INFLATION AND RELATIVE-PRICE VARIABILITY IN THE OPEN ECONOMY *

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The idea that the inflationary process is not neutral with respect to the structure of relative prices is receiving support from a growing number of studies. Glejser (1965) and Jaffee and Kleiman (1977) have found a positive correlation between differential inflation across sectors and the aggregate rate of inflation. Similarly, Vining and Elwertowski (1976) and Cukierman (1979) relate the variability of relative prices with that of the rate of inflation. Moreover, Parks (1978) has reported evidence indicating that it is mainly the amount of unexpected inflation which affects relative price variability.[1] Parks develops a model which results in an equation relating relative-price variability to unexpected inflation and real income. He tests the model using data for the United States, obtaining favorable results.[2]

An important feature of most previous studies is their abstraction from open economy considerations. In the case of an economy having close trade links with the rest of the world, there is a strong presumption that part of the domestic variability in relative prices will be due to foreign variability in relative prices. In fact, the existence of a mechanism of international transmission of relative price variability may lead to policy implications which differ from those of the existing literature. The purpose of this study is to develop and test an analytical framework for the determination of relative price variability in the context of a small open economy operating under a fixed exchange rate. Although we recognize that at any point in time relative prices may change in response to changes in the real side of the economy, our main focus here is on isolating the effects of inflation on the variability of relative prices.

In Section I we discuss the construction of measures of relative-price variability that are appropriate for the case of a two-sector (traded and non-traded) open economy. We then use these measures in order to assess the importance of relative-price variability within and between the sets of traded and nontraded goods. Section II begins with an examination of the relationship between inflation (expected, unexpected, and actual) in traded and nontraded goods and the relative-price variability of each.[3] A model of the open economy is then developed and its implications for relative price variability are derived. A key feature of the model is its emphasis on the specification of the market clearing mechanisms of a typical small open economy. While equilibrium requires zero domestic excess demand for nontraded goods, domestic excess demand for traded goods need not be zero, especially if the economy under study faces given international prices of traded goods. The model is then tested for the case of Mexico (1951-1976). Section III summarizes the implications of this study.

I. MEASUREMENT OF RELATIVE-PRICE VARIABILITY IN A TWO-SECTOR MODEL

The measure of relative-price variability adopted here is the one proposed by Theil (1967, Chapter 5):

$$(1) \qquad VP_t = \sum_i w_{it}(DP_i - DP)^2_t ,$$

where w_{it} is the share of expenditure on good i averaged over periods $t-1$ and t; DP_i is the rate of change in the price of good i between t-1 and t; DP is the average rate of inflation; and VP_t is our measure of relative-price variability. Since $(DP_i - DP)$ is the rate of change in relative price i, VP_t measures the nonproportionality of price movements. If all prices change at the same rate, VP_t will be equal to zero, and it increases with the dispersion of inflation rates across commodities.

Given n commodities, VP_t can be computed directly by (1). However, when subsets of the commodities considered differ in their economic characteristics, as in the case of traded and nontraded goods, the use of partial indexes of dispersion is more appropriate (see Theil, 1967, Chapter 7, Section 6). To derive such partial indexes for traded and nontraded goods, we express the aggregate rate of inflation as a weighted average of the inflation rates in each of the two sectors,[4]

(2) $$DP_t = \beta_t DP_{Tt} + (1 - \beta)_t DP_{NTt} ,$$

where T and NT index respectively traded and nontraded goods, and β is the share of traded goods in total expenditures.

With this specification, VP_t can be decomposed into three components,

(3) $$VP_t = VPR_t + \beta_t VP_{Tt} + (1 - \beta)_t VP_{NTt} ,$$

where VPR_t measures the <u>between-set</u> price variance, and VP_{Tt} and VP_{NTt} are the <u>within-set</u> measures, and correspond to equation (1) for T and NT separately (see notes to Table 1); the between-set measure is

(4) $$VPR_t = \beta_t (DP_T - DP)_t^2 + (1 - \beta)_t (DP_{NT} - DP)_t^2 .$$

We calculated Eq.(3) for Mexico using annual time series covering 1951-76. The VP_t index, which was constructed from data for prices and outputs in 47 sectors, is reported in Table 1, together with the partial variability indexes appearing at the right-hand-side of Eq.(3).[5] It is clear that the indexes fluctuate considerably: the values of VP_t range from a low of 0.56 (in 1968) to high values such as 8.33 (in 1953) and 6.10 (in 1974); VPR_t obtains values in the (0.0005, 1.079) range; and the values of VP_{Tt} and VP_{NTt} are in the ranges (0.42, 8.65) and (0.22, 8.55).

