of price changes across industries increasing with the inflation rate.

3.1. The model Sheshinski & Weiss (1983) with perfect anticipation

Let the optimization problem be defined as a maximization of discounted real profits V_0 , defined for a representative monopolist price setter as the discrete summation for all periods of real profits within each period $[t, t+1]$, where p_t is the nominal price prevailing for each period.

Then, we define:

$$V_0 = \sum_{\tau=0}^{\infty} \int_{t_{\tau}}^{t_{\tau+1}} F(P_{t_{\tau}} e^{-g\tau}) e^{-r\tau} dt - \beta e^{-r\tau+1}$$

where β is the cost of price adjustment, r is the real rate of interest, g is the rate of inflation and $F(\cdot)$ is a well-behaved function which defines real profits. The optimization problem then reduces to the choice of optimal sequences for t_{τ} and p_{τ} . The first order conditions are:

$$\frac{\partial V_0}{\partial t_{\tau}} = [-F(p_{\tau} e^{-g\tau}) + F(p_{\tau-1} e^{-g\tau}) + \beta r] e^{-r\tau} = 0$$

$$\forall \tau \geq 1$$

$$\frac{\partial V_0}{\partial p_{\tau}} = \int_{t_{\tau}}^{t_{\tau+1}} F'(p_{\tau} e^{-g\tau}) e^{-(r+g)t} dt = 0$$

The trivial case is $g = 0$, where the optimal choice is to keep the nominal price constant, and $\beta = 0$ where it is optimal to keep the real price constant continuously adjusting nominal prices. For $g \neq 0$ and $\beta > 0$, the first equation establishes that the differential in real profits from a nominal price change must equal the interest on the cost of price adjustment and the second equation establishes that the marginal profits should be equated to zero at each period. It can be shown that given independence of the optimal policy between periods, any solution to the initial period repeats itself, i.e., the optimal sequence for prices and t is fully determined after the first period of nominal price changes. The recursive form is:

$$p_{\tau} = p_{\tau-1} e^{ge}$$

So the optimal real price solution moves between two bounds, (S, s) where $S = se^{g\tau}$ is the upper bound for real price at the beginning of a given period, i.e., the real price prevailing immediately after a nominal price change, and s is the lower bound. It follows from the simple framework — basically full-anti-anticipation — that S and s are fixed bounds. They would also be fixed if one wants to introduce uncertainty by random demand shocks, resulting in random changes in the inflation rate, that are uncorrelated over time, and then, past shocks give no information regarding future changes in the inflation rate.⁶ It is important to observe that in this case the nominal adjustment in prices is given by the size of the (s, S) range. Since we are interested in the relationship

⁶ One can imagine that with some more structure, i.e., if shocks are somehow correlated, then the bounds itself could possibly change according to expectations regarding future demand movements.

between the inflation rate and the dispersion of relative prices, we have to make an assumption regarding the time structure of price decisions. Caplin (1985) has shown that deviations of relative prices from average are uncorrelated across firms, if the (s, S) rule is optimal. This strongly supports the argument that the timing of price decisions are dispersed among any given space of firms (industries).⁷ Therefore, to observe the above relationship, we can concentrate on the effect of the changes of the rate of inflation on the referred variance. The necessary conditions were differentiated with respect to g .

Let z be the real price in each period and there be a unique z^* s.t. $F'(z^*) = 0$, and so, if $z < z^*$, $F'(z) > 0$ and if $z > z^*$, $F'(z) < 0$. Then, differentiating yields:

$$\frac{ds}{dg} = \frac{rF'(S)}{B} A$$

$$\text{where } B = g^2 r F'(S) F'(s) (S^{r/g} - s^{r/g}) < 0$$

$$\text{and } A = \int_s^S F'(z) z^{r/g} \ln z \, dz < 0$$

therefore,

$$\frac{dS}{dg} > 0 \quad \text{and} \quad \frac{ds}{dg} < 0$$

so that an increasing inflation rate leads to larger price adjustments, i.e., larger amplitude of real prices changes. Sheshinski & Weiss (1983) show that even under a *stochastic* process of inflation a spread preserving increase in the expected rate of inflation leads to an increase in the size of the bounds in which real prices vary, i. e., larger price adjustments follow, if we assume that the variability of expected future prices is small.⁸ It is on this implication that the empirical work will focus.

4. Nash improving coordinating devices and relative prices

Before proceeding with the empirical estimation, it is worth to explore the question of relative prices setting as a problem of coordination using the simple example provided by Aumann's (1974) theoretical correlated equi-

⁷ There is no reason to believe that optimal decentralized price decisions can possibly be clustered.

⁸ Alternatively, empirical evidence could be looked for between the variance of overall inflation and the relative price changes variance, since it is also known that a mean-preserving increase in spread will under general circumstances lead to an increase in the amplitude of real price changes.

libria. Suppose the players of a game build a coordinating device or choose a coordinator so that they can correlate their actions and everybody in the game can be made better off compared with the possible Nash equilibrium. As an illustration, following Kreps (1990), suppose for player 1 and 2 there is a pay-off matrix such that:

		Player 2	
		t_1	t_2
Player 1	s_1	(4,4)	(3,5)
	s_2	(5,3)	(0,0)

Then, if there is a coordinator signaling⁹ to players 1 and 2 what moves to do, the solution for the game may be one in which both players are better off (s_1, t_1) compared to the Nash solutions, including the mixed strategy where players randomize and each get a pay-off of 3.75. The important point, in the example is that the choice regarding which move to make is still of each player, and the coordination just adds information, telling each player what probable move the other player will do.

