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ON MEASURES OF DISPERSION OF RELATIVE PRICES UNDER INFLATION

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ON MEASURES OF DISPERSION OF RELATIVE PRICES UNDER INFLATION.

BY

José Alberro, José Córdoba and Eytan Sheshinski *

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* The question discussed in this note was raised during joint work at the Dirección General de Política Económica y Social, S.P.P. México.

1. Instroduction.

Does Mexico, with a 60 per cent rate of inflation, have a - larger variability of relative prices than, say, Argentina - or Israel with a 500 per cent rate of inflation (until recen tly)?. The answer to this question is important when trying to lay the groundwork for a stabilization program. Such - measures of variability provide an index for the deviations of relative prices from their 'normal' level, i.e. their le- vel in the absence of inflation. Clearly, any stabilization program which involves a freeze on prices must take these - distortions into account.

In the following we show that two commonly used statistical - measures of dispersion of relative prices have no relation to the amplitude of real price fluctuations and hence are quite inappropriate as a policy tool. In the last section we pre - sent empirical evidence on Argentina, Israel and Mexico which suggests that the distribution of relative price is more dis- torted in the latter country.

2. A Measure of Relative Price Variability

Let $i \varepsilon 0,1$ be an index of good i and P_i (t) the log of the price of good i at time t. The period for which price observations are made (say, a month) is denoted by h. The standard measure of relative-price variability [*i.e.* Blejer (1981), (1983) ; Blejer and Leiderman (1982); Cukierman (1979)[1/]], denoted by V, is

$$V(t) = \int_0^1 W_i [\pi_i(t) - \pi(t)]^2 \, di \tag{1}$$

Where:

$\pi_i(t) = P_i(t) - P_i(t-h)$, is the rate of change of the price of good i between $t-h$ and t; $W_i \geq 0$ constant, $\int_0^1 W_i di = 1$

$\pi(t) = \int_0^1 W_i \, \pi_i(t) \, di$ is the aggregate rate of price change between $t-h$ and t, termed the 'rate of inflation' . Under a steady inflation rate, π, all π_i and V can be assumed to be constant, independent of t. It can be shown that, if firms change their prices periodically due to costs of price adjustments (Sheshinski and Weiss (1977) and (1983)), V has the following structure.

Suppose that all firms allow their real prices, *i.e*, $P_i(t) - P(t)$ [where $P(t) = \int_0^1 W_i \, P_i(t) di$ is the log of the price level], to fluctuate within given bounds (S,s) [2/] (Figure 1) .

This implies that the period of each price cycle, ε, is $\varepsilon = \frac{1}{\pi}(S-s)$. The bounds (S,s), and hence ε, depend on π. It has been shown that $\frac{d(S-s)}{d\pi} > 0$, *i.e.* the amplitude of real price variation increases with π and that under certain conditions, $\frac{d\varepsilon}{d\pi} < 0$, *i.e.* the frequency $(\frac{1}{\varepsilon})$ of price changes increases with π.

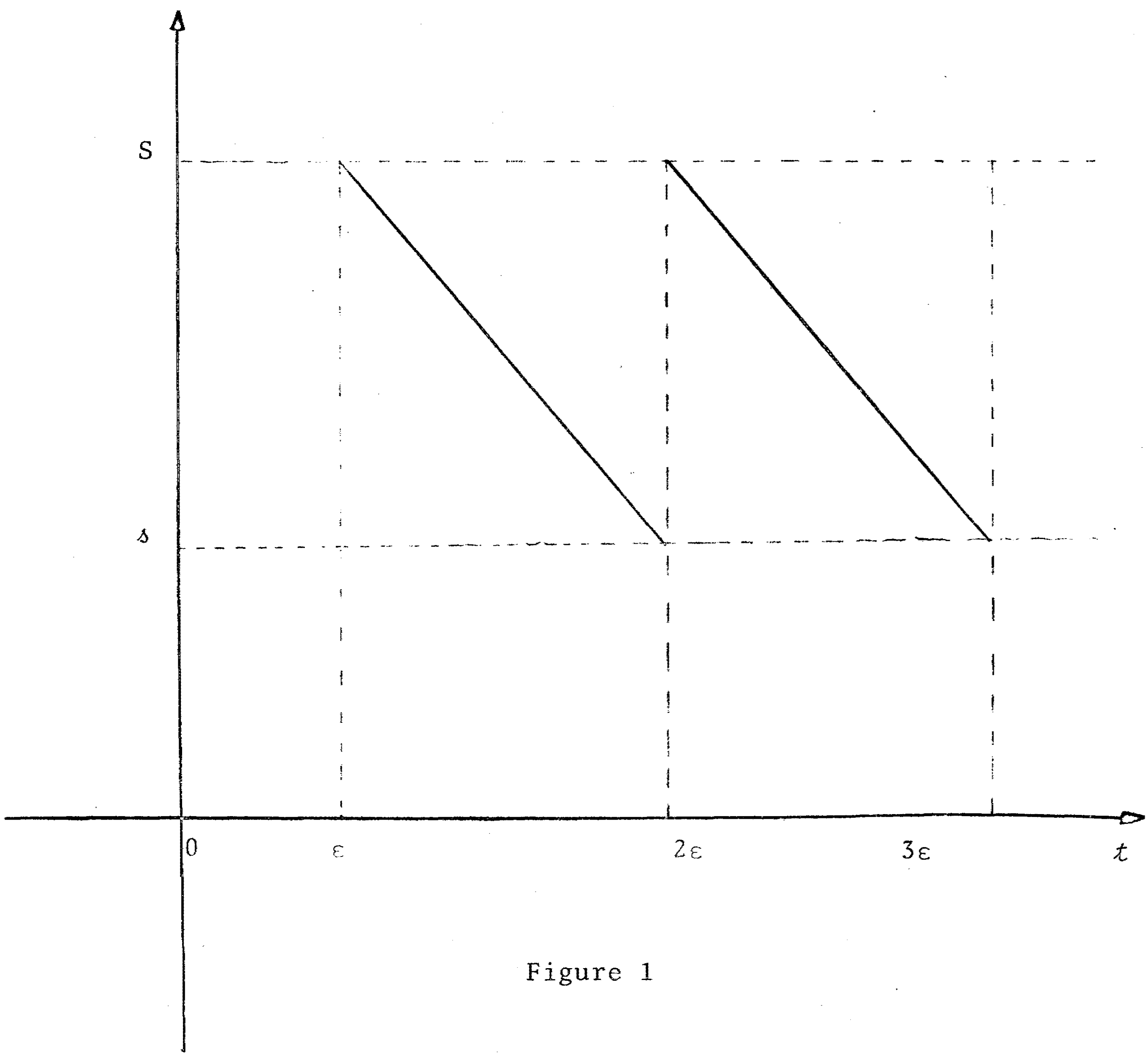


Figure 1

Given h and π (and hence ε), there exists a unique positive integer B, defined by

$$(B-1)\varepsilon \leq h \leq B\varepsilon, \tag{2}$$

such that each firm will change its price in any interval - $[t-h, t]$ either B-1 or B times, depending on its initial - price.

