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Inflation and Relative Price Variability in the Short and Long Run: New Evidence from the United States

IT IS NOW ESTABLISHED WISDOM that relative prices are not independent of the level of inflation. Studies by, for example, Mills (1927), Vining and Elwertowski (1976), Parks (1978), Fischer (1981), or more recently, Debelle and Lamont (1995) have documented a positive relationship between inflation and relative prices in the United States. Similar findings have been documented in an international context by Hercowitz (1981), using data from West Germany, Domberger (1987), for the United Kingdom, and Lach and Tsiddon (1992, 1993) for Israel. More recently Reinsdorf (1994), using data from the disinflationary early 1980s, finds a *negative* association between inflation and relative prices, suggesting possible nonlinearities in the relationship.

As emphasized by Reinsdorf, the issue is important if we are to add to our understanding of the inflation transmission mechanism, about how markets respond to inflationary shocks, and more generally, about the welfare costs of inflation. Models predicting a relative price–inflation linkage can be found in Lucas (1973), Barro (1976), Sheshinki and Weiss (1977), Cuikerman (1982, 1984), Cecchetti (1985), Benabou (1988), Tommasi (1992), and, Van Hoomissen (1988). Unfortunately, most of the theoretical models' predictions with respect to the inflation–relative price linkage are observationally equivalent. This equivalence hinders the interpretation of existing empirical findings. Ball and Mankiw (1995) are unique in the sense they argue the causation runs from relative prices to inflation, and they pro-

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vide some support for this view in the form of tests of their model's predictions concerning higher-order moments of the distribution of relative price changes.

Other obstacles to interpreting the existing empirical results can be categorized as data or econometric. First, Danziger (1987) notes that aggregation tends to mask the degree of relative price variability in the data, so that in fact the linkage may be stronger than commonly supposed. Alternatively, the observed variability may be due to product changes within the price indexes themselves. To the extent price indexes change to composition changes, the effect of inflation on relative price variability may be overestimated. Secondly, aggregation increases simultaneity biases; for example, an unobserved variable may be driving both inflation and relative price variability.[1] Third, Hartman (1991) suggests many existing results may be "definitional artifacts" resulting from the particular econometric specification chosen. Using a simple model of price changes, Hartman demonstrates that in the most common econometric specification (that is, a linear regression of relative price change variability on inflation and inflation squared or the absolute value of inflation), the OLS estimates will be positive as found in existing empirical studies. The result depends both on the assumed model of price change and the existence of both inflation and inflation squared terms in the regression. Additionally, many models make a distinction between expected and unexpected inflation, which necessitates a decomposition. Finally, Bomberger and Makinen (1993) argue some existing empirical findings may be due to important omitted variables from the econometric specifications used. In support of their contention, they re-do the estimations in Parks (1978) and find the results depend crucially on a single year, 1974.[2]

This paper has two primary objectives. First, the relationship between relative prices and actual inflation is studied using a data set and approach that, when combined, allow many of these empirical problems to be overcome.[3] Specifically, this study examines a panel of prices that allows dispersion to be studied at the commodity level from both the relative price and relative inflation perspectives. With this data set the issues of simultaneity and changes in price indexes due to changes in the composition of the bundle are mitigated. The estimation is done in a panel context where aggregate shocks which potentially affect both inflation and dispersion can be controlled for. We also implement a new test (utilizing the cross-sectional/time series detail of the data) of models relying on a relative/aggregate confusion motivation for the linkage between inflation and relative prices. The second objective of this study is to present evidence on the persistence of the relative price–inflation interaction. This question, that is, does a change in relative prices resulting from an inflammation shock diminish over time, is especially salient when considering the welfare implications of inflation.

1. See Lach and Tsiddon (1993) for another example focusing on the effects of aggregation.

2. Possible nonlinearities in the relationship, for example, Tsiddon (1993), Bertola and Caballero (1990), further hinder interpretation. I thank a reviewer for bringing this to my attention.

3. As pointed out by a referee, focusing on actual inflation camouflages some of the distinctions between various models predicting a relationship. For example, the Lucas (1973) model makes a prediction only about the effects of unexpected inflation. While this is true, focusing on actual inflation actually biases against finding a significant effect.

The main findings of this study are as follows: (1) Higher inflation (measured either at the product level or at the city level) is associated with greater cross-sectional dispersion of relative prices and of relative rates of inflation, both across cities and across products. These results do not depend on definitional artifacts (as in Hartman [1991], outliers (as in Bomberger and Makinen [1993]), or on aggregate shocks, or seasonal factors, that have not been explicitly controlled for in other studies; (2) As the ratio of aggregate to cross-sectional shocks (either across cities or products) rises, the strength of the dispersion/inflation linkage diminishes, thus supporting an implication of relative/aggregate confusion models; (3) The four time series of dispersion measures examined in this study all exhibit mean reversion, as do the measures of inflation. This implies that interactions among these variables are necessarily of a short-term nature. Indeed, examining that part of dispersion *not* associated with inflation yields virtually identical decay rates as the raw dispersion series; and finally, (4) an examination of the impulse response functions of these four dispersion measures provides additional evidence that the effect is short lived. The evidence suggests that much of the contemporaneous correlation between inflation and relative prices can be accounted for by the autoregressive properties the two series exhibit.

The outline for the remainder of this paper is as follows. Section 1 describes the data set and presents some summary information. Section 2 motivates the measures of dispersion to be examined and presents the first set of results on the inflation–relative price linkage. The persistence of these effects is studied in section 3, and section 4 concludes.

1. THE DATA SET

The price data have been collected from data published by the American Chamber of Commerce Researchers Association (ACCRA). The data are included in ACCRA's quarterly publication, *Cost of Living Index* (hereafter, *Index*). Each quarterly issue of the *Index* contains comparative price data for a large sample of urban U.S. areas, and a cost of living index computed by ACCRA from these price data. This study uses only the price data from which ACCRA's cost of living index is constructed.