The idea of income policies, namely, price controls/freezes, as coordinating devices is based on the assumption that it is feasible to cluster relative prices. On a purely theoretical ground, the existence of an improving correlated equilibria for a game with $n + 1$ agents does not imply that there can be in practical terms such a mechanism. If by income policies we mean price freezes/controls,¹⁰ then, the coordination device is not restricting each agents' play by expanding the agents' information set, but is merely constraining the individual decision making process.¹¹ In the next section we assess empirically the temporary effectiveness of recent income policies in Brazil.

⁹ The theoretical coordinating device, in this case, is a third part that rolls a die and tells each player what move to make according to the results of the rolling. If it comes 1 or 2, player 1 is told to play s_2 ; otherwise, s_1 . If it comes 1-4, player 2 is told to play t_1 ; otherwise, t_2 .

¹⁰ Still quoting Simonsen, "Income policies are an essential instrument to fight big inflation" and, "Income policies hardly can be imposed without price control", p. 526.

¹¹ Still Simonsen: "The central function of income policies is not to constrain individual decision making", p. 504. The constraint imposed by price freezes will invariably be binding if the monetary and fiscal policies are loose.

5. Intervention analysis^{12,13}

This section examines the impact of government induced shocks on the relative prices for the Brazilian economy. It is generally accepted that income policies are at least temporarily successful. In this section, we provide evidence that this may not be the case by looking at the behavior of the dispersion of relative prices during the implementation of the government induced shocks. The model developed is an Arima model on the variance of relative price changes (annual) and the idea is to assess the impact of the maxidevaluation that took place in Mar. 1983, the impact of price freezes of Feb. 1986, June 1987, Jan. 1989 and the demand shock of Mar. 1990. The variance itself is expressed in terms of rate of changes (monthly). The procedure allows inferences regarding the effect of government policies on the dispersion of prices for the Brazilian economy.¹⁴

5.1 Estimation of temporary effect of government induced shocks

Following Box & Tiao (1975), the transfer function-noise type of model for the relative price changes variance in terms of rate of change was defined:

$$\dot{S}_t^2 = \sum_{i=1}^k \left(\frac{\omega(L)}{\delta(L)} L^i \right) I_{it} + \epsilon_t$$

$$\text{where } \epsilon_t = \frac{\theta(L)}{\phi(L)} \eta_t$$

Where each I_{it} is defined as an extended pulse variable to model the policy on - policy off interventions, as defined below. The noise part of the model was identified as an AR(1). The entire period from Jan. 1980 to Sept. 1992 was used

¹² The great majority of the procedures outlined below are based on Granger & Newbold (1977) and Geweke (1984).

¹³ We refer the reader to the definition of the variables in the appendix.

¹⁴ The analysis of intervention, here, refers basically to the price freezes in effect after 1986. During the whole period of analysis a system of price controls was permanently in effect, for both agricultural and industrial products. The controls were somewhat relaxed in 1980 and tightened in 1983. However, as Baer (1989) acknowledges, the very nature of the system - firms had to justify price increases by cost schedules - was not one intended to force sectors to absorb price shocks, and therefore was not a price control in the widely used sense of the term.

for estimation. The sample autocorrelation for the rate of change in the variance does die out at high lags.¹⁵ This and the analysis of the plot of the noise part lead us to assume reasonable stationarity for the series. The analysis of the sample autocorrelation and partial autocorrelation strongly suggests an AR(1) process. Applying Hannan-Rissanen (1982) criterion to the series, the model selected was also first-order autoregressive, the order picked out by the AIC criterion.¹⁶

The effect of price shocks on the relative price changes variance is then analyzed with the following structure:

$$\dot{S}_t^2 = \sum_{i=1}^5 \omega_i I_{it} + (1 - \theta L)^{-1} \eta_t$$

The noise parameters follow the prior specification and the pulse inputs are specified such that the *instantaneous* effect of the price shocks/freezes can be estimated.¹⁷ I_i is the pulse function defined above. Define $i = 1, 2, 3, 4$ and 5 as follows:

To test the maxidevaluation that took place in 1983 (Feb., 21).

$I_1 = 1$ for Mar.-Sept. 1983, 0 otherwise

¹⁵ To test whether the null hypothesis that a particular sample autocorrelation is equal to zero we use the result by Bartlett (1946), modified by Davies & Newbold (1980).

¹⁶ Sample autocorrelations and partial autocorrelations for the relative price changes variance (expressed in terms of rate of change).

τ	1	2	3	4	5	6	7	8
$\hat{\rho}(\tau)$	0.37	0.15	-0.01	-0.01	0.12	0.04	-0.05	-0.11
$a_{\tau\tau}$	0.37	-0.02	-0.07	0.02	0.14	-0.05	-0.10	-0.05
τ	9	10	11	12	13	14	15	16
$\hat{\rho}(\tau)$	-0.19	0.08	-0.06	-0.06	-0.07	-0.13	-0.08	-0.10
$a_{\tau\tau}$	-0.12	0.02	-0.04	-0.04	-0.03	-0.08	-0.03	-0.08

Standard error of $\hat{\rho}(\tau) = 0.08$

Computable lags = 152

¹⁷ In this case,

$$\left(\frac{\omega(L)}{\delta(L)} L^b\right)_i = \omega_{0i}$$

The rate of change in the relative price changes variance in Jan. 1986 was treated as an outlier, given its magnitude and the fact that its value resulted from the management of prices by the government at that time.