Are fluctuations in VP mainly associated with fluctuations in a specific partial index? Although Table 1 is informative in this respect, its interpretation encounters the difficulty that the indexes are not directly comparable since the degree of aggregation of the data affects the different measures of variability (see Theil, 1967, pp.162-63). However, Eq.(3) can be used in order to assess the contribution of each index to the overall variability. Specifically, we calculate the shares of VP associated with between-sets variance (VPR/VP) and within-set variance [$\beta VP_T/VP$ and $(1-\beta)VP_{NT}/VP$]. These shares are reported in Table 2.

In column (1) of the table it can be seen that the share of total variability attributable to VPR (between-set variance) is for the most part negligible (and more or less constant). Most of VP, then, can be attributed to the dispersion of relative-price changes within each set. For this reason our analysis below focuses on the determination of relative-price variability within the traded and nontraded goods sets.

To assess the relative importance of within-set variability for the determination of overall relative price variability, we look at columns (2) and (3) of Table 2.

TABLE 1: RELATIVE PRICE VARIABILITY AND ITS COMPONENTS, MEXICO: 1951-76a/

	Total variability, VP_t	Between-sets variance, VPR_t	Within-set variance	
			Traded goods, VP_{Tt}	Nontraded goods, VP_{NTt}
1951	3.89	0.117	6.08	1.44
1952	3.09	1.079	2.15	1.88
1953	8.33	0.067	7.98	8.55
1954	3.17	0.140	4.29	1.71
1955	2.11	0.009	3.72	0.41
1956	1.89	0.006	2.07	1.69
1957	3.98	0.091	6.42	1.27
1958	3.12	0.678	3.94	0.89
1959	1.22	0.100	1.58	0.63
1960	1.68	0.356	1.46	1.18
1961	0.93	0.001	1.13	0.71
1962	0.90	0.001	0.42	1.41
1963	0.93	0.038	0.93	0.86
1964	1.11	0.092	1.02	1.02
1965	0.74	0.001	0.76	0.72
1966	0.88	0.245	0.94	0.31
1967	0.75	0.000	0.98	0.51
1968	0.56	0.033	0.68	0.37
1969	0.57	0.026	0.82	0.25
1970	0.73	0.013	1.17	0.22
1971	2.12	0.074	3.53	0.44
1972	0.86	0.022	1.34	0.32
1973	3.21	0.015	5.10	1.21
1974	6.10	0.160	8.65	3.12
1975	1.67	0.021	2.71	0.55
1976	2.42	0.037	3.86	0.88

Notes on following page.

Notes from Table 1.

a/ VP_t and VPR_t according to equations 1, 3, and 4. The within-set measures correspond to equation (1):

$$VP_{Tt} = \sum_{i=1}^{k} \alpha_{it} (DP_{iT} - DP_T)_t^2$$

and

$$VP_{NTt} = \sum_{i=k+1}^{n} \alpha_{it} (DP_{iNT} - DP_{NT})_t^2 .$$

There are k traded and n - k nontraded commodities; α_{it} is the average share of commodity i in expenditure on traded goods (i = 1, ..., k) or nontraded goods (i = k + 1, ..., n). DP are first-differences of logs. The indexes computed are here multiplied by 1000.

TABLE 2: PROPORTION OF VP ACCOUNTED FOR BY EACH COMPONENT [a/]

