INFLATION AND INTRAMARKET PRICE VARIABILITY: EMPIRICAL EVIDENCE FROM U.S. FOOD PRODUCTS

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food products, inflation, intermarket price variability, intramarket price variability

Abstract
The objective of this paper is to examine the response of relative price variability on U.S. food markets to food price inflation to identify whether such inflation influences the structure of relative prices between different food products. Results show that changes in food price inflation rate have a strong positive effect on the structure of relative prices across food products. In addition, the expected rate of inflation is found to be more important than the unexpected components as a determinant of relative price variability.

1. Introduction

A large number of studies have analyzed the relationship between the level of inflation rate and changes in relative prices of particular products and/or markets in the U.S. and other countries (Vining and Elwertowski 1976; Parks 1978; Cukierman 1979; Hercowitz 1981; Domberger 1987; Lach and Tsiddon 1992; Debelle and Lamont 1997; Fielding and Mizen 2000; Bakhshi 2002). Within this literature, change in relative prices has been measured by relative price variability and has been identified as an indicator of the real costs of inflation.

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More specifically, an important function of the price system is to efficiently transmit the information that economic agents need in order to allocate resources efficiently. Given that the information required is contained in relative prices, the noise coming from inflation can make it difficult to optimally utilize the information. As such, inflation can induce welfare-diminishing resource misallocation by directly affecting relative price dispersion (Friedman 1977; Fischer 1981). In fact, many empirical studies have found evidence of the positive relationship between inflation and relative price variability.

Studies linking inflation and relative price variability tend to fall into one of two categories (Domberger 1987). The first type of study concentrates on the analysis of cross-sectional variation of price movements in different (between) markets around the mean rate of price change for the whole economy, which is referred to as the analysis of *intermarket* price variability (Vining and Elwertowski 1976; Parks 1978; Cukierman 1979; Chang and Cheng 2000). Under this classification, relative price variability is defined as the dispersion of the products' own inflation rates (or industry averages) around an aggregate rate of inflation. By emphasizing that industries and sectors may differ in their speeds of adjustments to inflation shocks, the second type of study focuses on the analysis of the dispersion of prices *within* markets, which is known as the analysis of *intramarket* price variability (Domberger 1987; Lach and Tsiddon 1992). According to this classification, relative price variability refers to the dispersion of an individual product's own inflation rate around an industry average rate of inflation.

Until recently, however, empirical literature on the dispersion of relative prices in agricultural economics mostly concentrates on the analysis of the *intermarket* price variability in the U.S. and other countries (i.e., European countries). Lapp and Smith (1992), for example, use a measure of relative price variability across a set of agricultural commodities (47 commodities) in the U.S.; they find that these relative prices are more volatile when inflation rates are higher. Similarly, Reziti (2005) constructs a measure of relative price variability across a set of Greek agricultural products (53 products); he shows that changes in inflation rate have a strong positive effect on the dispersion of prices. Accordingly, empirical work on the *intramarket* price variability has received relatively little attention. To my knowledge, Lach and Tsiddon (1992) and Loy and Weaver (1998) are the only two studies that have tackled this issue. These two studies adopt disaggregated price data in examining the effects

of inflation on the dispersion of agricultural prices; they find evidence that a positive relationship between relative price variability and the rate of inflation holds for an intramarket measure. However, their analyses focus on the Israeli and Russian agricultural markets.

Furthermore, given the rapid spike in U.S. food prices during the period of 2007-2008, it is very interesting to explore the effect of food price inflation on changes in relative price structure across food products within the U.S. agricultural market. More specifically, since the early 1980s, consumer food prices in the U.S. have been stable in an overall sense. During 1982-2006, for example, the Consumer Price Index for all food (food CPI) has increased an average rate of 2.9% annually, nearly identical to the Consumer Price Index for all items (overall CPI) (3.3%). Moreover, the food CPI has never increased above the average annual rate of 4% over the past 15 years (Figure 1). Since the summer of 2007, however, this trend has changed dramatically as consumers in the U.S. have begun to face higher food prices at supermarket checkout lines. During the second half of 2007, for example, the food CPI, led by prices

