Relative Agricultural Price Changes in Different Time Horizons

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Abstract:

Using a monthly data covering from 1974:1 to 2002:12, this paper explores the linkage between changes in macroeconomic variables (real exchange rate and inflation rate) and changes in relative agricultural prices in different time horizons (1, 12, 24, 36, 48, and 60 months). Controlling for factors likely to determine the long run trend of relative agricultural prices, the results show that long-term changes in real exchange rate has had a significant negative correlation with the long-term changes in relative agricultural prices. Conversely, changes of the general price have a role in explaining short-term changes in relative agricultural price at best.

Keywords: Relative agricultural price, exchange rates, inflation rates

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1. Introduction

Changes in relative commodity prices in an economy occur continuously in response to changes in many real factors of demand and supply of different commodities. Although this proposition has received substantial theoretical support, many economists have tried to discern the potential effect of macroeconomic factors on relative commodity prices. In their rational expectation model based on the assumption of imperfect information, Lucas (1973) and Barro (1976) show that unanticipated inflation can create a "misconception" of absolute and relative price changes, which implies that the inflation leads to a dispersion of prices among different commodities. A positive relationship between the inflation rate and relative price dispersion is also discovered in the menu cost models, which assumes the existence of a menu cost when changing prices (Shehinski and Weiss, 1977; Ball and Romer, 1989; Ball and Mankiw, 1995).

In agricultural economics, many studies (e.g., Frankel, 1986; Grennes and Labb, 1986; Robertson and Orden, 1990; Saghaian et al., 2002) have examined the potential effect of monetary shock on changes in relative agricultural prices. However, they have concentrated on the issue of the speed of price adjustment, while studies examining the issue of the size of adjustment are sparse. Moreover, because these studies have focused only on the effect of monetary shock on the relative commodity prices holding long-run money neutrality hypothesis[1], concentration is limited on short-term changes of relative agricultural prices.

We insist in this paper that, in an open economy, the U.S. dollar movements have been important role in explaining the *long-term* movements of relative agricultural price because the changes can cause different degrees of supply shock on different industry sectors. In fact, under the assumption of well-working foreign exchange markets, the possibility of U.S. dollar movements influencing relative agricultural prices might not be easily accepted because the macroeconomic shocks only cause a temporal overshooting problem of the nominal exchange rate (e.g., Dornbusch 1976). However, many empirical studies about foreign exchange markets suggest that there is some degree of inefficiency in the markets (e.g., Frankel and Froot, 1987, 1990; Froot and Frankel, 1989; Ito, 1990; Frankel and Rose, 1995). As a result, there have been large and persistent deviations of the nominal exchange rate from its monetary fundamentals (Dornbusch, 1987; Rogoff, 1996)[2]. Therefore, the large and persistent fluctuation of the U.S. dollar, which cannot be explained by monetary variables, can cause different degrees of supply shocks in different industry sectors. Moreover, due to the persistency, unlike to the monetary shock, it causes relative *long-term* variation of the relative commodity price.

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[1] An exception is Saghaian et al. (2002). They find that the money is not neutral to relative agricultural price even in the long-run.

[2] This is called nominal exchange rate misalignment problem in the relevant literature.

There are some reasons we believe why U.S. agricultural prices are expected to be more sensitive to the U.S. dollar movement, relative to the prices of other industries' and non-tradable goods. First, the agricultural sector is heavily involved in international trade (more than 30 percent of domestic agricultural products is exported). Second, the demand for agricultural products is inelastic compared to other manufacturing products (Kileson and Poole, 2000)[3]. As a result, supply shock induced by exchange rate movements can more sensitively affect the agricultural prices than other manufacturing and non-tradable good's prices.