To test the effect of price freezes, respectively:

$I_2 = 1$ for Mar.-Nov. 1986, 0 otherwise - 1986 shock (Feb., 28)

$I_3 = 1$ for Jun.-Sept. 1987, 0 otherwise - 1987 shock (Jun., 12)

$I_4 = 1$ for Jan.-July 1989, 0 otherwise - 1989 shock (Jan., 14)

To test the demand shock of 1990 (Mar., 15)

$I_5 = 1$ for Apr. 1990-Sep. 1991, 0 otherwise

Final Parameters: $\hat{\sigma}^2 = 1504.7$; AIC = 1564.8

	θ	I_1	I_2	I_3	I_4	I_5
Coeff.	0.33	60.59	-20.52	-8.65	10.25	-29.31
T-stat	(4.30)	(3.04)	(-1.18)	(-0.35)	(0.51)	(-2.20)

The results obtained with different specifications of the variables are very close to the ones reported here.¹⁸ The adequacy of the model was checked on the basis of the autocorrelations of the residual.¹⁹

¹⁸ Using the standard deviation as the dependent variable similar results were obtained:

Final parameters: $\hat{\sigma}^2 = 1504.7$;

AIC = 1564.8

	θ	I_1	I_2	I_3	I_4	I_5
Coeff.	0.25	25.33	-9.67	-1.51	5.44	-15.25
T-stat	(3.17)	(3.05)	(-1.36)	(-0.14)	(0.65)	(-2.82)

Using the rate of change of the log of the variance we obtain the following results:

Final parameters: $\hat{\sigma}^2 = 13.54$

AIC = 14.08

	θ	I_1	I_2	I_3	I_4	I_5
Coeff.	0.12	5.83	-2.42	0.24	1.09	-2.86
T-stat	(1.51)	(3.77)	(-1.86)	(0.12)	(0.70)	(-2.93)

¹⁹ We tested each individual autocorrelation for the residuals ϵ_t on the basis that they should constitute a white noise process. Since there are 152 observations $T^{-1/2} = 0.08$. The autocorrelations of the residuals below were not significant at any lag. We also use the Box-Ljung statistic to test the jointly autocorrelations of the residuals. For the variance series the portmanteau statistic was calculated from the residual autocorrelations for the lags $\tau = 8$, and 16, respectively 6.86 and 12.98. The computed statistics were compared with values of the $\chi^2_{(8)}$ and $\chi^2_{(16)}$ statistics. We do not reject the null hypothesis of no autocorrelation.

Sample autocorrelation of the residuals

τ	1	2	3	4	5	6	7	8
$\hat{\rho}(\tau)$	-0.04	0.04	-0.11	-0.07	0.13	0.03	0.01	0.00
τ	9	10	11	12	13	14	15	16
$\hat{\rho}(\tau)$	-0.15	0.06	0.02	0.04	0.03	-0.06	-0.05	0.00

Standard error of $\hat{\rho}(\tau) \approx 0.08$

The results lead us to conclude that while the maxidevaluation of 1983 greatly affected the structure of relative prices, the price freezes were not effective in controlling the dispersion of these prices and the demand shock induced by the Collor Plan had a more significant influence in that direction. These results should be interpreted carefully. From the simple observation of the behavior of the dispersion of relative prices one can see that the relative price changes variance increased immediately before all the price freezes, indicating that the controls were established at a moment of adjustment of relative prices. This definitely can help to explain the low values of I_3 and I_4 . This also points to the fact that the rate of inflation itself may have given signals to private agents of the possibility of the implementation of new price freezes. The occurrence of the first price freeze signaled to agents that such a shock could be part of the set of policy instruments for the government. A government would resort to it in the case agents perceived the lack of monetary and fiscal adjustment. That is to say, price freezes were a good "bet" whenever the inflation reached a given "high" level. It is interesting to test if there was some anticipation of a price freeze looking at the movements of the variance of relative prices.

5.2 Estimation of anticipation of price freezes

Using the same procedure as before, the following structure was estimated:

$$\dot{S}_t^2 = \beta I + \beta_a I_a \beta_b I_b + \beta_c I_c + (1 - \theta L)^{-1} \varepsilon_t$$

I_i is a pulse function, $i = a, b$ defined to test the anticipation of price shocks:

$I = 1$ for Sept. 1985-Jan. 1986, 0 otherwise

$I_a = 1$ for Dec. 1986-May 1987, 0 otherwise

$I_b = 1$ for Sept. 1988-Dec. 1988, 0 otherwise

$I_c = 1$ for Oct. 1989-Feb. 1990, 0 otherwise

If the intervention — price freeze — was anticipated it must have had an effect on the variance of relative prices, prior to the intervention effectively taking place.²⁰ The anticipation is plausible in this case given that a certain level of inflation is known among agents in the economy as signaling a high

²⁰ It is as if the delay parameter b in the transfer function-noise model could be positive for interventions.

probability of a price freeze. This is an indirect effect of the first price freeze by *introducing the possibility* of income policies as such in the economy. It is therefore specially interesting to compare the difference in the magnitudes of the coefficients that represent anticipation before the first price freeze, the price freezes that took place later in the 1980s and the demand shock induced by the Collor Plan.