FIGURE 1. Overall CPI and food CPI in the United States (% change year ago)
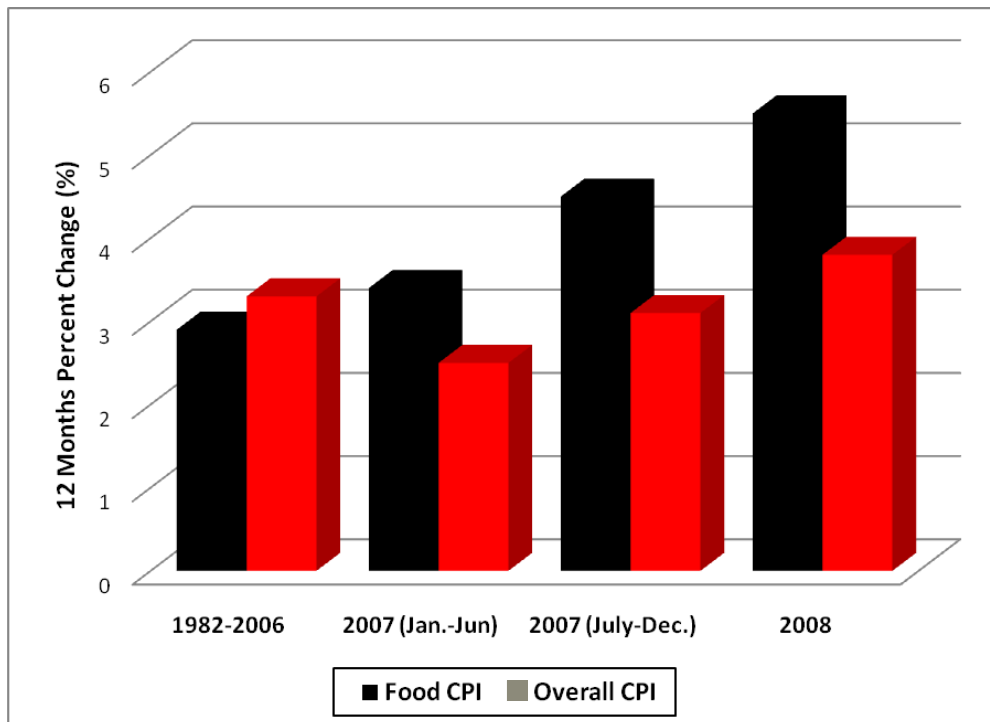
for beef (6.5%), poultry (8.0%), eggs (13.6%), and dairy products (21.7%), rose by 4.5%, the highest increase since 1990 (Figure 2). As a result, the food CPI increased much faster than the overall CPI (3.1%) during the same period.

FIGURE 2.   Food CPI and retail food prices in the United States, 2007
(% change year ago)
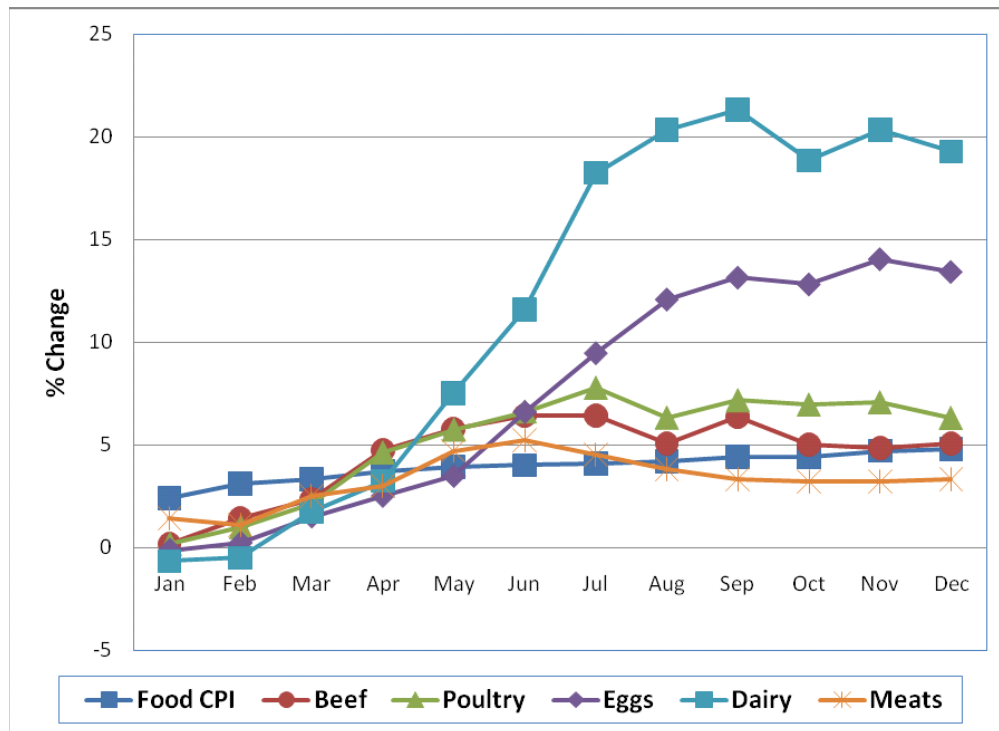
In this paper, therefore, I attempt to extend the scope of previous work by assessing the real costs of inflation processes in the U.S. within the context of *intramarket* price variability. The empirical focus is on the examination of the response of relative price variability on U.S. retail food markets to food price inflation to identify whether such inflation affects the structure of relative prices between different products. Since both expected and unexpected inflation could affect the dispersion of prices, I also analyze the role of the two different characteristics of inflation as determinants of changes in relative price structure and determine what components of inflation is affecting changes in relative prices and hence welfare costs of inflation. To that end, I use monthly data for ten food products in the U.S. over the period from January 1982 to December