In steady-state, firms will be uniformly distributed over - time (Caplin and Spulber (1985)). Hence, the fraction of firms that will change their price B times over any interval $[t-h,t]$ is $\frac{h - (B-1)\varepsilon}{\varepsilon} = \frac{h}{\varepsilon} + 1-B$

while the rest, $B - \frac{h}{\varepsilon}$, will change B-1 times. Each price - change increases the log of price by $S-s = \pi\varepsilon$. The aggregate price increase over any interval $[t-h, t]$ is πh thus, - in view of (1),

$$\begin{aligned} V &= (B\pi\varepsilon-\pi h)^2 \left(\frac{h}{\varepsilon} + 1 - B\right) \\ &+ [(B-1)\pi\varepsilon - \pi h]^2 \left(B-\frac{h}{\varepsilon}\right) \\ &= \pi^2 (B\varepsilon - h) [h-(B-1)\varepsilon] \end{aligned} \tag{3}$$

Clearly, whenever ε is an integral number of h, $V=0$. Furthermore, for a given value of B,

$$\frac{dV}{d\pi} = 2\pi\ (B\varepsilon - h)\ [h-(B-1)\varepsilon] + \pi^2\ [(B\varepsilon - h)\ (1-2B) + B\varepsilon]\ \frac{d\varepsilon}{d\pi} \qquad (4)$$

By assumption $\frac{d\varepsilon}{d\pi} < 0$. Hence, whenever $B\varepsilon - h \overline{\sim} 0$, $\frac{dV}{d\pi} < 0$,and whenever $h-(B-1)\varepsilon \overline{\sim} 0$ then, $\frac{dV}{d\pi} > 0$, (See figure 2).

Obviously, V is monotone neither in π nor in B and therefore cannot serve as a measure of relative price dispersion.[3/]

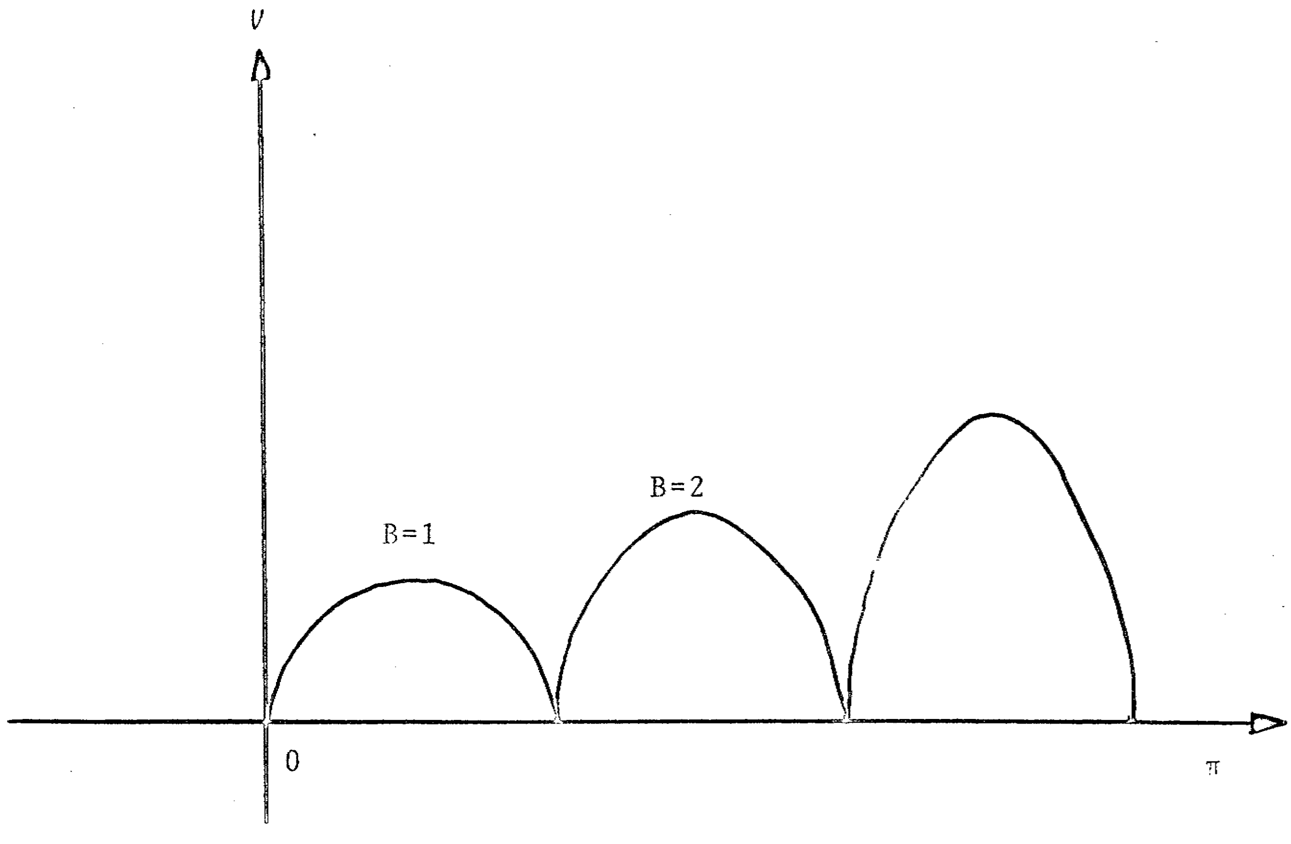


Figure 2

3. Another Inappropriate Measure[4/]

An appropriate measure of the dispersion of the real price of any good is $\pi\varepsilon = S - s$, *i.e.* the amplitude of real price variation. Hence we can construct the following measure:

$$W = \int_0^1 W_i \,(S_i - s_i)\, di = \pi \int_0^1 W_i \,\varepsilon_i \, di \qquad (5)$$

Since $\pi\varepsilon_i$ is monotone in π, W is also increasing in π. The calculation of W requires information on the frequency of price adjustments ε_i for each good i. Alternatively, an appropriate measure of real price dispersion is, say, the variance of real prices (which is constant over time in steady-state).

However, if the real price of a good reported at a certain point in time is actually, by statistical collection (sampling) procedures, an average of real prices over a certain period, say h, then the dispersion of real prices thus defined is an underestimate of the true dispersion.