The estimation procedure and the same diagnostic tests were performed.²¹

Final Parameters: $\hat{\sigma}^2 = 1521$; AIC = 1593.1

	θ_1	I	I_a	I_b	I_c
Coeff.	0.35	-8.89	44.62	40.55	46.70
T-stat	(4.66)	(-0.37)	(2.01)	(1.65)	(1.99)

The coefficients I_a and I_c are significant at the 5% and I_b at the 10% level respectively, indicating that some anticipation did indeed took place and only after the first price freeze. These results are robust to different specifications of the above model.²²

The main results of this section are compatible with optimal behavior from the part of agents that forecast expected change in policy and as a result of the anticipation of the policy effects, the very own policies become less effective. This is suggestive about the effects of learning by the part of private agents and the question of reputation by the part of the government. It is evidence of the process by which private agents perceive and anticipate the lack of fiscal

²¹ The adequacy of the model was checked as specified for model 1.

²² Alternative specifications of the model tested the anticipation of each one of the price shocks at a time. The following results were obtained.

Final parameters: $\hat{\sigma}^2 = 1660.5$ AIC = 1682.4			Final parameters: $\hat{\sigma}^2 = 1633$ AIC = 1654.8		
	θ_1	I		θ_1	I_b
Coeff.	0.38	-8.44	Coeff.	0.36	42.79
T-stat	(5.19)	(-0.34)	T-stat	(5.0)	(1.62)
Final parameters: $\hat{\sigma}^2 = 1616.8$ AIC = 1638.1			Final parameters: $\hat{\sigma}^2 = 1627.3$ AIC = 1648.8		
	θ_1	I_a		θ_1	I_c
Coeff.	0.38	46.97	Coeff.	0.34	44.14
T-stat	(5.17)	(2.01)	T-stat	(4.52)	(1.90)

adjustment and monetary control which result in the ineffectiveness of the income policies in Brazil.

6. Causality²³

This section examines the expected relation between the variability of relative prices and the mean of the inflationary process as it is implied by the model on optimum pricing policy of section 3.²⁴ The importance of the causality analysis resides on the fact that income policies attempt to cluster relative price decisions implicitly assume that the dispersion of prices is not caused with or without statistical observable delay by the rate of change in the inflation rate. Otherwise, the dispersion of relative prices could not be considered a *control* variable. If we can establish, in this context, that there is strong evidence for the causality of the rate of inflation on the relative prices change variance, then, we should not try to control inflation by controlling relative prices. The relevance of discussing causality in this context cannot be exaggerated. The policymaker must distinguish among those variables that he can immediately manipulate. As Skyrms (1988) puts it the distinctions among acts, states and consequences are causal distinctions and if a policymaker were to ignore them, he would be in danger of practicing "voodoo economics".

The empirical analysis models *the rate of change* of both the dispersion of relative prices and the inflation rate. There are two reasons for that. One of them is the requirement of stationarity and the other is the empirical fact that changes in the inflation and not the inflation *per se* are more interesting for analysis in a country with high and persistent history of inflation. The period was restricted to Dec/ 1979-Feb. 1986, given that there is not enough observations between shocks to carry out the analysis for in-between periods after 1986. We started the analysis by simply observing the cross-correlation between \dot{s}_t^2 and $\dot{\pi}_t$. Both series are stationary and were transformed to be zero mean. For $\tau = -1, 0, 1$ the

²³ The usual and more general definition of causality found in the literature, Zellner (1979) is predictability according to a law or set of laws. Let $F(X_{t+\tau}/\Omega)$ be the conditional distribution function of X given Ω . If $F(X_{t+\tau}/\Omega) = F(X_{t+\tau}/\Omega - Y_t)$ for all $\tau > 0$. Then, Y_t does not cause X_t . If $F(X_{t+\tau}/\Omega_{t+1} - X_{t+1}) = F(X_{t+1}/\Omega_{t+1} - X_{t+1} - Y_{t+1})$. Then, Y_t does not instantaneously cause X_t . In the statistical work basically we try to find evidence against the proposition that the rate of inflation has no predictive value for the dispersion of relative prices in the linear sense.

²⁴ A recent objection to the usefulness of the investigation of the empirical relation between the mean and the dispersion of the inflationary process was raised by Hartman (1991). The paper, however, is inconclusive since one of the main results points to the fact that there are, *a priori*, no general implications for the sign of coefficients in a simple linear regression relating the dispersion of prices to the mean and therefore the correlation does not follow from a definitional artifact.

cross correlation is statistically significant.²⁵ However, as illustrated by Box & Newbold (1971), the raw cross-correlogram may be misleading.²⁶ Therefore we considered prewhitening the input:

$$\text{let } \dot{s}_t^2 = \alpha_0 \dot{\pi}_t + \alpha_1 \dot{\pi}_{t-1} + \dots + \zeta_t$$

$$\text{and } \dot{\pi}_t = (1 - \theta L)^{-1} (1 + \phi L) \epsilon_t, \text{ then}$$

$$\epsilon_t = (1 + \phi L)^{-1} (1 - \theta L) \dot{\pi}_t$$

$$\text{define } v_t = (1 + \phi L)^{-1} (1 - \theta L) \dot{s}_t^2$$

$$\text{then, } v_t = \alpha_0 \epsilon_t + \alpha_1 \epsilon_{t-1} + \dots + (1 + \phi L)^{-1} (1 - \theta L) \zeta_t$$

$$\text{thus, } \text{cov}(v_t, \epsilon_{t-\tau}) = \alpha_\tau \text{var}(\epsilon_{t-\tau})$$

As a result, the coefficients α_τ are proportional to the cross-correlation between v_t and $\epsilon_{t-\tau}$, and the latter is useful in determining the direction of causality. Thus, to calculate the cross-correlation of the whitened series we estimate the following two processes:²⁷

²⁵ The sample cross-correlations were computed:

τ :	-12	-11	-10	-9	-8	-7	-6	-5	-4	-3	-2	-1
$\hat{\rho}(\tau)$:	-0.12	-0.12	-0.04	-0.09	0.06	-0.03	-0.01	0.09	0.03	0.12	0.17	0.28

τ :	0	1	2	3	4	5	6	7	8	9	10	11	12
$\hat{\rho}(\tau)$:	0.51	0.32	0.11	0.17	0.19	0.18	0.23	0.07	-0.13	-0.08	-0.02	-0.11	-0.27

Standard error of $\hat{\rho}(\tau) \approx 0.13$. The correlation at lags 6 and 12 with opposite sign is put down as being non economic meaningful and may be due to sampling error.

