Inflation and Relative Price Variability

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The paper develops a natural measure of the amount of relative price variability. The variance of relative price change is shown to be correlated with the rate of change in the price level using data for consumer goods in both the Netherlands and the United States. This association has been noted in other data for a variety of countries. Using a multisectoral supply-and-demand framework, the paper goes on to show how changes in relative prices and ultimately the variance of relative price changes are related to supply conditions changes in real income and the amount of unanticipated inflation. The model is used as the basis for an analysis of movements in the prices of consumer goods in the United States for the period 1929–75. The amount of unanticipated inflation (measured as the difference between the actual rate and a time-series predictor) is a more important determinant of relative price variability than the rate of inflation.

I. Introduction

The proposition that changes in real economic variables, in particular the supply of output, are independent of fully anticipated movements in nominal variables has received substantial theoretical support, but attempts to test this natural rate hypothesis have encountered formidable difficulties, and to a large extent its validity remains an open question. This paper offers a straightforward test of the proposition that an unanticipated inflation has an effect on movements in relative prices, an effect that is absent in the case of a fully anticipated inflation.

The paper proposes a convenient measure of the amount of relative price change occurring in an economy or in a subsector of an economy

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during a given period. This measure is found to move in a predictable way in response to certain real forces in the economy, but it is also affected by the amount of unanticipated inflation. This finding thus documents in a new way certain nonneutral effects of unexpected inflation.

In an interesting recent paper, Vining and Elwertowski (1976) present evidence of a relationship between the variance of relative prices and what they call "general price change instability" using the detailed price data that underlie both the wholesale and consumer price indexes. They also show that the shape of the distributions of relative price changes shifts in a systematic way with general price-change instability. The authors fail to provide a precise definition of their notion of general price-change instability[1] or to provide a microeconomic link to explain the transmission of the general instability into the relative price variance. Both of these tasks are addressed in the present paper.

Changes in relative prices occur continuously in response to change in real income, family composition, and many other determinants of demand on the one hand and to changes in technology, resource availability, and other determinants of supply on the other hand. It is widely agreed that to the extent that monetary expansion and the resulting inflation are fully anticipated, they would have no effect on the movements in relative prices and the underlying resource allocation. It has long been recognized by theorists, however, that unanticipated inflation can result in changes in relative prices together with changes in the allocation of resources. Mises, for example, in *The Theory of Money and Credit* (1953), discusses these nonneutral effects of inflation. More recently, explanations of Phillips-curve phenomena in labor markets have relied on the effect of temporary mistakes in suppliers' price expectations brought about by unanticipated inflation. Lucas and Rapping (1969) and Lucas (1973) employ models containing these features. Gordon and Hynes (1969) show how a given change in aggregate demand can produce quite different responses in different markets because of differences in the information content of the change. These authors focus on the behavior of suppliers and show that, depending on the past experience of producers in different sectors, a nominal price change can be interpreted primarily as an increase in relative prices, to which they respond with increased supply, or as a largely inflationary change to which there is little supply response. Finally, Bordo (1976) has rediscovered the interesting early work of Cairnes (1873) and has attempted to retest Cairnes's hypotheses and extend the explanation of how changes in the quantity of money are transmitted into price and quantity changes in different sectors of the economy.

[1] For descriptive purposes, a measure similar to that proposed by Klein (1975), a multiyear moving standard deviation, might serve the authors' purposes.

Section II of the paper proposes a useful way of measuring the overall degree of relative price change in a given period. Annual time-series data on personal consumption expenditure by major type in the Netherlands and in the United States, together with the corresponding price series, are then used to construct a series for this measure of price variability as well as a series that measures the overall rate of inflation for the set of consumption goods. There is a fairly weak but statistically significant association between the measure of price variability and the squared rate of price change, a suggestion that the amount of relative price change increases when the rate of price change increases. Simple theories suggest, however, that it may be the amount of *unanticipated* inflation rather than the rate itself that is important in affecting the amount of relative price change. Section III begins with a brief survey of several explanations of price movements and their effects. Then a simple multiple-market framework is used to explore the determinants of price and quantity changes in a context where suppliers lack complete information about movements in the general price level. The variance of relative price changes is related to changes in real income, to real factors on the supply side, and to a measure of "surprise," that is, the difference between actual and anticipated rate of change in the general price level. A test of this relationship using annual data for the United States from 1929 through 1975 provides confirmation for the hypothesis that although fully anticipated price changes are neutral in their effect, unanticipated changes in the rate of inflation have a significant effect on the amount of relative price change in the economy and result in nonneutral real effects.