Consider first the following example: take two goods whose (log of) real prices at time t are $P_1(t)$ and $P_2(t)$, respectively. If both prices follow the same (S,s) policy, then in steady-state

$$P_1(t) - P_2(t) = \frac{\pi\,\varepsilon}{2} \quad \text{(or symmetrically, } -\frac{\pi\varepsilon}{2}\text{)}$$

At $t+h$

$$P_1(t+h) = P_1(t) + [(B-1)\,\varepsilon - h]\pi \qquad (5)$$

$$P_2(t+h) = P_2(t) + (B\varepsilon - h)\pi$$

Obviously, $P_1(t+h) - P_2(t+h) = -\frac{\pi\varepsilon}{2}$, *i.e.* the spread of prices at $t+h$ is the same spread as at t, only the ran - king of prices has been reversed (see figure 3).

Taking the average price for each good at t and at $t+h$, - denoted $\overline{P}_1$ and $\overline{P}_2$ respectively,

$$\overline{P}_1 = \frac{1}{2}[P_1(t) + P_1(t+h)] = P_1(t) + \frac{1}{2}[(B-1)\,\varepsilon - h]\pi$$

$$(6)$$

$$\overline{P}_2 = \frac{1}{2}[P_2(t) + P_2(t+h)] = P_2(t) + \frac{1}{2}(B\varepsilon - h)\pi$$

Hence, $\overline{P}_1 - \overline{P}_2 = 0$, *i.e.* the "prices" $\overline{P}_i$ indicate no - - variation in relative prices.

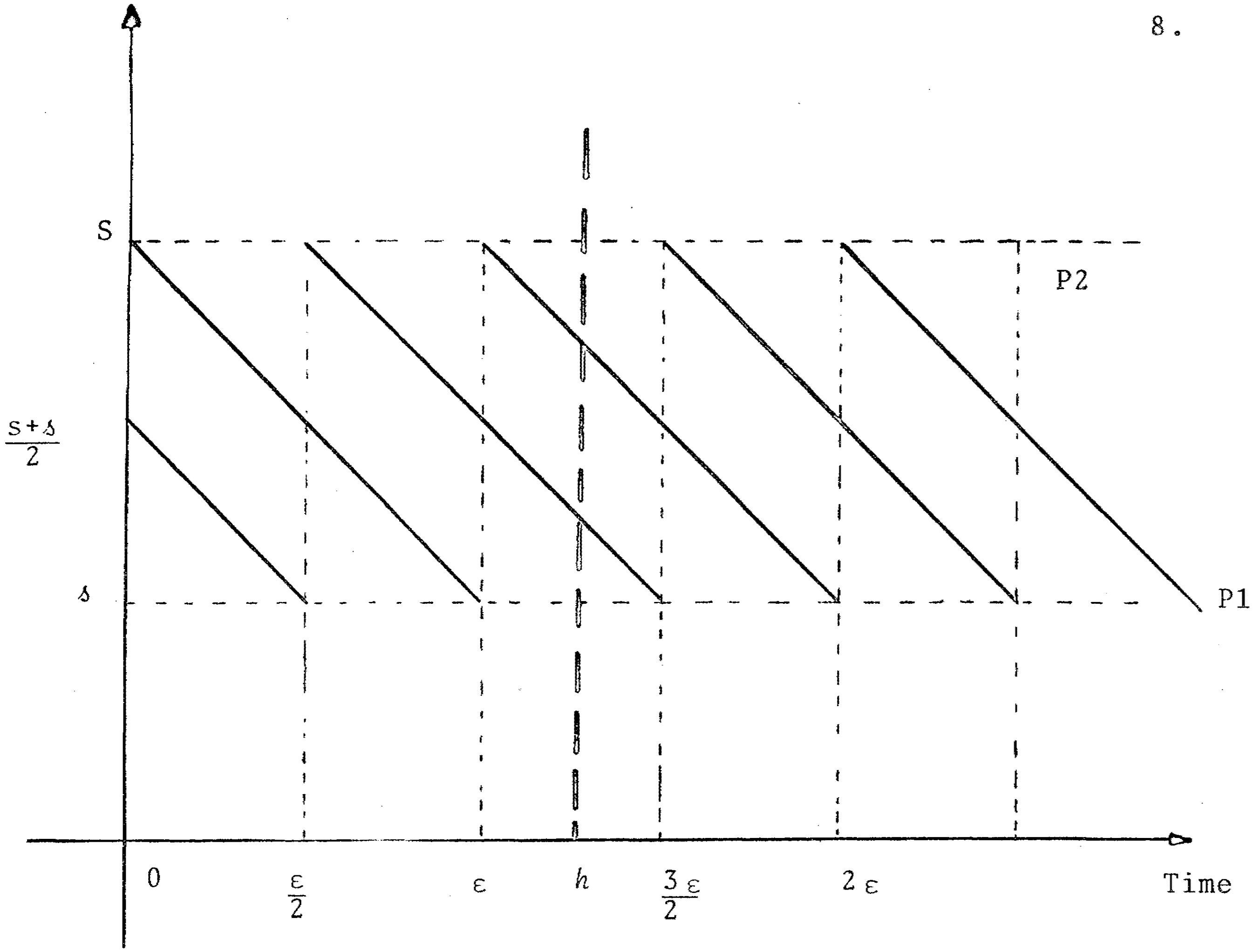


Figure 3

We shall now prove the following general proposition:

Proposition. In steady-state, when all firms follow an (S,s) pricing policy, if reported real prices are an average of real prices at different points in time, then the reported mean real price is the time mean, while the reported variance of real prices is smaller than the time variance.

Proof. Assume without loss of generality that B=1 ($h<\varepsilon$), *i.e.* firms change their price at most once within each interval - $[t, t+h]$. Suppose further that prices are simpled twice - within each period, at the beginning and at the end.

Let us index firms by i so that the initial real price of firms i is $S-i$ and i is uniformly distributed over $[0, S-s]$.