²⁶ Two independent AR(1) series may exhibit rather high values of R^2 , simply depending on the value of the coefficients for each process. As Granger & Newbold (1977) suggest, if the AR coefficients are the same and equal 0.9, $E[R^2] = 0.47$, since $\text{var}(R) = (1 + \theta \phi) / T(1 - \theta \phi)$. So we should be careful in analyzing cross-correlations of integrated processes even in the case that stationarity holds. Box & Newbold showed that in the case that the series are not even stationary, as in the case of random walk series, not only the cross-correlations are quite large but also the correlogram exhibits a smooth pattern not useful at all for statistical inferences. The smoothness results from the sample cross covariance process to follow approximately

$$(1-L)^2 \hat{\lambda}_\tau = \xi_\tau, \text{ where } \xi_\tau \text{ is some white noise process.}$$

²⁷ According to the same standard model building strategy worked out in the initial part of the paper. We whitened the rate of change of the inflation rate with an ARMA (1,1) and then applied the parameter estimates to estimate \hat{v}_t .

Final parameters:

	AR ₁	MA ₁
Coeff.	0.81	0.49
T-stat	(5.90)	(2.44)

$$\hat{\varepsilon}_t = (1 + \hat{\phi}L)^{-1} (1 - \hat{\theta}L) \dot{\pi}_t$$

$$\hat{v}_t = (1 + \hat{\phi}L)^{-1} (1 - \hat{\theta}L) \dot{s}_t^2$$

Below we compute the sample cross-correlation between \hat{v}_t and $\hat{\varepsilon}_{t-\tau}$.²⁸ For $\tau = 0$ the cross-correlation is statistically significant and large. For $\tau=6,8$ and 12 it is also quite large, although in the bounds of two standard deviations. The low values for the cross-correlations at negative lags strongly suggest the case of unidirectional causality. The large value without time delay may be a good indication of instantaneous causality, a qualification further discussed below.

6.1 On co-integration

As noted in Granger (1988), co-integration implies some causality in means between the variables. Before we perform any causality test we then check for co-integration between the inflation rate and the variability of relative prices using the variables prior to the transformations that will allow us to obtain stationary series. The first lag autocorrelation for both series are 0.935 and 0.953, respectively. Let $z_t = s_t^2 - \alpha \pi_t$ be $I(0)$. The estimation results in $\hat{\alpha} = 1.5908$ for the co-integration regression (variables log transformed). The augmented Dickey-fuller test based on the residual series z_t of the co-integration regression was performed for different lags specifications.

$$\Delta \hat{z}_t = -\rho \hat{z}_{t-1} + \sum_{j=1}^p \beta_j \Delta \hat{z}_{t-j} + e_t$$

We selected the following two on the basis of the AIC criteria.

Lag p	2	6
Coeff. $\hat{\rho}$	-0.23	-0.34
t-stat	(-3.04)	(-4.03)
Number of observations	= 74	

²⁸ The sample cross-correlation between \hat{v}_t and $\hat{\varepsilon}_{t-\tau}$

τ :	-12	-11	-10	-9	-8	-7	-6	-5	-4	-3	-2	-1
$\hat{\rho}(\tau)$:	-0.08	-0.08	0.01	-0.09	0.11	-0.05	-0.06	0.05	-0.08	0.00	-0.02	0.07

τ :	0	1	2	3	4	5	6	7	8	9	10	11	12
$\hat{\rho}(\tau)$:	0.40	0.20	-0.14	-0.02	0.06	0.05	0.19	0.05	-0.20	-0.05	-0.07	-0.03	-0.25

Standard error of $\hat{\rho}(\tau) \approx 0.13$

We reject the null hypothesis of no co-integration for $p = 6$. This is some indication that the relationship between the two variables depicted by the co-integration regression is stable. It is worth remarking that this does not establish a relationship between the mean and the variance of the same random variable, since the variance analyzed here, as defined in the appendix, is the unconditional variance of relative price changes, and not the unconditional variance of inflation.

