The Behavior of Prices and Inflation: An Empirical Analysis of Disaggregated Price Data

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This paper analyzes the effects of inflation on the dispersion of prices, as well as other aspects of price behavior, using disaggregated data on prices of foodstuffs in Israel during 1978–84. We find that the effect of expected inflation on intramarket price variability is stronger than the effect of unexpected inflation. We show that even in times of high inflation, price quotations are not trivially short and price changes are not synchronized across firms. These facts, taken together, confirm that there is some staggering in the setting of prices. We find that the distribution of real prices is far from being uniform, as many menu cost–based models assume or conclude. In fact, as inflation increases to very high levels, this distribution is not even symmetric. When the annual inflation rate reaches 130 percent, there are equal chances of finding real prices above or below the market average, but upward deviations in the real price are further away from zero than downward ones. Furthermore, as the annual rate of inflation more than doubles from 60 to 130 percent, real prices are pushed toward both tails of the distribution.

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I. Introduction

In the past decade, many empirical studies on the dispersion of prices reached the conclusion that price dispersion is positively correlated with the rate of inflation. Different mechanisms accounting for this correlation have been advanced in the literature. The approach of Lucas (1973), based on imperfect information, emphasizes the role of unexpected inflation and inflation variability in generating intermarket price dispersion (Hercowitz 1981; Cukierman 1984). Another approach builds on the presumption that nominal price changes are costly; that is, they are subject to menu costs. The optimal policy, in this case, is to set prices discontinuously according to an (S, s) pricing rule (Sheshinski and Weiss 1983). In choosing the target and the threshold, a price setter contemplates the future evolution of inflation. Therefore, the distribution of inflation and, in particular, expected inflation affects the dispersion of prices. Still another theory that leads to a positive relation btween inflation and price dispersion is based on costly consumer search (Bénabou 1988). Stigler and Kindahl (1970) argue that the search process by itself is not sufficient to reduce price dispersion when consumers' information about prices erodes as a result of inflation. Van Hoomissen (1988*a,* 1988*b*) posits that information's obsolescence due to inflation reduces the optimal stock of price information that consumers wish to hold, thereby leading to greater price dispersion.

The high rates of inflation in the past two decades have prompted many economists to believe that there is something fundamentally wrong with inflation, beyond its intrinsic potential to surprise. Recent developments have shown that price dispersion or, more generally, its distribution is linked to a central question in macroeconomics: Does the aggregate price level lag behind money supply in such a way that it can be exploited by the monetary authority? Can expected changes in the money supply affect output? A series of recent papers that tackle this issue examine the relationship between the behavior of the aggregate price level and the distribution of prices (Caplin and Spulber 1987; Tsiddon 1988; Caballero and Engel 1989*b*). It follows, then, that different approaches posit different channels by which inflation affects the dispersion of prices and, implicitly or explicitly, emphasize different aspects of the inflationary process as the primary cause of such an effect. In addition, it is now recognized that the pattern of this dispersion may have important macroeconomic implications.

In this study we analyze these interrelated issues using disaggregated data on prices of foodstuffs in Israel during 1978–84. Inflation was not a new phenomenon in the Israel of 1978, and it persisted at very high rates until 1985. The price data were drawn from the sample used in the monthly computation of the consumer price index by the Central Bureau of Statistics. These are micro-level data, collected at the store level, and therefore most closely resemble the data envisioned by the cost of adjustment theory: price quotations at the level of the price setter.

We find that *expected inflation* has an important effect on price dispersion. We show that the effect of the expected component of inflation on intramarket price variability is stronger than the effect of the unexpected part of inflation. As the data show that over 80 percent of total price variability comes from the intramarket price variability, the effect of expected inflation on overall price variability is substantial.

We show that, even in times of high inflation, price quotations are not trivially short and price changes are not synchronized across firms. These facts, taken together, confirm that there is some staggering in the setting of prices. We find that the distribution of real prices is far from being uniform, as many menu cost–based models assume or conclude. In fact, as inflation increases to very high levels, this distribution is not even symmetric. Our results point toward the following description: when the annual inflation rate reaches 130 percent, there are equal chances of finding real prices above or below the market average, but upward deviations in the real price are further away from zero than downward ones. On the other hand, *extreme* deviations from the market average are more likely to be downward. Furthermore, as the annual inflation rate more than doubles from 60 to 130 percent, real prices are pushed toward both tails of the distribution.

The paper is organized as follows. Section II describes the data, and Section III presents the results on the relationship between price dispersion and inflation. Section IV analyzes other aspects of the behavior of prices: the duration of price quotations and the size of price changes, the synchronization in the timing of price changes across firms, and the distribution of real prices. Section V presents conclusions.

II. Description of the Data

The data set consists of price quotations on 26 food products reported by a sample of stores. The data were collected by the Central Bureau of Statistics (CBS) for the purpose of computing the consumer price index (CPI). The 26 products satisfy the following criteria: they are homogeneous, they did not change substantially either in quality or in the structure of their markets, and their prices were not controlled by the government during the period investigated.

The periods for which most of the data are available are 1978–79, 1981–82, and the first nine months of 1984 (before the first stabilization program was launched). The data for 1980 and 1983 practically disappeared from the CBS archives. We have 1983 data for only three fish products and for rice; in the case of beef for soup, the data for 1978 were missing.

The data were collected by monthly visits to stores always in the same week of the month. Each week therefore reflects a different set of stores, and the same stores reappear every four weeks. Chain stores are not included in the sample, and the only information available, besides the price quotation, is the store's code number and the week, month, and year in which it was sampled.

Table 1 presents basic statistics. For each product there usually is a different number of reporting stores (col. 1), and of course the same store may report prices of several products. Unfortunately, the data set is not balanced: not only is the number of nonmissing observations per store different across stores, but the calendar period in which they were sampled is also different. In half of the 57 months, the number of stores with nonmissing data exceeds the figure in column 2; the median store coverage ratio is well above 50 percent for almost all products. Similarly, half of the reporting stores have data for at least the number of months appearing in column 4. Dividing the numbers in column 4 by 57 yields the median monthly coverage ratio appearing in column 5: with a few exceptions, 50 percent of the stores appear in well above 55 percent of the sample period. At least for the first 20 products, these figures suggest that although there are some changes over time in the identity of the sampled stores, there is a sizable core of stores constituting a more or less balanced set of data.

III. Relative Price Variability and Inflation

As mentioned in the Introduction, a number of not necessarily competing hypotheses have been advanced to explain the observed positive association between relative prices and some aspects of inflation. These theories usually emphasize the role of different characteristics of inflation—such as expected or unexpected inflation or its variability—in explaining this positive correlation.

At this stage it is convenient to clarify what we mean by relative price variability and price dispersion. The first concept refers to the tendency of relative prices to change over time and is usually proxied by the cross-sectional (i.e., across stores in this paper) variance of the

TABLE 1

BASIC STATISTICS FOR 1978–79, 1981–82, AND 1984:1–1984:9

Product	Number of Reporting Stores (1)	Median Number of Stores per Month (2)	Median Store Coverage Ratio (Col. 2 ÷ Col. 1) (3)	Median Number of Months per Store (4)	Median Monthly Ratio (Col. 4 ÷ 57) (5)
1. Tea bags	31	22	.71	47	.82
2. Fresh beef	56	40	.71	47.5	.83
3. Frozen goulash	50	43	.86	54	.95
4. Challah bread	21	14	.67	49	.86
5. Cocoa powder	30	26	.87	54	.95
6. Fish fillet	30	20	.67	42	.74
7. Arrack	31	21	.68	47	.82
8. Buri fish	29	18	.62	39	.68
9. Codfish	28	19	.68	47.5	.83
10. Frozen beef liver	45	30	.67	40	.70
11. Fresh beef liver	45	26	.58	33	.58
12. Chicken breast	54	38	.70	48	.84
13. Chicken liver	53	36	.68	39	.68
14. Rice	26	22	.85	54	.95
15. Turkey breast	43	24	.56	37	.65
16. Steak	48	29	.60	38.5	.67
17. Beef for soup*	50	28	.56	28.5	.63
18. Chicken legs	48	29	.60	35.5	.62
19. White vermouth	27	11	.41	18	.32
20. Liquor	21	9	.43	12	.21
21. Champagne	22	11	.50	18	.32
22. Vodka	29	16	.55	30	.53
23. Red wine	31	19	.61	32	.56
24. Rosé wine	28	17	.61	35	.61
25. Hock wine	33	19	.58	33	.58
26. Sweet red wine	34	18	.53	34	.60

* Does not include data for 1978, so col. 5 equals col. 4 divided by four.

rate of change of price. We emphasize that relative prices here refer to the comparison of prices across the store dimension, with the product held fixed. On the other hand, price dispersion is a more static concept that looks at the cross-sectional variance of price *levels*. Clearly, these two concepts are different objects. In fact, they may even move in opposite directions. At the risk of creating a false dichotomy between the different models mentioned in the Introduction, we chose to elaborate on two, and by no means the only, particular approaches. The first approach underlies the positive effect of expected inflation on the dispersion of prices and is associated with menu cost models. The presence of fixed costs of adjustment in nominal prices induces the firm to change its nominal prices in a discontinuous fashion. If the inflation rate is nonnegative and if some other conditions are satisfied, the optimal policy is to change the nominal price only when the real price hits a lower threshold, s (Caplin and Sheshinski 1987). The nominal price is changed so that the new real price equals a higher return point, S. One feature of this optimal policy is that the characteristics of the inflationary process affect both the threshold and the return points. In particular, the distance between S and s increases with the *expected* value of inflation, and since, under further restrictions (Tsiddon 1988), such an increase is associated with a greater dispersion of prices, the relationship between them is positive. These models are usually concerned with the price-setting behavior of sellers of a single product. Hence, they have direct implications for *intramarket* dispersion of prices.

The second approach is based on the Lucas-type confusion between aggregate and relative shocks and emphasizes the positive effect of *unexpected* inflation on relative price variability. These models are constructed in such a way that the relevant dimension along which prices are compared is the product dimension. One way of carrying on this comparison in a meaningful fashion is to compare the rates of change in the products' price. That is, the relevant concept is the dispersion of the products' own inflation rates around an aggregate rate of inflation. According to the terminology adopted in this paper, this is the *intermarket* relative price variability. This effect has been confirmed empirically by several studies (Parks [1978], Fischer [1981], and Hercowitz [1981]; for a dissenting view, see Jaffee and Kleiman [1977]).