II. A Measure of Relative Price Variability

The measures of inflation with which we are most familiar are price indexes, such as the consumer price index (CPI), the wholesale price index (WPI), and the various implicit price deflators from the National Income and Products Accounts of the United States.[2] These common indexes all rely on the Laspreyres formula in their construction, although the weights are revised periodically. These and other indexes attempt to provide a measure of either the average level of prices or the average rate of change of prices.

I introduce the following notation for prices and other related variables. Let p_{it} represent the price of the ith commodity in time period t. Depending on the context, the commodity may be a very narrowly defined commodity on which we can obtain a direct observation of the price, or it may itself be an index associated with a set of closely related commodities.

[2] For an extensive discussion of these and other more specialized indexes together with a consideration of problems associated with their construction, see Triplett (1975).

The rate of change in the ith price between periods $t-1$ and t will be denoted Dp_{it} and is defined as the difference in the natural logarithm of prices in the two periods. That is, $Dp_{it} = \ln p_{it} - \ln p_{it-1}$. An aggregate price level for the set of commodities, $i = 1, \ldots, n$, is denoted P_t, and the rate of change in this index is defined as the weighted average of the rates of change for the individual commodities. This is a standard Divisia price-index formulation that is widely used in econometric work.[3] Thus

$$DP_t = \sum_{i=1}^{n} w_{it}^* Dp_{it},$$

where w_{it}^* is the average expenditure share on the ith commodity in years $t-1$ and t. The weights are nonnegative and sum to one.

A simple measure of the degree of relative price change between period $t-1$ and t is given by the weighted sum of squared deviations of the individual rates of price change around the average. I define this variance measure[4] as

$$VP_t = \sum_{i=1}^{n} w_{it}^* (Dp_{it} - DP_t)^2.$$

The factor $Dp_{it} - DP_t$ is the rate of change in the ith relative price, that is, the logarithmic difference in the relative price p_{it}/P_t. The average rate of change for the relative prices is zero, and VP_t is the variance. Thus VP_t is seen to be a measure of nonproportionality of the price movements, for if all prices change by the same rate, the common DP_t, then the variance measure will be zero; the measure will be larger the more nonproportional the price changes become.

Theil (1967) computed the values of DP_t and VP_t for data constructed by A. P. Barten on per capita consumption and prices in the Netherlands during the years 1921–39 and 1948–64 for 16 classes of consumer goods. Thus it was possible to examine very cheaply the behavior of the measure of relative price variability. Table 1 gives the Dutch time series of DP_t and VP_t. Table 2 includes results of simple regressions of VP_t and DP_t.

[3] See, for example, Jorgenson and Griliches (1967) and Theil (1967, chaps. 5–7).

[4] Theil (1967, p. 155) includes a measure of the variance of the log changes in the individual prices in his decomposition of the variance of the log changes in the expenditure shares. He calls it a "measure of the change in the *price structure*" and notes its properties. Theil's interest there is primarily in demand behavior, and he does not pursue the association studied here. The variance measure of relative price change will be affected by the degree of aggregation of the data. The more disaggregated the data the larger the variance measure will be (see Theil 1967, pp. 162–63). Vining and Elwertowski use an unweighted average in defining both the average rate of price change and the variability of relative prices. The use of weights makes sense from a sampling point of view, for if we were to draw at random a dollar's worth of expenditure from total expenditure, the probability that we would observe price p_{it} is given by the expenditure share w_{it}. In light of the effect of disaggregation on the variance measure, their use of different numbers of commodities in different years presents difficulties.