6.2 Testing causality

The following hypothesis is tested:

$$H_0 = \dot{\pi}_t \text{ does not cause } \dot{s}_t^2$$

The tests that follow were computed and used to infer whether the causal inferences were at all sensitive to the choice of test statistic employed.

a) Granger Test

OLS estimations of the following constrained and unconstrained model were done:

$$\dot{s}_t^2 = \sum_{i=1}^n \gamma_i \dot{s}_{t-i}^2 + v_t \quad (t = 1, \dots, T)$$

$$\dot{s}_t^2 = \sum_{i=1}^n \alpha_i \dot{s}_{t-i}^2 + \sum_{i=1}^m \beta_i \dot{\pi}_{t-i} + \varepsilon_t \quad (t = 1, \dots, T)$$

n was set by AIC criteria among *all* possible lag specifications for $n = 6$ to 18: the AIC choice among a lower lag specification and among a higher lag specification was picked for analysis, respectively $n = 6$ and $n = 12$. The null hypothesis corresponds to $\beta_i = 0$ for $i = 1, \dots, m$. The total number of observations is 72. In all cases below, $m = n$. For $m = 6, m = 8$ and $m = 15$ the computed F^{29} statistics are 1.2392, 1.7483, 1.1396. Comparing these with values of the $F_{[6.54]}, F_{[8.48]}, F_{[15.27]} > 2$ statistics we do not reject the null hypothesis of no causality in any case. The models specified were indeed highly parameterized, but no one of the regressions considered were spurious in terms of the general F test.

²⁹ $F = [RSS_e - RSS]/m/[RSS/(t-m-n)]$

A variant of this test proposed by Geweke, Meese & Dent (1983) is to calculate the following large sample statistic that asymptotically:

$$T(\hat{\sigma}_c^2 - \hat{\sigma}^2)/\hat{\sigma}^2 \sim \chi^2_{[n]}$$

where $\hat{\sigma}_c$ and $\hat{\sigma}$ are respectively the variance of residuals for the constrained and unconstrained equations defined above. For $n = 6$, $n = 8$ and $n = 15$ the computed χ^2 statistics are 2.7643, 11.0915, 24.4670. Comparing these with values of the $\chi^2_{[6]}$, $\chi^2_{[8]}$ and $\chi^2_{[15]}$, we observe the same results as before and do not reject the null hypothesis of no causality.

b) Pierce & Haugh Test

We test causality alternatively through the residuals of the whitened rates of dispersion and inflation. We fit the following model using the Arima fitted before for the transfer function model.

$$(1 - \theta L)\dot{\pi}_t = (1 + \phi L)\epsilon_t$$

In this case, we do not apply the fitted model, but rather fit an extra model:

$$(1 - \Psi L)\dot{s}_t^2 = v_t$$

The null hypothesis in this case corresponds to

$$E[v_t, \epsilon_{t-\tau}] = 0, \forall \tau > 0$$

The test of causality is based on the sample cross correlations — given below — between the residual series for the fitted Arima models. The only lag that the cross-correlation is significantly different from zero is lag zero, suggesting the non-existence of causality, at least in terms of temporal causality.³⁰

³⁰ Sample cross-correlations of the residuals

τ	-12	-11	-10	-9	-8	-7	-6	-5	-4	-3	-2	-1
$\hat{\rho}(\tau)$:	-0.09	-0.08	0.03	-0.08	0.13	-0.02	-0.03	0.10	-0.02	0.08	0.08	0.12

τ	0	1	2	3	4	5	6	7	8	9	10	11	12
$\hat{\rho}(\tau)$:	0.41	0.14	-0.12	0.00	0.06	0.02	0.14	0.00	-0.23	-0.09	0.02	-0.05	-0.26

Standard error of $\hat{\rho}(\tau) = 0.13$

c) Instantaneous causality

If the variance of relative price changes is better forecast using the present rate of change in the inflation rate rather than using just the lags of the latter, then, this is quite a good indication of instantaneous causality. A possible test for instantaneous causality would be to observe the cross-correlations for the prewhitened series — each by its own filter. However, in this case, one cannot infer anything about the direction of causality. The following test of instantaneous causality should avoid the problem of determining direction of relationship. Comparing the above equation with the second equation of the

$$\dot{S}_t^2 = \sum_{i=1}^n \alpha_i^2 \dot{S}_{t-i}^2 + \sum_{i=0}^m \beta_i^2 \dot{\pi}_{t-i}^2 + \varepsilon_t^2 \quad (t = 1, \dots, T)$$

Granger Test, the null hypothesis corresponds to $\beta_0 = 0$. For $m = 6, 8, 15$ the computed F^{31} statistics are 7.6061, 5.4389, 6.3884. Comparing these with values of the $F_{[1.53]}, F_{[1.47]}, F_{[1.26]} < 5$, we *reject* the null hypothesis of no causality, respectively at the 1% level, 5% level and also 5% level.

The evidence gathered so far greatly appoints to instantaneous causality. The question regarding the direction of the tested causality is supported by the theory specified in section 3 so that *a priori* one could feel comfortable regarding the non-existence of a time gap. The more appealing explanation to support the finding of instantaneous causality from the rate of change in the inflation rate to the relative price changes variance is the fact that the finite time delay between cause and effect in the case analyzed is small compared to the time interval over which the data was collected [Granger (1988)]. The apparent causation in this case is due to temporal aggregation. This indeed could very well be the case since weekly observations could possibly lead to the existence of a time gap.

The possibility of true instantaneous causality — when no measurable time delay should be expected — is a possibility that does not render itself interesting in the sense that *there must be a time gap, between the decision of fixing relative prices and the observation of an accelerating inflation, no matter how small it is*. We have reasons to believe that during the more recent periods of fast acceleration of inflation this gap may indeed be even smaller and not measurable in the statistical sense and yet existent in the true economic system where the decision lags may be smaller than the ones implied by the lag structure of the data.

³¹ $F = (RSS_c - RSS) (T-m-n)/(1 RSS)$

It is also possible that a jointly causal variable have been excluded. In this case further research could search in some other direction beyond the one that this work explores, investigating the relationship with a jointly causal variable — possibly a control variable — for the two variables examined.³² It is suggestive that Taylor (1981) reports some evidence that the degree of monetary accommodation is correlated with aggregate inflation variability across countries. However, as Granger (1980) remarks, we should be careful with this argument, since the missing variable argument is always a possible one.