Over time, ACCRA's sample of cities and commodities varies. At the beginning of the sample period (1975.1) there were 166 cities and 44 items priced. The number of cities steadily increased to 297 in 1992.4; however, each report contains a distinct sample of cities. The sample used in this study consists of 48 cities which appeared in roughly 90 percent of the quarterly surveys.

The goods and services sampled, however, are much less variable, though there have been additions to and subtractions from the list. As noted, in 1975.1 ACCRA's sample consisted of forty-four goods and services. For this study, fourteen of these original items are dropped since some ultimately dropped out of ACCRA's sample and some, for example, public utilities, were clearly regulated by public authorities. Another two commodities were omitted because they were arguably not stan-

TABLE 1

PERCENTAGES OF NEGATIVE AND POSITIVE PRICE CHANGES

	Nominal		Real
Goods	< 0	> 0	< 0
Grocery Items:			
Steak	43.9	53.0	50.9
Hamburger	43.7	51.3	52.2
Bacon	46.0	50.7	52.0
Whole Chicken	46.7	48.5	52.9
Milk	32.9	50.5	60.2
Eggs	45.6	49.4	52.8
Margarine	44.3	46.2	55.6
Potatoes	45.0	52.4	49.4
Bananas	43.6	48.3	52.5
Lettuce	44.9	50.7	49.8
Bread	42.4	48.2	53.2
Cigarettes	16.7	67.5	51.8
Coffee	47.2	47.6	57.6
Sweet Peas	41.8	47.0	53.5
Tomatoes	40.3	45.4	56.6
Peaches	34.7	53.9	52.8
Tissue	39.2	48.7	54.8
Washing Powder	37.6	54.2	53.2
Shortening	44.1	47.8	60.4
Orange juice	39.7	47.7	55.3
Baby food	21.0	36.3	62.7
Soft Drink	40.4	46.2	56.8
Nongrocery Items:			
Gasoline	38.4	54.8	52.9
Hospital Room	8.7	52.4	55.6
Doctor	18.7	42.2	61.1
Dentist	22.8	40.8	62.1
Man's Haircut	18.9	39.2	64.2
Dry Cleaning	16.5	51.6	58.6
Appliance Repair	22.7	42.9	64.0
Movie	12.0	27.3	73.5
Bowling	14.3	36.4	66.9
Liquor	25.1	41.8	69.3

dardized products: for example, the monthly rent on a two-bedroom apartment, and the monthly payment (principal and interest) on a new house with eighteen hundred square feet of living area. Four additional commodities were subsequently added to the sample bringing the total to thirty-two standardized commodities, predominantly grocery items, that appeared in the quarterly publications. Some commodities did, however, change over the sample, typically as a result of a change in manufacture packaging. This change was accounted for by assigning a missing value to the last quarter prior to the change.

The reported prices were obtained by ACCRA as an average over a small number of sellers in the city, on the Thursday, Friday, or Saturday of the first week of each quarter. Unless taxes are included in the quoted price they are excluded, and coupons and other discounts are not valid unless provided and available at the point of purchase. Since 1982, the reported price of any item in the ACCRA publication represents the average over a minimum of five and a maximum of ten sellers. Because ACCRA is focusing on midlevel managers, the prices represent those occurring in areas within the city where these managers are likely to buy the item, and the sellers are sampled based on market share as determined by ACCRA's representatives.

The actual data collection is done by the local chamber of commerce staff or volunteers for the chamber. Explicit instructions and data forms are provided for each data collector. Some prices are obtained by phone and usually the respondents do not know it is for a survey. Once collected, the data is sent to one of nine different regional coordinators for checking. Finally, the data is sent to Houston where it is transferred to computer and subjected to both computer and visual checks for outliers. Publication occurs approximately five and one-half months after the original data are collected.

The cities and commodities chosen are described in Appendix Tables 1 and 2. Table A1 highlights the fact that the coverage in terms of cities is quite wide, with twenty-seven states represented. In terms of commodities, twenty-two are grocery items while the remaining ten add considerable diversity in coverage and are mostly services. Table A2 gives more explicit definitions of the commodities. For this study all data were input by hand from ACCRA's *Index* and have been extensively checked for errors. Table 1 presents some summary information on the price flexibility pattern in the data. The first two columns show that in terms of the number of price changes, prices do not always rise; indeed, column 3 shows that in real terms (relative to the CPI), the chance of a price decline is about equal to that of a price rise in any given period. In addition to differences in price flexibility, the table also demonstrates the importance of idiosyncratic shocks among these goods and services, even at relatively low levels of inflation.

2. INFLATION AND RELATIVE PRICE VARIABILITY

This section discusses the variability measures examined, the empirical framework adopted, and the estimation results. In the next section the persistence of these effects is discussed.

The structure of the data set enables variability to be measured relative to a product average, or relative to an average of the cities, and it is possible to consider relative price variability and relative price change, or inflation, variability. Consequently, four variability measures are discussed corresponding to two measures of relative price variability and two measures of relative inflation variability. Relative inflation variability is what is most commonly examined in empirical studies in this area since they lack actual price data; however, as noted by Beaulieu and Mattey (1994), relative price variability is more relevant from a theoretical point of view.