TABLE 1

THE RATE OF INFLATION, DP_t, AND THE VARIANCE OF RELATIVE PRICE CHANGES, VP_t: THE NETHERLANDS, 1921–63 (Excluding War Years)

Year	DP_t	VP_t
1921–22	−.1019	.004945
1922–23	−.0426	.002010
1923–24	.0026	.001731
1924–25	.0069	.001698
1925–26	−.0475	.002240
1926–27	−.0111	.000874
1927–28	.0052	.000454
1928–29	−.0042	.001302
1929–30	−.0444	.002826
1930–31	−.0651	.003625
1931–32	−.0861	.004626
1932–33	−.0376	.004549
1933–34	−.0153	.001230
1934–35	−.0343	.000399
1935–36	−.0400	.001953
1936–37	.0445	.001770
1937–38	.0215	.001785
1938–39	.0088	.001737
1948–49	.0459	.001950
1949–50	.0803	.003052
1950–51	.1018	.001963
1951–52	.0017	.005671
1952–53	−.0094	.000481
1953–54	.0379	.001548
1954–55	.0153	.000690
1955–56	.0188	.002759
1956–57	.0450	.001445
1957–58	.0145	.001457
1958–59	.0092	.000217
1959–60	.0176	.001547
1960–61	.0169	.000268
1961–62	.0220	.001131
1962–63	.0274	.000586

SOURCE.—Theil 1967, tables 5.7 and 5.8.

Given this body of data, the relationship must account for episodes of both inflation and deflation. This can be done with regressions of the form $VP_t = \alpha + \beta|DP_t| + u_t$, or $VP_t = \alpha + \beta DP_t^2 + u_t$. Alternatively, we can allow for different degrees of response for positive price changes than for negative price changes by the use of regressions of the form $VP_t = \alpha + \beta_+|DP_t^+| + \beta_-|DP_t^-| + u_t$, or $VP_t = \alpha + \beta_+(DP_t^+)^2 + \beta_-(DP_t^-)^2 + u_t$, where DP_t^+ (or DP_t^-) represents the product of DP_t and a dummy variable that takes the value one when DP_t is positive (or negative) and zero otherwise.

The real testing will be done below using U.S. data, but first we can take advantage of the Dutch data for a preliminary look at the behavior of relative price variance. Most of the prewar observations show a decline in the average price level, whereas almost all of the postwar observations show increases in the average price level. There does appear to be an

TABLE 2

Regression Results for the Netherlands, 1921–64, Dependent Variable, VP_t (Excluding War Years)

Constant	$(DP_t)^2$	$(DP_t^+)^2$	$(DP_t^-)^2$	$\bar{R}^2$
.0015 .... (5.85)	.254 (3.31)	...	...	.238
.0015 .... (6.32)	...	.094 (0.97)	.387 (4.33)	.344

Note.—Student t-values are in parentheses.

association between the rate of inflation and the variance in the relative price changes. Somewhat surprisingly, the relationship is stronger in the case of price declines than it is for the price increases. The best-fitting regression includes both a positive and negative branch with different slope coefficients.

These regressions do not attempt to separate the effects of anticipated and unanticipated inflation. Thus the 1951–52 observation seems a major outlier involving the largest observed variance in the sample with almost no change in the average price level. As such it very much weakens the relationship for the positive branch. This observation and several that are less dramatic suggest, however, that the large variance may be associated with unanticipated change in the rate of inflation. The 1951–52 change follows 3 years of rapid and accelerating inflation reaching 10.18 percent in 1950–51. The surprise associated with a change from 10 percent inflation to almost none in 1 year may account for some of the changes in relative prices. Other episodes are suggestive as well. There seem to be cases in which change to a new level of inflation is accompanied by large variance in the change in relative prices, but if the inflation rate remains constant at the new level there is a tendency for the variance to decline. In the following section I shall attempt to formalize the notion of surprise, or unanticipated price change, and relate it to the measure of price variability.

In order to check the primitive notions developed using the Dutch data, we can look at similar data for the United States. They consist of the expenditure and implicit deflator series for "Personal Consumption Expenditure by Major Type." I have used the 12-commodity breakdown annually for the years 1929–75.[5] Table 3 presents the series of DP_t and VP_t for the United States. Table 4 gives several simple regressions based on the entire sample as well as both the prewar and postwar subperiods.

[5] Expenditure and price data are from tables 2.3 and 7.11 of the National Income and Product Accounts of the United States. The data for 1946–75 appear in U.S. Department of Commerce *Survey of Current Business* (1976). For 1929–45 the Office of Business Economics kindly provided me with the unpublished preliminary tables for the recent benchmark revisions.