2007. Food prices used for the analysis are (1) cereal and bakery; (2) beef; (3) pork; (4) poultry; (5) eggs; (6) dairy; (7) fruits and vegetables; (8) nonalcoholic beverages; (9) other food at home; and (10) food away from home. These products comprise approximately 93% of total household expenditure shares in the U.S. (Table 1). Given the significance of these markets, therefore, this timely analysis will shed light on the welfare consequences of inflation by identifying the linkage between inflation processes and relative prices across products.

The remainder of this paper is organized as follows. Section II briefly introduces the theoretical considerations underlying relative price variability. Section III discusses the measurement of relative price variability and the data used for the empirical analysis. In section V the results for intramarket price variability are reported and discussed. Finally, section V makes some concluding remarks.

TABLE 1. Estimated U.S. food expenditure shares, as of December 2007

Product	Expenditure share (%)
Cereal and bakery	7.4
Beef	3.8
Pork	2.4
Poultry	2.3
Eggs	0.9
Dairy products	6.4
Fruits and vegetables	8.4
Nonalcoholic beverages	6.7
Other food at home	9.9
Food away from home	44.6
Sub-total	92.8
Total	100.0

Source: Economic Research Service (ERS), USDA

2. The theoretical consideration on relative price variability

To derive a general expression for relative price variability (V_t), following Domberger (1987), I first define a simple unweighted aggregate price index (P_t) as follows:[1]

$$P_t = \frac{1}{\sum_i n_i} \sum_i \sum_j P_{ijt} \tag{1}$$

where $i = 1,...,m$ counts over markets; $J = 1,...,n$ counts over products; and P_{ijt} is a measure of the price of product j in market i in period t. Following the theorem of variance decomposition (Johnston and Dinardo 1996), the variance of P_{ijt} around the overall mean P_t can be expressed as follows:

$$\sum_i \sum_j (P_{ijt} - P_t)^2 = \sum_i \sum_j (P_{ijt} - P_{it})^2 + \sum_i n_i (P_{it} - P_t)^2 \tag{2}$$

where P_{it} is the mean price index in the i th market and equals $(1/n_j) \sum_j P_{ijt}$. Thus, the general expression for relative price variability (V_t) can be defined as follows:

$$V_t = \sum_i \sum_j (P_{ijt} - P_{it})^2 + \sum_i n_i (P_{it} - P_t)^2 \tag{3}$$

Equation (3) demonstrates how overall price variability in a given period can be decomposed into two separate components such as *within*-market (*intramarket*) and *between*-market (*intermarket*) components. More specifically,

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[1] In practice, such index is generally constructed using weighted means of individual price relatives. In the U.S., for example, the consumer price index (CPI) is a base-weighted index comprising 211 individual indices, one for each relevant market/industry. Each of these indices in turn is on the basis of a large sample of individual commodity prices. In this section, my attention is limited only to unweighted price index for analytical purpose; however, I will return to this issue when discussing measurement of relative price variability for my analysis (see endnote [4]).

the first term on the right-hand side of equation (3) measures the variability of an individual product indices (r_{ijt}) around an industry (market) average rate of index (r_{it}). The second term captures the variability of an industry (market) averages (r_{it}) around the overall mean (r_t). Again, since the literature on agricultural economics has mostly concentrated on the analysis of the second term, this study attempts to fill this gap by analyzing the first term.

The literature cited previously has found the positive correlation between relative price variability and inflation. Given the emphasis that places on the role of different characteristics of inflation such as *expected* and *unexpected* inflation, the theories explaining this positive correlation tend to fall into one of two categories: (1) imperfect (limited) information models and (2) menu cost models (Keynesian sticky price models). Imperfect information models, based on a rational expectation framework, assume that agents can obtain information on the equilibrium price in their own market much more quickly than on the general price level (Lucas 1973; Barro 1976). Under this circumstance, only unexpected demand shocks (i.e., unanticipated changes in the money stock) lead to a temporary inflation-unemployment trade-off as a result of confusion (misconceptions) between relative and general price level fluctuations. As such, the models predict a positive relationship between *unexpected* inflation and the dispersion of the industry (market) averages around the overall rate of inflation.[2] According to the terminology used in this approach, this coincides with the *intermarket* price variability.