The main topic of this section is the empirical relationship between the variability of relative prices of a given product across stores, the *intramarket* price variability for short, and the expected and unexpected components of inflation.[1]

[1] Domberger (1987) also draws the distinction between intermarket and intramarket price variability. Until recently, however, empirical work on the latter was rarely con-

Before we discuss the empirical results, some comments on our measure of price variability and expected inflation are in order. The unit of time used in the analysis is 1 month. We therefore aggregate our weekly data into a monthly measure of price variability.[2] Let P_{ijt} be the price of product i in store j during month t and N_{it} the number of stores quoting prices of product i in month t. As mentioned above, relative price variability is usually measured by the cross-sectional variance of the rates of change of price (Domberger 1987; Van Hoomissen 1988*a*), and we do the same. Denote the rate of change in the price of product i in store j between month $t - 1$ and month t by $DP_{ijt} = \ln P_{ijt} - \ln P_{ijt-1}$. Their standard deviation across stores is

$$\mathrm{SDP}_{it} = \left[\frac{1}{N_{it} - 1} \sum_{j} (DP_{ijt} - DP_{i.t})^2\right]^{1/2}, \tag{1}$$

where $DP_{i.t}$ is product i's monthly average inflation rate and equals $(1/N_{it}) \sum_j DP_{ijt}$.[3]

Some of the advantages of working with rates of change are that they eliminate possible store effects in price levels (which enter multiplicatively), they may also help in taking care of possible nonstationarities in price levels, and they facilitate aggregation over products and thus make the results of our analysis comparable to other results obtained at more aggregate levels.

On the other hand, menu cost models have direct implications on the relationship between inflation and intramarket price dispersion. In fact, as Danziger (1987) points out, in a model with homogeneous firms and a deterministic environment, relative price variability is not a monotone function of the rate of inflation. In his model, based on

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ducted, presumably because of the lack of adequate data. Domberger (1987) and Van Hoomissen (1988*a*) are recent exceptions. The latter paper studies intramarket price variability in Israel using the same type of data we use, but for a smaller number of products during a longer period, 1971–84. Both papers, however, focus on issues different from the ones analyzed here.

[2] There are two reasons for doing this. First, the data on inflation rates provided by the CBS come on a monthly basis, and, more important, the number of price quotations per week is quite low: its average ranges between 2.5 and 10.7 per product. On a monthly basis, however, the averages range between 10.3 and 42.8 price quotations per product.

[3] Notice that between-week differences in inflation rates were not removed from (1). Were we measuring price dispersion, i.e., dispersion of price *levels*, this would constitute a problem since it implies a built-in positive correlation between the monthly variability measure and inflation. Thus the between-week effects of inflation have to be removed from measures of price dispersion such as the cross-sectional coefficient of variation of price levels. Regarding (1), between-week differences in inflation *rates* are not that significant, and it is not obvious in which direction these differences move as inflation changes over time. It should also be noted that this measure incorporates stores' inflation rates that were sampled at about 21 days apart, the length of time between the first and last weeks in the month.

Rotemberg's (1983) approach, an increase in the inflation rate increases SDP at low levels of inflation but decreases it when inflation is too high. This occurs despite the fact that the *dispersion* of prices is always positively related to the inflation rate.

This suggests that one should look for the effects of expected inflation on a measure of price dispersion such as the cross-sectional coefficient of variation of price *levels*. The problems we run into while analyzing the relationship between this measure of price dispersion and inflation convinced us that we need a more structured and empirically oriented model to tackle this issue. Hence we do not present results on the effects of the components of inflation on the cross-sectional dispersion of price levels.[4]

While it is true that menu cost models do not imply a univocal relationship between price variability and inflation, is there still anything to be learned from the empirical finding that this relationship is positive? Our answer is also positive for two reasons. First, even in Danziger's stylized model, the association between SDP and inflation implied by menu cost models is positive provided that the period between successive price adjustments exceeds 1.5 months (Danziger 1987, p. 707). In our sample, the average duration of a price quotation is 1.9 months in 1978–79 and decreases to 1.5 months during 1982. From our empirical perspective, therefore, the nonmonotonicity of the SDP-inflation relationship may be relevant only in a small portion of the data (see n. 8). Second, the stores in our data are, of course, not homogeneous as the Rotemberg model assumes. In the Rotemberg model the distance of the (S, s) band is $k\pi^{1/3}$, where k is a constant whose value depends on the parameters of the model and π is the rate of inflation. One manifestation of store heterogeneity is, presumably, in heterogeneous values for k. In this case, the difference in the magnitude of the (S, s) band between two stores, $(k_1 - k_2)\pi^{1/3}$, is increasing in the inflation rate. Moreover, as shown in Appendix A, in most empirically relevant circumstances, DP_{ijt} is approximately equal to the length of the (S, s) band of product i sold by store j at some time between months $t - 1$ and t. Thus the preceding argument implies that the cross-sectional variance of DP_{ijt} is increasing in the inflation rate. Store heterogeneity opens up a new channel by which

[4] The Durbin-Watson statistics from the regression of the coefficient of variation (CV) on expected ($E\Pi$) and unexpected ($U\Pi$) inflation indicated the presence of AR(1) disturbances in all products. In half of them, however, this was a result of "dynamic misspecification"; i.e., the nonlinear restriction imposed by the AR(1) model on the regression of CV against its own lagged value and current and lagged values of $E\Pi$ and $U\Pi$ was rejected by the data. In these cases, CV_{-1} was highly significant (t-values above 10), and $E\Pi$ and $U\Pi$ became insignificant. In the other half of the products, the evidence for positive effects of expected and unexpected inflation was not that strong either: only about half of them had significant positive estimates.

inflation affects price variability, a channel that is missing in theoretical models that assume a homogeneous environment.

In sum, while the direct implications of menu cost models are better cast in terms of the relationship between price dispersion as defined here and inflation, we have argued that, for our empirical purposes, a similar implication holds for the relationship between relative price variability and the rate of inflation.

The details on the construction of the expected and unexpected inflation series are provided in Appendix B. For each product, expected inflation is taken to be the predicted value from a regression of current own inflation rate on past own inflation rates, past aggregate inflation rates, and time-related variables. It should be pointed out that although the CPI inflation rate was very high in the sample period, its behavior was not particularly erratic. In fact, the inflationary process in Israel was characterized by "steps" (Liviatan and Piterman 1986). This feature gives meaning to the notion of (linearly) predicting the inflation rate. Since the model used for prediction is a very simple and intuitive one (see App. B), we are confident that the decomposition into expected and unexpected components is not an artificial one.[5]

For each of the 26 products, we constructed time series of the measure of intramarket relative price variability, SDP, and of expected and unexpected inflation. We estimated sets of regressions with SDP_{it} as the dependent variable, where $i = 1, \ldots, 26$ indexes products. The regression results are provided in table 2. In one set, SDP was regressed on the expected ($E\Pi_{it}$) and unexpected ($U\Pi_{it}$) inflation rates of each product (cols. 1 and 2); in the other case, the regressor was the actual inflation rate of each product, now denoted by Π_{it} (col. 6).[6] These regressions are to be interpreted as a summary of correlations or reduced-form associations only since we are not estimating a structural model. Since we are dealing with foodstuffs, which are strongly affected by seasonality, government policy, and other "common" shocks, the natural framework for estimation is a system of seemingly unrelated regressions (SUR), which allows for

[5] Lach (1991) shows that errors in the decomposition of inflation into expected and unexpected components that, by construction, cancel each other do not affect the consistency of the ordinary least squares (OLS) estimator of the parameter of expected inflation but may under- or overestimate the parameter of the unexpected part of inflation depending on the sign of $\beta_1 - \beta_2$ in eq. (2). The latter term's sign can be consistently estimated since the plim of the OLS estimator of $\beta_1 - \beta_2$ is proportional to the true difference $\beta_1 - \beta_2$. Hence, we can still draw unbiased inferences on the relative effects of expected vs. unexpected inflation based on OLS estimators.

[6] The inflation rate Π_{it} is conceptually the same as $DP_{i,t}$. Since the former variable is taken directly from the CBS publications, it is not exactly equal to $DP_{i,t}$. See App. B for details.

TABLE 2

INTRAMARKET PRICE VARIABILITY

Dependent Variable: SDP

Product	$E\Pi_i$* (1)	$U\Pi_i$* (2)	R^2 (3)	$E\Pi_i$† Contribution (4)	$U\Pi_i$† Contribution (5)	Π_i (6)
1	.68	.63	.36	.35	.08	.67
	(6.2)	(3.0)				(6.5)
2	.28	.17	.32	.29	.05	.24
	(7.3)	(3.4)				(7.5)
3	.21	.40	.09	.03	.07	.30
	(2.1)	(3.4)				(4.0)
4	.50	.46	.60	.41	.17	.49
	(8.8)	(5.7)				(11.5)
5	.54	.59	.77	.73	.19	.55
	(12.4)	(7.0)				(13.5)
6	.34	.31	.45	.31	.15	.34
	(7.1)	(5.3)				(8.9)
7	.40	.44	.33	.20	.12	.43
	(6.2)	(5.4)				(8.6)
8	.49	.26	.48	.40	.06	.39
	(8.2)	(3.4)				(8.5)
9	.41	.47	.59	.32	.32	.43
	(7.5)	(8.5)				(10.8)
10	.54	.23	.37	.37	.01	.49
	(8.8)	(1.9)				(8.5)
11	.36	.17	.55	.53	.03	.31
	(10.4)	(2.6)				(10.2)
12	.67	.50	.20	.16	.04	.60
	(5.4)	(3.2)				(6.6)
13	.46	.33	.46	.42	.15	.41
	(10.8)	(7.2)				(12.0)
14	.63	.44	.17	.14	.03	.56
	(4.8)	(2.4)				(5.9)
15	.49	.28	.42	.31	.04	.43
	(8.8)	(3.5)				(10.5)
16	.51	.22	.58	.56	.05	.39
	(11.0)	(3.7)				(9.7)
17	.27	.16	.38	.34	.06	.24
	(4.6)	(1.9)				(4.7)
18	.08	.22	.15	.02	.11	.13
	(1.5)	(3.7)				(3.5)
19	.46	.28	.36	.32	.05	.38
	(8.2)	(3.7)				(8.2)
20	.49	.06	.24	.24	.00	.41
	(6.4)	(.5)				(6.1)
21	.40	.40	.43	.36	.12	.40
	(6.8)	(4.2)				(7.8)
22	.36	.55	.41	.18	.22	.43
	(6.1)	(7.6)				(9.4)
23	.46	.60	.51	.24	.29	.52
	(8.3)	(10.5)				(12.6)
24	.28	.16	.17	.12	.05	.18
	(4.0)	(3.1)				(4.6)

TABLE 2 (*Continued*)

Dependent Variable: SDP

Product	$E\Pi_i$* (1)	$U\Pi_i$* (2)	R^2 (3)	$E\Pi_i$† Contribution (4)	$U\Pi_i$† Contribution (5)	Π_i (6)
25	.23	.43	.03	.01	.01	.38
	(.9)	(1.1)				(1.8)
26	.68	.58	.50	.46	.17	.68
	(11.9)	(8.5)				(14.3)
Average‡	.43	.36				.41
	(.15)	(.16)				(.14)

NOTE.—The equations for each product were estimated jointly using the SUR procedure, with the exception of product 17, for which 1978 data were missing. *t*-values are in parentheses.

* Expected and unexpected inflation are defined in App. B.