TABLE 3

THE RATE OF INFLATION, DP_t, AND THE VARIANCE OF RELATIVE PRICE CHANGES, VP_t: UNITED STATES, 1930–75

Year	DP_t	VP_t
1930	−.0286	.001502
1931	−.1084	.004932
1932	−.1246	.004425
1933	−.0383	.003487
1934	.0693	.006971
1935	.0236	.001631
1936	.0112	.000128
1937	.0351	.000201
1938	−.0184	.001747
1939	−.0093	.000193
1940	.0122	.000145
1941	.0677	.000893
1942	.0972	.015121
1943	.0926	.002172
1944	.0579	.001514
1945	.0375	.000689
1946	.0787	.006389
1947	.1018	.001237
1948	.0565	.000373
1949	−.0054	.001235
1950	.0166	.000306
1951	.0647	.000560
1952	.0220	.000545
1953	.0180	.000812
1954	.0095	.000565
1955	.0075	.000449
1956	.0200	.000155
1957	.0329	.000159
1958	.0224	.000309
1959	.0187	.000366
1960	.0188	.000129
1961	.0113	.000055
1962	.0137	.000109
1963	.0152	.000067
1964	.0127	.000052
1965	.0168	.000110
1966	.0276	.000299
1967	.0252	.000129
1968	.0395	.000133
1969	.0440	.000173
1970	.0450	.000258
1971	.0424	.000307
1972	.0349	.000228
1973	.0538	.001155
1974	.1012	.003942
1975	.0741	.000391

SOURCE.—Based on National Income and Product Accounts implicit deflators for personal consumption expenditure by major type of product.

The regression results given in Table 4 again show a significant relationship between the variance in relative prices and the rate of inflation or deflation in all three cases. The regression residuals have much higher variance in the period 1930–45 than in the period 1947–75. The regression relation is, however, considerably stronger for the latter period.

TABLE 4

Regression Results for the United States, 1930–75, Dependent Variable VP_t

Period	Constant	DP_t^2	$\bar{R}^2$
Postwar (1948–75)	.00010 (0.91)	.258 (5.97)	.562
Prewar (1930–41)	.00118 (1.66)	.288 (2.51)	.347
Entire sample excluding war (1930–41, 1948–75)	.00030 (1.40)	.324 (5.97)	.476

Note.—Student t-values are in parentheses.

Up to this point I have undertaken only to present evidence that the amount of relative price variability exhibits some systematic patterns of behavior that may be related to the rate of price change or to a measure of unanticipated price change. Glejser (1965) presents cross-section evidence of a similar relationship. Vining and Elwertowski (1976) find an association between the variance of relative prices and the degree of general price-change variability. Although operationally they provide only a graphical specification for this variable, it corresponds with the notion of variance in the price level (in logs) around its anticipated level. The apparent association between periods of high rates of either inflation or deflation and price-change instability accounts for the possibility of detecting an influence on relative price variance with either variable. In the next section I shall survey various ideas that have been advanced to explain differential price movements and begin to formulate a simple-market model within which we may be able to specify more precisely the factors that could be expected to affect the amount of relative price change.

III. Determinants of Movements in Relative Prices

A. Explanations of the Effects of Monetary Stimulus

A number of distinct theories have been offered to help explain differences in the rate of response of prices to monetary stimulus. The Cairnes (1873) explanation cited by Bordo (1976) emphasizes differences in the underlying supply-and-demand elasticities for various commodities together with a less well-developed notion of differences in the origin and sequence of expenditures arising from the stimulus. Cairnes's predictions seem to depend largely on his beliefs concerning supply elasticities in the short run. Mises's (1953) discussion is quite similar to that of Cairnes, especially in its attention to the origin and propagation of the stimulus. Bordo contributes several additional factors that may help to differentiate fast-responding sectors from slow responders. These factors include the effect of inventories, the degree of market organization, and the degree of nonmarket intervention. He cites the conflicting views with respect to market concentration and adds an interesting hypothesis concerning price predictability and the prevalence of long-term contracts.

The papers of Gordon and Hynes (1969), Klein (1973), and Lucas (1973) emphasize informational aspects of market adjustment. Although these papers are firmly grounded in a market context, they attempt to develop the implications of a world in which the economic agents, particularly suppliers, lack complete information about relative prices. Agents must make decisions concerning output levels on the basis of information about the general price level that is based on its past behavior but without knowledge about its current level. In this context it is possible for suppliers to be fooled temporarily into believing that a given price movement reflects a change in relative prices. Although they cannot be fooled continuously, the temporary error can produce real output effects.