Menu cost models stipulate that the presence of fixed costs of adjustment in nominal prices (known as fixed menu costs) induces the firm to change its nominal prices intermittently rather than continuously according to an (*S*, *s*) pricing rule (Sheshinski and Weiss 1983). More specifically, since nominal price changes are costly, firms adjust nominal prices whenever the real prices of their goods fall to a lower bound, *s*, at which time they should raise nominal prices so that real prices equal the upper bound, *S*. As such, the models predict that, as inflation is *expected* to increase, the difference between the optimal *s* and *S* increases, thereby resulting in a greater dispersion of prices. These models are usually associated with the price-setting behavior of sellers of a single

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[2] In other words, since the imperfect information model is derived from the so-called Lucas-type confusion between aggregate and relative shocks, only *unexpected* inflation affects relative price variability.

product.[3] In combination, this approach thus has direct implications for a link between *expected* (anticipated) component of inflation and *intramarket* price variability (Lach and Tsiddon 1992). Note that, since menu cost models incorporate the rational expectations framework in addition to price rigidity, this approach relates to the *unexpected* component of inflation as well (Lapp and Smith 1992).

3. Measurement of relative price variability and data

In the literature, relative price variability is usually measured by an unweighted standard deviation of price changes around the mean price change (Vining and Elwertowski 1976; Domberger 1987), and I do the same.[4] More specifically, let P_{jt} represent the food price index of j th product in time period t[5] The rate of change in the food price index of j th product between periods $t-1$ and t is denoted ΔP_{jt} and is expressed as the difference in the natural logarithm of price indices in the two periods:

$$\Delta P_{jt} = \ln P_{jt} - \ln P_{jt-1} \tag{4}$$

Similarly, ΔP_t denotes the mean rate of price changes for the set of food products, $j = 1,...,n$ as follows:

$$\Delta P_t = (1/n) \sum_j (\ln P_{jt} - \ln P_{jt-1}) \tag{5}$$

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[3] According to the menu cost approach, since higher average inflation tends to increase relative price variability, this approach relates price variability to *trend inflation* rather than unexpected inflation or the change in inflation rate (Reziti 2005).

[4] Since the main objective of this paper is to analyze the intramarket price variability in U.S. food markets, my empirical focus is given to the first term in equation (3). In addition, the analysis of intermarket variability generally uses both weighted and unweighted measures of standard deviation of the rates of price change, while the analysis of intramarket variability exclusively adopts the latter. As Domberger (1987) points out, however, the fundamental relationship under investigation is insensitive to the specific weighting procedure used (p. 553).

[5] For consistency P_{ijt} is a more correct expression as used in equations (1)-(3). Since this paper deals with food products within the U.S. food market, however, I exclude a subscript i (that is, food market) and prefer to use P_{jt} instead of P_{ijt} for simplicity

I emphasize that the definition of relative price variability in this study refers to the dispersion of food product price movements around the food market average and thus, the terminology adopted here is the *intramarket* price variability. As such, ΔP_t used here is (approximately) equal to the Consumer Price Index (CPI) for all food (food CPI). From equations (4) and (5), therefore, the relative price variability (SD_t) measured by standard deviation of the individual rate of price change around the food market average can be expressed as:

$$SD_t = \left[\frac{1}{n-1} \sum_{j}^{n} (\Delta P_{jt} - \Delta P_t)^2 \right]^{1/2} \tag{6}$$

The term ($\Delta P_{jt} - \Delta P_t$) is the rate of change in the j th relative price; that is, the logarithmic difference in the relative price (P_{jt} / P_t). If all prices change at the same rate, then SD_t is equal to be zero. If the dispersion of price changes across products is nonproportional, on the other hand, then SD_t becomes larger. As such, SD_t is said to be a measure of nonproportionality of the price movements (Parks 1978).

The data used in this paper consist of monthly prices of ten food products — cereal and bakery, beef, pork, poultry, eggs, milk, fruits and vegetables, nonalcoholic beverages, other food at home and food away from home — covering 26 years from January 1982 to December 2007. All price series are quoted in U.S. consumer price indices for each of ten products (2000=100) and are collected from the Bureau of Labor Statistics (BLS) in the U.S. Department of Labor (DOL).

4. Empirical methods and results

4.1. Identifying relationship between intramarket price variability and inflation

Since my analysis involves the intramarket price variability, following equation (6), I construct time-series of the measure of relative price variability (SD_{jt}) for each of the 10 products. I then estimate sets of regressions with SD_{jt} as the de-

pendent variable, where $j = 1,2,...,10$ indices products. The model to be estimated takes the following form:

$$SD_{jt} = \alpha_i + \beta_i \pi_{jt} + \varepsilon_{jt} \tag{7}$$

where π_{jt} is the inflation rates of each product, where $j = 1,2,...,10$ indices products; and ε_{jt} is 10×1 vectors of disturbances each of which is independently and identically distributed with zero mean and a possible non-diagonal covariance matrix.