Let P_{ijt} be the price of the ith commodity in the jth city at time t, and $\bar{P}_{it}$ be the average price across cities. Then define the first measure of dispersion, V_{1it}, to be the sum of the squared deviations of relative prices (R_{ijt}) around the average (unweighted) relative price for product i, ($\bar{R}_{it}$). This yields

$$V_{1it} = \left[\frac{1}{n-1} \sum_{j=1}^{n} (R_{ijt} - \overline{R}_{it})^2 \right]^{.5}, \quad (1)$$

that is, $R_{ijt} = \ln(P_{ijt}/\overline{P}_{it})$, and, $\overline{R}_{it} = 1/\tilde{n}\Sigma_{j=1}^{\tilde{n}} R_{ijt}$, where there are $\tilde{n}$ non-missing observations, with a maximum number of cities (n) equal to 48.[4] Similarily, taking the sum of the squared deviations of R_{ijt} over the m commodities, around city averages, yields a second measure of dispersion.

$$V_{2jt} = \left[\sum_{j=1}^{m} w_i(R_{ijt} - \overline{R}_{jt})^2 \right]^{.5} \quad (2)$$

where $\overline{R}_{jt}$ is the average of the R_{ijt} across the m products within a given city, that is, $\overline{R}_{jt} = 1/\tilde{m}\ \Sigma_{i=1}^{\tilde{m}} R_{ijt}$, and the average is taken over the $\tilde{m}$ non-missing observations, with a maximum number of products equal to 32. For this study, all prices have been normalized to index form as in the U.S. consumer price index. That is, the city price index is computes as $P_{jt} = \Sigma_{i=1}^{n} w_i P_{ijt}$, where the commodity weights (w_i) were taken from the *BLS Handbook of Methods* to correspond to those used in the U.S. consumer price index, and $\Sigma_{i=1}^{n} w_i = 1$.

Thus V_{1it} is a measure of the within-product relative price variability and V_{2jt} measures the within-city variability. If we believe that resources are more mobile within a given city, then distortion at the city level (V_{2jt}) is the appropriate indicator of potential real resource misallocation. V_{1it} measures the dispersion of individual prices (holding the product constant across cities) around the overall average of that product. If specific (but mobile) capital characterizes the production process, then this measure best reflects the potential for inflation to affect resource allocation. Intuitively, this latter possibility seems more relevant to the short run.

As noted above, most empirical studies have focused on the variability of relative inflation. Hence, for comparison, the third measure of dispersion is

$$V_{3it} = \left[\frac{1}{n-1} \sum_{j=1}^{n} (\Pi_{ijt} - \overline{\Pi}_{it})^2 \right]^{.5} \quad (3)$$

where the summation is over the n cities in the sample. Π_{ijt} = the rate of change in the price of the ith product in city j, and $\overline{\Pi}_{it}$ is the average rate of change in product i's price (Π_{ijt}) over all – (non-missing) cities. Finally, define

$$V_{4jt} = \left[\sum_{i=1}^{m} \mathrm{w}_i(\Pi_{ijt} - \overline{\Pi}_{jt})^2 \right]^{.5} \quad (4)$$

4. Clearly other weighting schemes (for example, economic weight) could be proposed. The unweighted average across cities is appealing by virtue of its simplicity.

TABLE 2

SUMMARY STATISTICS

	V_{1it}	V_{3it}	Inflation
Number of Observations	2289	2234	2283
Number of Cross-Sections	32	32	32
Cross-sectional Mean	2.50	2.43	.0006
Cross-sectional S.D.	.403	.551	.005
Skewness	−.787	−1.02	.337
Kurtosis	.824	1.47	.525
	V_{2jt}	V_{4jt}	Inflation
Number of Observations	3070	2917	2955
Number of Cross-sections	48	48	48
Cross-sectional Mean	2.13	2.29	.0012
Cross-sectional S.D.	.140	.356	.117
Skewness	.342	.104	−8.10
Kurtosis	2.48	1.23	67.37

NOTES: all variability measures are in logs, and inflation is defined as the first difference in the log of the average price.

where Π_{jt} is the rate of inflation within a given city. The interpretations of V_{3it} and V_{4jt} are similar to those for V_{1it} and V_{2jt}. That is, V_{3it} measures the dispersion of inflation rates, holding the product constant around the overall product inflation rate for the product. This measure corresponds most closely to that in Lach and Tsiddon (1992). V_{4jt} measures the dispersion of product-level inflation rates in a given city around the average for that city. This measure of variability is similar to that used in Parks (1978) except that here the dispersion is measured across actual goods and services rather than across broad subaggregates of the Personal Consumption Deflator.

Table 2 presents some summary statistics on the variability and inflation measures studied. In the top panel of the table measures V_{1it} and V_{3it} are grouped together to emphasize that the variability is measured across products. Similarly, in the lower panel measures V_{2jt} and V_{4jt} are grouped together since there are forty-eight cross-sections (cities) over which dispersion is calculated. From the table, two things are worth mentioning. First, note the large number of observations. These sample sizes encourage a belief that the inflation–relative price linkage can be estimated with a high degree of precision. Second, note that the mean and cross sectional standard deviation of both variability and inflation are comparable, though variability is slightly larger (both mean and standard deviation) when measured across products, and inflation is slightly higher and more variable when measured across cities.

Columns 1–4 of Table 3 present results from a fixed-effects regression model of the form:

$$V_t = \sum_{i=1}^{n} \lambda_i + \sum_{t=1}^{68} \eta_t + \sum_{k=1}^{3} \theta_k + \beta \,\overline{\Pi}_t + u_t \tag{5}$$

where λ_i is a product/city effect and n is the number of cross-sections (32 for V_{1it} and V_{3it}, and 48 for V_{2jt} and V_{4jt}), θ_k are quarterly dummies, η_t are time dummies. Π_t is average (across cities, or products) inflation, and u_t is a regression disturbance.