Then if a firm does not change its price in $[t, t+h]$ its reported real price, $\overline{P}_i$, is

$$\overline{P}_i = S-i - \frac{\pi h}{2} \quad . \tag{7}$$

If a firm changed its price (to a real price of S) then its reported real price is 5/

$$\overline{P}_i = \frac{3}{2} S-i - \frac{s}{2} - \frac{\pi h}{2} \tag{8}$$

Let $m = 1 - \frac{h}{\varepsilon}$, $0 \leq m \leq 1$. It is easy to see that

$$\overline{P}_i = \begin{cases} S-i - \frac{\pi h}{2} & \text{if } 0 \leq i < m(S-s) \\ \frac{3}{2} S - i - \frac{s}{2} - \frac{\pi h}{2} & \text{if } m(S-s) \leq i \leq S-s \end{cases} \tag{9}$$

Hence, the mean real price, $\overline{P}$, is given by

$$\overline{P} = \frac{1}{S-s} \int_0^{S-s} \overline{P}_i \, di = \frac{1}{S-s} \{ \int_0^{m(S-s)} (S-i - \frac{\pi h}{2}) \, di$$

$$+ \int_{m(S-s)}^{S-s} \left(\frac{3}{2} S - i - \frac{s}{2} - \frac{\pi h}{2}\right) di \qquad (9)$$

Integrating (9), we find that

$$\bar{P} = \frac{S+s}{2} + \frac{(S-s)}{2} (1-m) - \frac{\pi h}{2} \qquad (10)$$

Using that $m = 1 - \frac{h}{\varepsilon}$, and $\pi\varepsilon = S-s$, we see that $\frac{(S-s)}{2}(1-m) = \frac{\pi h}{2}$. Hence, $\bar{P} = \frac{S+s}{2}$, the true mean of real prices.

Now, with a uniform distribution of real prices over $[s,S]$, the variance, V , is known be.

$$W = \frac{(S-s)^2}{12}$$

Let us calculate the variance of reported real prices, $\bar{P}_i$, denoted $\bar{W}$:

$$\bar{W} = \frac{1}{S-s} \left\{ \int_0^{m(S-s)} \left(S - i - \frac{\pi h}{2}\right)^2 di + \int_{m(S-s)}^{S-s} \left(\frac{3}{2} S - i - \frac{s}{2} - \frac{\pi h}{2}\right)^2 di \right\} - \frac{(S+s)^2}{4} =$$

$$= \frac{1}{S-s} \left\{ \int_0^{S-s} \left[\left(S - \frac{\pi h}{2}\right)^2 - 2 \left(S - \frac{\pi h}{2}\right) i + i^2 \right] di \right. \tag{12}$$

$$\left. + \int_{m(S-s)}^{S-s} \left[\left(S - \frac{\pi h}{2}\right)(S-s) + \left(\frac{S-s}{4}\right)^2 - (S-s) i \right] di \right\} - \left(\frac{S+s}{4}\right)^2$$

Performing the integration, (12) becomes :

$$\overline{W} = \left(S - \frac{\pi h}{2}\right)^2 + \frac{1}{3} (S-s)^2 - \left(S - \frac{\pi h}{2}\right) (S-s) \quad m +$$

$$+ \left(\frac{S-s}{4}\right)^2 (1-m) - \left(\frac{S-s}{2}\right)^2 (1-m^2) \quad . \tag{13}$$

Again, substitutig $m = 1 - \frac{h}{\pi}$ and $\pi\varepsilon = S - s$ in (13), we find, after some manipulations, that

$$\overline{W} = K \left(\frac{S-s}{12}\right)^2 = K W \tag{14}$$

Were $K = 1 - 3\ (1-m)\ m$. Clearly

$K = 1$ when $m = 0$ $(h = \varepsilon)$ or $m = 1$ $(h = 0)$.

That is, there is no bias when observations are instantaneous or if the length of the reporting period coincides with the - length of the firms price cycle. It is easy to verify that - K atains a unique minimum at $m = \frac{1}{2}$, where $K = \frac{1}{4}$.

thus, $\frac{1}{4} \leq K \leq 1$ and $\overline{W} \leq W$ ||.

Obviously the bias in $\overline{W}$ could be quite significant: up to 75 per cent of the true variance.

4. Empirical Evidence

Clearly if a firm changed its price more than once within each time interval $[t, t+h]$, the argument would remain the same and equation (8) would become

$$\bar{P}_i^2 = 2S - i - \frac{\pi h}{2} - s \qquad \text{in the case of two changes}$$

$$\bar{P}_i^3 = \frac{5S}{2} - i - \frac{\pi h}{2} - \frac{3s}{2} \qquad \text{in the case of three}$$

and

$$\bar{P}_i^n = \frac{(2+n)S - ns}{2} - i - \frac{\pi h}{2} \qquad \text{in the case of } n.$$

Obviously the difference between the average prices in the case of one and n changes is

$$\delta = \frac{n-1}{2}(S-s)$$

reflecting the n price cycles (a price cycle occurs when the real price goes from S to s).

Hence the important element here is not the number of full price cycles but the fractile (F) left over after an integer number of nominal price changes. As before, the fifference between measared and real price will be largest when $F = \frac{1}{2}$

Consider, on the other hand, the case in which different commodities are being produced, each with its own bounds - $[S_j, s_j]$ ($j = 1,2,$). Suppose the distribution of Fi - - across commodities is $N(0,\sigma_F^2)$; the measured variance of - the distribution of relative prices underestimates the actual variance but that bias does not change with n , the number of price changes. Therefore our empirical estimates of that statistic across countries allows us to make valid comparisons independently of the level of inflation, which affects the - number of price changes per unit of time.

Figure 1 graphs estimates of $V(t)$ for Argentina, Israel and Mexico over the 1983-1986 period [6]. Figures 2 and 3 - include measures in tune with the discussion so far : they are moving averages of the weighted mean of the absooute - value of the difference between the actual real price of a product and it's long run average value. Hence the measures are:

$$J(t) = \sum_{i=0}^{4} \sum_{j=1}^{\ell} \alpha_j \left| \{P_j(t-i) - P(t-i)\} - \frac{S_j + s_j}{2} \right|$$

where ℓ = 61 in the case of Argentina, ℓ = 104 in the case - of Israel and ℓ = 301 in the case of Mexico. The differences between the two reflect the difficulty in empirically - - estimating $\frac{S_j + s_j}{2}$. Figure 2 takes the long run average - real price to be a centered moving average of five - - - observations:

$$\frac{1}{5} \sum_{i=0}^{4} P_j(t+2-i) - P(t+2-i)$$

until the price and wage control date and a straight six - period moving average thereon 7/. Figure 3, on the other - hand takes $\frac{S_j + s_j}{2}$ to be a straight six period moving - - average throughout the period:

$$\frac{1}{6} \sum_{i=0}^{5} P_j(t-i) - P(t-i).$$ Obviously the two - measures (V, J) yield significantly different results: V indicates the dispersion of relative prices in Mexico is equivalent to Israel's even though its rate inflation is considerably lower. The regression results of Table I - indicate that, proportionately, the caos in price formation is much worse in Mexico than in the other two countries.