A system of seemingly unrelated regressions (SUR)－also, called joint generalized least squares (JGLS) or Zellner estimation－is used to estimate equations (7). The SUR is a generalization of ordinary least squares (OLS) for multi-equation systems. Like OLS, the SUR method assumes that all the regressors are independent variables, but it uses the correlations among the errors in different equations to improve the regression estimates (Zellner 1962). In our case, for example, since inflationary shocks are likely to influence the price movements in each product to a greater or lesser extent, one may suspect there to be contemporaneous error correlations across different products. Under this circumstance, OLS is not the minimum variance estimator. To overcome this problem, therefore, I adopt the SUR method and estimate all 10 equations as a single system.

The results of the estimates of the system of equation (7) are reported in Table 2. As seen in column 1, for one case, I regress SD_{jt} on the actual inflation rate of each product (π_j).[6] For the other case, the regressor is the mean rate of price changes for the set of food products, measured by the food CPI (π_t) (column 2).[7] All the coefficients on both inflation variables (π_j) and

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[6] Conceptually, the inflation rate of each product, π_j, is the same as ΔP_{jt}. Since the former is collected directly from the BLS, however, it is not exactly equal to the calculated ΔP_{jt}. For this reason, the results of these regressions can be interpreted as a summary of correlations or reduced-form associations between the two variables only because I am not estimating a structural model (Lach and Tsiddon 1992).

[7] For this case, equation (7) is estimated separately by OLS, since the regressors are identical across equations. For this reason, I do not estimate both π_j and π_t in one equation using the SUR method, although SD_{jt} can be explained by own inflation rate and also by aggregate food inflation rate.

TABLE 2. Regression results for relative price variability and inflation variables

Product	Own inflation rate $(\tilde{\jmath})$		Food inflation rate $(\jmath\iota)$	
	Coefficient	DW	Coefficient	DW
Cereal and bakery	0.012 (1.99)**	1.87	0.068 (4.92)**	1.79
Beef	0.176 (5.03)**	1.98	0.227 (2.09)**	2.02
Pork	0.113 (5.17)**	1.91	0.332 (2.46)**	1.91
Poultry	0.068 (3.22)**	1.96	0.468 (3.66)**	2.03
Eggs	0.127 (5.76)**	1.80	1.138 (2.47)**	2.02
Dairy products	0.229 (6.16)**	1.78	0.257 (2.71)**	1.98
Fruits and vegetables	0.079 (3.04)**	1.80	0.688 (2.89)**	1.83
Nonalcoholic beverages	0.226 (7.04)**	1.70	0.206 (1.86)*	1.97
Other food at home	0.047 (2.78)**	1.91	0.156 (2.68)**	1.89
Food away from home	0.233 (2.96)**	1.79	0.175 (5.01)**	1.85

Note: ** and * denote significance at the 5% and 10% levels, respectively. The 5% critical bound for the Durbin-Watson statistics is (1.76, 1.78). The equations for the own inflation rate of each product are estimated jointly using the SUR procedure. The equations for the aggregate inflation rate of each product are estimated separately by OLS, since the regressors are identical across equations. To save space, the estimated constant terms are not reported.

$(\jmath\iota)$ are found to be positive and significant at least at the 10% level. This indicates that higher inflation leads to higher intramarket variability across products, thereby inducing the welfare cost of inflation.[8] In addition, the results show that the estimated coefficients of $\tilde{\jmath}$ are generally larger than those of $\jmath$. For example, the estimated coefficient of averages 0.372 over all ten products, which is much higher than the average value of the coefficient of $\jmath$,

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8 Following previous studies (e.g., Domberger 1987; Lapp and Smith 1992), I use both linear and quadratic functional forms in estimating equation (7). The former outperforms the latter in terms of goodness of fit and parsimony. Hence, my presentation focuses on the results derived from the linear regression.