TABLE 3

THE RELATION BETWEEN INFLATION AND DISPERSION

	Dispersion Measure			
	V_{1it}	V_{2jt}	V_{3it}	V_{4jt}
Log Inflation	.3028 (.0538)	.0966 (.0481)	.7751 (.0908)	2.060 (.2470)
Quarterly dummies	yes	yes	yes	yes
Commodity dummies	yes	no	yes	no
Time dummies	yes	yes	yes	yes
City dummies	no	yes	no	yes
Adjusted R-Squared	.63	.66	.11	.31
Std. Error of regression	.2811	.2300	.3420	.2777
Number of observations	2281	2951	2232	2915

NOTES. Heteroskedasticity consistent standard errors in parenthesis All variability measures are in logs
The individual regressions run were of the form:

$$\log V_t = \sum_{i=1}^{n} \lambda_i + \sum_{t=1}^{68} \eta_t + \sum_{k=1}^{3} \theta_k + \beta\,\overline{\Pi}_t + u_t$$

where λ_i is a product effect, n is the number of cross-sections (32 for V_{1it} and V_{3it}, and 48 for $V_{2jt}$and V_{4jt}), η_t are time dummies, θ_k are quarterly dummmies, $\overline{\Pi}_t$ is average (across cities, or products) inflation, and u_t is a regression disturbance

This specification allows effects that are specific to particular goods (for example, changes in product-specific taxes, or systematic measurement differences), cities, seasons (quarters), and aggregate macroeconomic effects to be controlled for explicitly. Failure to control for these effects when they are correlated with the included regressors implies OLS, or GLS, estimates are biased and inconsistent (for example, Hausman and Taylor 1981). Thus, this specification greatly increases the amount of information obtained relative to univariate estimations.[5] Similarly, it is possible to infer the effects of aggregation on the estimates by comparing estimates here with typical aggregate results.

The conclusion from Table 3 is that relative price variability and relative inflation variability are positively associated with inflation. However, looking across the columns it is apparent that the effect varies widely according to how variability is measured. Secondly, the inflation-variability linkage appears much stronger for relative inflation variability than for relative price variability, though the relatively low explanatory power of the regression suggests some important omitted effects. These results support those of earlier studies which find a positive relationship, and are opposite those in Reinsdorf (1994), though Reinsdorf's results may be due to his early 1980s' sample period. Since the relationship holds for all four variability measures and since aggregate, seasonal, and product/city effects have been controlled for, these results demonstrate the generality of the inflation-variability linkage. That is, it cannot be claimed the results depend on outliers, as in the case of the Parks (1978) results, on seasonal effects, aggregation, or due to specific cities or commodities. In addition to highlighting the wide range of estimates of the magnitude of the

5. Note also that this specification avoids the "definitional artifact" critique pointed out by Hartman (1991), since the regression includes only an inflation term, and not the two terms: inflation, and inflation squared.

relationship, it should be noted that these results do not allow us to discriminate among the various models predicting a relationship. Next, some evidence is presented testing an application specific to models relying on misperceptions, or relative/aggregate confusion.

According to Lucas (1973) an increase in the variability of aggregate shocks reduces firms' output response since with higher variability of aggregate shocks there is a higher probability that an observed change in P_{ijt} is purely nominal. That is, firms react less to any given shock when aggregate variability is higher because firms are not as likely to be fooled. This implies that, in a regression of variability on inflation, the coefficient on inflation will be negatively related to the ratio of nominal to real shocks.

Existing studies have ignored this implication, presumably due to the difficulty of identifying the sources of the various shocks. Below, this implication is tested using a proxy for relative-aggregate confusion that uses both the time series and cross-section variation in the data set. Consider the ratio of aggregate inflation variability to cross-section inflation variability. This ratio identifies real shocks as those particular to the product, or market; nominal shocks, then, are a residual once these (product or market) specific shocks have been accounted for. Under the null hypothesis, an increase in this ratio is associated with lower relative price variability since firms become less surprised by any given increase in aggregate inflation.

To implement the test of the hypothesis that the coefficient on inflation is itself a function of the ratio of the variance of nominal to real shocks, this ratio is interacted with the inflation variable in the estimated equations. The resulting proxy, Ψ, is equal to the ratio of the variance of aggregate shocks (σ_η^2) to cross-section shocks (σ_ϵ^2 or σ_ν^2) times the relevant inflation rate. The relative-aggregate confusion model predicts the sign on the estimated coefficients to be negative.

The empirical counterparts for σ_η^2, σ_ϵ^2, and σ_ν^2 are constructed as follows. First, assume σ_η^2 can be approximated by the conditional variance of the aggregate inflation (Π_t) process, which for this data is modeled as an ARCH(1) process, that is,[6]

$$\Pi_t = \alpha_1 + \beta_1\Pi_{t-1} + u_t,\ u_t \sim N(0, \sigma^2_{\eta,t}), \text{ and } \sigma^2_{\eta,t} = \alpha_2 + \beta_3\sigma^2_{\eta,t-1} \tag{6}$$

where Π_t is equal to the unweighted average of the city inflation rates. The instruments for σ_ϵ^2 and σ_ν^2 are also estimated directly from the data, that is,

$$\sigma^2{}_\nu = \frac{1}{m-1}\sum_{t=1}^{m} (\Pi_{it} - \overline{\Pi}_{it})^2$$

$$\sigma^2{}_\epsilon = \frac{1}{n-1}\sum_{j=1}^{n} (\Pi_{jt} - \overline{\Pi}_{jt})^2 \tag{7}$$

6. The GARCH(1, 1) model has been applied to the U.S. GNP deflator by Bollerslev (1986). For the data used in this paper a χ^2 test for the presence of ARCH(1) effects yields the value .0142, with a significance level of .9047, while tests for GARCH(1, 1), and GARCHM(1, 1) effects were easily rejected.