7. Concluding remarks and policy implications

Recent income policies in Brazil have attempted to cluster relative price decisions based on the need to coordinate individual behavior. The equilibrium under a decentralized price setting is shown to be improvable by means of an appropriate coordination mechanism on a theoretical level. This is the backbone of income policies in Brazil. Section 4 shows the inadequacy of interpreting income policies as such a coordinating device. They generally involve constraining the control problem of agents that would otherwise follow an optimal rule for setting their relative prices.

The empirical evidence corroborates the ineffectiveness of such coordinating devices even on a temporary basis. Based on the dispersion of relative price changes we fail to reject the hypothesis that price clustering occurs as a result of temporary price freezes. This is robust evidence against the essence of recent income policies in Brazil. The lack of impact on the dispersion of relative prices of such policies is apparent when compared to the statistically significant effect the maxidevaluation of 1983 had on this same dispersion. The first result of the paper is the temporary ineffectiveness of all the price freezes. It is an empirical argument against the somewhat settled idea that income policies are temporarily successful.

The second result of the paper is the evidence on the anticipation of the price shocks after agents learned in the sample that such price shocks could take place with certainty. In a model of relative prices setting under stochastic inflation this anticipation would most probably take the form of a higher risk premium with increased inflation and dispersion of relative prices. Some sort of assurance that government induced shocks will not take place may require changes of legal nature. Otherwise, it would take a long time to build a nonreneging commitment of the government not to implement new alternative versions of past income policies.

³² Granger (1988) discusses this point and the related issue of co-integration.

Finally, according to the empirical evidence, we fail to reject the change in the dispersion of relative price as being Granger caused by the change in the inflation rate. We find, however, evidence of instantaneous causality, some degree of co-integration and significant instantaneous cross-correlation between the variables. All these results point to the fact that³³ a jointly causal variable have been excluded. Further investigation should pursue the relationship with an appropriate jointly causal variable, a control variable other than relative prices. It is suggestive that Taylor (1981) reports some evidence that the degree of monetary accommodation is correlated with aggregate inflation variability across countries. This also points to the policy result that the use of an adequate control variable such as a *sustainable fiscal adjustment* cannot be replaced by attempts to control even temporarily relative prices.

Therefore, the past observed income policies in Brazil have either not performed or have performed inadequately the role of coordinator of relative prices, in the sense used by Blanchard (1983). The government policies tried to cluster relative price decisions when the relative prices deviations signal that these decisions cannot be clustered by their own nature. They are the natural outcome of a decentralized price setting economy. It is argued that as a result of the non-existence of a practical feasible device to cluster relative price decisions other than constraining the choice of agents, there should be no control of relative prices apart from the one that results from controlling the causes of the underlying inflation. This refutes the position that income policies may needed to coordinate individual behavior and can substitute an adequate and sustained fiscal adjustment.

³³ It could very well be the case that the finite time delay between cause and effect in the case analyzed is small compared to the time interval over which the data was collected. The apparent causation in this case is due to temporal aggregation. This indeed could very well be the case since weekly observations could possibly lead to the existence of a time gap.

Appendix I

Table 1
Evolution of real relative prices
(cumulative percentage change during selected periods)

<i>Industries</i>	<i>Price freezes¹</i>	<i>Demand shock²</i>	<i>Other periods</i>	<i>Jan. 80 - Sept. 92</i>
Vegetables	87.48	-32.33	83.94	133.36
Grains	9.90	-19.76	-7.98	-18.85
Fibers	34.70	-34.63	60.39	41.23
Oilseeds	119.83	-67.79	-50.42	-64.90
Roots	69.18	-40.40	123.23	125.09
Dairy	82.73	-40.19	-15.15	-7.27
Coffee	-5.30	-3.92	-49.25	-53.82
Other agric.	36.89	-48.87	1.17	-29.19
Minerals	-19.29	28.54	-44.60	-42.53
Silicates	-10.68	4.12	65.20	53.64
Iron/steel	-16.53	24.65	-10.39	-6.75
Non-ferrous	-15.51	10.62	-24.19	-29.15
Agric. equipment	6.32	38.95	25.20	84.97
Indus. equip.	-12.92	3.81	13.91	2.97
Other equip.	-4.62	15.01	8.67	19.21
Natural fabrics	-20.17	-7.22	-20.45	-41.08
Artificial fabrics	-22.94	21.40	-38.22	-42.21
Hosiery	28.74	-19.10	-57.94	-56.19
Apparel	104.31	1.29	-57.82	-12.70
Shoes	-5.12	2.41	-53.04	-54.37
Alcoholic bev.	-27.02	7.36	90.87	49.56
Non-alcoh. bev.	-12.41	19.70	NA	NA
Tobacco	0.36	10.63	41.94	57.60
Cereals	-20.42	26.87	97.03	98.91
Sugar	-39.84	30.35	1.79	-20.18
Oils	-16.99	15.98	1.86	-1.93
Cocoa	43.39	3.53	NA	NA
Foodstuffs	-62.27	3.65	NA	NA
Meat	48.28	-23.15	-11.18	1.21
Dairy	7.54	24.32	-4.60	27.55
Feedstuffs	12.28	6.32	-20.97	-5.65
Appliances	-46.47	11.20	-23.47	-54.44
Motors	-17.72	22.28	-8.95	-8.39
Electric equip.	-24.49	-8.07	43.15	-0.62
Motor vehicles	-35.75	8.29	91.56	33.28
Other vehicles	-10.02	29.14	2.24	18.80
Lumber	76.79	-7.16	-24.68	23.62
Wood furniture	8.91	10.69	13.06	36.29
Metal furniture	-10.64	37.16	-5.95	15.27
Other furniture	-4.00	-3.58	6.92	-1.03
Paper	-36.10	11.28	46.78	4.37
Rubber	-25.50	14.38	21.21	3.29
Leather	34.58	-22.34	-6.67	-2.46
Fuels	2.45	28.95	-25.42	-1.47
Paints	-17.37	20.95	95.77	95.65
Plastics	-29.89	6.32	44.99	8.08
Fertilizers	38.54	-14.00	-34.35	-21.78
Chemicals	-25.71	10.63	30.54	7.28