Table I

	Argentina	Israel	México
R^2	.1844	.0201	.4576
Estimate of a	.3113	.0834	.4760
Standard Error	.0672	.1098	.0628
t Statistic	4.6340	.7597	7.5744

The estimated regresion is $X = a + bY + e$ with

X = logarithm of the monthly inflation rates

Y = logarithm of V

In Mexico "controlled prices"[8/] account, directly and indirectly for about a third of the Consumer Price Index. Figure 4 compares an estimate of J(t) which excludes them, with Argentina's and Israel's; in turn, Figures 5 and 6 compare the behaviour - of that dispersion index for Mexico in both cases when - - - controlled prices are included and when they are excluded - from J(t) for the two definitions of average real prices - - previously defined. It can be seen that, since the curve that includes all products is always above the one that excludes - controlled prices, price setting by the government adds to the caos in price formation and that it may do so to a considerable extent (about 35%). Furthermore, the increase in both measures from July 1985 to its maximum in January 1986, underlines the wisdom of not having implemented a price freeze during that - period.

5. CONCLUSIONS

In this paper we have criticized two of the commonly used - - statistical measures of dispersion of relative prices on the grounds that they have no relation with the amplitude of real price fluctuations. We estimated the bias of the measured - variance of the real price vis a vis the time variance and -

found it could be large.

Finally our empirical estimates underline the need to use the appropiate measure since the calculated relative price variability in Argentina, Israel and Mexico differ - - - considerably depending on the variable that is used; we - also showed the importance of taking controlled prices - into account.

FIGURE 1

VAR. POND. DE TASAS DE INFL. MENSUAL

(SUAV. 5C)

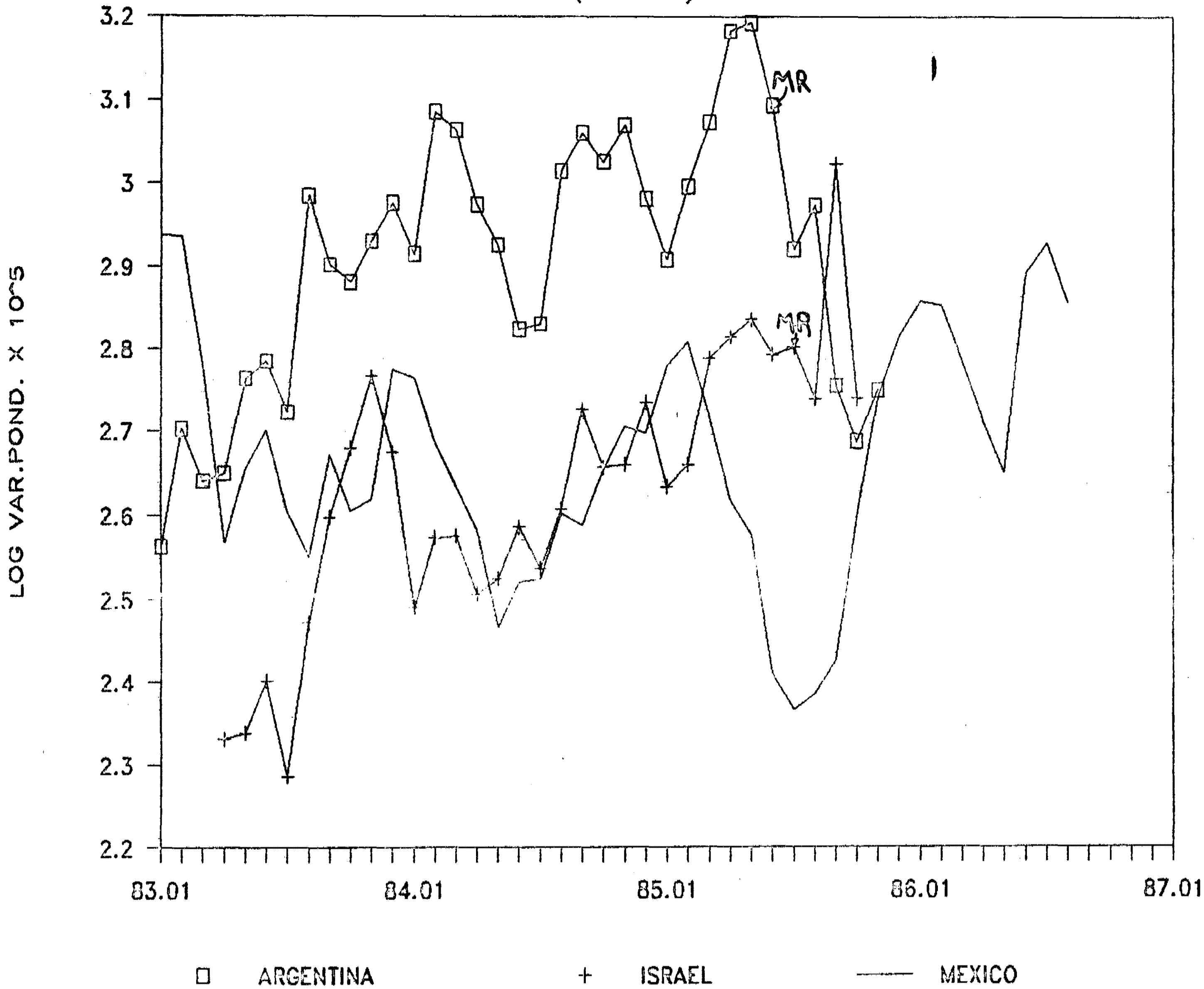


FIGURE 2

Med.pond. de |P.r.−P.obj| suav. mc5

(P.obj.=p.m.c5 hasta RM−3,p.m.r6;Mc5)