† From eq. (2), the total variation of SDP_t around its sample mean, with the subscript *i* omitted, is given by

$$\sum_t (SDP_t - SDP_.)^2 = b_1^2 S_{ee} + b_2^2 S_{uu} + 2b_1 b_2 S_{eu} + \sum_t \hat{\epsilon}_t^2 + 2b_1 \sum_t \hat{\epsilon}_t E\Pi + 2b_2 \sum_t \hat{\epsilon}_t U\Pi,$$

where $SDP_.$ is the sample mean of SDP_t; S_{ee} and S_{uu} are the sum of squared deviations around their means of expected and unexpected inflation, respectively; $S_{eu} = \Sigma_t(E\Pi_t - E\Pi_.)(U\Pi_t - U\Pi_.)$; $\hat{\epsilon}_t$ is the generalized least squares residual from eq. (2); and b_1 and b_2 are the generalized least squares estimators appearing in cols. 1 and 2. Then R^2 is defined by

$$R^2 = \frac{b_1^2 S_{ee} + b_2^2 S_{uu} + 2b_1 b_2 S_{eu}}{\sum_t (SDP_t - SDP_.)^2},$$

and the contributions of $E\Pi$ and $U\Pi$ are $b_1^2 S_{ee}/\Sigma_t(SDP_t - SDP_.)^2$ and $b_2^2 S_{uu}/\Sigma_t(SDP_t - SDP_.)^2$, respectively.

‡ Averages across products of individual estimates and their standard errors (in parentheses).

contemporaneous correlation across different products.[7] On average, the best results, in terms of goodness of fit and parsimony, were obtained when the explanatory variables entered linearly. Quadratic specifications, although significant in some individual products, do

[7] This two-step procedure, first running actual inflation on a set of exogenous and predetermined variables and then using the predicted and residual values as regressors in an OLS regression of the SDP equation, provides consistent and efficient estimators of β_1 and β_2 in eq. (2), at least in the context of a single equation (Pagan 1984). The only problem lies in the computation of the variance of β_1, which is underestimated by standard OLS procedures. Notice, however, that this bias disappears under the null hypothesis: $\beta_1 = \beta_2$. Using an instrumental variable procedure to surmount this problem (Pesaran 1987) essentially produces the same numerical estimates as the two-step procedure, so that the downward bias in the variance estimator of β_1 is numerically insignificant. It is easy to show that the systemwide SUR estimators generated by the two-step procedure are still consistent. The main reason for not selecting the instrumental variable procedure is that the two-step procedure allowed us to use complete data from 1977:1 to 1984:9 to construct the expected and unexpected inflation and to run separate regressions for 1977–79 and 1980–84. This is technically more difficult to accomplish with the instrumental variable procedure since data for SDP in 1977, 1980, and 1983 are missing. For more details, see App. B.

not change the overall picture and are not presented here.[8] Hence, the models being estimated in each case are

$$\mathrm{SDP}_{it} = \beta_{i0} + \beta_{i1} E\Pi_{it} + \beta_{i2} U\Pi_{it} + \epsilon_{it} \tag{2}$$

and

$$\mathrm{SDP}_{it} = \alpha_{i0} + \alpha_{i1} \Pi_{it} + \mathbf{u}_{it}, \tag{2'}$$

where ϵ_t and $\mathbf{u}_t$ are 26×1 vectors of disturbances each of which is independently and identically distributed with zero mean and a possible nondiagonal covariance matrix.[9]

Table 2 clearly shows that in most products, expected inflation has a positive effect on price variability. In 24 out of the 26 products this coefficient is positive and significant. For unexpected inflation, b_2 is positive and significant in 22 products. In 17 products the estimated coefficient is higher in $E\Pi$ than in $U\Pi$, and this difference is statistically significant in nine cases.

In order to assess the explanatory power of the regression, we decompose the variation of SDP for each product into explained and unexplained parts. Since expected and unexpected inflation are orthogonal by construction, the explained part consists of $b_1^2 S_{ee} + b_2^2 S_{uu}$, where S_{ee} and S_{uu} are, respectively, the sum of squared deviations around their means of the expected and the unexpected inflation rates. Since the residuals are not orthogonal to the regressors, their covariance with $E\Pi$ and $U\Pi$ times their respective coefficient estimates is included in the unexplained part (see n. † in table 2). We define a goodness-of-fit measure for each product as the ratio of the explained sum of squares to the variance of SDP, which may exceed one. This is denoted by R^2 and is presented in column 3. Furthermore, this explained part can be attributed either to expected $(b_1^2 S_{ee})$

[8] In estimating the SUR equations, we did not allow for different numbers of observations in each product. The four products for which 1983 data exist were jointly estimated in a separate run. Their estimates are basically unchanged. The results presented in the tables correspond to the periods 1978–79, 1981–82, and 1984:1–1984:9. Quadratic specifications may be important because of Danziger's point discussed above. Regressing SDP against linear and quadratic expected inflation shows that the quadratic term is significantly negative in only three out of the 26 products. A simple calculation indicated that the estimated slope becomes negative in only a few months in late 1984, maybe because of the phenomenon pointed out by Danziger. Since this is a very minor portion of our data, we did not pursue this issue further, on the presumption that the overall results will not be affected.

[9] There is a group of products (3, 5, 9, 10, 12, and 14) that exhibit first-order serial correlation and fourth-order autoregressive conditional heteroscedasticity of their error terms. Excluding them from the SUR procedure does not alter the results of the remaining products whose disturbances are serially uncorrelated and conditionally homoscedastic. Since even in the case in which the estimated variances of these six products are doubled the main conclusions are not affected, we chose to ignore this issue.

or to unexpected ($b_2^2 S_{uu}$) inflation. A comparison of columns 4 and 5 reveals that whatever the explanatory power of the regression is, this additional explanation is due to the expected component of inflation rather than to the unexpected one.[10]

The estimated coefficient of expected inflation averages 0.43 (standard error 0.15) over all products,[11] which is 20 percent higher than the average value of the coefficient of the unexpected rate of inflation, 0.36 (0.16), and very similar to the average coefficient of actual inflation, 0.41 (0.14).[12]

Two additional sets of regressions were run using different measures of inflation. Their results are provided in table 3. Columns 1–3 indicate that aggregate inflation matters. Aggregate inflation is defined by the rate of change of the CPI (see App. B). The estimates indicate a pattern similar to the one found when the product's own inflation rates were used: expected aggregate inflation ($E\Pi_{\text{cpi}}$) has a positive effect on intramarket price variability. On average, across products, the estimated coefficient of $E\Pi_{\text{cpi}}$ is 0.50 and that of $U\Pi_{\text{cpi}}$ is 0.36.[13] A comparison of the estimated coefficients across products in tables 2 and 3 indicates that these averages mask a more variable response to aggregate inflation across products than to own inflation. The estimated coefficient of $E\Pi$ is larger than that of $U\Pi$ in 16 products, but significantly so in only five products. Columns 4 and 5 pre-

[10] The reason that cols. 4 and 5 do not add up to col. 3 is that the orthogonality of the regressors is not satisfied because of our use of a sample period in these regressions different from the sample period used in the construction of these variables (see App. B). In most cases, the discrepancy is less than 1 percent of the total.

[11] Here and in the rest of the paper, the standard error refers to the standard deviation of individual product estimates around their average.

[12] Restricting the same coefficients over all products results in almost identical estimates: 0.43 (0.012) for $E\Pi$ and 0.35 (0.016) for $U\Pi$. These restrictions are rejected by the data (F-value of 5.4 with [48, 1,275] degrees of freedom). As explained in the text, SDP may move in a direction opposite to that of a measure of the dispersion of price levels, and this happens because there are both positive and negative changes in prices. To partially cope with this problem, SDP was computed only for stores that had positive changes in prices (UPSDP). In this case (not presented here), the average coefficients of predicted and unpredicted inflation are quite similar, 0.24 (0.18) and 0.23 (0.19), respectively. Because of the large number of stores that do not change prices, especially at the beginning of the sample period, UPSDP is sometimes missing and sometimes based on a small number of observations. Weighing the observations by the latter did not change the results, and the large number of nonoverlapping missing values made it inappropriate to try to improve efficiency by SUR methods. Hence, these results are based on individual OLS regressions and are significantly less accurate than the results in table 2. Because of the large standard errors, only 15 products have significantly positive coefficients for $E\Pi$, and $U\Pi$ is significantly positive in only nine products. In 14 products the estimated coefficient is larger for $E\Pi$ than for $U\Pi$, but this difference is significant in only three products.

[13] These estimates were generated by the two-step procedure explained in Pagan (1984). An instrumental variable estimation method delivers estimates that are, numerically, very similar. See n. 7 for an assessment of this method.

TABLE 3

INTRAMARKET PRICE VARIABILITY

Dependent Variable: SDP

Product	$E\Pi_{cpi}$* (1)	$U\Pi_{cpi}$* (2)	R^2 (3)	$E\Pi_i$† (4)	$\lvert U\Pi_i \rvert$† (5)
1	.56	.20	.12	.59	−.45
	(2.7)	(.39)		(5.0)	(−1.3)
2	.38	.33	.36	.18	.22
	(5.3)	(1.9)		(3.9)	(2.3)
3	−.13	1.01	.08	.12	−.01
	(−.58)	(1.94)		(1.1)	(−.1)
4	.43	1.23	.21	.49	−.30
	(2.59)	(2.95)		(7.2)	(−1.6)
5	.53	.19	.30	.43	.30
	(4.64)	(.65)		(7.6)	(1.75)
6	.44	−.07	.30	.38	−.3
	(5.18)	(−.41)		(6.2)	(−2.5)
7	.55	−.39	.34	.38	−.07
	(4.68)	(−1.3)		(4.9)	(−.6)
8	.23	.15	.04	.56	−.14
	(1.64)	(.49)		(7.6)	(−.9)
9	.26	−.45	.11	.40	−.23
	(2.22)	(−1.8)		(5.4)	(−2.0)
10	.96	1.71	.50	.47	−.18
	(6.25)	(4.43)		(7.4)	(−1.0)
11	.41	.23	.48	.34	−.07
	(6.81)	(1.57)		(9.8)	(−.7)
12	.83	.84	.29	.63	−.37
	(4.42)	(1.79)		(4.3)	(−1.3)
13	.45	.58	.44	.39	−.15
	(5.93)	(3.04)		(7.1)	(−1.9)
14	1.61	.33	.35	.54	.30
	(5.79)	(.57)		(3.6)	(.9)
15	.64	.85	.45	.42	.63
	(5.98)	(3.2)		(7.4)	(5.2)
16	.52	.59	.44	.44	.38
	(6.08)	(2.76)		(9.5)	(4.2)
17	.24	.37	.21	.25	.15
	(2.83)	(1.84)		(4.0)	(1.2)
18	.17	.32	.09	.09	.09
	(1.85)	(1.44)		(1.7)	(1.0)
19	.50	.46	.16	.42	−.1
	(3.01)	(1.11)		(6.8)	(−.7)
20	.31	.21	.05	.46	−.04
	(1.55)	(.43)		(5.1)	(−.2)
21	.20	−.05	.03	.31	.02
	(1.24)	(−.14)		(4.6)	(.1)
22	.46	−.27	.18	.31	.35
	(3.07)	(.71)		(4.6)	(3.1)
23	.55	.20	.21	.27	.37
	(3.72)	(.55)		(3.7)	(3.3)
24	.42	.43	.14	.20	.04
	(2.81)	(1.14)		(2.5)	(.5)
25	.89	.27	.10	.16	.04
	(2.34)	(.28)		(.6)	(.1)