In a recent paper Sheshinski and Weiss (1976) consider a model in which there are real costs associated with nominal price changes and study the effects of a steady anticipated inflation on the frequency and size of price changes. In their model there is no uncertainty about the rate of inflation, but because of the adjustment costs, price adjustments occur at discrete intervals. During the interval between price changes, the real price of the good declines. One apparent implication of their model is that "if the timing of firms' price adjustments is independent, then we would observe a variance of price changes across products or firms which increases with the rate of inflation" (p. 31). Differences among firms or industries in the costs of price adjustment could lead to different rates of response to monetary stimulus.

B. A Multimarket Model of Relative Price Change

The multimarket model presented here provides a framework for understanding how the basic determinants of supply and demand combine with expectations about the rate of inflation to affect the amount of relative price change resulting from changes in nominal income. In order to focus on the effect of price expectations, the basic supply-and-demand apparatus has been kept as simple as possible.

The quantity of the ith commodity supplied in period t is given by the supply function

$$\ln q_{it} = \alpha_i + \beta_i \ln (p_{it}/P_t^*) + \gamma_i t, \qquad i = 1, \dots, n, \tag{1}$$

where α_i, β_i, and γ_i are supply function parameters. The parameter β_i is the supply elasticity with respect to changes in real price of good i, and γ_i is the trend rate of growth of output; P_t^* denotes the anticipated level of the general price index.[6] The quantity demanded is given by the function

$$\ln q_{it} = \eta_{ii} \ln p_{it} + \eta_{io} \ln m_t + \theta_i, \tag{2}$$

where η_{ii} is the own-price elasticity and η_{io} the income elasticity of demand, and where m_t denotes nominal income. It will be convenient to ignore cross-price effects but to preserve homogeneity, so that we will assume $\eta_{ii} + \eta_{io} = 0$.[7]

The supply-and-demand functions can be written in terms of the logarithmic differences. Using the notation introduced in Section II, we can write

$$Dq_{it} = \beta_i(Dp_{it} - DP_t^*) + \gamma_i, \tag{1'}$$

$$Dq_{it} = \eta_{ii}(Dp_{it} - Dm_t). \tag{2'}$$

The assumption that markets clear, given the rate of change in income and given the suppliers' expectations concerning the general rate of inflation, results in a reduced form

$$Dp_{it} = \frac{1}{\beta_i - \eta_{ii}} (-\eta_{ii}Dm_t + \beta_i DP_t^* - \gamma_i), \tag{3}$$

$$Dq_{it} = \frac{1}{\beta_i - \eta_{ii}} (-\beta_i\eta_{ii}Dm_t + \beta_i\eta_{ii}DP_t^* - \eta_{ii}\gamma_i). \tag{4}$$

Assuming normally shaped supply-and-demand functions with $\beta_i > 0$, $\eta_{ii} < 0$, the model provides standard predictions concerning the effect of income change. A rise in income will increase both price and quantity, with the amounts depending on the elasticities of supply and demand. The Cairnes predictions are based on the notion that the smaller the elasticity of supply, the more the price change and the less the quantity change associated with a given increase in nominal income. In addition, we see that the anticipated rate of inflation affects the price and quantity response as we would expect. The higher the anticipated inflation, the

[6] This supply function can be rationalized by assuming that the production function for the ith good involves as inputs all n goods. The supply function of a profit-maximizing firm can then be written in terms of the ith good's output price and the set of input prices. For certain choices of production function, the input prices will appear as an index. In the case of the Cobb-Douglas function, for example, the input prices will appear as a geometric average of the input prices. It would be straightforward to follow Lucas (1973) and assume that suppliers formulated their price expectations by combining their view of the prior distribution of DP_t with current information about price changes in their own sector. The resulting formulation would give rise to an equation like (8) for predicting relative price variance; therefore, the extra calculations are not carried along.

[7] The assumption that cross-price effects are negligible greatly simplifies the following derivations. If cross-price derivatives are introduced into the demand equations (2), the reduced-form solution (3) and (4), as well as the subsequent decompositions of the change in relative prices (6) and the variance of relative price changes (8), all retain their basic form. The coefficients, however, all become more complicated and include the various cross effects.

higher the resulting price change and the lower the quantity change. Finally, a positive trend in supply reduces the price change and increases the quantity change.