	Between sets, VPR_t/VP_t	Within-set	
		Traded goods $\beta_t VP_{Tt}/VP_t$	Nontraded goods $(1-\beta)_t VP_{NTt}/VP_t$
	(1)	(2)	(3)
1951	0.03	0.78	0.19
1952	0.34	0.34	0.32
1953	0.00	0.48	0.52
1954	0.04	0.69	0.27
1955	0.00	0.90	0.10
1956	0.00	0.55	0.45
1957	0.03	0.82	0.15
1958	0.22	0.64	0.14
1959	0.08	0.67	0.25
1960	0.21	0.45	0.34
1961	0.00	0.62	0.38
1962	0.00	0.24	0.76
1963	0.04	0.51	0.45
1964	0.08	0.47	0.45
1965	0.00	0.53	0.47
1966	0.28	0.55	0.17
1967	0.00	0.67	0.33
1968	0.06	0.62	0.32
1969	0.05	0.74	0.21
1970	0.02	0.83	0.15
1971	0.03	0.87	0.10
1972	0.03	0.79	0.18
1973	0.00	0.81	0.19
1974	0.03	0.72	0.25
1975	0.01	0.82	0.17
1976	0.02	0.80	0.18

a/ β and $1 - \beta$ are respectively the shares of traded and nontraded goods in total expenditure. 0.00 indicates less than 1 per cent.

In almost all cases the variability within traded goods accounts for a much larger fraction of the total than the variability within nontraded goods. An important implication of this finding is that insofar as domestic economic variables affect mostly the variability within the set of nontraded goods, while foreign (exogenous) factors affect primarily the variability within traded goods, it is apparent that a large fraction of relative-price variability in the open economy is attributable to variables that are beyond the direct control of the domestic authorities. In the next section we examine the extent to which domestic and foreign variables (including inflation) differ in their effect on the measures of within-set relative-price variability.

II. DETERMINANTS OF RELATIVE-PRICE VARIABILITY IN THE OPEN ECONOMY.

Simple Tests of the Effects of Inflation on Relative-Price Variability.

Before considering a more complete macroeconomic model, we investigate the direct relationship between inflation and the variability of relative prices within each group of commodities. To do so, we estimated equations of the form:

$$(5) \qquad VP_{jt} = a_j + b_j (DP_j)_t^2 \qquad j = T, NT$$

Table 3 reports the results for Eq.(5). While relative-price variability in the traded-goods sector appears to be positively and significantly affected by actual inflation in that sector, no significant relationship is found for the nontraded sector.

As mentioned in the introduction, however, it has been recently asserted that inflation affects relative-price variability only when it is not anticipated by economic agents. This hypothesis can be tested by estimating equations of the form:

TABLE 3: INFLATION AND RELATIVE-PRICE VARIABILITY, MEXICO: 1951-76.[a/]

Coefficients	Traded Goods (T)		Nontraded goods (NT)	
	Equation (5)	Equation (6)	Equation (5)	Equation (6)
Constant	0.002 (0.0004)	0.002 (0.0005)	0.001 (0.0004)	0.0005 (0.0004)
Inflation				
Actual $(DP_{jt})^2$	0.111 (0.031)		0.017 (0.032)	
Expected $(EDP_{jt})^2$		0.026 (0.057)		0.037 (0.049)
Unexpected $(DP_{jt} - EDP_{jt})^2$		0.325 (0.097)		0.291 (0.071)
R^2	0.362	0.365	0.013	0.441
D.W.	1.427	1.883	1.691	2.055

a/ The dependent variable is VP_{jt} (j = T, NT); Numbers in parentheses are the standard errors of the coefficients.

$$VP_{jt} = a_j + b_j(EDP_j)_t^2 + c_j(DP_j - EDP_j)_t^2 , \quad j = T, NT \tag{6}$$

where the actual rate of inflation has been decomposed into an expected part, denoted by EDP, and an unexpected part, denoted by DP - EDP. In order to estimate Eq.(6), measures of expected inflation in each sector are required. Here we have fitted simple first-order autoregressive processes of the form $DP_{jt} = d_{0j} + d_{1j}DP_{jt-1}$ + residual; (j = T, NT). With this simple specification, the best predictor for DP_{jt} is $d_{0j} + d_{1j}DP_{jt-1}$; we have used this predictor to approximate the expected inflation variable, EDP_{jt}.[6]

The results of estimating (6) are also presented in Table 3. We find that for both sets of goods (traded and nontraded) only unexpected inflation has a significant and positive effect on the pertinent measures of relative-price variability; the effects of expected inflation are not significant.

Formulation and Testing of a Multimarket Model of the Open Economy.