TABLE 3 (*Continued*)

Product	$E\Pi_{cpi}$* (1)	$U\Pi_{cpi}$* (2)	R^2 (3)	$E\Pi_i$† (4)	$\|U\Pi_i\|$† (5)
26	.68	.19	.24	.52	−.17
	(4.03)	(.44)		(7.3)	(−1.4)
Average‡	.50	.36	.24	.37	.01
	(.32)	(.48)	(.15)	(.15)	(.26)

NOTE.—t-values are in parentheses.
* The equations were estimated separately by OLS since the regressors are identical across equations. There are 54 observations for all products except products 6, 8, 9, and 14. These have 66 observations corresponding to the additional year 1983. Expected and unexpected CPI inflation are defined in App. B.
† The equations for each product were estimated jointly using the SUR procedure, with the exception of product 17, for which 1978 data were missing. Expected and unexpected own-product inflation are defined in App. B.
‡ Averages across products of individual estimates and their standard errors (in parentheses).

sent estimates of the coefficients in a regression in which the explanatory variables are $E\Pi_{it}$, as in table 2, and the absolute value of $U\Pi_{it}$, denoted by $|U\Pi_{it}|$. The motivation behind this specification is that $|U\Pi_{it}|$ may be used as a proxy for the variance of the inflation process.[14] In only four products is the coefficient of $|U\Pi_{it}|$ numerically larger than the coefficient of expected inflation. On average, $|U\Pi_{it}|$ has no discernible positive effect on relative price variability.

These results show that expected inflation affects *intramarket* relative price variability. They also suggest that the expected component of inflation has more explanatory power than the unexpected one. This, in turn, implies that using actual inflation data as an explanatory variable captures mostly the effect of expected inflation and not of Lucas-type confusion effects. In this sense, the results presented here are consistent with the predictions of menu cost models.

IV. The Behavior of Prices

There are at least two reasons why the characterization of the behavior of prices is extremely important for economic analysis: The first reason has to do with the welfare costs of inflation, and the other is concerned with the constraints put by the behavior of rational price setters on an active monetary policy. The second issue will guide us in our description of the behavior of prices. It can be defined more precisely by the following questions: Do prices lag behind money in a way that can be used by the government to pursue an active monetary policy? Can expected changes in the money supply affect output? In the following subsections we describe four aspects of price behavior that have a bearing on these questions: the duration of price quota-

[14] We thank an anonymous referee for suggesting this point.

tions and the magnitude of price changes (subsection A), the synchronization of price changes (subsection B), and the shape of the distribution of real prices (subsection C). Although each aspect is important in itself, it is the interaction between them that is crucial for the proper understanding of nominal rigidities.

A. *Duration of Price Quotations and the Size of Price Changes*

Over all products and stores, the duration of a price quotation in 1978–79 was 1.9 months, and in 1981–82 it was 1.6 months.[15] That is, prices lasted approximately 20 percent less in 1981–82 than in 1978–79. This is not surprising if one takes into account that the CPI monthly inflation rate rose from 4.9 percent in 1978–79 to 6.6 percent in 1981–82, a 35 percent increase. In addition, this empirical observation is implied by theory: Menu costs models, assuming a not too asymmetric profit function, imply that prices change more frequently when expected inflation increases, even though the (S, s) band itself increases (Rotemberg 1983; Tsiddon 1991).

Table 4 presents several results on the direction and size of price changes and on the duration of price quotations for each product in two periods, 1978–1979:6 and 1982.[16] First, we ask the following question: Within each period, what proportion of all price quotations are changing from month to month? And, in which direction are they changing? For each product, we classified all nonmissing observations, corresponding to each store-month combination, into one of three categories according to the direction in which the price changed from the previous month: it either decreased, remained constant, or increased. In columns 1, 2, 7, and 8, we present the percentage of all

[15] The duration of a price quotation is defined as the number of months that elapsed between two different price quotations, provided that there are no missing values in between.

[16] These periods correspond to the two extremes of the sample in terms of the level of inflation. Restricting the first period to the year 1978 reduces the number of observations considerably, since in 1978 there were significantly fewer price changes. Hence, it was extended till June 1979, which corresponds to the same inflationary step. On the other hand, adding 1981 or parts of it to 1982 considerably increases the standard errors, presumably because of the reduction in duration from 1981 to 1982. In 1984 the duration recorded for most prices was 1 month. This is, of course, a result of the discrete feature of the data. In all likelihood, prices changed more than once a month. Although it may be possible to deal effectively with this type of truncation, we chose to exclude the observations for 1984 from the analysis in this section.

observations falling into the no change (No) and upward changes (Up) categories. The first striking observation is that during 1978–1979:6, in most products, prices are quite stable in the sense that in more than 50 percent of the opportunities recorded (number of stores times 18 months), stores opted not to change prices. Such stability is less common in 1982 but is still not negligible. The average over products of cases in which the price was not changed is 61 percent in the first period and decreases to 39 percent in 1982. These nontrivial figures can be taken to be evidence favorable to the menu cost approach. It is interesting to note that even when inflation is always positive there are instances in which prices are reduced in nominal terms. This could be taken to be indicative of the presence of idiosyncratic shocks affecting the stores (Tsiddon 1988). This, however, does not happen very often: 5 percent of all price quotations in both periods.

In computing the duration of a price quotation, we are faced with the problem that the observed values of durations are truncated at 1 month, the period of time between observations. Hence, all inferences on duration are made on the basis of its truncated distribution. In order to unravel the nontruncated moments, we would have to appeal to strong distributional assumptions. We do not proceed this way; instead we assume that the truncated distribution shifts between periods in the same way as the underlying nontruncated distribution. We believe that our inferences on the changes in the distribution of durations between periods would not be subject to such a severe truncation bias. In this fashion, we hope to capture the qualitative changes between periods of low and high inflation. Since, in both periods, the observed median duration is around one, this results in a very nonsymmetric distribution of observed durations. Hence, in columns 3–5 and 9–11 we present the second and third quartiles and the mean of the truncated distribution of durations. In both periods and in all products the first quartile is 1 month. A close look at the figures reveals that, although there is heterogeneity across products in the duration of prices, in almost all products, the high tail of the distribution is squeezed down. The average duration of a price quotation across products and stores was reduced by 40 percent between 1978–1979:6 and 1982, from 2.2 to 1.5 months, as can be seen in columns 5 and 11. This reduction reflects a decline in the average duration in each product, sometimes by as much as 40 percent (products 1, 4, 13, 14, 20, 24, and 26) and by as little as 12 percent (product 6).

Out of the 10 products whose price quotations had an average duration of over 2.5 months in 1978–1979:6, seven of them are

TABLE 4

DURATION AND CHANGE OF PRICE QUOTATIONS

	A. 1978–1979:6						B. 1982					
	CHANGES (%)*		DURATION (Months)			(S, s) BAND (%):†	CHANGES (%)		DURATION (Months)			(S, s) BAND (%):
PRODUCT	No (1)	Up (2)	50% (3)	75% (4)	Mean (5)	MEAN (6)	No (7)	Up (8)	50% (9)	75% (10)	Mean (11)	MEAN (12)
1	76	16	3	6	4.1	4.4	45	49	1	2	1.6	10.5
2	43	53	1	2	1.6	7.1	37	61	1	2	1.3	11.1
3	60	33	1	3	2.0	6.0	29	67	1	2	1.4	10.4
4	85	14	3	4	3.0	14.0	35	62	1	2	1.5	11.8
5	53	42	2	3	2.0	8.8	45	50	1	2	1.5	10.6
6	48	50	1	2	1.7	7.2	39	54	1	1	1.5	10.0
7	69	26	2	3	2.2	8.0	47	50	2	2	1.75	12.7
8	41	44	1	2	1.6	6.8	20	76	1	1	1.2	10.3
9	51	39	1	2	1.8	7.5	38	54	1	2	1.4	11.8
10	43	51	1	2	1.6	7.8	32	61	1	2	1.4	10.5
11	53	44	1	2	1.9	8.1	40	57	1	2	1.5	10.9

12	39	55	1	2	1.5	7.2	21	75	1	1	1.2	12.1
13	56	39	1	3	2.2	7.1	26	72	1	2	1.3	11.2
14	70	25	1	4.5	2.7	6.4	47	43	1	2	1.6	6.9
15	39	55	1	2	1.4	7.7	25	73	1	1	1.2	11.4
16	48	49	1	2	1.7	7.5	37	59	1	2	1.4	10.4
17‡	53	45	1	2	1.5	7.8	33	59	1	1	1.3	9.9
18	40	49	1	2	1.6	6.2	18	79	1	1	1.2	10.8
19	72	21	2	3	2.6	14.7	53	44	1	2	1.8	14.4
20	80	17	2	4	2.8	10.9	49	49	1	2	1.2	12.5
21	76	19	2	4	2.4	11.7	53	43	1	3	1.9	13.0
22	77	20	2	3	2.6	11.8	49	48	1	2	1.8	12.1
23	76	22	2	3.5	2.5	12.6	51	46	2	2	1.7	14.3
24	78	19	3	4	3.1	14.4	43	54	1	2	1.6	12.6
25	76	20	2	4	2.5	12.1	52	47	1.5	2	1.7	14.1
26	78	19	2	4	2.8	14.0	44	53	1	2	1.6	12.8
Average§	61	34	1.6	3	2.21	9.1	39	57	1.1	1.85	1.48	11.5
	(16)	(14)	(.7)	(1.1)	(.65)	(3.0)	(11)	(11)	(.3)	(.5)	(.21)	(1.6)

* Percentage of all observations in which price quotations remained constant (No) or changed upward (Up).
† Arithmetic mean of all nonzero DP_{ijt}, by product.
‡ Since data for 1978 are not available, figures are for January to June 1979.
§ Average across products of individual estimates and, in parentheses, their standard errors.

wines and liquors.[17] These are also the products that exhibited the largest reduction in the average durations of their price quotations. This group of products seems to behave differently than the other food items. This will be seen more clearly shortly, when the magnitude of the price changes is analyzed. However, to understand the reasons that make these products different, a more thorough study is called for.

Since the monthly inflation (CPI) during 1978–1979:6 was 3.9 percent, roughly speaking, the real price of the product eroded by 8.6 percent before the nominal price was changed. Correspondingly, with a CPI increase of 7.3 percent in 1982, the real price decreased, on average, by 11 percent before the store decided to revise it. These figures are quite large, especially when one takes into account that most products in the analysis are basic foodstuffs, products usually believed to have many substitutes. Yet their real price can change by about 8.5–11 percent before their nominal price is adjusted upward. The large magnitude of the real price erosion does not seem to be an isolated phenomenon of the particular data set utilized. Cecchetti (1986) reaches a similar conclusion using price data on American magazines: real prices erode by as much as 25 percent before they are adjusted upward. These results strongly suggest that nominal rigidities are a common phenomenon.