The main objective of the paper is to identify the important macroeconomic factors which explain the changes of relative agricultural price in different time horizons (1, 12, 24, 36, 48, and 60 months). To examine the issue, we develop a time-series model which identifies unobservable real factors affecting variation of relative agricultural prices in the long run. Unlike the short-term overshooting problem of the agricultural price induced by monetary shock, we found the long-term changes of relative agricultural price in the U.S. have been strongly correlated with the U.S. dollar movements. The paper is organized as follows. Section 2 includes a brief discussion about exchange rate movements under the floating system. In Section 3, we present empirical models to examine this issue. The main empirical findings are presented in Section 4. The paper is summarized in Section 5.

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[3] Recently, Kim and Koo (2002) find that the U.S dollar movements affect the performance of U.S agriculture exports differently than other industry sectors, which implies possibly different degrees of domestic supply shocks induced by the U.S dollar movements.

2. Exchange Rate Movements under the Floating System

Before examining this issue empirically, we discuss the question of whether U.S. dollar movements can affect the long-term variation of relative agricultural price on theoretical grounds. In fact, the possibility might not be easily supported by standard macroeconomic models. To explain this point, consider the theory of purchasing power parity (PPP), which is one of the fundamental assumptions in the flexible price monetary model. Under the assumption of a fully integrated world goods market, the model assumes the following PPP condition:

(2.1) $$\frac{P_t}{P_t^*} = \theta \cdot S_t \quad \text{or} \quad \frac{P_t}{P_t^* S_t} = R_t = \theta,$$

where P_t and P_t^* are the aggregate price levels in the home and foreign countries, respectively; S_t and R_t are the nominal and real exchange rates between the home and foreign countries (i.e., units of home currency required to buy one unit of foreign currency); and θ represents factors that cause the nominal exchange rate to deviate from the PPP, such as transportation costs and trade barriers, which are assumed to be constant in the long run. An important point to consider is that real exchange rates (R_t) are assumed to be a fixed constant (θ) in both the short-run and the long-run. Therefore, from this point of view, we do not expect that movements of real exchange rate can cause any cyclical long-term variations in real variables, such as agricultural export and real agriculture prices.

Under the sticky price monetary model (Dornbusch, 1976), short-run deviation of nominal exchange rates from the PPP can be explained by stickiness of nominal wages and prices. In other words, the speed of adjustment of the financial market in response to a nominal shock is assumed to be faster than that of commodity market, which could possibly cause temporary overshooting of nominal exchange rates. However, according to the model, the nominal shock should cause only temporal overshooting of nominal exchange rates, which might cause short-term volatility rather than long-term cyclical variations in real exchange rates.

However, recent empirical evidence in international macroeconomics and finance appears to contradict the Dornbusch model by revealing the strong possibility of a persistent deviation of nominal exchange rates from PPP. This would indicate that there is a possibility of long-term fluctuation of real exchange rates. Empirical evidence also demonstrates the possibility of some degree of inefficiency in the foreign exchange market. The most compact form of explanation for these deviations is the possibility of a rational speculative bubble[4]. If the nominal exchange rate moves, it will drift in the same direction for a long time unless an important economic event changes the direction of expectations held by foreign exchange market participants (Frankel and

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[4] Speculative bubbles are defined as a phenomenon of nominal exchange rate movements that are not based on economic fundamentals, but rather are based in self-confirming expectations (Frankel and Rose, 1995). Although many economists believe that speculative bubbles are one of the important sources of unexpected movement of nominal exchange rates during the post-Bretton Woods era, there is not a universally accepted reason for what starts a bubble or what causes them to burst. Potential explanations for the sources of speculative bubbles are the influential effect of 'noise' traders in foreign exchange markets (De Long et al., 1990, 1991); heterogeneous beliefs of economic agents (Hart and Kreps, 1987); and systematic forecasting error (Froot and Frankel, 1989). More detailed discussion of this issue is summarized in Frankel and Rose (1995), and Taylor (1995).

Rose, 1995; Frankel 1996). Under both the sticky price model and the rational speculative bubble model, we can define the real exchange rate as:

$$\frac{P_t}{P_t^* S_t} = R_t = \theta + f_t \,, \qquad (2.2)$$

where f_t relates to unobservable stochastic factors that cause fluctuation in the real exchange rates.