0.131.[9] This implies that intramarket price variability is more responsive to market average inflation (food CPI) across products than to own inflation.[10]

It is worth mentioning that, when dealing with time-series data, the possibility of nonstationarity in a series raises issues about parameter inference and spurious regression (Wooldridge 2000). Before estimating equation (7), the existence of a unit root in the series in the two models − calculated relative price variability (V_{jt}), own inflation rates of each product (π) and aggregate inflation rate ($\tilde{\pi}$) − is thus investigated using the Dickey-Fuller generalized least squares (DF-GLS) test (Elliot et al. 1996). All the series (42 series) are found to be stationary. Thus, standard econometric methods can be applied to produce parameter estimates with desirable asymptotic properties.[11]

4.2. Determinants of intramarket price variability

In this section, to assess the role of different characteristics of inflation as determinants of intramarket price variability and to determine what components of inflation is affecting the dispersion of prices and how much, I first build time-series of expected and unexpected inflation for each of 10 products. Then, I regress SD_{jt} on the expected and unexpected inflation rates of each product. The model to be estimated takes the following forms:

$$SD_{jt} = \sigma_{j0} + \sigma_{j1}\pi_{jt} + \sigma_{j2}\pi_{jt}^{\cdot} + u_{jt} \tag{8}$$

where π_{jt} and $\pi_{jt}^{\cdot}$ are expected and unexpected inflation rates of each product, where $j = 1,2,...,10$ indices products; and u are 10×1 vectors of disturbances each of which is independently and identically distributed with zero mean and a possible non-diagonal covariance matrix.

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[9] These are averages of individual estimates of 10 products in Table 2.

[10] It should be noted here that given the fact that the inflation rate π_j is conceptually the same as Δr_{jt}, price variability of each product (j) is basically higher than the CPI (π_p), defined as the mean rate of price changes for the set of food products. As such, the results obtained from equation (7) may be strongly affected by the ways of constructing the two different indices, rather than by characteristics of different types of food products/markets; this interpretation should thus be viewed with caution.

[11] For brevity, however, the results of unit root tests are not reported.

It should be noted that to construct time series of expected and un-expected rates of inflation for each of 10 products, I rely on a forecasting model developed by Larch and Tsiddon (1992). In this model, product *j*'s rate of inflation (   ) is estimated for the lags of its own inflation rates (ˇ) and of aggregation inflation (   ), as measured by changes in the monthly food CPI as follows:

$$\pi_{jt} = \hat{\alpha} + \sum_{k=1}^{p} \hat{\varepsilon}_k \pi_{jt-k} + \sum_{k=1}^{q} \hat{\varphi}_k \pi_{fp,t-k} \qquad (9)$$

where *p* and *q* are lag order. In order to estimate equation (9), it is necessary to determine the lag lengths for the forecasting model since it is sensitive to changes in the lag structure. To make the selection of lag order manageable, I restrict myself to choose from zero to nine own lags and food CPI lags, respectively. The optimal lag lengths for equation (9) are then determined by the Schwarz (SC), Hannan-Quinn (HQ), and Akaike (AIC) information criteria using likelihood ratio (LR) tests (Doornik and Hendry 1994). The R^2 values for these regressions lie between 0.6 and 0.8 with exception of poultry (0.47). In the serial correlation and heteroskedasticity tests using the *F-* form of the Lagrange Multiplier (LM) test, the null hypothesis of no serial correlation and no heteroskedasticity cannot be rejected at the 5% level. Given the reasonable forecasting power and white noise residuals, the estimates are reasonable prox-ies for expected and unexpected inflation rates. Finally, the predicted values from the OLS regressions are taken to be expected inflation for each product, while the residuals are taken to be unexpected inflation.[12]

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[12] In addition to this forecasting model, I also employ two alternative measures of the expected and unexpected inflation series, since the results may be sensitive to the measure used (Lapp and Smith 1992). The measures are: (1) a univariate ARMA (1, 1) time-series model of *ᴶ*; and (2) an AR (1) model of ˇᴶ. Although there are many possible formulations for the construction of the expected and unexpected in-flation rates, these two measures are the most widely used formulations in the liter-ature (e.g., Parks 1987; Lapp and Smith 1992; Reziti 2005). The results generated by these models indicate patterns similar to the ones found when the forecasting model developed by Larch and Tsiddon (1992) is used. To save space, therefore, I do not report the results here.

The results of the estimates of the system of equation (8) are summarized in Table 3. For expected inflation (π_j), all estimated coefficients except other food at home are found to be positive and significant at the 5% level (column 1), implying that expected inflation has a positive effect on the extent of price variability. For unexpected inflation (π_j), on the other hand, the coefficient is found to be positive and significant in 6 out of the 10 products (column 2). In addition, the results show that the estimated coefficient of expected inflation is mostly larger than that of unexpected inflation; on average, for example, the estimated coefficient of π_j is 0.170 across products and that of π_j is 0.113. These findings suggest that expected component of inflation has more explanatory power than the unexpected one in explaining intramarket price variability across different products in the U.S. agricultural market.