As shown in Appendix A, DP_{ijt} is approximately equal to the (S, s) band $(= S - s)$ of product i sold by store j at some time between t and $t - 1$, in percentage terms. Hence, averaging over all price changes provides an estimate of the average length of the (S, s) band, which should not differ much from the figure of 8.5–11 percent mentioned above. In columns 6 and 12 of table 4 we present this average change. We see, in fact, that the length of the (S, s) band was on average 9.1 percent in 1978–1979:6, when monthly inflation was 3.9 percent, and 11.5 percent in 1982 with a 7.3 percent monthly inflation rate. Most products, except wines and liquors, show substantial increases in their (S, s) band lengths. In the latter group, the average change

[17] In nine of these 10 products, the median duration is greater than or equal to 2 months. All wines and liquors are made in Israel, and therefore their prices are not directly related to the exchange rate. The other three products are tea bags, sweet challah bread, and rice. One possible cause of this may be the different market structures of these products. If there are reasons to believe that, because of the greater possibility for product differentiation among wines and liquors, competition in these markets may be less stringent than in the markets for fresh chicken or beef, the reported differences in the duration of price quotations are not unexpected. Menu cost models imply that the less concave the profit function around the optimal price (i.e., the more monopoly power a price setter has), the wider the range of real prices he accepts without changing the nominal price (i.e., the longer is the duration of a price quotation).

of price did not change as much and actually decreased in some items (they were very high to begin with in 1978–1979:6). The picture that emerges from the data shows that the response to inflation of stores selling wine and liquors is mainly to shorten the duration of the nominal price quotations; in other products, the elicited response is more like an increased magnitude of the nominal price change, the (S, s) band.

The preceding analysis did not differentiate between increases and decreases of the nominal price. Restricting ourselves to the upward changes only, which constitute 95 percent of the sample, we obviously find that the average change in price is larger in both periods. However, the increase in the average size of the change between the periods is somewhat smaller. On average, the increase in price amounted to 12.3 and 12.9 percent in each period, respectively. As before, wines and liquors behave differently, and, in fact, the average increase in their price was actually reduced between 1978–1979:6 and 1982.

Finally, one may ask whether the data indicate the presence of systematic differences across stores in the size of the price change and in their durations. In other words, are there store effects? Intuitively, we would expect to find that either there are no store effects at all, neither in the size of the price change nor in their duration, or, if there are any store effects in the size of the price change, they would have to be positively correlated with corresponding effects in the duration of nominal price quotations. For each product and in each period, we separately tested the equality of store-specific means of price change and of their duration, $X_{ij.} = (1/T_{ij}) \Sigma_t X_{ijt}$, where X is either the change in price (DP) or the duration of the nominal price quotation and T_{ij} is the number of nonmissing observations of product i in store j. Regarding the magnitude of the price change, the F-test cannot reject the null hypothesis of equality among stores in basically all products.[18] Note that this is still consistent with possible store effects in the price *level*. The F-test results concerning the equality among stores in the average duration of their price quotations are more ambiguous. In 1978–1979:6, the null hypothesis is rejected in seven products, whereas in 1982 it is rejected on 12 occasions. These rejections should be taken with a grain of salt because they may be caused by a small group of stores that differ too much from the others.[19] In any case, the simple correlation between the mean change

[18] In 1978–1979:6, in four products the null is rejected at the 5 percent but not at the 1 percent significance level. In 1982, the null hypothesis is rejected in only one case, product 8.

[19] This is conceivable since the long tail of the truncated distribution of durations implies the presence of some extreme values. This means that, given a period of fixed

in price with the mean duration per store over all products is significantly different from zero: .33 in 1978–1979:6 and .60 in 1982.[20]

B. The (A)Synchronization of Price Changes

The preceding subsection reported that a substantial number of prices remain constant for at least 2 months. At the existing inflation rates, this is not a trivially short period of time. Thus whether the timing of price changes is synchronized across firms or not is an important issue for the analysis of the real effects of monetary policy.

Suppose that there is a monetary expansion that affects the demand faced by each store and that, in response, stores increase their prices. Suppose, further, that stores do not act simultaneously; that is, price changes are not fully synchronized across stores. Then on changing its price, each store must take into account the fact that other stores have not yet done so. Were the store to ignore this in setting its price, its price would be too high and it would lose customers. If price changes were fully synchronized, stores would still take into account the behavior of rival stores; in equilibrium, however, the aggregate price will immediately increase so as to leave real balances unchanged.[21]

In many models it is this lack of synchronization in the pricing decision across stores that makes the lag in the aggregate price level longer than the duration of an average price quotation.[22] This result is certainly not specific to the menu cost model; it can be equally generated by other forms of nominal rigidities.

In trying to assess the degree of synchronization in price changes with our data, we are faced with the problem that the period of

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length (18 or 12 months), the averages computed for the stores that have these long durations are based on a small number of observations. Hence, these extreme values may not be averaged out over time. The null hypothesis is rejected at the 5 percent significance level in products 1, 2, 3, 5, 9, 16, and 17 in 1978–1979:6 and in products 5, 7, 8, 12, 14, 16, 18, 19, and 21–24 in 1982. However, it cannot be rejected at the 1 percent significance level in products 14, 16, 21, and 22 in 1982.

[20] At the individual product level, in the first period, in only two cases is this correlation negative, but not significantly so. In another eight products, the correlations are positive but also not significantly so. In 1982, all but two correlations are significantly positive.

[21] Under ideal conditions, of course. It is implicitly assumed that a monetary expansion has an instantaneous effect on demands. As a referee pointed out, it takes time for money to change hands, and thus money supply may have real effects on relative prices.

[22] See Fischer (1977). This statement is not always correct. In Caplin and Spulber (1987), e.g., there is no synchronization at all, and aggregate price adjusts immediately to changes in the money supply. Their assumption on the distribution of prices plays a key role in deriving this result (see subsection *C*).

observation is 1 month. As inflation increases, from 1981 onward, the duration of a price quotation is reduced and may be less than 1 month. This implies that we would observe a large proportion of firms changing prices each month, giving the impression of a high degree of synchronization across stores. For an annual rate of inflation of 130 percent in 1982, this intepretation is problematic.[23] We therefore proceed to analyze the relatively low-inflation period of 1978–79, which still had a 60 percent yearly inflation. For each month and for each product, we compute the proportion of reporting stores that changed their prices in that month.[24] These proportions appear in figure 1. For the 23 months between 1978:2 and 1979:12, this proportion averages .395, with a standard deviation of .113. This implies that each store changes its price every 2.5 months or that each month, on average, 40 percent of the stores change their prices. If price changes were fully synchronized, we would expect these proportions to vary considerably over time. In the extreme case, we should observe a sequence of zeros (no price changes) followed by a one (all stores change prices simultaneously), another sequence of zeros, and so on. A simple calculation shows that the standard deviation implied by such a pattern is four times as large as the estimated one.[25] Thus the relatively small variance in the estimated proportions is indicative of less than perfect synchronization; in other words, price changes are staggered. The pattern seems to be cyclical, perhaps because of seasonalities (observe the large November effect); starting in mid-1979, there also seems to have been, not surprisingly, a positive trend.

Another way of checking the degree of synchronization is by looking at the serial correlation among the estimated proportions. When prices change every 1.9 months on average, full synchronization implies some negative autocorrelation of order less than three; perfect staggering would imply no autocorrelation at all. None of the estimated autocorrelations is statistically significant, perhaps because of

[23] It is not the same to have all firms changing their prices within a month when the monthly inflation is 1 percent as when it is 8 percent. This is not to say that synchronization does not increase with inflation. There are many reasons why this must be so. It is true of both menu cost–based (Tsiddon 1988) and information-based (Van Hoomissen 1988*b*) models. We merely say that the finding that most stores change prices within the month, as it occurs in the 1980s, should be interpreted with care.

[24] We concentrate on only upward changes since we want to relate to the effects of an aggregate shock and downplay the role of idiosyncratic factors. Note, also, that we are not dealing with the interesting question of comovements in the prices of different products sold by the same store.

[25] Suppose that prices change every 3 months (eight times in 2 years); the average proportion is then .333 with a standard deviation of .481.

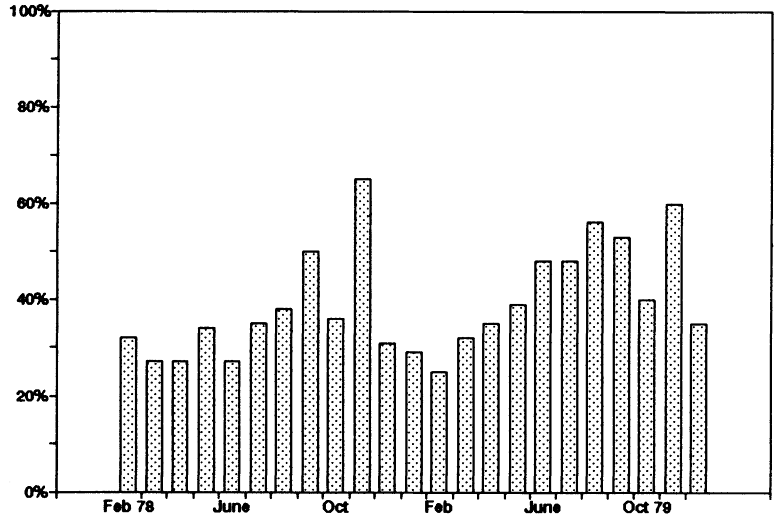

FIG. 1.—Proportion of price changes: (a)synchronization of price changes (February 1978 to December 1979).

the small number of observations.[26] Nevertheless, the first three autocorrelations are positive, indicating that price changes are not fully synchronized.

Yet another confirmation of the lack of synchronization in price changes comes from the cross-correlations between changes in the money supply and the proportion of price changes. Since we do not know the speed at which changes in the money supply are transmitted to the food market, we look at cross-correlations between these two series, up to four lags. Intuitively, if there is full synchronization, some of these cross-correlations should be positive, but we find no evidence of this: all four cross-correlations are negative or very close to zero.[27]

These results indicate that in food markets, characterized by inexpensive and relatively homogeneous products, and in a period in which yearly inflation was 60 percent, price changes do not cluster. Stores do not coordinate the timing of their price changes, not even implicitly.

[26] The estimates of the first five autocorrelations are .29, .39, .03, −.11, and −.29, with long standard errors on the order of .25.

[27] The estimates for these cross-correlations are .09, −.44, −.26, −.05, and −.11 for the contemporaneous changes and four lags of changes in the money supply. Standard errors are approximately .20.