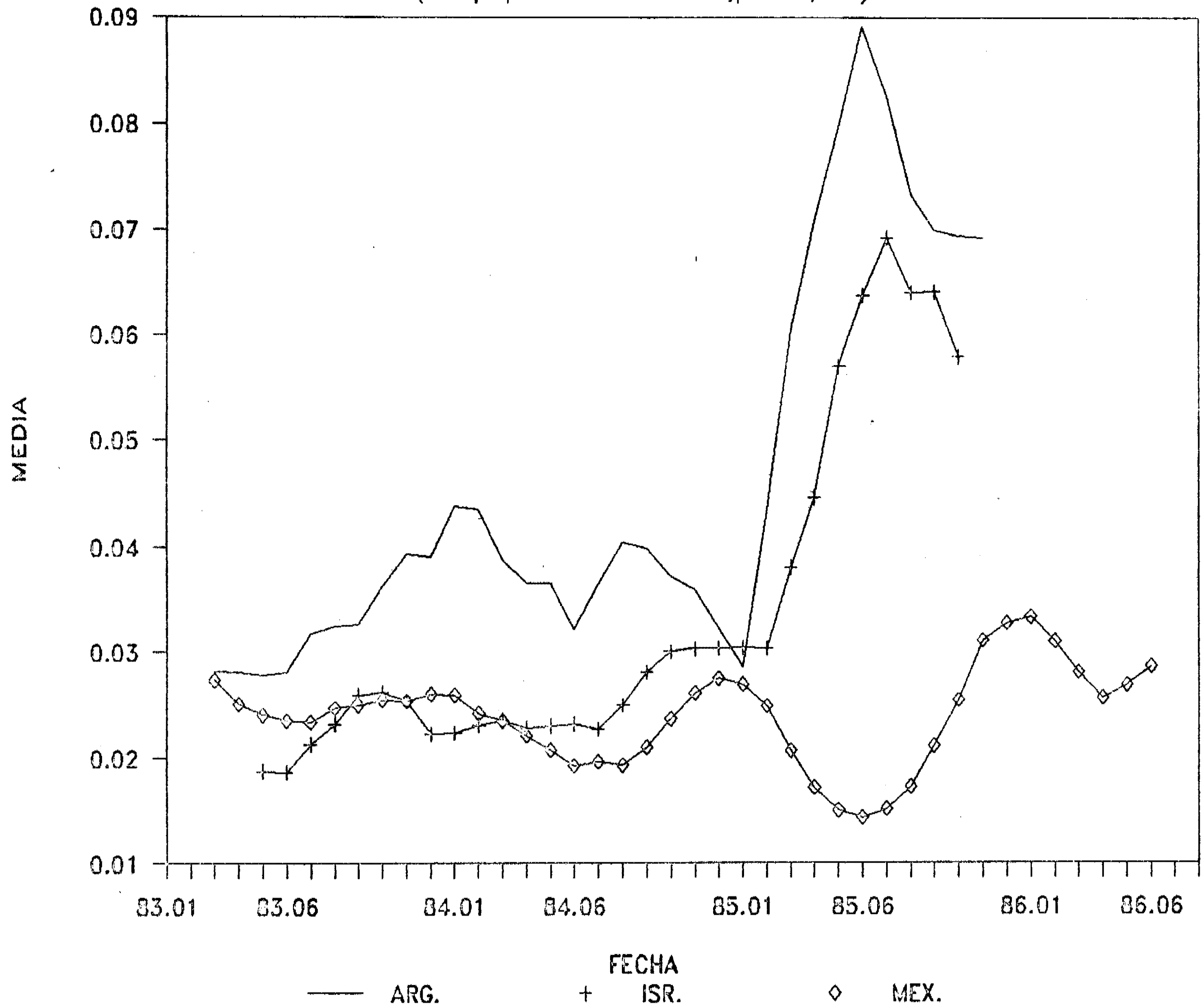


FIGURE 3

Media pond.de |P.r. − P.obj.| suav.(5c)

(P.obj.=prom movil rez. de 6 obs.)

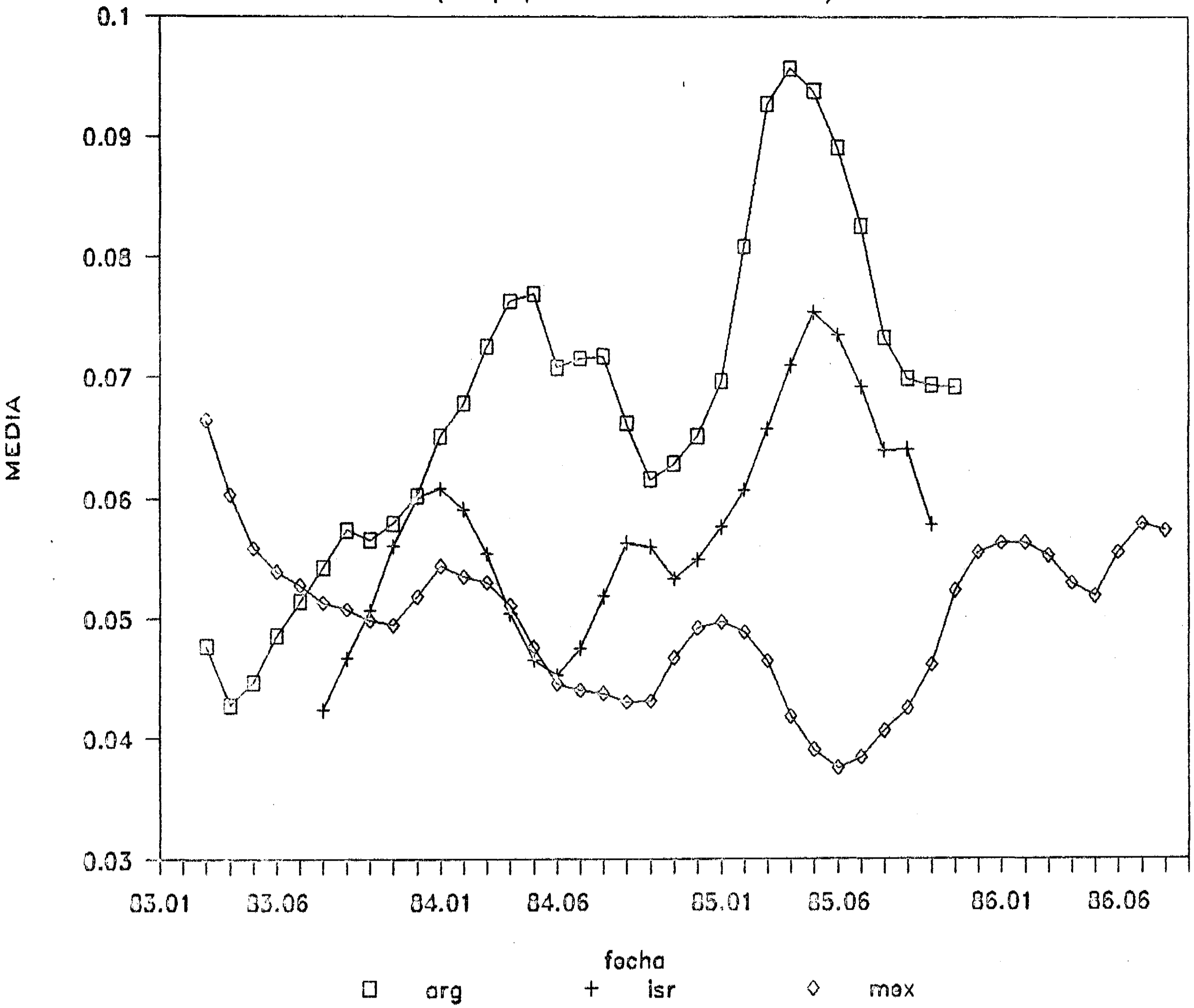


FIGURE 4

Media pond.de |P.r. – P.obj.| suav.(5c)

(P.obj.=prom movil rez. de 6 obs.)