The results presented above shed light on the relationship between inflation and relative price variability for traded and nontraded goods. However, these results are derived from simple,model-free, specifications that abstract from other macroeconomic determinants of relative price variability than inflation. Our purpose here is to develop and test a multimarket model of the open economy that will enable us to analyze the role of inflation in conjunction with that of other macroeconomic variables in the generation of relative-price changes. As the market clearing mechanism operates differently for traded and nontraded goods, each sector is modeled separately.

Nontraded Goods: The basic model consists of supply and demand equations expressed in rates of change for each nontraded good:

(7) $$DQ^s_{it} = \gamma^s_i + \varepsilon^s_i (DP_i - DP^*)_t$$

(8) $$DQ^d_{it} = \gamma^d_i - \varepsilon^d_i (DP_i - DP^*)_t + \lambda_i (DM - DP_i)_t \ ,$$

where DQ^s and DQ^d are the rate of change of respectively commodity supply and demand, DP_i is the rate of change in the price of commodity i, DP^* is the expected rate of inflation, DM is money growth, and $i = k + 1, \ldots, n$ (i.e., i indexes the $n - k$ nontraded goods). Eq.(7) asserts that the supply of each commodity has a trend-growth component (γ^s_i), which presumably captures the supply effects of secular changes in technology, resource availability, and in other underlying determinants of supply. In addition, it is postulated that the supply of commodity i is an increasing function of its perceived relative price; i.e., ε^s_i should be positive.[7] Eq.(8) asserts that the quantity demanded of commodity i depends on three variables: (a) γ^d_i, which is a trend-growth component that represents the demand effects of secular changes in permanent real income, family composition, and other underlying determinants of demand; (b) the perceived relative price of commodity i: whenever agents perceive an increase in the price of i relative to other commodity prices, they reduce their demand for i; (c) $(DM - DP_i)$, which stands for a real-balance ffect on demand. When the government increases its transfers of money to the public, part of the additional transfers is translated into an increase in the demand for commodities as reflected by the λ_i coefficient $(\lambda_i > 0)$.[8]

Assuming that the market for each nontraded good clears, the equilibrium rate of change of the price of each good is given by

$$(9) \qquad DP_{it} = \left(\frac{1}{\varepsilon_i + \lambda_i}\right)[\varepsilon_i DP^*_t + \lambda_i DM_t + \gamma_i],$$

where $\varepsilon_i \equiv \varepsilon^d_i + \varepsilon^s_i$ and $\gamma_i = \gamma^d_i - \gamma^s_i$. The model results in standard predictions: increases in money growth and in expected inflation result in increases in prices. Similarly, an increase in underlying excess-demand growth increases prices.

The rational-expectations assumption in conjunction with the definition of the price level [Eq.(2)] implies that the expected rate of inflation is a weighted average of expected inflation in traded and nontraded goods:

$$(10) \qquad DP^*_t = EDP_t = \beta_t EDP_{Tt} + (1 - \beta)_t EDP_{NTt} \quad ,$$

where E is the mathematical expectation conditional on a given information set that will be specified below.

Substituting Eq.(10) into (9), subtracting DP_{NTt} from both sides, and using the identity $DP_{NT} \equiv EDP_{NT} + (DP_{NT} - EDP_{NT})$ yields

$$(11) \qquad (DP_i - DP_{NT})_t = \left(\frac{1}{\varepsilon_i + \lambda_i}\right)[\varepsilon_i \beta_t (EDP_T - EDP_{NT})_t - \varepsilon_i (DP_{NT} - EDP_{NT})_t + \lambda_i (DM - DP_{NT})_t + \gamma_i].$$

Eq.(11) expresses the actual change in the relative price of nontraded commodity i as a function of the expected change in the traded/nontraded price ratio, unexpected inflation in nontraded goods, and growth in real money balances. As previously defined, relative price variability within nontraded goods is measured by