TABLE 4

THE IMPORTANCE OF RELATIVE-AGGREGATE CONFUSION

	Dispersion Measure			
	V_{1it}	V_{2jt}	V_{3it}	V_{4jt}
Log Inflation	.4461 (.0753)	.9021 (.2078)	.7705 (.0916)	2.042 (.2484)
Ψ_{in}	−.8441 (.0623)		−2.306 (2.544)	
Ψ_{jn}		−1.077 (.0436)		−5.288 (1.680)
Quarterly dummies	yes	yes	yes	yes
Commodity dummies	yes	no	yes	no
Time dummies	·yes	yes	yes	yes
City dummies	no	yes	no	yes
Adjusted R-Squared	.62	.66	.11	.32
Std. Error of regression	.2820	.2287	.3431	.2783
Number of observations	2219	2871	2200	2873

NOTES: Heteroskedasticity consistent standard errors in parenthesis. All variability measures are in logs. The individual regressions run were of the form.

$$\log V_t = \sum_{i=1}^{n} \lambda_i + \sum_{t=1}^{68} \eta_t + \sum_{k=1}^{3} \theta_k + \beta\,\overline{\Pi}_t + \delta\Psi_{\eta t} + u_t$$

where $\Psi_{\eta t} = \sigma_\eta^2/\sigma_\nu^2(\overline{\Pi}_{it})$ in regressions for V_{1it} and V_{3it}, and, $\Psi_{\eta t} = \sigma_\eta^2/\sigma_\epsilon^2(\overline{\Pi}_{jt})$, in regressions for V_{2jt} and V_{4jt}. λ_i is a product effect, n is the number of cross sections (32 for V_{1it} and V_{3it}, and 48 for V_{2jt} and V_{4jt}), η_t are time dummies, θ_k are quarterly dummmies, $\overline{\Pi}_t$ is average (across cities, or products) inflation, and u_t is a regression disturbance.

where Π_{it} is the first difference of the log of product i's average (across cities) price and $\overline{\Pi}_{it}$ is the average of the Π_{it} across all products. Π_{jt} and $\overline{\Pi}_{jt}$ are defined similarly. Thus $\sigma^2{}_\nu$ represents the variability of product specific shocks, and $\sigma^2{}_\epsilon$ is a measure of the variability of city-specific shocks. The estimations equation in (5) become

$$V_t = \sum_{i=1}^{n} \lambda_i + \sum_{t=1}^{68} \eta_t + \sum_{k=1}^{3} \theta_k + \delta\Psi_{\eta t} + \beta\overline{\Pi}_t + u_t\,, \tag{5'}$$

where $\Psi_{\eta t} = \sigma_\eta^2/\sigma_\nu^2\ (\Pi_{it})$ in regressions for V_{1it} and V_{3it}, and $\Psi_{\eta t} = \sigma_\eta^2/\sigma_\epsilon^2\ (\overline{\Pi}_{jt})$ in regression for V_{2jt} and V_{4jt}. All other variables are defined as in equation (5). Again, according to relative-aggregate confusion models, $\delta < 0$. The evidence presented in Table 4 is strongly in favor of this hypothesis. In all four cases the estimate of δ is negative, and the estimate is statistically significant at the 1 percent level in three out of the four cases.

To summarize, Tables 3 and 4 have provided evidence that the variability of relative prices and of relative inflation is positively related to inflation. This result holds both when variability is measured within a city (across products) and across cities. This result is not a "definitional artifact," does not depend on outliers, and holds even after controlling for specific time, seasonal, and city/commodity effects. Finally, the evidence supports the hypothesis that as the variance of nominal shocks increases relative to real shocks, the elasticity of relative price variability with respect to inflation diminishes. The next section asks whether the effects of inflation on relative prices and relative inflation rates persist.

3. PERSISTENCE EFFECTS

In this section the persistence of the inflation–relative price linkage is examined from two perspectives. First, the persistence characteristics of the univariate time series dispersion measures are compared to the residual, or unexplained (that is, after accounting for the effect of inflation) dispersion. The results here indicate that it is possible to reject the unit root null hypothesis for both series, implying the relationship between inflation and dispersion is short run. The second, more agnostic, perspective is gained by examining the impulse response functions derived from a vector autoregressive model of variability and inflation. The motivation for this approach is the lack of agreement in the literature over the functional form of the relationship and even over which variable is the appropriate right-hand-side variable (for example, Ball and Mankiw 1995).

In Table 5 the persistence of the inflation–relative price linkage is studied by comparing unit root tests on the dispersion measures themselves to analogous tests on dispersion after controlling for the impact of inflation. It is unlikely that this test would be conclusive were it not for the large amount of data available in the current panel data set. Levin and Lin (1992) show that the power for rejecting the unit root null hypothesis increases rapidly in a panel setting. However, as in the usual setting, the t-statistic converges asymptotically to a (noncentral) normal distribution. In order to determine the appropriate critical values for this particular data set, five thousand Monte Carlo simulations are undertaken. In each simulation, an artificial data set is generated under the null hypothesis. The error is assumed to be i.i.d. normal (0, 1). All observations that are missing in the actual data set are also deleted in the artificial data set before computing the t-statistic. As a result, the sample size in the simulations is exactly the same as the actual data set. Similar simulations are performed to compute the actual power of the test against various alternatives (see Appendix Table A3 for details). Finally, these simulations were performed for both dispersion measures V_{1it} and V_{2jt} since the two cross-sections are different, that is, V_{1it} is a cross-section of goods and V_{2jt} is among cities.

In the top panel of Table A3 the simulations involving V_{1jt} are presented. The t-statistics under the null are indeed skewed to the left of zero in the simulations, with a mean of -9.41. The 1 percent and 5 percent critical values for rejecting the null are -10.13 and -9.96, respectively. Even with these large critical values the unit root null hypothesis is easily rejected at the 1 percent level for V_{1it}. According to the power calculations, at the 1 percent level we would reject the null only 4.8 percent of the time if the alternative $\beta = .001$ were actually true. However, the power increases dramatically (to 98 percent) if the alternative $\beta = .10$ were actually true. The 1 percent and 5 percent critical values from simulations involving V_{2jt} (in