¹Mar.-Nov. 86; Jun.-Sept. 87. Jan.-Jul. 89.

²Apr.-Sept. 90.

Appendix II

Definition of variables and data set

The period of analysis covers 1979/12 to 1990/10. The following measures are used throughout the paper:

1. Variance of relative price changes

$$S_t^2 = \sum_{k=1}^n w_k (\pi_{kt}^a - \pi_t^a)^2$$

2. Variance of inflation

$$S_t^2(\pi) = \sum_{i=1}^{12} (\pi_{t+1-i}^m - \bar{\pi}_t^m)^2 / 12$$

3. Variance of relative prices³⁴

$$S_{p_k}^2 = \sum_{k=1}^n w_k ((p_k/p_t) - 1)^2$$

Notation: a = annual, m = monthly, w = weight, k is a specific industry and $n = 46$

π_t = Nominal rate of change of the wholesale price level

π_{kt} = Nominal rate of change of the relative price k

p_t = General wholesale price index - aggregate supply³⁵

p_{kt} = Relative wholesale price for industry k

The indexes and (weights) used for analysis are the 48 components of the wholesale price index³⁶

³⁴ The variance of relative price changes is a measure of the behavior of relative prices at every point in time. The variance of inflation is a measure of the cumulative behavior of relative prices since, specifically for this measure the base of the index was set 1977 = 100.

³⁵ The indexes are published every month by Fundação Getulio Vargas and they compose the IPA-OF (wholesale prices). The base period is Dec. 1989 = 100, except for the plots of real relative prices that the analysis of cumulative changes required the base period 1977 = 100.

³⁶ The weights may not add up due to rounding.

Farm products $n = 8$

- Vegetables and fruits (3.3)
- Grains and pulses (14.6)
- Plant fibers (0.6)
- Oilseeds (0.5)
- Roots and tubers (2.0)
- Livestock, poultry and dairy (5.9)
- Cocoa (2.4)
- Others (1.4)

Industry $n = 40$

- Crude minerals (5.6)
- Lime and silicates (2.4)
- Metallurgy
 - Iron and steel (7.3)
 - Non-ferrous metals (1.6)
- Machinery
 - Agricultural equip. (0.7)
 - Industrial equip. (1.3)
 - Others (0.7)
- — Electrical Equipment
 - Household appliances (1.3)
 - Motors and generators (0.6)
 - Others (1.0)
- Foodstuffs
 - Cereals and bakery (2.9)
 - Sugar (1.4)
 - Oils and fats (0.5)
 - Meat and poultry and fish (3.7)
 - Milk and dairy (0.6)
 - Others (1.4)
 - Salt and feedstuffs (0.7)
- Tobacco products (1.5)
- Lumber and wood products (2.8)
- Coffee (1.4)
- Vehicles equipment
 - Motor vehicles (2.9)
 - Others (0.2)
- Furniture
 - Wood furniture (1.5)
 - Steel furniture (0.4)
 - Others (0.3)
- Chemicals
 - Fuel and lubricating oils (6.7)
 - Paints and varnishes (0.5)
 - Plastics (1.5)
 - Fertilizers (1.8)
 - Others (3.6)
- Textiles
 - Natural fabrics (3.7)
 - Artificial fabrics (1.2)
 - Hosiery (0.4)
 - Apparel (0.5)
 - Shoes (1.2)
- Beverages
 - Alcoholic (0.7)
 - Non-alcoholic (0.6)
- Leather and skins (1.2)
- Paper and paperboard (1.6)
- Rubber products (0.8)

Resumo

O estudo examina políticas de preço ótimas em um ambiente inflacionário e o efeito sobre a dispersão de preços relativos no Brasil durante a última década. O estudo estima, através da variabilidade de preços relativos, primeiro, a efetividade temporária dos choques macroeconômicos induzidos pelo governo, dentre os quais, a maxidesvalorização de 1983, os três congelamentos de preços sucessivos, e o choque de demanda produzido pelo plano Collor; em seguida o estudo testa a antecipação dos choques pelos agentes; e finalmente, testes de causalidade entre inflação e a dis-

persão de preços relativos são realizados. A evidência empírica indica que as políticas de renda recentes não desempenharam adequadamente o papel de coordenação de preços relativos e que houve antecipação do segundo e do terceiro congelamentos. Apenas causalidade instantânea entre a média e a dispersão do processo inflacionário é observável na amostra considerada. O estudo argumenta que a falta de um mecanismo prático para ordenar decisões de preços relativos sem restringir a escolha dos agentes torna inadequado o controle de preços relativos que não seja *resultado* do combate às causas primárias do processo inflacionário.

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