C. The Distribution of Real Prices

We have shown that stores maintain constant their nominal prices for relatively long periods of time and that when they do update them they do not do so simultaneously. Do these two characteristics of firm behavior carry any macroeconomic implications? Specifically, suppose that there is a change in the money supply. Does the presence of lags in the adjustment of the nominal price at the firm level imply a lag in the adjustment of the aggregate price level? In order to draw any conclusions on the behavior of the aggregate price level, we must know the distribution of real prices. This is the topic investigated in this subsection.

In using our price data to determine the shape of the distribution of real prices, we must first define what we mean by real prices. The price data, after being transformed into natural logarithms, are denoted by lowercase letters. The real price of product i in store j during the tth month, Y_{ijt}, is the deviation of the log of the nominal price, p_{ijt}, from the log of the (geometric) mean of product i's prices across all stores sampled that month: $Y_{ijt} = p_{ijt} - p_{i.t}$, where $p_{i.t} = (1/N_{it}) \Sigma_j p_{ijt}$. Thus Y_{ijt} is approximately equal to the percentage deviation of the store's price from the (geometric) mean price in month t.

The distribution of real prices in which we are interested is the cross-sectional (monthly) distribution of Y_{ijt}, for each product i. Suppose, first, that we concentrate on a single month for a particular product. Using data on Y_{ijt} to analyze its distribution may be problematic owing to the presence of store heterogeneity in the price level.[28] Heterogeneity of optimal prices across stores may result from different demand and supply conditions in different locations or, to a lesser extent, from the systematic differences in sampling dates over the month. It is first assumed that this heterogeneity is reflected only in a time-invariant mean real price for each store, $EY_{ijt} = \mu_{ij}$. Hence, we can write $Y_{ijt} = \mu_{ij} + \epsilon_{ijt}$, where $\epsilon_{ijt} = Y_{ijt} - EY_{ijt}$.[29] If theoretical models on pricing behavior ignore store heterogeneity, then their conclusions may be interpreted as saying something about the distribution of ϵ. In this case, one would like to recenter the data and analyze the deviations of each store's real price from its expected value, or the within-store real price variation, $Y_{ijt} - \mu_{ij}$. On the other hand, if one's goal is to compare distributions in different periods,

[28] Recall that in the previous subsection we did not find significant store effects in price *changes*.

[29] Since Y is the log of nominal price, this formulation implies that, for a given realization of ϵ, the price of product i in store j deviates by a fixed μ_{ij} percent from the (geometric) mean price. In terms of the nominal price (not in logs), we can write $P_{ijt} = \exp(p_{i.t} + \mu_{ij} + \epsilon_{ijt})$.

these store effects do not have to be removed provided, of course, that the set of stores does not change much between the periods. This procedure has its advantages because the store effect formulation may not be satisfactory or because the store effects are believed to be partly endogenous and therefore not independent of the rate of inflation. In this case the data should be left uncentered.[30] We look at both possibilities.

Next, we address the time-series dimension of the data. The average number, over all products, of price quotations per month ranges from 10 to 44. The first nine months of 1984 are omitted from the analysis, so that only 48 months remain, corresponding to the years 1978–79 and 1981–82. This presents two problems: first, there are not enough data in each month for the results to be robust; second, the analysis and interpretation of 48 cross-sectional distributions for each of the 26 products are untractable. Thus we aggregated the data into the two subperiods analyzed previously corresponding to low and high inflationary steps, 1978–1979:6 and 1982. Here, aggregation means that in each case we look at the sum of frequency counts over 18 or 12 months. However, even if the distribution of $Y_{ijt} - \mu_{ij}$ has the same zero mean over time, it may not have the same variance over time if the latter is affected by the rate of inflation. Since there is still considerable variation in the inflation rates within each subperiod, it may be the case that the variance of Y_{ijt} is not constant over time. To handle this problem, we assume that, for each product i, the variance of Y_{ijt} changes over time but that in every month all stores have the same variance. Hence, for all j in N_{it}, $\text{var}(Y_{ijt}) = \sigma_{it}^2$. Thus to make sense of the time aggregation under this circumstance, we should look at the distribution of the following variable:

$$Z_{ijt} = \frac{Y_{ijt} - \mu_{ij}}{\sigma_{it}}. \qquad (3)$$

Notice that the expected value and variance of the standardized real price, Z, are the same across products. Provided that these two moments determine all other higher moments of the distribution of Z, one can go a step further and argue that aggregating again over products does not change the distribution.[31] Aggregation now means adding the frequency counts of Z over different products. This is the procedure adopted here.

For example, if the store's real price Y_{ijt} has expected value μ_{ij} and variance σ_{it}^2 and comes from a uniformly distributed population, then Z_{ijt} is also uniformly distributed on the interval $(-\sqrt{3}, \sqrt{3})$ for each

[30] We thank a referee for pointing out this issue.

[31] This is satisfied, for example, by the normal and uniform distributions.

product. Under this hypothesis, a plot of the frequency counts of all data points in all products should have an approximately rectangular shape over that interval; otherwise this would constitute evidence against the null hypothesis that real prices are uniformly distributed.

Since the parameters μ_{ij} and σ_{it}^2 are unknown, they have to be estimated. It can be shown that, as T and N tend to infinity, the estimators in equation (4) are consistent for μ_{ij} and σ_{it}^2:

$$Y_{ij.} = \frac{1}{T_{ij}} \sum_t Y_{ijt}, \quad \mathrm{STD}_{it} = \left[\frac{1}{N_{it}} \sum_j (Y_{ijt} - Y_{ij.})^2 \right]^{1/2}, \tag{4}$$

where T_{ij} is the number of nonmissing price quotations of product i in store j.[32] Hence, the time-series dimension of the data is used to estimate the store effects and the cross-section dimension to estimate the variance of real prices.

In analyzing the distribution of real prices, we shall take the uniform distribution as a benchmark case since it has been the focus of recent research in aggregate (S, s) behavior. Caplin and Spulber (1987) derive the result that money is neutral in a model in which all price setters follow a constant (S, s) rule, the initial distribution of (the log of) real prices is uniform, and shocks to the money supply are nonnegative and have no idiosyncratic component. At any instant, the number of firms that change their prices, as they pass the threshold, multiplied by the change in their real price is just sufficient to compensate for the deterioration of real prices in all firms that do not change their prices. The average (aggregate) price level, therefore, equals the (normalized) quantity of money at every instant. This conclusion strongly depends on the distributional assumption on (the log of) real prices. This result was later strengthened by Caballero and Engel (1989*b*), who show that adding nonnegative idiosyncratic shocks makes the uniform distribution the unique stable equilibrium distribution, thereby dispensing with the arbitrary assumption of uniformity of the initial distribution.[33] In equilibrium, then, discrete price adjustments at the firm level do not carry over and generate a lag in the response of the aggregate price level to changes in the money supply. If these models, or their assumptions, are approximately correct, then plots of the frequency counts of the standardized

[32] For Y_{ij} to be consistent, it is sufficient that $\{Y_{ijt}, t = 1, 2, \ldots\}$ be a stationary sequence with mean μ_{ij}, satisfying $\gamma(T_{ij}) \rightarrow 0$ as $T_{ij} \rightarrow \infty$, where γ is the autocovariance function of Y. For STD_{it} to be consistent, it is sufficient to require that $\{(Y_{ijt} - \mu_{ij})^2, j = 1, 2, \ldots\}$ be a stationary ergodic sequence with finite mean or that its elements be asymptotically uncorrelated.

[33] They also qualify their result by showing that, when out of equilibrium, firms' lags are transmitted to the aggregate price level and that heterogeneity across firms speeds up convergence to equilibrium (Caballero and Engel 1989*a*).

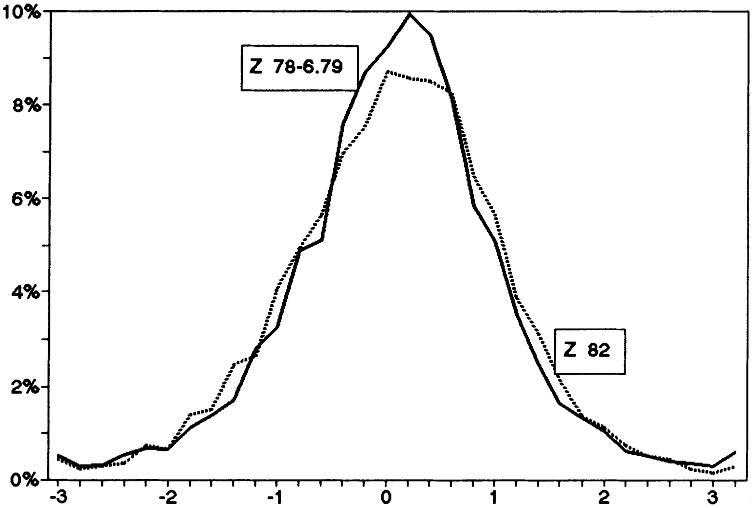

FIG. 2.—Distribution of real prices—Z

real price should be approximately rectangular over the interval $(-\sqrt{3}, \sqrt{3})$.

Figure 2 presents histograms of standardized real prices, Z_{ijt}, for all 26 products taken together, in each of the two periods (table 5 presents summary statistics). Recall that, by construction, Z_{ijt} has zero mean and unit variance. The histograms give us the impression that the distributions of standardized real prices are unimodal. One may even conjecture that they are symmetric or slightly skewed to the left. The apparent symmetry of the estimated quantiles and the low negative values of the skewness coefficients tend to confirm this. Formally, we use the observed moment ratio $b_1 = m_3/m_2^{3/2}$, where m_k is the kth sample moment about the sample mean, as a base for a symmetry test. The standardized value of b_1 has a limiting standard normal distribution, and its value is denoted by S (Gupta 1967). The results of the test imply that the data reject the hypothesis that the distribution is symmetric about zero in favor of the alternative that it is skewed to the left in 1982 but not in 1978–1979:6. Apparently, even the small observed differences in the quantiles cannot be supported under symmetry given the large number of observations used.[34] The following features of the distribution of real prices are worthwhile mentioning. First, the estimated median, for all practical purposes, coincides with the zero mean. Second, the left skewness

[34] One should be cautious with these formal tests since they assume independently and identically distributed variables.

TABLE 5

PROPERTIES OF THE DISTRIBUTION OF REAL PRICES—Z

Period	N	Skewness	Kurtosis	Max	95%	90%	75%	50%	25%	10%	5%	Min	S
1978–1979:6	10,965	−.06	1.58	6.27	1.61	1.16	.56	.028	−.56	−1.21	−1.67	−5.32	1.09
1982	7,080	−.13	.81	5.29	1.58	1.20	.63	.03	−.61	−1.25	−1.68	−4.72	2.90

NOTE.—$P > D_N < .01$ and $P > D_U < .01$.

means that there is a mass concentration, a hump, above zero that seems to occur between the 75 and the 95 percent quantiles. In other words, the upper tail is thinner than the lower tail. Note also that, in general, the third quartile is further away from zero than the first quartile. All this points toward the following: There are equal chances of finding real prices above or below the market average, but upward deviations in the real price are further away from zero than downward ones. On the other hand, *extreme* deviations from the market average are more likely to be downward. These features better describe the distribution of prices in 1982 than in 1978–1979:6. Visual inspection of the histograms indicates that, for all products taken together and in all periods, the distribution of (standardized) real prices is *not* uniform. Also, it is *not* normal as the Kolmogorov-Smirnov test results confirm (their respective tail probabilities are denoted by $P > D_u$ and $P > D_n$). Finally, the results of separate analysis of the histograms and related statistics for each of the 26 products for the four years 1978–79 and 1981–82 are consistent with the aggregate picture.