TABLE 5

THE PERSISTENCE OF DISPERSION AND SHOCKS TO DISPERSION

	Dispersion Measure			
	V_{1it}	V_{2jt}	V_{3it}	V_{4jt}
Regression 1:				
Lagged dispersion	−.2859	−.4572	−.8750	−.8167
	(.0156)	(.0158)	(.0220)	(.0190)
Quarterly dummies	yes	yes	yes	yes
Commodity dummies	yes	no	yes	no
Time dummies	yes	yes	yes	yes
City dummies	no	yes	no	yes
Adjusted R-Squared	.62	.66	.11	.32
Std. Error of regression	.2820	.2287	.3431	.2783
Number of observations	2219	2871	2200	2873
Half life	2.06	1.13	.33	.41
	Unexplained Dispersion Measure			
	υ_1	υ_2	υ_3	υ_4
Regression 2:				
Lagged υ	−.2759	−.4601	−.8587	−.8042
	(.0152)	(.0161)	(.0221)	(.0190)
Quarterly dummies	yes	yes	yes	yes
Commodity dummies	yes	no	yes	no
Time dummies	yes	yes	yes	yes
City dummies	no	yes	no	yes
Adjusted R-Squared	.09	.20	.39	.40
Std. Error of regression	.1964	.1928	.3398	.2691
Number of observations	2145	2818	2192	2784
Half life	2.15	1.12	.35	.42

NOTES: Heteroskedasticity consistent standard errors in parentheses. All variability measures are in logs.
The specification for regression 1 was.

$$\Delta \log V_t = \beta \log V_{t-1} + \text{quarter, time, and, city or good, dummies}$$

For regression 2, replace V, with unexplained dispersion—defined as the residual series from the regressions in Table 3.

the lower panel of Table A3) are even larger in absolute value (−15.01 and −14.36, respectively), yet the null hypothesis is still comfortably rejected for V_{2jt}. If we assume the distributions for V_{3it} and V_{4jt} are similar to those for V_{1it} and V_{2jt}, then the unit root null is rejected for these variability measures as well. These results imply that dispersion is stationary, and therefore no long-run relationship exists between inflation and dispersion.

Table 5 also presents the implied half-lives of deviations in these dispersion measures. According to the table, shocks to dispersion dissipate very quickly; in the worst case, after two quarters half of the disequilibrium is eliminated. Inflation rates apparently adjust even faster. The half-life of deviations to the dispersion of inflation rates is less than half a quarter.

In the bottom panel unit root tests on the "residual" dispersion series (that is, the residual from a regression of dispersion on inflation and time, product, and seasonal dummies) are reported.[7] The results are illustrative. Note the magnitude of the coef-

7. In some sense, this regression is redundant since the two series cannot be cointegrated.

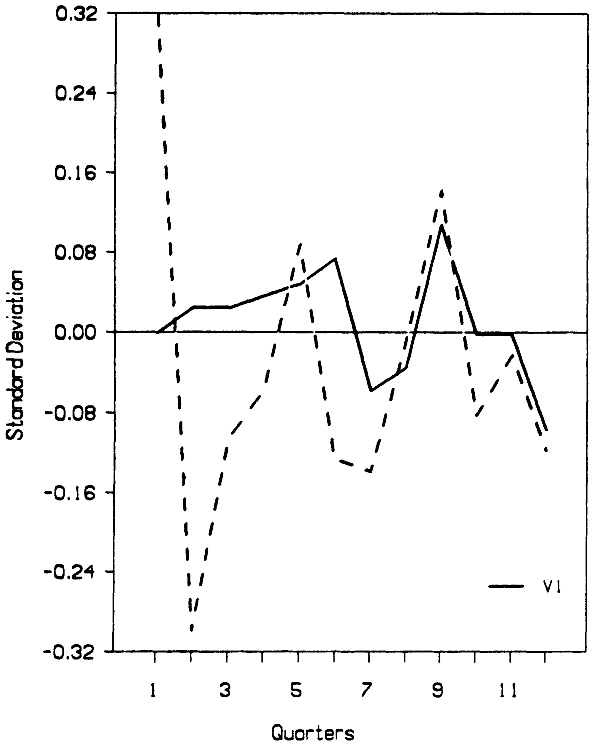

FIG. 1. Impulse Responses (V1 and Inflation)

ficients on the lagged residual series are virtually identical to those in the top panel. Since inflation effects have been filtered out of the residual dispersion series, these results confirm those in the top panel that the major force driving dispersion is not inflation.

An alternative way to view the dynamic interaction between the two series is through their impulse response functions. The impulse response functions plotted in Figures 1 through 4 show the dynamic paths of each endogenous variable to a one-standard-deviation shock to the inflation equation. Impulse response functions were obtained from estimating a two-equation VAR allowing for eight lags of each of the endogenous variables.[8] The plots support the results from the univariate analysis.

In particular, the most persistent effect of inflation seems to be on V_{1it}, and the least persistent effects seem to be for measures V_{3it} and V_{4jt}. Visually, the graphs suggest that once the mutual interdependence of inflation and variability are accounted for, a one-time shock to inflation is dissipated fairly quickly. That is, with the exception of the impulse response for V_{1it}, variability returns to its preshock state in a little over a year. Moreover, the effects implied by these impulse response

8. Estimates from a VAR with twelve lags produced virtually identical impulse response functions.

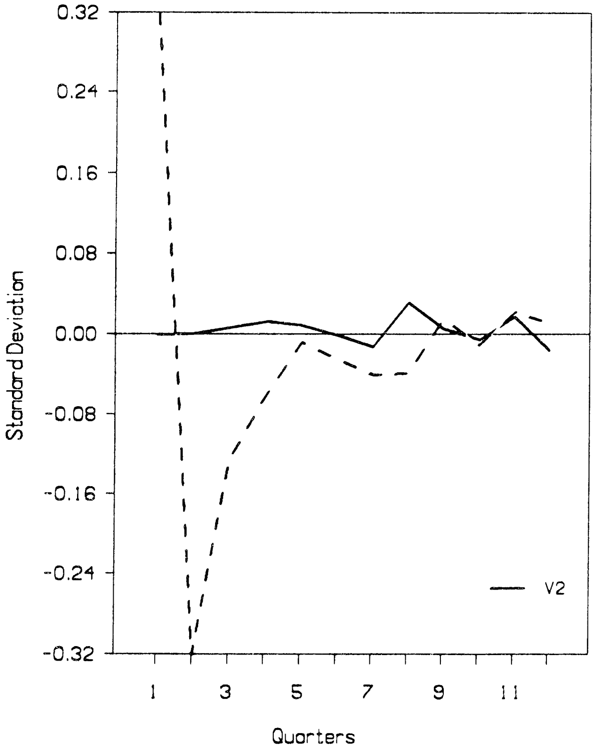

FIG. 2. Impulse Responses (V2 and Inflation)

functions are much smaller than the contemporaneous effects estimated from single-equation regression analysis. That is, the dominant movement in the graphs is the reversion in the inflation series, and the effects on the variability series are generally two orders of magnitude smaller than the initial shock to inflation. Combined with the evidence in Table 5, this suggests that much of what is typically attributed to the effect of inflation on relative prices actually reflects simultaneity, or mutual dependence, in the two series coupled with strong mean reversion tendencies in both series.