(12) $$VP_{NTt} = \sum_{i=k+1}^{n} \alpha_{it}(DP_i - DP_{NT})_t^2 .$$

Combining expressions (11) and (12), we obtain

(13) $$\begin{aligned} VP_{NTt} = {} & f_0 + f_1(EDP_T - EDP_{NT})_t^2 + f_2(DP_{NT} - EDP_{NT})_t^2 + \\ & + f_3(DM - DP_{NT})_t^2 + f_4(EDP_T - EDP_{NT})_t + f_5(DP_{NT} - EDP_{NT})_t \\ & + f_6(DM - DP_{NT})_t + Z_t , \end{aligned}$$

where Z_t includes further interaction terms,

$$\begin{aligned} Z_t = {} & f_7(EDP_T - EDP_{NT})_t(DP_{NT} - EDP_{NT})_t + \\ & + f_8(EDP_T - EDP_{NT})_t(DM - DP_{NT})_t + f_9(DP_{NT} - EDP_{NT})_t(DM-DP_{NT})_t , \end{aligned}$$

and the f coefficients are defined as follows:

$$\phi_i = \frac{\alpha_{it}}{(\varepsilon_i + \lambda_i)^2} \qquad i = k+1,\ldots,n$$

and

$$f_0 = \sum\phi_i\gamma_i^2 \qquad f_2 = \sum\phi_i\varepsilon_i^2$$

$$f_1 = \sum\phi_i(\varepsilon_i\beta_t)^2 \qquad f_3 = \sum\phi_i\lambda_i^2$$

$$f_4 = 2\sum\phi_i\varepsilon_i\beta_t\gamma_i \qquad f_7 = -2\sum\phi_i\varepsilon_i^2\beta_t$$

$$f_5 = -2\sum\phi_i\varepsilon_i\gamma_i \qquad f_8 = 2\sum\phi_i\beta_t\varepsilon_i\lambda_i$$

$$f_6 = 2\sum\phi_i\lambda_i\gamma_i \qquad f_9 = -2\sum\phi_i\varepsilon_i\lambda_i \quad .$$

According to Eq.(13), the variability of relative prices within the subset of nontraded goods responds to changes in the expected price ratio between traded and nontraded goods, unexpected inflation in nontraded goods, changes in real money growth, and excess-demand shifts. Notice that the f coefficients are functions of the underlying nontraded supply-and-demand parameters. Our previous assumptions imply that $f_0, f_1, f_2, f_3, f_7, f_8, f_9$ should be positive. The sign of f_4, f_5, f_6 is ambiguous and depends on the assumption made about the signs of the γ_i, the trend-growth of excess demand for each nontraded good. If the γ_i have the same sign for all commodities, then

$$\left.\begin{array}{l} f_4, f_6 \gtrless 0 \\ f_5 \lessgtr 0 \end{array}\right\} \text{ according to } \gamma_i \gtrless 0 \text{ (for all } i\text{)}.$$

If γ_i differ across commodities, the signs of these f coefficients will be indeterminate. However, f_4 and f_5 will be opposite in sign, while the sign of f_6 depends on the particular configuration of the appropriate weights.

In order to test the model on the basis of equation (13), three additional assumptions were made. First, we assumed that the coefficients are stable, so that the equation can be estimated on the basis of standard regression methods. Second, an assumption about the mechanism of formation of the expectations EDP_T and EDP_{NT} is required. As before, the expectational variables were constructed from the fitted values of estimated first-order autogressions of inflation in traded and nontraded goods (see note (6)). Third, we have assumed that the contribution of Z_t to VP_{NTt} is of a negligible order of magnitude and is ignored in the estimated equations.[9]

Eq.(13) was estimated using the Cochrane-Orcutt technique, obtaining the results shown in column (1) of Table 4.[10] The equation appears to explain movements of VP_{NTt} quite satisfactorily. The variables measuring the traded/nontraded price ratio, unexpected nontraded goods inflation, and real money growth enter with the hypothesized sign and are significant in squared form. In linear form, only the price-ratio variable is not significant. As expected, f_4 and f_5 have opposite signs; f_6 turns out to be negative.

Overall, this pattern of empirical results indicates that the model is quite compatible with the sample information. In particular, there are significant effects of unexpected inflation in nontraded goods on the amount of relative-price variability within these goods. Notice that expected inflation in nontraded goods is not included in the estimated equation. This is so because the model embodies the hypothesis that, other things being equal, expected inflation will not affect relative-price variability. To test this implication of the model, we re-estimated the equation with the inclusion of expected inflation in nontraded goods in both squared and linear forms, obtaining the results in column (2) of Table 4. The expected-inflation variables are not significant. Moreover, with only one exception the coefficients of the other variables are not much affected by the inclusion of the expected inflation variables; however, the estimates are less precise. All in all, then, the findings reported in this and the previous section indicate that inflation in nontraded goods affects the degree of relative-price variability in these goods only when it is not anticipated by economic agents.