TABLE 3. Regression results for relative price variability and expected and unexpected inflation of each product

Product	Expected inflation (π_j)	Unexpected inflation (π_j)	DW
Cereal and bakery	0.018 (2.32)**	0.015 (0.31)	1.87
Beef	0.163 (2.78)**	0.155 (3.52)**	1.98
Pork	0.154 (2.74)**	0.134 (3.23)**	1.92
Poultry	0.132 (2.38)**	0.009 (0.26)	1.96
Eggs	0.260 (2.66)**	0.004 (0.10)	1.79
Dairy products	0.269 (4.01)**	0.185 (4.15)**	1.99
Fruits and vegetables	0.092 (2.51)**	0.079 (1.81)*	1.80
Nonalcoholic beverages	0.151 (2.66)**	0.258 (6.58)*	1.82
Other food at home	-0.040 (-0.79)	0.093 (2.44)**	1.92
Food away from home	0.498 (4.01)**	0.099 (0.99)	1.79

Note: ** and * denote significance at the 5% and 10% levels, respectively. The 5% critical bound for the Durbin-Watson statistics is (1.76, 1.78).The equations are estimated jointly using the SUR procedure. t-values are in parentheses.

For completeness, I also run an addition set of regressions using an aggregate inflation rate, as defined by the rate of change of the food CPI.[13] The estimates are reported in Table 4. The results show that expected inflation (π_j) is found to have a positive effect on price variability in 9 out of the 10 cases (column 1). For unexpected inflation (u_j), on the other hand, the estimated coefficients turn out to be significant and positive only in 3 out of the 10 cases (column 2). In addition, as seen in Table 3, the estimated coefficient of expected aggregation inflation is larger than that of unexpected one; on averages across products, for example, the estimated coefficient of π_j is 0.475 and that of u_j is 0.201. The finding provides further confirmation that the effect of expected inflation on intramarket price variability is much stronger than that of unexpected inflation. In this sense, my findings here are more consistent with the menu cost model than the imperfect information model in explaining the relationship between inflation and the variability of relative prices within the U.S. agricultural sector; that is, as inflation is *expected* to increase, a discrete (rather than continuous) and different price adjustment across products due to fixed menu costs results in a change in the structure of real prices, thereby inducing the inefficiency of resource allocation and thus welfare loss.[14] In addition, a comparison between Tables 3 and 4 indicates that the estimated co-

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[13] To that end, following Larch and Tsiddon (1992), the expected (unexpected) aggregate inflation rate is estimated using the predicted value (residual) of π taken from a forecasting model in which current π^p is regressed by its past nine lags and monthly seasonality, which is determined using likelihood ratio tests; the resulting residuals are found to behave as white noise. In addition, the results generated by this model are consistent with those obtained from such two alternative measures as (1) a univariate ARMA (1, 1) time-series model of π^p; and (2) an AR (1) model of π^p. Finally, for this case, equation (8) is estimated separately by OLS, since the regressors are identical across equations.

[14] Again, under the menu cost theory, inflation costs arise because the presence of fixed costs of adjustment in nominal prices induces the firm to change its nominal prices slowly and differentially across products. Due to the fact that this model incorporates rational expectation like imperfect information models, the consideration of unexpected inflation is relevant here as well (see the theoretical consideration on relative price variability section). Finally, it should be pointed out that I did not consider other factors such as macroeconomic variables and/or competition and market structure in the U.S. food markets in my analysis; hence, this conclusion also should be viewed with caution.

efficient of aggregate inflation is generally higher than that of own inflation, implying that intramarket price variability has a more variable response to aggregate inflation across products than to own inflation.

TABLE 4. Regression results for relative price variability and expected and unexpected aggregate inflation

Product	Expected inflation (π_j)	Unexpected inflation (π_j)	DW
Cereal and bakery	0.030 (2.54)**	0.096 (1.17)	1.88
Beef	0.232 (1.72)*	0.221 (1.20)	2.02
Pork	0.545 (3.55)**	-0.049 (-0.21)	1.94
Poultry	0.462 (2.90)**	0.062 (0.26)	2.01
Eggs	1.694 (2.06)**	1.275 (1.85)*	1.85
Dairy products	0.421 (2.14)**	0.242 (1.04)	1.99
Fruits and vegetables	0.737 (2.99)**	-0.400 (-0.46)	1.84
Nonalcoholic beverages	0.114 (0.89)	0.398 (2.04)**	1.97
Other food at home	0.159 (2.62)**	0.077 (0.86)	1.92
Food away from home	0.358 (2.24)**	0.091 (5.86)**	1.89

Note: ** and * denote significance at the 5% and 10% levels, respectively. The 5% critical bound for the Durbin-Watson statistics is (1.76, 1.78). The equations are estimated separately by OLS, since the regressors are identical across equations.