4. SUMMARY

This study's contributions fall into two categories. First, this paper presents new evidence that a positive association exists between inflation and relative prices and relative inflation rates in very disaggregated data for the United States over the period 1975 through 1992. There is also evidence that the response of relative prices and relative inflation rates to inflation varies inversely with the information content

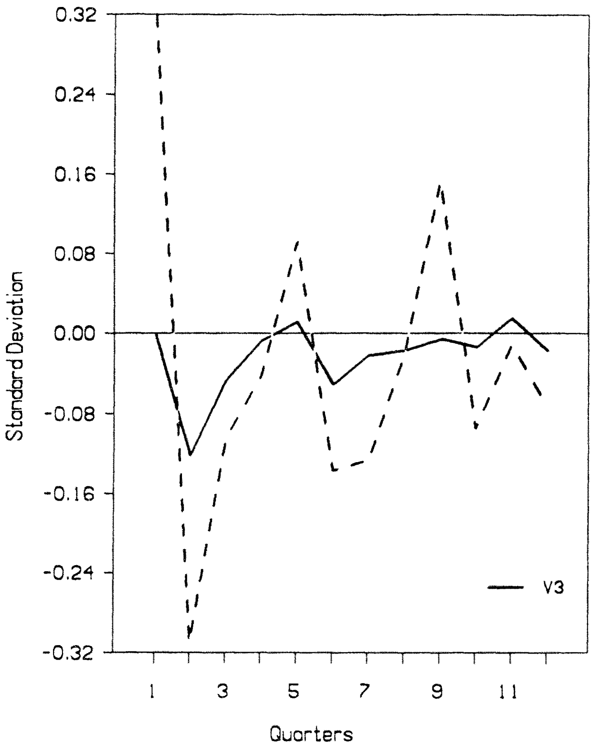

FIG. 3. Impulse Responses (V3 and Inflation)

of a given shock to inflation. The relationship is studied from two cross-sectional perspectives using individual price series collected from forty-eight U.S. cities.

The findings are free of many of the criticisms of earlier empirical studies. The data studied are prices of individual goods and services that have undergone very little quality change over the sample period. Also, the data are sampled from a broad cross-section of U.S. cities over a reasonably long time period. Importantly, since individual good, time, and quarter effects have been controlled for separately, the results cannot be driven by a single product, or year, as found by Bomberger and Makinen (1993) for the earlier results of Parks (1978).

The second objective of this paper was to present evidence on the persistence of the effects of inflation on relative prices. Results from this part of the study challenge the traditional interpretation of the relationship between inflation and relative prices. Several conclusions can be drawn from this analysis. First, the effect of inflation on relative prices is not long run, that is, the two series are not cointegrated. Results from vector autoregressions imply the effect is much smaller than indicated by typical estimates. Combined, these imply that the welfare implications due to relative price effects of inflation are minor. Secondly, estimated impulse response

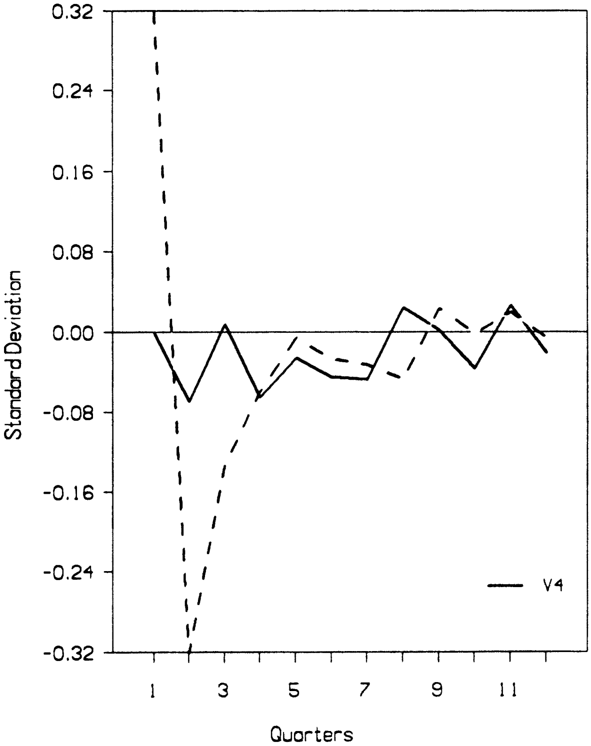

FIG. 4. Impulse Responses (V4 and Inflation)

functions indicate the effect of inflation on relative prices is greater than the effect of inflation on relative inflation rates. Loosely, this result is in accord with menu cost models since following a shock, these models predict some prices change and not others. The result is that relative price dispersion increases. It is less clear from these models why relative rates of inflation should exhibit inertia.