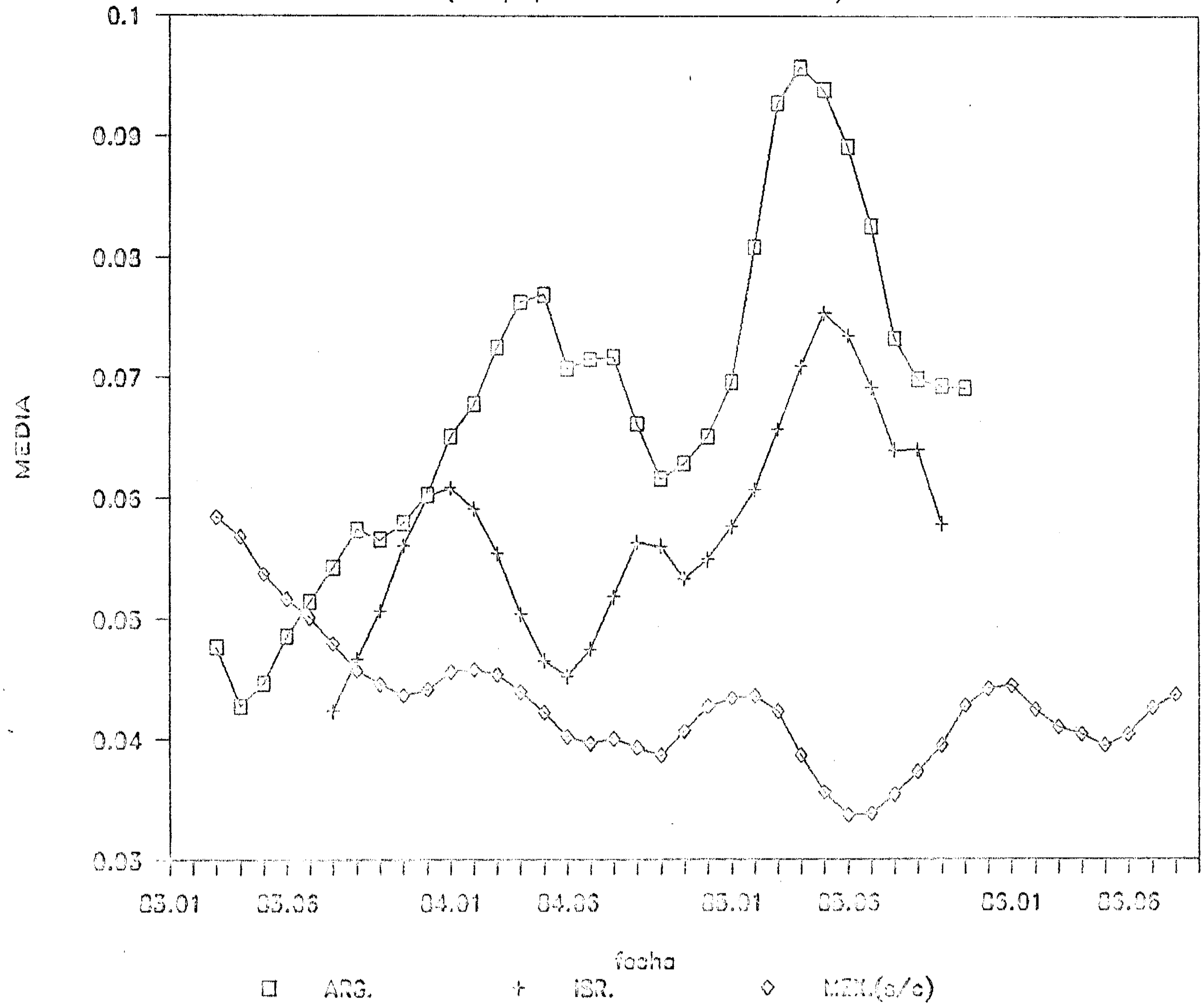
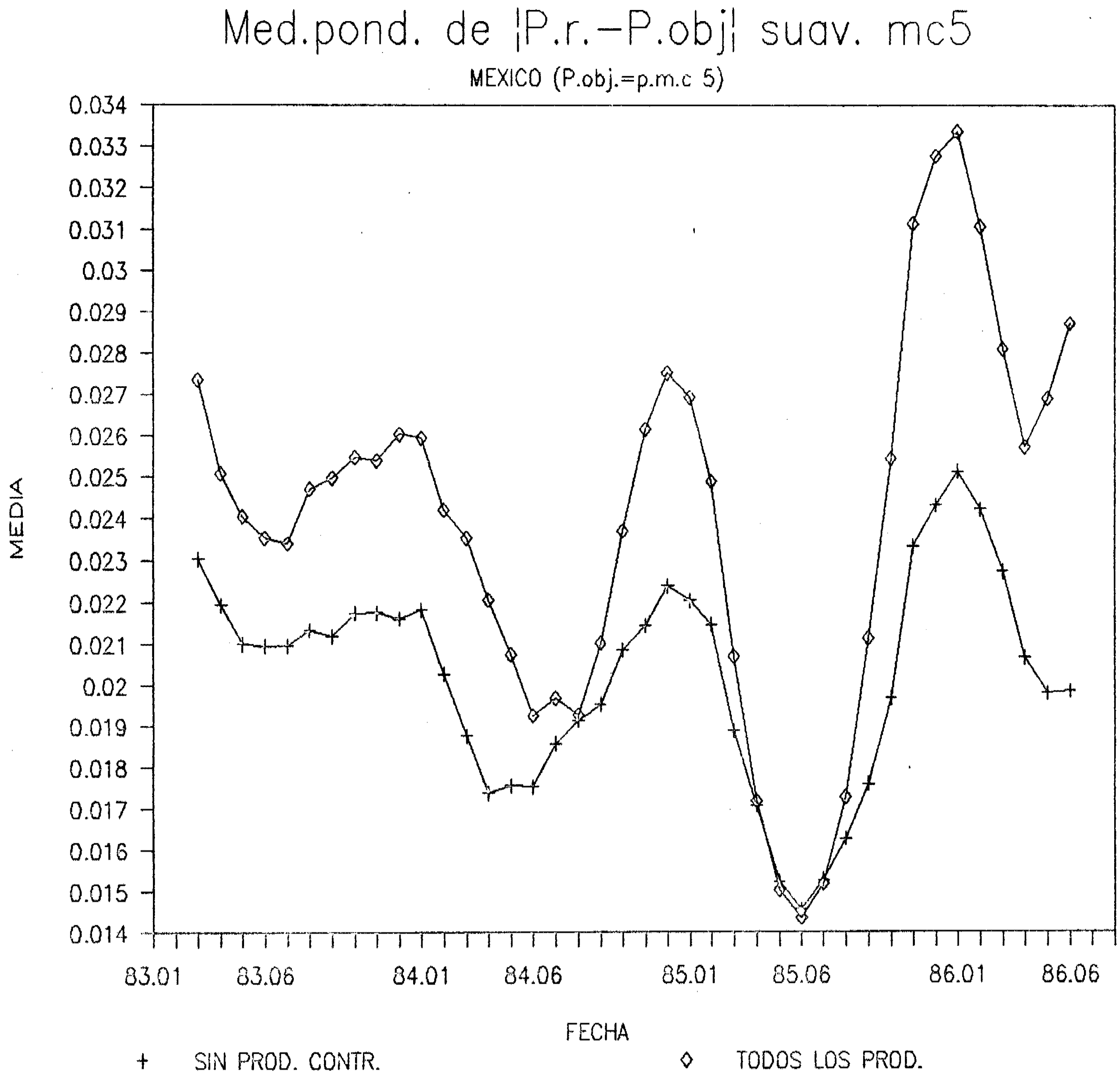