Recall that the actual average rate of price change is defined as

$$DP_t = \sum_{i=1}^{n} w_{it}^{*} Dp_{it}. \tag{5}$$

Using (3) we can write the actual change in the ith relative price as

$$Dp_{it} - DP_t = \frac{1}{\beta_i - \eta_{ii}} [-\eta_{ii}(Dm_t - DP_t) - \beta_i(DP_t - DP_t^{*}) - \gamma_i]. \tag{6}$$

As defined above, the variance of the relative price changes is

$$VP_t = \sum_{i=1}^{n} w_{it}^{*}(Dp_{it} - DP_t)^2. \tag{7}$$

From (6) it can be seen that the change in relative prices can be decomposed into components that involve the basic supply-and-demand parameters together with terms that involve the rate of change in real income $(Dm_t - DP_t)$ and the unanticipated inflation rate $(DP_t - DP_t^{*})$.

Combining the expression for changes in relative prices (6) with the definition of the variance (7), we can obtain the following expression for the determinants of the relative price variance:

$$\begin{aligned} VP_t = A_0 &+ A_1(Dm_t - DP_t)^2 + A_2(DP_t - DP_t^{*})^2 \\ &+ A_3(Dm_t - DP_t)(DP_t - DP_t^{*}) + A_4(Dm_t - DP_t) \\ &+ A_5(DP_t - DP_t^{*}), \end{aligned} \tag{8}$$

where the coefficients, A_i, are given by

$$A_0 = \sum_i \frac{w_{it}^{*}\gamma_i^2}{(\beta_i - \eta_{ii})^2},$$

$$A_1 = \sum_i \frac{w_{it}^{*}\eta_{ii}^2}{(\beta_i - \eta_{ii})^2},$$

$$A_2 = \sum_i \frac{w_{it}^{*}\beta_i^2}{(\beta_i - \eta_{ii})^2},$$

$$A_3 = 2\sum_i \frac{w_{it}^{*}\beta_i\eta_{ii}}{(\beta_i - \eta_{ii})^2},$$

$$A_4 = 2\sum_i \frac{w_{it}^{*}\gamma_i\eta_{ii}}{(\beta_i - \eta_{ii})^2},$$

$$A_5 = 2\sum_i \frac{w_{it}^{*}\gamma_i\beta_i}{(\beta_i - \eta_{ii})^2}.$$

The relative price variance can thus be written as a combination of the effects of changes in real income, unanticipated inflation, and supply shifts where the weights involve underlying supply-and-demand parameters in a systematic way.

If the coefficients are stable, we may test the formulation by using (8) as the basis for a regression for predicting movements in VP_t.

Although the coefficients are compounded from supply-and-demand parameters from all the sectors, the standard assumptions imply that A_1 and A_2 should be positive and A_3 negative. The supply trend terms may have either sign, although there is a presumption that they are positive for most of the period covered. If they are all positive, then the constant A_0 as well as A_5 would be positive and A_4 negative. In order to carry out a test of the formulation, however, we must provide an operational definition for the unanticipated rate of inflation.

C. A Measure of Unanticipated Inflation and Regression: Results for the United States

So far I have not specified a mechanism for generating the anticipated rate of inflation, DP_t^*, or the surprise or unanticipated rate, $DP_t - DP_t^*$. There are many possible formulations that we might try. The one that I have examined here is based on time-series analysis of the inflation-rate series DP_t alone. Implicitly I have also assumed that suppliers in different sectors use the same information and arrive at the same prediction. In contrast, Lucas's (1973) model permits suppliers in a given sector to use current price information from their own but not other markets in developing a price predictor.

Analysis of the inflation-rate series (DP_t), using methods and programs outlined in Nelson (1973), shows that the series itself is nonstationary but that the differenced series ($DP_t - DP_{t-1}$) appears random. A particularly simple time-series model of the form $DP_t = DP_{t-1} + \mu + \varepsilon_t$ is suggested, and the best predictor for DP_t is, therefore, $D\hat{P}_t = \hat{\mu} + DP_{t-1}$. Using this predictor as the anticipated rate of inflation, DP_t^*, we also get a simple measure of surprise or unanticipated inflation for use in regressions based on expression (8).