Traded Goods: In modeling relative-price variability in the traded goods sector, we consider two alternatives. First, it can be hypothesized that the underlying factors that determine relative-price variability in nontraded goods are also relevant for traded goods. Alternatively, to the extent that the law of one price applies and the economy faces exogenous international prices of traded goods, the relative-price variability of traded goods is exogenous to the economy under consideration.

In principle, it is possible to use the same model as before. Indeed, the only modification needed is to define i as indexing the k traded commodities. If this is done, and if we substitute (10) into (9), subtract DP_{Tt} from both sides, and use the condition $DP_{Tt} \equiv EDP_{Tt} + (DP_{Tt} - EDP_{Tt})$, we obtain

(11')
$$(DP_i - DP_T)_t = \left(\frac{1}{\varepsilon_i + \lambda_i}\right)[\varepsilon_i(1 - \beta)_t(EDP_{NT} - EDP_T)_t - $$
$$- \varepsilon_i(DP_T - EDP_T)_t + \lambda_i(DM - EDP_T)_t + \gamma_i] ,$$

where $i = 1,\ldots,k$. Recall that relative-price variability within traded goods is defined by

(12')
$$VP_{Tt} = \sum_{i=1}^{k} \alpha_{it}(DP_i - DP_T)^2_t .$$

TABLE 4: COEFFICIENTS FOR EQUATION (13) FOR NONTRADED GOODS WITH AND WITHOUT EXPECTED INFLATION (DEPENDENT VARIABLE, VP_{NTt}) MEXICO, 1951-76 a/

	Coefficients	
	(1)	(2)
Constant	0.001 (0.0003)	-0.002 (0.0015)
$(EDP_T - EDP_{NT})^2_t$	0.899 (0.472)	0.560 (0.994)
$(DP_{NT} - EDP_{NT})^2_t$	0.270 (0.063)	0.340 (0.069)
$(DM - DP_{NT})^2_t$	0.312 (0.077)	0.252 (0.096)
$(EDP_T - EDP_{NT})_t$	0.021 (0.015)	0.019 (0.048)
$(DP_{NT} - EDP_{NT})_t$	-0.017 (0.005)	-0.020 (0.005)
$(DM - DP_{NT})_t$	-0.037 (0.009)	-0.025 (0.013)
$(EDP_{NT})^2_t$		-0.262 (0.273)
EDP_{NTt}		0.054 (0.042)
R^2	0.864	0.896
D.W.	2.08	2.30
ρ	-0.601 (0.163)	-0.571 (0.168)

a/ Column (1) is the estimate of equation (13), omitting Z_t. The column (2) estimate includes also expected inflation in nontraded goods. Numbers in parentheses are standard errors; ρ is the estimated first-order autocorrelation coefficient of the residuals.

As before, we can now combine (11') and (12') to obtain Eq.(13') for VP_{Tt},

$$VP_{Tt} = f_0 + f_1'(EDP_{NT} - EDP_T)_t^2 + f_2(DP_T - EDP_T)_t^2 + f_3(DM - DP_T)_t^2 +$$

(13')

$$+ f_4'(EDP_{NT} - EDP_T)_t + f_5(DP_T - EDP_T)_t + f_6(DM - DP_T)_t + Z_t' \, ,$$

where the f coefficients are defined as before, except that now $i = 1,\ldots,k$ and that $f_1' = \sum\phi_i\varepsilon_i^2(1-\beta)_t^2$, $f_4' = \sum\phi_i\varepsilon(1-\beta)_t\gamma_i$, i.e., $(1-\beta)_t$ replaces β_t; Z_t' is the traded-goods counterpart of Z_t and here too this term is ignored in the analysis.