Figure 3 provides information on the nonstandardized real price Y_{ijt} (summary statistics appear in table 6). Examining Y_{ijt} serves two purposes: First, it reinforces our previous conclusions on the distribution of real prices in the sense that they are not a result of data manipulations. Second, it has the advantage that it lends itself to straightforward interpretation: 50 percent of the real prices lie within 8–10 percent of the average price, and 80 percent lie within 20–25 percent of it. These are nonnegligible differences. For a product with an average price of one, we have a 50 percent probability of paying less than 0.9 or more than 1.1. The summary statistics also confirm that the dispersion of real prices increases with inflation. First, the variance of the distribution increases by 50 percent. Second, in 1978–1979:6, with 3.9 percent monthly inflation, 25 percent of the prices are 8 percent below the mean and 25 percent are 9 percent above the mean. In 1982, when inflation per month is 7.3 percent, the first and third quartiles are -10 and 12 percent, respectively. In 1978–1979:6, the distribution is symmetric, but in 1982 it is skewed to the left. This means that, while in the low-inflation period deviations from the mean price in both directions have the same probability, by the time we reach 1982 with an inflation rate that is almost doubled, the probability of being 10 percent above the mean price is higher than that of being 10 percent below. In both periods, the distributions are neither uniform nor normal.

The empirical results show that real prices are not uniformly distributed. This conclusion is robust to different transformations of the

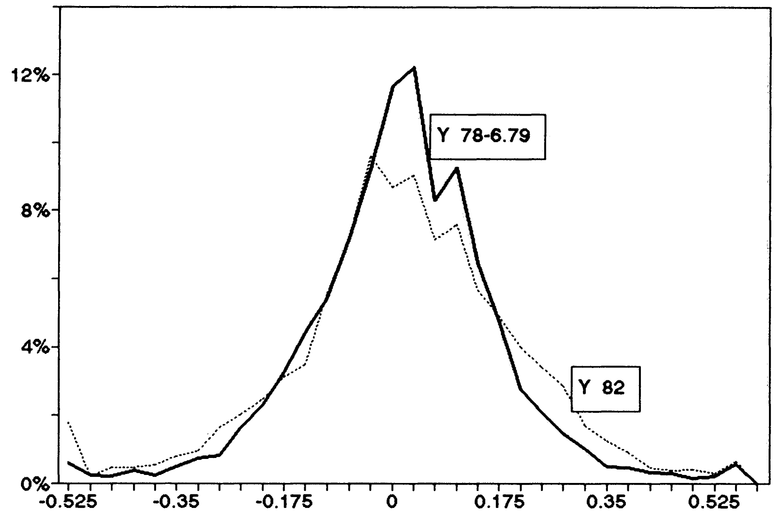

FIG. 3.—Distribution of real prices—*Y*

raw price data.[35] To conclude, however, that it indicates that (*S*, *s*) behavior is not supported by the data is not warranted. The uniform distribution is the result of aggregate (*S*, *s*) behavior only under very specific assumptions. If these are not met, then the conclusion that the distribution of real prices is uniform does not have to follow. In fact, Tsiddon (1988) shows that if idiosyncratic and aggregate shocks follow a Brownian motion, implying positive and negative values, then the stationary equilibrium distribution is not uniform. If, in addition, shocks do not have zero mean, the distribution of real prices is not symmetric.[36] Two-sided menu cost models such as this provide additional predictions that can be tested by the data.

First, as expected inflation increases, real prices spend more time closer to the lower threshold, implying that we should observe a decrease in the lower quantiles, that is, a shift in the mass of the distribution toward the lower tail. Second, in these models, the distribution of the durations of price quotations is directly related to the distribution of real prices. In a sense, they are mirror images of each other: if the distribution of real prices becomes more humped to the left as

[35] For additional results, see Lach and Tsiddon (1990).

[36] If shocks are two-sided, then there are two thresholds and one return point. The additional threshold emerges when the real price increases too much in the opposite direction to that of the trend. The shape of the stationary distribution in the case of Brownian motion without drift is triangular.

TABLE 6

PROPERTIES OF THE DISTRIBUTION OF REAL PRICES—Y

Period	N	Var	Skewness	Kurtosis	Max	95%	90%	75%	50%	25%	10%	5%	Min	S
1978–1979:6	10,965	.027	.06	6.86	2.33	.24	.17	.09	.002	−.08	−.18	−.26	−1.07	.23
1982	7,080	.041	−.58	2.96	1.01	.3	.24	.12	.002	−.10	−.23	−.33	−1.16	6.7

NOTE.—$P > D_N < .01$ and $P > D_U < .01$.

inflation increases, so does the distribution of durations, and vice versa.[37]

Can we detect a shift in the distribution of real prices between 1978–1979:6 (3.9 percent monthly inflation) and 1982 (7.3 percent monthly inflation) in the direction predicted by the theory? The answer is a qualified yes. Figure 2 indicates that the interquartile range is larger in 1982 than in 1978–1979:6, meaning that prices were pushed to *both* tails of the distribution. The change in the distribution of Y in figure 3 seems to obey the same pattern as for Z. However, these changes in the shape of the distribution do not seem to be dramatic, and although not inconsistent with the theory, they are probably not strong enough to support any decisive answer. At this stage, therefore, we find the evidence on this issue only suggestive.

The other prediction from two-sided menu cost models is related to the distribution of durations. The theoretical results were derived for positive changes in price (Tsiddon 1988). Thus figure 4 and table 7 present statistics on the distribution of durations between upward changes of nominal prices only.[38] We address two issues: the asymmetry of the distribution and the degree to which it changes between 1978–79 (low inflation) and 1981–82 (high inflation). Looking first at the four years together, we see that the distribution is clearly skewed to the right; it has a positive skewness coefficient. Comparing the two subperiods, we observe an increase in the coefficient of skewness from 0.86 to 1.34. This is a direct consequence of the flattening in the upper tail and the increase in the lower tail of the distribution: the upper 10 percent of the durations exceeded 7 months in 1978–79 but exceeded only 3 months in 1981–82; the lower 25 percent were up to 2 months in 1978–79 but only up to 1 month in 1981–82.[39] Thus the duration data are consistent with the prediction from two-sided menu cost models, and if one is willing to accept the theoretical

[37] Specifically, if the evolution of the log of the real price can be represented by a Brownian motion (with drift) and if nominal prices are always increasing, then the duration of a price quotation is (asymptotically) distributed as an inverse Gauss variable. This distribution is humped to the left, and as the average real price deteriorates faster (i.e., as the rate of inflation increases), so does the asymmetry of this distribution.

[38] The figures correspond to a group of 11 products (1, 4, 7, and 19–26) that are homogeneous in terms of the average durations (and variance) of their nominal prices. The nominal prices in these products lasted, on average, 3 months in 1978–79 but only 2 months in 1981–82. There are two reasons for concentrating on this subset of products: First, because of their homogeneity, we avoid the use of an arbitrary normalization to aggregate heterogeneous products, and, second, their long duration minimizes the left-censoring problem inherent to duration data.

[39] It should be pointed out that a decrease in the mean duration does not necessarily imply an increase in the skewness of the distribution. Actually, a decrease in the mean duration can be accomplished by an increase in the symmetry of the distribution; i.e., average duration and skewness decline together.

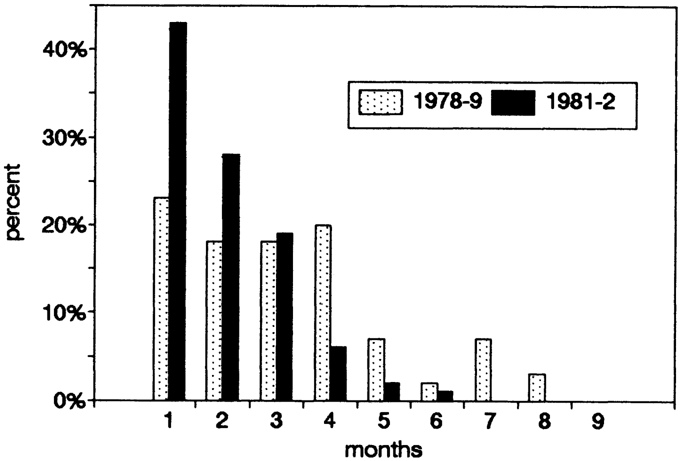

FIG. 4.—Duration of price quotations, products 1, 4, 7, and 19–26

results at face value, they imply that the distribution of real prices gets more humped to the left as the rate of inflation increases.

V. Conclusion

This paper deals with the relationship between real prices and the rate of inflation using disaggregated price data on 26 Israeli foodstuffs, between January 1978 and September 1984. Inflation was not a new phenomenon in these markets, and although rates of inflation were high, this was not, for the most part, a period of hyperinflation.

We have shown that relative price variability is directly linked to expected inflation. Specifically, intramarket price variability is mostly affected by the expected component of inflation. Since our data show that intramarket variability contributes around 80 percent of total price variability, expected inflation is a more important variable in explaining price variability than its unexpected counterpart. We interpret this strong effect through the lens of menu cost models, but even if one objects to this kind of reasoning, the fact remains that the price system is not invariant to expected inflation.

The paper discusses other aspects of price adjustments as well. When monthly inflation is around 3.9 percent in 1978–1979:6, real prices erode by 9 percent on average before they are changed. As the monthly rate of inflation climbs to 7.3 percent in 1982, the erosion in real prices is larger, 11.5 percent on average. This is accompanied by a shortening of the time between price adjustments. In fact, the average duration of a price quotation drops from 2.2 to 1.5 months

TABLE 7

SUMMARY STATISTICS FOR FIGURE 4

Period	N	Mean	Skewness	10%	25%	50%	75%	90%	Range
1978–79, 1981–82	2,077	2.4	1.55	1	1	2	3	4	10
1978–79	645	3.2	.86	1	2	3	4	7	10
1981–82	1,432	2.0	1.34	1	1	2	3	3	7

between 1978–1979:6 and 1982. Fixed time-dependent rules, therefore, are not empirically validated by the data.

We also show that price changes are not synchronized across stores. At the store level, asynchronization must mean that the magnitude of the adjustment in the nominal price (when adjusted) is smaller, since other stores have not changed their prices yet. At the macro level, this asynchronization suggests that the aggregate price level does not adjust immediately to changes in the money supply. This suggestion is reinforced by the shape of the distribution of real prices. Models based on (S, s) pricing rules posit a relationship between the response of the aggregate price level to monetary shocks and the distribution of real prices across stores. Recall that the distribution of real prices is unimodal but not always symmetric about zero. It is neither normal nor uniform. A nonuniform distribution suggests, at least within the framework of discrete (S, s) pricing rules, that firms' discontinuous price adjustments *do* slow down the response of aggregate price to monetary shocks. Thus individual firm lags are transmitted into lags of the aggregate price level. The dynamic behavior of the aggregate price level is, therefore, affected by the dispersion of real prices.