Table 5 presents the regression results based on (8) for the United States during the postwar, prewar, and combined periods. The regression equations based on the model explain movements in the variance of relative price changes rather well in all three samples. In each case, real income and the measure of surprise or unanticipated inflation enter significantly in both squared and linear form, and their coefficients have the predicted sign. The only sign reversals occur in the coefficient of the real income × surprise interaction variable; and in both cases where a

TABLE 5

THE VARIANCE OF RELATIVE PRICE CHANGES: REGRESSION RESULTS FOR THE UNITED STATES, 1930–75, BASED ON EQUATION (8)

Coefficient and Variable	Postwar (1948–75)	Prewar (1930–41)	Combined (1930–41, 1948–75)
A_0 (constant)	.0011 (3.17)	−.0002 (0.34)	.0005 (2.86)
$A_1(Dm_t - DP_t)^2$	.747 (2.70)	.268 (2.95)	.209 (4.12)
$A_2(DP_t - DP_t^*)^2$	.243 (2.21)	.582 (6.55)	.465 (9.67)
$A_3(Dm_t - DP_t)(DP_t - DP_t^*)$	−.529 (1.70)	.103 (0.86)	.089 (1.03)
$A_4(Dm_t - DP_t)$	−.054 (2.68)	−.015 (2.93)	−.019 (5.58)
$A_5(DP_t - DP_t^*)$	.015 (2.04)	.008 (1.71)	.006 (1.92)
D-W	1.82	2.33	1.49
$\bar{R}^2$	.667	.896	.850

NOTE.—Student t-values are in parentheses.

reversal occurs, the coefficient is insignificant. The results are generally consistent with supply trends that are positive.

The results show clearly the distinct effect of unanticipated inflation on changes in relative prices. In this form of the model an inflation that is fully anticipated has no effect on the changes in relative prices that occur as the result of real factors affecting supply and demand, whereas unanticipated inflation increases the amount of relative price variance. The relative size of the contributions of real income and unanticipated inflation to the explanation of the price-change variance is different in the prewar and postwar periods. In the earlier period, the unanticipated inflation variable has a stronger effect on the variance than does a change in real income, whereas in the later period the real income variable becomes more important.

Although the market model developed above suggests no separate role for the actual (or anticipated) rate of inflation in determining relative price variance, several previous studies have suggested that the rate of inflation should in fact appear. We have already shown directly that the rate of inflation is positively associated with variance. Vining and Elwertowski (1976) show a relationship between price-variance and price-change instability, but they also note that their measure of instability is correlated with the rate of inflation. Glejser's (1965) cross-country analysis finds a relationship between relative price variance and the average rate of inflation. An obvious question that arises is whether the rate

TABLE 6

THE VARIANCE OF RELATIVE PRICE CHANGES: REGRESSION RESULTS FOR THE UNITED STATES, 1930–75, BASED ON EQUATION (8) WITH THE ADDITION OF $A_6\ DP_t^2$

Coefficient and Variable	Postwar (1948–75)	Prewar (1930–41)	Combined (1930–41, 1948–75)
A_0 (constant)	.0006	−.0002	.0001
	(1.31)	(0.34)	(0.54)
$A_1(Dm_t - DP_t)^2$	.501	.199	.125
	(1.64)	(1.28)	(2.41)
$A_2(DP_t - DP_t^*)^2$	.191	.555	.444
	(1.72)	(5.26)	(10.28)
$A_3(Dm_t - DP_t)(DP_t - DP_t^*)$	−.527	.030	−.005
	(1.76)	(0.17)	(0.06)
$A_4(Dm_t - DP_t)$	−.032	−.011	−.008
	(1.37)	(1.28)	(1.86)
$A_5(DP_t - DP_t^*)$	.013	.008	.006
	(1.72)	(1.69)	(2.09)
$A_6(DP_t^2)$	.107	.081	.148
	(1.63)	(0.58)	(3.16)
D-W	1.99	1.96	1.44
$\bar{R}^2$	.691	.881	.882

NOTE.—Student t-values are in parentheses.

of inflation has an effect on variance when entered together with the real variables and the unanticipated inflation variable suggested by the model. The natural rate hypothesis, at least in its pure form, denies such a role for the rate of inflation, but several theoretical considerations provide a basis for its inclusion. As mentioned above, Sheshinski and Weiss (1976) show that if there are real resource costs associated with price changes and if these costs vary across sectors, then there will be an association between the variance of price changes and the (fully anticipated) rate of inflation. Furthermore, the duration of contracts would be affected by different rates of inflation, and the choice would in turn influence the movement of relative prices. Although these factors have not been integrated fully into the market model, a simple empirical experiment will allow some preliminary judgment about the possible importance of the rate of inflation as a separate variable.