5. Concluding remarks

Although the empirical studies on the analysis of the intermarket price variability in the agricultural sector have been widely conducted, relatively little attention has been paid to the analysis of the intramarket price variability. With

the recent sharp acceleration in food prices in the U.S., this study thus contributes to the literature by analyzing the real costs of such inflationary processes in the U.S. in a framework of intramarket price variability. To that end, using monthly price data on 10 food products in the U.S. over the past 25 years, this study examines the effect of inflation on the dispersion of relative prices, as well as the determinants of price behavior between different products within the U.S. agricultural market.

The results show a positive relationship between relative price variability and the rate of inflation for an intramarket measure; that is, through disrupting the structure of relative prices between food products within the U.S. agricultural sector, food price inflation affects changes in intramarket price variability, thereby inducing welfare-diminishing resource misallocation. I also find that the effect of the expected component of inflation on intramarket price variability is much stronger than the effect of the unexpected part of inflation. This finding is thus consistent with the menu cost models in predicting the sources of the link between inflation and relative price variability; that is, inflation costs arise as a result of sticky prices that slowly and differentially adjust across different products.

Finally, it is worth mentioning that, although this paper examines the relationship between the level of inflation rate and relative price variability in U.S. food market, my approach can be readily applied to all countries, particularly experiencing fast-rising food prices. As bad weather conditions this summer were considered to be the major culprits behind the recent deterioration of the fresh food market (i.e., cabbage and other vegetables for kimchi, the Korean national dish), for example, consumers in Korea have been suffering from the rapid spike in food prices; annual cabbage price inflation increased by 400% between August and September in Seoul, Korea's capital city, and year-on-year inflation jumped from 2.6% to 3.6% during the same period. Under the circumstances, therefore, the approach adopted here can be used to empirically assess the hypothesis that a positive relationship between inflation rate and price variability in the recent food market in Korea.

References

Bakhshi, H. 2002. Inflation and relative price variability. Economics Letters 76(1): 27-33.

Barro, R. 1976. Rational expectations and the role of monetary policy. Journal of Monetary Economics 2: 1-32.

Chang, E. C., and J.W. Cheng. 2000. Further evidence on the variability of inflation and relative price variability. Economics Letters 66(1): 71-77.

Cukierman, A. 1979. The relationship between relative prices and the general price level: a suggested interpretation. American Economic Review 69(3): 444-447.

Debelle, G., and O. Lamont. 1997. Relative price variability and inflation: evidence from U.S. cities. Journal of Political Economy 105(1): 132-152.

Domberger, S. 1987. Relative price variability and inflation: a disaggregated analysis. Journal of Political Economy 95(3): 547-566.

Elliott, G., T.J. Rothenberg and J.H. Stock. 1996. Efficient tests for an autoregressive unit root. Econometrica 64: 813-836.

Fielding, D., and P. Mizen. 2000. Relative price variability and inflation in Europe. Economica 67(1): 57-78.

Fisher, S. 1981. Relative shocks, relative price variability, and inflation. Brookings Papers on Economic Activity 2: 381:431.

Friedman, M. 1977. Nobel lecture: inflation and unemployment. Journal of Political Economy 85: 451-472.

Hercowitz, Z. 1981. Money and the depression of relative price. Journal of Political Economy 89(2): 328-356.

Johnston, J., and J. Dinardo. 1996. Econometric Methods. 4th edition. New York: McGraw-Hill

Lach, S. and D. Tsiddon. 1992. The behavior of prices and inflation: an empirical analysis of disaggregated price data. Journal of Political Economy 100(2): 349-389.

Lapp, J.S., and V.H. Smith. 1992. Aggregate source of relative price variability among agricultural commodities. American Journal of Agricultural Economics 74(1): 1-9.

Loy, J.P.F., and R.D. Weaver. 1998. Inflation and relative price volatility in Russian food markets. European Review of Agricultural Economics 25(3): 373-394.

Lucas, R.E. 1973. Some international evidence on output-inflation tradeoffs. American Economic Review 63(3): 326-334.

Parks, R.W. 1978. Inflation and relative price variability. Journal of Political Economy 86(1): 79-95.

Reziti, I. 2005. The relationship between macroeconomic variables and relative price variability in Greek agriculture. International Advances in Economic Research 11(1):

111-119.

Sheshinski, E., and Y. Weiss. 1983. Optimum pricing policy under stochastic inflation. Review of Economic Studies 44(2): 287-303.

Vining, D.R., and T.C. Elwertowski. 1976. The relationship between relative prices and the general price level. American Economic Review 66(4): 699-708.

Wooldridge, J.M. 2000. Introductory Econometrics: a Modern Approach. Mason, Ohio: South-Western College Publishing.

Zellner, A. 1962. An efficient method of estimating seemingly unrelated regressions and tests of aggregation bias. Journal of American Statistical Association 57(2): 500-509.