This paper is mostly descriptive and does not attempt to formally test discrete price adjustment rules. We plan to use our detailed data set to investigate this and other aspects of firms' pricing behavior in future work.

Appendix A

Interpretation of DP_{ijt}

Suppose that time is measured by days, indexed by $v = 1, 2, \ldots$, and that a month always has 30 days. Thus beginnings of months correspond to values of v equal to 1, 31, 61, 91, and so on. This subsequence is generated by $1 + 30(t - 1)$, where $t = 1, 2, \ldots$ is understood to represent the beginning of the tth month. Since we deal with a single product, we omit the subscript i from this Appendix. The observed nominal price quotation of store j in month t is denoted by P_{jt}. It corresponds to the price observed when the store was sampled during month $t - 1$, say, at day c, with $1 + 30(t - 2) < c < 1 + 30(t - 1)$. It is, by definition, the *last* nominal price chosen by store j *prior* to being sampled during month $t - 1$. Suppose that this nominal price was chosen at day $v = \tau$, that is, $P_{jt} = P_{j\tau}$. At the time of the change in price, τ, the store's real price was $P_{j\tau}/P_{.\tau}$, where $P_{.\tau}$ is the average price across stores at time τ. If the store follows an (S, s) pricing policy, this real price is store j's upper bound at time τ, $S_{j\tau}$. On the other hand, just before the nominal price is adjusted, the real price in store j reaches the lower bound $s_{j\tau}$. This lower bound equals the previous nominal price divided by the average price level at time τ, $s_{j\tau} = P_{j\tau'}/P_{.\tau}$, where τ' is the date of the *last* change in the price level before time τ (i.e., the price remained unchanged between τ' and

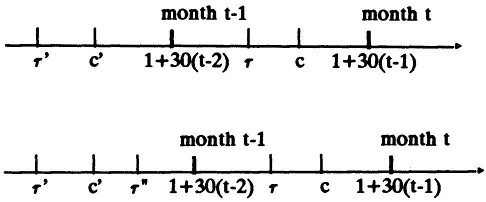

FIG. A1.—Time lines

τ). A measure of the length of the (S, s) band at time τ is then given by $\log(S_{j\tau}/s_{j\tau}) = \log(P_{j\tau}/P_{j\tau'})$.

How do observed prices in months t and $t - 1$ relate to prices in days τ and τ'? Suppose that store j changes its price *once* between two consecutive sampling dates, $c' < c$. Then it must be the case that τ' is less than c', which is less than $1 + 30(t - 2)$; that is, the last change before day τ occurred before the sampling date in month $t - 2$. Thus the events occur in the following sequence: $\tau' < c' < 1 + 30(t - 2) < \tau < c < 1 + 30(t - 1)$, as shown in the first time line in figure A1. Under these assumptions we get, in terms of observed prices, $P_{jt-1} = P_{j\tau'}$ and $P_{jt} = P_{j\tau}$. Hence, $DP_{jt} = \log(P_{jt}/P_{jt-1})$ is a measure of $\log(S/s)$ at day τ during month $t - 1$.

If the store did not change its price between c' and c, we cannot hope to measure the length of the (S, s) band from events (none) that occurred during month $t - 1$, and, in fact, DP_{jt} will be zero. If, on the other hand, the store changes its price more than once between c' and c, the matching with observed prices is not so clear-cut. Let τ' denote now the last time the nominal price was changed before the beginning of month $t - 1$ and prior to being sampled during month $t - 2$. This implies that the observed price in month $t - 1$ is $P_{j\tau'}$. Let τ'' be the first time the price is changed *after* the sampling date in month $t - 2$, and let τ be as before. That is, $\tau' < c' < \tau'' < 1 + 30(t - 2) < \tau < c < 1 + 30(t - 1)$, as shown in the second time line in figure A1. Then $DP_{jt} = \log(P_{jt}/P_{jt-1}) = \log(S_{j\tau}/P_{j\tau'}) + \log P_{.\tau} = \log(S_{j\tau}/s_{j\tau''}) + \log(P_{.\tau}/P_{.\tau''})$ since, by the same reasoning as before, $s_{j\tau''} = P_{j\tau'}/P_{.\tau''}$. Thus if prices change more than once between sampling dates, DP_{jt} is measuring the difference between the (S, s) bounds in different days that are at most 4 weeks apart and the inflation rate between these days. The bias is $DP_{jt} - \log(S_{j\tau}/s_{j\tau}) = \log(s_{j\tau}/s_{j\tau''}) + \log(P_{.\tau}/P_{.\tau''})$. The second term is usually positive but should be empirically negligible because the difference in the price level within a month is small compared to the size of DP_{jt}. If the lower bound decreases with expected inflation, as some models predict, the first term is negative. Hence, the approximation might not be too bad.

The outcome of all this is that the standard deviation of DP serves as an approximation to the standard deviation of the length of the (S, s) band in percentage terms.

Appendix B

Inflation Forecasts

In this Appendix we explain how the monthly inflation forecasts were constructed for each of the 26 products and for the aggregate inflation rate.

We restrict ourselves here to a 1-month-ahead forecast and ignore other possibilities.

The data utilized to calculate the inflation rates are the monthly price indexes for each product in 1977:1–1984:9. These indexes were computed by the Central Bureau of Statistics from virtually the same data used by us throughout our analysis. The difference stems mainly from the exclusion from our working sample of stores that were infrequently sampled.

The different behavior of inflation between the earlier and later years of the sample period dictated that we perform separate forecasts for 1977:1–1979:12 and 1980:1–1984:9. The expected inflation of product i's price in month t is given by

$$E\Pi_{it} = \hat{\alpha} + \sum_{\tau=1}^{q} \hat{\delta}_\tau \Pi_{it-\tau} + \sum_{\tau=1}^{p} \hat{\beta}_\tau \Pi_{\mathrm{cpi},t-\tau} + \text{time}.$$

That is, product i's rate of inflation, Π_{it}, is forecasted by the last q values of its own inflation rates, $\Pi_{it-\tau}$; by the last p observations on aggregate inflation, as measured by changes in the monthly CPI, $\Pi_{\mathrm{cpi},t-\tau}$; and by time-related variables. These time-related variables were either monthly dummies, a linear time trend, or a dummy variable for the year 1984, the year in which inflation climbed to a higher step. The choices among these options followed F-test results. To make the choice of lag length manageable, we confine ourselves to choose from among zero, three, or six own lags and zero or three CPI lags. Again, on the basis of goodness-of-fit criteria, values of q were chosen from among these options (see cols. 1 and 2 in table B1), while p was set at three. The estimates in the expected inflation equation were obtained from OLS regressions run on each product separately.[40] Their R^2 values are presented in columns 3 and 4 of table B1.

In the first period, monthly dummies were added in all regressions. The R^2 statistics lie between .6 and .9 with two exceptions, products 6 and 9. The regression results, not presented here, indicate that the fit of the model is significantly diminished when the own inflation lags are excluded from the regression.

In the second period, the 1984 dummy variable coefficient was in general significantly positive. Not surprisingly, the predictive power in this period, as measured by the R^2, is between one-half and two-thirds of that in the first period.

Each forecasting regression was tested for serial correlation using the Lagrange multiplier test suggested by Breusch (1978) and Godfrey (1978) for testing the existence of autoregressive or moving average errors of order p = 1, 2 and for autoregressive conditional heteroscedasticity (ARCH) as in Engle (1982). In general, the null hypotheses of no serial correlation and conditional homoscedasticity cannot be rejected by the data.[41]

[40] We also experimented with data on aggregate money balances (M1) and with more sophisticated methods of estimation (autoregressive integrated moving average [ARIMA] estimation). Inclusion of M1 does not provide additional information when we use the CPI, and the ARIMA model did not yield significantly better results. The ARIMA procedure did not allow us to break the sample down into two subperiods because of the small number of observations.

[41] Both in 1977–79 and in 1980–84, the null hypothesis of uncorrelated errors against the alternative AR(1)–MA(1) (AR(2)–MA(2)) cannot be rejected in 24 (23) products, at a 5 percent significance level. Engle's (1982) ARCH(p) disturbance model

TABLE B1

INFLATION FORECASTS

	LAGS OF OWN INFLATION*		R^2	
PRODUCT	1977–79† (1)	1980–84 (2)	1977–79 (3)	1980–84 (4)
1	6	3‡	.90	.48
2	3	3§	.64	.45
3	3	3‡§	.86	.27
4	6	3§	.74	.19
5	3	3†‡§	.69	.79
6	3	3§	.43	.54
7	3	3§	.83	.30
8	6	3§	.85	.42
9	3	3†§	.46	.43
10	6	6†§	.83	.71
11	3	6†§	.61	.66
12	3	3‡	.57	.38
13	3	3†	.83	.35
14	6	6§	.84	.60
15	3	6§	.70	.52
16	3	3§	.75	.39
17	3	3§	.71	.45
18	3	3‡§	.67	.32
19	3	3†§	.81	.44
20	3	3†§	.91	.41
21	3	6†§	.90	.49
22	6	3	.94	.20
23	3	3	.92	.13
24	3	3	.92	.09
25	3	3†	.90	.46
26	3	3†	.87	.42

* In addition, three lags of aggregate inflation, Π_{cpi}, were used in each product.
† Monthly dummies were added.
‡ A linear time trend was added.
§ A dummy for 1984 was added.

The expected aggregate inflation rate ($E\Pi_{cpi}$) was computed as the predicted value of Π_{cpi} given by a regression of current Π_{cpi} on its previous six lags and on monthly dummies from 1977:1 to 1984:9:

$$\begin{aligned} E\Pi_{cpi,t} = {} & -.02 + .36\Pi_{cpi,t-1} + .08\Pi_{cpi,t-2} + .30\Pi_{cpi,t-3} \\ & + .04\Pi_{cpi,t-4} + .09\Pi_{cpi,t-5} + .24\Pi_{cpi,t-6} \\ & + \text{monthly dummies}; \quad R^2 = .78, T = 86. \end{aligned}$$

The resulting residuals are well behaved: the Lagrange multiplier test statistics for testing the existence of autoregressive or moving average errors of order $p = 1, 2, 3, 6$, and 12 are, respectively, 0.67, 3.10, 3.6, 4.36, and 17.28,

is rejected in all products for $p = 1, 2$, and 4 in the 1977–79 regressions. In 1980–84, however, conditional homoscedasticity is rejected in two to three products in favor of the ARCH(p) model.

which are significantly less than the χ_p^2 critical value at 5 and 10 percent significance levels. In addition, Engle's ARCH(p) disturbance model is also rejected by the data for $p = 1, 2, 4$, and 12: the corresponding test statistics (χ_p^2) are 0.04, 0.21, 0.68, and 7.20.

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