The distinguishing features of the two models are as follows. Under the sticky price monetary model, f_t might be serially correlated, but the coefficient of autoregression is far less than one, so that the deviation of real exchange rates from an arbitrary constant θ would die out within a short-time period. By contrast, under the assumption of a rational speculative bubble, f_t could possibly have a unit root or near unit root. Therefore, under the assumption of the existence of a rational speculative bubble, real exchange rate movements can have an explanatory power for long-term variation of relative agricultural price via different degrees of domestic supply shocks.

3. Empirical Model Derivation

In the relevant literature (e.g., Vining and Elwertowski, 1976; Parks, 1978; Fisher, 1981; Lach and Tsiddon, 1992; Bomberger and Makinen, 1993; Debelle and Lamont, 1997), economists have examined the issue of the effect of nominal shocks (inflation rate) on *changes* in different commodity prices. The empirical question in this case is whether the changes in the general price level are correlated with the variability of relative price changes in an economy. The relationship between price change dispersion among different commodities (or inter-market price change dispersion) and general inflation rates is typically estimated with the following model,

$$RPD_t = \alpha + \beta \cdot \Delta \ln p_t + \gamma \cdot \Delta \ln z_t + \eta_t \tag{3.1}$$

where RPT_t is a measure of price change dispersion of different commodity groups; $\Delta \ln p_t$ is a rate of general inflation; and $\Delta \ln z_t$ are rates of changes of other relevant variables. Inter-market price change dispersion is usually measured by a variation (or standard deviation) of changes of relative prices compared to general inflation rates such as[5]:

$$RPD_t = \frac{1}{N}\sum_{1}^{N}(\pi_{it} - \pi_t)^2 \tag{3.2}$$

where $\pi_{it} (= \ln p_{it} - \ln p_{it-1})$ is the rate of change of the *ith* commodity group; $\pi_t (= \ln p_t - \ln p_{t-1})$ is an inflation rate for the period; and *N* is the number of the commodity groups.

Although the empirical model specification (3.1) with the measure (3.2) is appropriate to examine the effect of the general inflation rate on the relative price dispersion problem at the macroeconomic level, it is not appropriate to examine the issue of the relative price change of a specific commodity group compared to other commodity groups. Because we are concentrating on the price variation problem of a specific commodity group (agricultural products) in comparison to other groups in

[5] This definition is used in Parks (1978) and Fisher (1981).

different time horizon, we should derive the empirical model to fit our economic question.

To do that, we first assume that there are long-run relationships between general price level and the price level of each commodity group. For instance, consider there are only two goods in an economy, agricultural and non-agricultural products.

(3.3) $$\ln p_t^a = \alpha_0 + \beta_0 \cdot \ln p_t + \eta_t$$

(3.4) $$\ln p_t^{na} = \alpha_1 + \beta_1 \cdot \ln p_t + \mu_t$$

where p_t is the general price level; p_t^a is the price level of agricultural goods; p_t^{na} is the price level of non-agricultural goods; $p_t = w^a p_t^a + w^{na} p_t^{na}$ and $w^a + w^{na} = 1$; w^a and w^{na} are weights of the components of the deflator for each commodity group; and η_t and μ_t are observable and unobservable stochastic components including macroeconomic shocks, and idiosyncratic shocks for each commodity group, which can affect the real price movements of each group. In the long-run, the relationship between the general price level and the price of a commodity group *i* are determined by relative supply and demand conditions between the groups of the commodities (Kileson and Poole, 2000). For instance, by Engel's law, if the income elasticity of the agricultural goods is less than that of the non-agricultural group, the coefficient β_0 is expected to be less than β_1. Subtracting (3.4) from (3.3), we have

(3.5) $$\ln p_t^a - \ln p_t^{na} = (\alpha_0 - \alpha_1) + (\beta_0 - \beta_1) \cdot \ln p_t + (\eta_t - \mu_t)$$

If we do not consider the long-run coefficients β_0 and β_1, which are expected to be determined by unobservable real factors (e.g., different degrees of income elasticities), we might easily obtain a statistically significant effect of general price level on relative price movement of different commodity groups. However, this result suffers from an omitted variables problem, and it is difficult to conclude that the general price movements are related to the relative price movements.