Eq.(13') is the expression for relative-price variability of traded goods. Using the same model and estimation technique as for Eq.(13) we get

$$VP_{Tt} = \underset{(0.001)}{0.003} + \underset{(1.180)}{1.689}(EDP_{NT} - EDP_T)_t^2 + \underset{(0.244)}{0.310}(DP_T - EDP_T)_t^2 -$$

(14)

$$- \underset{(0.255)}{0.069}(DM - DP_T)_t^2 - \underset{(0.044)}{0.023}(EDP_{NT} - EDP_T) +$$

$$+ \underset{(0.016)}{0.004}(DP_T - EDP_T)_t - \underset{(0.038)}{0.008}(DM - DP_T)_t$$

$R^2 = 0.559$

D.W. $= 1.37$

$\rho = \underset{(0.200)}{-0.181}$

The results yield no support for the specification of Eq.(13'). Only the constant term in the equation is significantly different from zero. Moreover, the F statistic for testing the significance of the regression as a whole is $F(6,17) = 3.59$, which is below the 1 percent critical level. This should be compared with the results for the nontraded sector [Table 4, column (1): the F statistic is $F(6,17) = 18.03$, which is much greater than the 4.10 (1 percent) critical value]. Thus, the multimarket model of the previous section appears to be applicable only to the nontraded sector of the small open economy under study.

With this finding in mind, we now turn to a second approach in modeling the price behavior in the traded good sector. This approach is based on the idea that in a small open economy operating under fixed exchange rates and well-integrated with the rest of the world, the behavior of domestic prices of traded goods will follow closely that of the corresponding international prices. At the extreme, when no change occurs in domestic tariffs or transportation costs, the following condition is implied

$$DP_{iT} = DP^{W}_{iT} , \tag{15}$$

that is, the domestic rate of change of the price of traded commodity i will be equal to its rate of change in international markets (DP^{W}_{iT}). In this case, relative-price variability in traded goods, Eq.(12'), can be expressed as

$$VP_{Tt} = \sum_{i=1}^{k} \alpha_{it}(DP^{W}_{i} - DP^{W}_{T})^{2}_{t} \tag{16}$$

that is to say, VP_{Tt} is externally given to the small open economy.

Clearly, the validity of this approach must rest on a direct test of the hypothesis embodied in Eq.(15). This could be done by looking at disaggregated data, using a methodology similar to Isard's (1977). However, this would be far beyond the scope of this paper. Instead, we directly test the hypothesis that domestic relative price variability closely follows its foreign counterpart. This can be done by using the VP variable constructed by Hercowitz (1980) for the United States as a proxy for foreign variability, and relating it to the measures of domestic variability constructed by us. The following regression results were obtained:[11]

Traded Goods.

(17)
$$VP_{Tt} = \underset{(0.005)}{0.015} + \underset{(0.675)}{2.625}\, VP_{USt}$$

$R^2 = 0.459$
$DW = 1.93$
$F(1,23) = 15.09$
$\rho = 0.194$

Nontraded Goods.

(18)
$$VP_{NTt} = \underset{(0.004)}{0.009} + \underset{(0.623)}{0.722}\, VP_{USt}$$

$R^2 = 0.086$
$DW = 1.99$
$F(1,23) = 1.34$
$\rho = 0.179$

The estimates indicate that U.S. relative price variability only affects the variability within Mexican traded goods. While the coefficient of VP_{US} is highly significant in the traded-goods equation, it is insignificant in the nontraded goods equation (see also the difference in the F-statistic obtained in each case). These findings support the notion that the relative price variability of Mexican traded goods is mainly determined by factors external to the Mexican economy; a result that differs from the one obtained above for nontraded goods.

III. SUMMARY AND CONCLUDING REMARKS

This paper has explored the determination of relative-price variability in a fixed-exchange-rate open economy, Mexico (1951-76). Once the distinction is made between traded and nontraded goods, it is appropriate to decompose overall relative-price variability into three components: the variability within each set of goods and between the two sets. The results of Section I suggest that in the Mexican economy an important part of the overall variability is explained by the within traded-goods variance, and that the between-sector variance accounted for a negligible share. Accordingly, the analysis of Section II was confined to the determination of the within-sector variance. When the actual rate of inflation is decomposed into expected and unexpected components, only the unexpected components appear to have strong and significant effects on the within-sector variability of relative prices. These findings emerged from the simple examination of the relationship between variability within traded and nontraded goods and their respective rates of inflation as well as from the results of testing the multimarket model presented in Section II. This model results in an equation relating relative-price variability within a set of goods to unexpected inflation in these goods, the expected rate of change of the traded/nontraded price ratio, and real money growth. While the model provided a satisfactory explanation of variability within nontraded goods, evidence was presented in support of treating the within-traded-goods variability as exogenous to the Mexican economy.