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APPENDIX TABLE A1

CITIES AND COMMODITIES INCLUDED

Cities		Goods	
C1	Birmingham AL	G1	Steak
C2	Mobile AL	G2	Hamburger
C3	Blythe CA	G3	Bacon
C4	Indio CA	G4	Whole Chicken
C5	Palm Springs CA	G5	Milk
C6	Denver CO	G6	Eggs
C7	Lakeland FL	G7	Margarine
C8	Boise ID	G8	Potatoes
C9	Champaign, Urbana IL	G9	Bananas
C10	Peoria IL	G10	Lettuce
C11	Ft. Wayne IN	G11	Bread
C12	Indianapolis IN	G12	Cigarettes
C13	Cedar Rapids IA	G13	Coffee
C14	Lexington KY	G14	Sweet Peas
C15	Louisville KY	G15	Tomatoes
C16	Baton Rouge LA	G16	Peaches
C17	Lafayette LA	G17	Tissue
C18	New Orleans LA	G18	Washing Powder
C19	Benton Harbor MI	G19	Shortening
C20	Traverse City MI	G20	Frozen Orange Juice
C21	Columbus MS	G21	Baby Food
C22	St. Joseph MO	G22	Soft Drink
C23	St. Louis MO	G23	Gasoline
C24	Falls City NE	G24	Hospital Room
C25	Hastings NE	G25	Office Visit Doctor
C26	Omaha NE	G26	Office Visit Dentist
C27	Reno, Sparks NV	G27	Man's Haircut
C28	Newark NJ	G28	Dry Cleaning
C29	New York NY	G29	Major Appliance Repair
C30	Hickory NC	G30	Movie
C31	Columbus OH	G31	Bowling
C32	Altoona PA	G32	Liquor
C33	Rapid City SD		
C34	Vermillion SD		
C35	Chattanooga TN		
C36	Knoxville TN		
C37	Abilene TX		
C38	El Paso TX		
C39	Ft. Worth TX		
C40	Houston TX		
C41	Lubbock TX		
C42	Salt Lake City UT		
C43	Charlseton WV		
C44	Appleton WI		
C45	Eau Claire WI		
C46	Madison WI		
C47	Oshkosh WI		
C48	Casper WY		

APPENDIX TABLE A2

DESCRIPTIONS OF COMMODITIES INCLUDED

Grocery Items:
Steak, lb. round steak (75.1–80.3); T-bone steak (80.4–92.4) USDA Choice
Ground Beef or Hamburger, lb.
Bacon, lb., national brands
Whole Chicken, lb. Grade A Frying
Milk, 1/2 gal. carton
Eggs, doz. large Grade A
Margarine, lb.
Potatoes, 10 lbs. white or red
Bananas, lb.
Lettuce, each
Bread, 24 oz. (75.1–80.2); 20 oz. (80.3–92.4)
Cigarettees, carton Winston king-size
Coffee, 2 lbs. (75.1–80.2); 1 lb. (80.3–88.3); 13 oz. (88.4–92.4); Maxwell House, Hills Brothers, Folgers
Sweet Peas, #303 can 15–17 oz. (75.1–85.4); 17 oz. (86.1–92.4); Del Monte or Green Giant
Tomatoes, #303 can 15–17 oz. (75.1–85.4); 14.5 oz. (86.1–92.4); Del Monte or Green Giant
Peaches, #2 1/2 can approx. 29 oz. (75.1–85.4); 29 oz. (86.1–92.4); Del Monte or Libby's halves or slices
Tissue, 1 roll (75.1–79.1); 4 rolls (79.2–80.2); Kleenex brand 175 count box (80.3–92.4)
Washing Powder, 49 oz. (75.1–88.4); 42 oz. (89.1–92.4); Giant Tide, Bold, or Cheer
Shortening, 3-lb. can all vegetable, Crisco brand
Frozen orange juice, 6-oz. can (75.1–85.4); 12 oz. can (86.1–92.4)
Baby food, jar, 4 1/2 oz. strained vegetables
Soft Drink, 1 qt. Coca-Cola (75.1–79.2); 2-liter (79.3–92.4)

Nongrocery Items:
Gasoline, regular including taxes national brand full service (75.1–79.1); unleaded: full service (79.2–81.4); self-service (82.1–92.4)
Hospital Room, semiprivate cost per day
Office Visit Doctor, general practitioner routine exam of existing patient
Office Visit, Dentist, teeth cleaning and inspection, no x-ray or fluoride treatment
Man's Haircut, no styling
Dry Cleaning, man's two-piece suit
Major Appliance Repair, service call excl. parts, color TV (75.1–79.1); Washing machine (79.2–92.4)
Movie, first-run indoor evening price
Bowling, price per line evening price
Liquor, 750 ml. bottle Seagram's 7 Crown (75.1–88.3); J&B Scotch (88.4–92.4)

APPENDIX TABLE A3

SIZE AND POWER SIMULATIONS OF TEST FOR H_0: $\beta = 0$

DISPERSION MEASURE: V_{1it}

Size		Power against alternatives (%) H_a: β =		
Size	Critical Value	−.001	−.01	−.1
1%	−10.13	1.8	26.1	97.8
5%	−9.96	5.8	38.1	98.3
10%	−9.86	11.0	44.2	98.4

DISPERSION MEASURE: V_{2jt}

Size		Power against alternatives (%) H_a: β =		
Size	Critical Value	−.001	−.01	−.1
1%	−15.01	1.3	5.9	64.2
5%	−14.36	5.8	13.7	72.6
10%	−14.03	11.0	18.6	76.5

NOTES In each simulation, artificial data are generated according to the following specification for each product (city in the bottom panel), and time period.

$$\log V_t = (1 + \beta) \log V_{t-1} + \sum_{i=1}^{n} \lambda_i + \sum_{t=1}^{68} \eta_t + \sum_{k=1}^{3} \theta_k + u_t .$$

The coefficients for product (city), time, and quarter dummies are assumed to be the same as the estimated values in the regressions in Table 5. The initial values are assumed to be equal to be the actual values in 1975.1. Before computing the test statistics to test H_0 $\beta = 0$, all observations missing in the actual data set are deleted. Therefore, each artificial data set has the same number of observations (for example, 2,219 for V_{1it}). Twenty-five hundred simulations are performed for each panel. The error term is assumed to be an i.i.d standard normal variate.