Table 6 presents regression results based on an equation that includes the variables that appear in equation (8) but with the square of the rate of inflation, DP_t^2, added as well.

The $\bar{R}^2$ increases slightly with the addition of DP_t^2 in the postwar period and in the combined prewar and postwar sample, but it decreases slightly in the prewar sample. The DP_t^2 variable is not individually significant in either separate period, although it is significant in the combined sample. The significance and size of the coefficients of both real and surprise variables are reduced by the inclusion of the rate of inflation. The empirical problem of determining whether there is a separate role for the rate of inflation is complicated by the apparent association between the rate of inflation and the movements in both real income and in the unanticipated inflation.[8] Thus although the simpler market model presents a clearer picture of the determinants of relative price variance, at least for the body of data considered here, there is some evidence supporting a separate effect for the rate of inflation. Its magnitude is much smaller, however, than that of unanticipated inflation, from one-third to one-sixth the size, depending on the sample.

IV. Further Implications of the Results

The empirical finding of Vining and Elwertowski (1976) of a positive relationship between the variance of relative price changes and the amount of price-change instability, for which they offer no explanation, can be interpreted rather directly in terms of the model presented above. I have found a linear relationship between the variance of relative price change and the squared measure of surprise or unanticipated inflation. Thus $VP_t = \alpha_0 + A_2(DP_t - DP_t^*)^2$, where α_0 represents the remaining terms of expression (8). If we were to average over the observations of a particular subperiod or over observations for a particular country, we would have a linear relationship between the variance of relative price changes for the period or country and the variance of price changes around their expected rate. This latter variance is a more precise representation of the Vining and Elwertowski notion of general price-change instability. It also corresponds with Lucas's (1973) notion of the variance of the prior distribution of the (changes in) the price level. The model and the regression results thus provide a foundation for the descriptive finding of Vining and Elwertowski and at the same time point to a misspecification in Lucas's formulation that is also noted by Vining and Elwertowski.

Lucas specifies the price structure in the following way: The nominal price change for the *i*th good can be decomposed into the relative price change plus the change in the general price level. (Lucas formulates his model in terms of levels, but the same argument can be made in terms of rates of change.) He then asserts that the relative price change and the general price-level change are both normally distributed random variables, independent and each having constant variances. The findings presented here and those of Vining and Elwertowski show that there is

[8] Logue and Willett (1976) find cross-section evidence of a relation between the rate and variability of inflation.

in fact a relationship between the two variances. Some forms of relationship might present difficulties for Lucas's prediction and empirical test, but the linear relationship preserves the sign of the prediction on which his test rests. Thus although misspecified in terms of the variance relationship, the Lucas model's key prediction is robust to the generalization implied by the relationship between the variances.

Certain observations that Glejser (1965) found anomalous can be explained in a straightforward way by the present model. In looking at the data for two periods for France, 1950–56 and 1950–60, he finds it surprising that "the magnitude of the relative price changes was smaller in the latter period albeit the inflation rate was larger and the productivity increase about the same" (p. 77). Although there seems to be a tendency for the variance of the rate of inflation around its expected level to increase with the rate of inflation, the later period in France seems to have been one in which the higher rate of inflation also involved greater predictability. The reduction in the amount of unanticipated inflation predicts lower variance in the relative price changes using the relationship implied by the model.

It would be useful to extend the present study in several different directions. Is the relationship found here affected by the degree of commodity aggregation and by the length of the time period used? The Vining and Elwertowski findings go part of the way in answering the question for a broader array of commodities, but it would be useful to try to extend the regression results to the larger sample of prices. It would be desirable to introduce expenditure weights and to try to maintain a consistent set of commodities for such an exercise.

Verifying the relationship for data from countries other than the United States would also seem useful. The Glejser study, although interesting for the empirical relationship he finds between variance and the rate of inflation, does not attempt to test any satisfactory explanation of the relationship, and from the point of view of the present findings, the relationship that he finds is not the one of primary importance but rather stems from the apparent association between the rate of inflation and the general price-change instability.

It would also be interesting to test alternative ideas about the formation of expectations. The rather naive time-series predictor used here might be replaced by an interest-rate predictor. Some experiments along these lines are currently under way.

Finally, it would be useful to have a means for evaluating the costs and consequences of the increased price variability associated with unpredicted inflation. Jaffee and Kleiman (1976) attempt to develop a welfare measure of the costs of uneven inflation that represents a useful start on this problem.

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