In general, our results imply that total relative-price variability in Mexico is significantly affected by expected changes in the traded/nontraded price ratio, real money growth, unexpected inflation, and the external variability within traded goods. These findings are consistent with those of previous studies. However, by explicitly considering open-economy specifications, we have identified additional variability effects that operate through the international-trade sector of the economy. Some of these effects may be exogenous to the small open economy.

An important implication of the findings is that there appears to be a mechanism of international transmission of relative-price variability, at least under fixed exchange rates. Thus, a significant positive correlation between the amount of relative-price variability across countries can be expected to emerge from the data. If this is so, then small open economies would have to put up with the welfare implications of 'imported' relative-price variability.[12]

Finally, there are at least two promising directions for future research. First, since domestic relative-price variability may depend closely on foreign variability, it would seem appropriate to study the desirability and feasibility of domestic policies aimed at insulating the economy from such external variability. Second, it would be interesting to extend our analytical framework to the case of a flexible-exchange-rate regime. This would increase the number of countries for which this type of analysis can be performed.

FOOTNOTES

* This paper was prepared at the Center for Latin American Development Studies, Boston University, where both authors are Research Associates. The authors are grateful to S.Freund and Z.Hercowitz for their many helpful suggestions.

1 Interestingly, a relationship such as that found by Parks was also observed by Graham (1930, pp.175-76) in his study of the German hyperinflation.

2 Another relevant study in this context is the one by Hercowitz (1981) who investigates the relationship between unanticipated money growth and relative price variability for the case of the German hyperinflation. His findings show a statistically significant correlation between unanticipated money growth and the degree of price dispersion.

3 Examination of the relationship for expected vs. unexpected inflation is also of interest from the standpoint of current controversies in the macro literature. Models such as Parks' yield the implication that inflation affects relative price variability only when it is unexpected. On the other hand, other models emphasize the existence of differential (across sectors) real costs associated with price and contract adjustment, which may result in non-negligible effects of anticipated inflation on relative price variability (see Sheshinski and Weiss (1978), and Taylor (1980)).

4 Eq.(2) is derived from a standard Divisia price-index formulation.

5 The basic data used are published by the Bank of Mexico (1969, 1977). Nontraded goods comprise construction and housing, transportation and communications, commerce, public services, and other services, and are subdivided into ten sectors.

6 The estimated equations are

$$DP_{Tt} = \underset{(0.013)}{0.022} + \underset{(0.169)}{0.667}\, DP_{Tt-1} \qquad R^2 = 0.394 \quad D.W. = 1.889 \quad h = 0.50$$

$$DP_{NT_t} = \underset{(0.016)}{0.032} + \underset{(0.186)}{0.597}\, DP_{NTt-1} \qquad R^2 = 0.301 \quad D.W. = 2.034 \quad h = -0.26$$

Numbers in parentheses are standard errors of the coefficients; h is Durbin's statistic for testing serial correlation in autoregressive models.

7 This supply function can be rationalized in terms of a specification that assumes that the production function for good i uses all n goods as inputs. See Parks (1978). Alternatively, it can be assumed that while labor demand by sector i depends on the nominal wage deflated by P_i, labor supply is a function of the nominal wage deflated by P^*. Assuming that production varies mainly as a function of labor employment and that the labor market clears, a supply function of this form can be obtained.

8 Hercowitz (1981) includes a similar real-balance effect in his aggregate-demand specification.

9 This assumption was mainly made in order to avoid problems of multicollinearity. Note also that Z_t does not have a straightforward interpretation.

10 The data used here are described in Note 8. In addition, data on money growth were taken from the M_1 series of International Financial Statistics.

11 VP_{US} is the index of relative price variability calculated by Hercowitz (1980) using disaggregation of the wholesale price index. The estimations below used the Cochrane-Orcutt Technique.

12 On the welfare implications of relative-price variability, see Jaffee and Kleiman (1977) and Fernandez (1980). On its effects on real economic variables, see Keynes (1924, pp.35-36), Friedman (1977). Empirical evidence on this issue is presented by Blejer and Leiderman (1980).

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