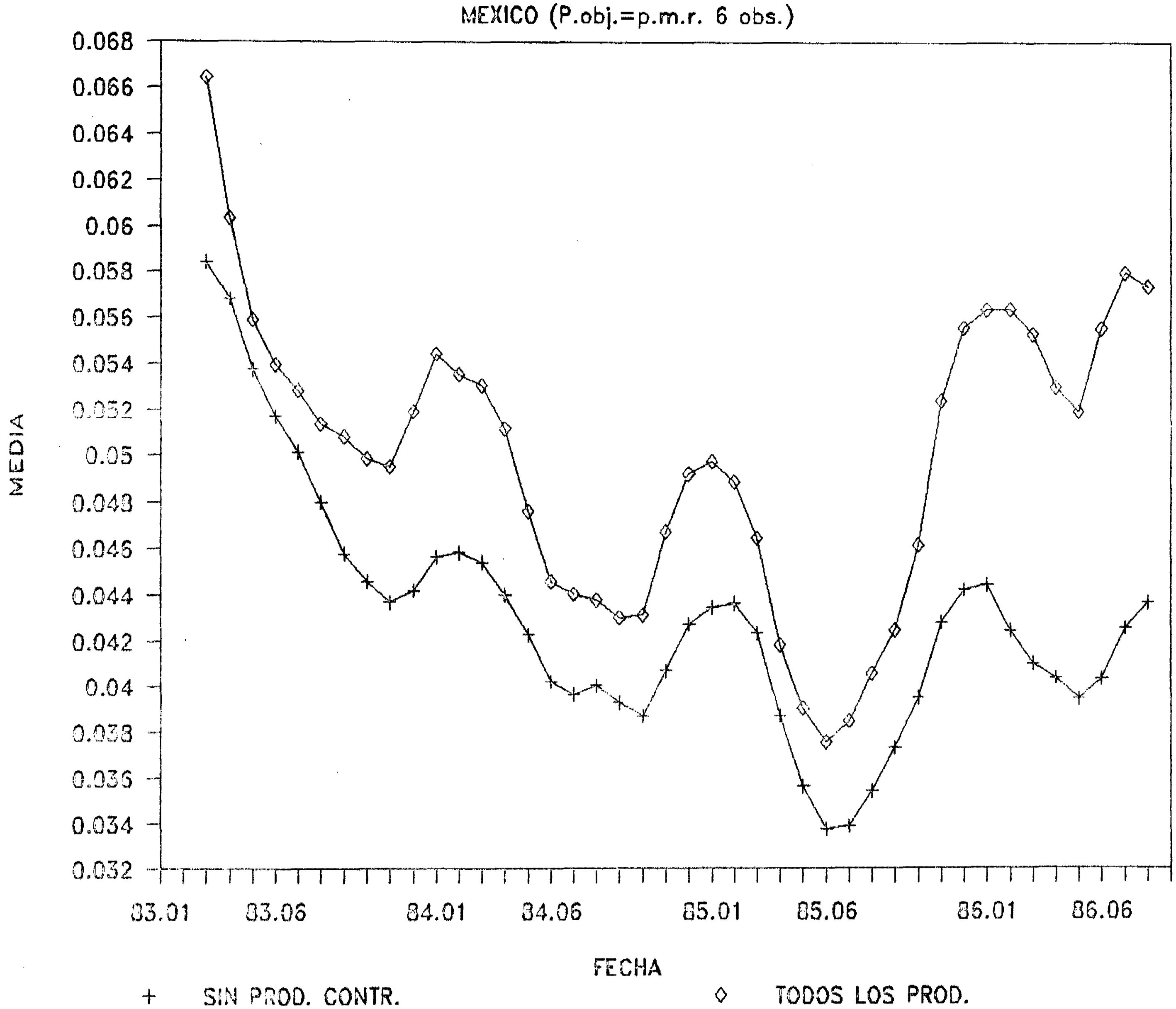
FIGURE 5



Med.pond. de |P.r.-P.obj| suav. mc5
MEXICO (P.obj.=p.m.c 5)
MEDIA
0.034
0.033
0.032
0.031
0.03
0.029
0.028
0.027
0.026
0.025
0.024
0.023
0.022
0.021
0.02
0.019
0.018
0.017
0.016
0.015
0.014
83.01
83.06
84.01
84.06
85.01
85.06
86.01
86.06
FECHA
+ SIN PROD. CONTR.
◊ TODOS LOS PROD.

FIGURE 6

Media pond. de |P.r. - P.obj.| suav. 5c



FOOTNOTES

1/ For surveys of this literature see Fischer (1981) and Marquez and Vining (1982).

2/ In log terms.

3/ Obviously neither the coefficient of variation

$$K = \frac{V}{\pi}$$

nor $K^2 = \frac{K}{V}$ are aceptable.

4/ This section was motivated by the statistical procedures for price collection in Mexico. See Alberro (1986)

5/ Suppose the firm changed its price at $T_i \varepsilon [t, t+h]$.

Then $\bar{P}_i = \frac{1}{2} [S - i + S - \pi (h - t_i)]$. But

$t_i = \frac{1}{\pi} (S - i - s)$, hence (8).

6/ In each case, the relative price variability was - measured across components of the Consumer Price - Index.

7/ The introduction of a price and wage control program would undully affect the estimates of the long run - real price.

8/ These include both governmentally produced goods and services (oil water, electricity, steel, etc..) and - privately produced commodities whose price is - - controlled (corn, tortillas, bread, rice, etc..)

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