To eliminate this possibility, we rewrite the equation (3.3) and (3.4) as

(3.6) $$\ln p_t^a - \beta_0 \cdot \ln p_t = \alpha_0 + \eta_t$$

(3.7) $$\ln p_t^{na} - \beta_1 \cdot \ln p_t = \alpha_1 + \mu_t .$$

Subtracting equation (3.7) from equation (3.6), we have

(3.8) $$(\ln p_t^a - \beta_0 \cdot \ln p_t) - (\ln p_t^{na} - \beta_1 \cdot \ln p_t) = (\alpha_0 - \alpha_1) + (\eta_t - \mu_t) .$$

If we decompose the stochastic term η_t and μ_t as the macroeconomic shocks, such as inflation and exchange rate shocks, and the unobservable commodity group specific idiosyncratic shocks ($\eta_t = \theta_0 + \lambda_0 \ln p_t + \pi_0 \ln r_t + \omega_t$; $\mu_t = \theta_1 + \lambda_1 \ln p_t + \pi_1 \ln r_t + \varepsilon_t$ where r_t is real exchange rate, and ω_t and ε_t are unobservable commodity group specific idiosyncratic shocks), we have

(3.9) $$(\ln p_t^a - \beta_0 \cdot \ln p_t) - (\ln p_t^{na} - \beta_1 \cdot \ln p_t) = \kappa + \gamma \cdot \ln p_t + \delta \cdot \ln r_t + \zeta_t$$

where $\kappa = \alpha_0 - \alpha_1 + \theta_0 - \theta_1$; $\gamma = \lambda_0 - \lambda_1$; $\delta = \pi_0 - \pi_1$; and $\zeta_t = \omega_t - \varepsilon_t$. By differencing equation (3.9) with lag length *k*, we develop our final empirical model as

(3.10) $$\Delta_k (\ln p_t^a - \beta_0 \cdot \ln p_t) - \Delta_k (\ln p_t^{na} - \beta_1 \cdot \ln p_t) = \kappa + \gamma \cdot \Delta_k \ln p_t + \delta \cdot \Delta_k \ln r_t + \zeta_t$$

where $\Delta_k \ln z_t = \ln z_t - \ln z_{t-k}$ for any variable z_t.

With the model specification (3.10), we examine the question of which macroeconomic factors cause more deviation in the price of a commodity group from its long-run equilibrium level in comparison to other commodity group. If food and agricultural prices are more (less) sensitive to the changes in general price level than prices of other commodity groups, the estimated coefficient γ is expected to be positive (negative). If U.S. dollar appreciation causes more supply shock in the domestic food and agricultural markets than other commodity groups, the expected sign of δ is negative.

Different lag lengths are important in examining the main hypothesis of the paper. If we believe that the inflation rate causes only short-run effects on the changes in relative price differences, the significance of the estimated coefficients should be die out where k is large enough. This means that changes in general price level cannot explain the changes of relative prices between different commodity groups in the long-run. The real exchange rate, however, can explain relatively *long-term* changes of relative prices. As we discussed before, the misalignment problem of the U.S. dollar is prolonged and persistent; once the U.S. dollar appreciates (depreciates), it continues the trend for several years in a row. Therefore, we expect that the supply shocks generated by U.S. dollar movement also continue for several years, which can better explain relative price changes in a longer time period than inflation rate. In practice, we use two-step estimation procedure. In the first stage, we estimate the cointegration vector, which explains long-run relationship between the general price level and the price level of each commodity group. In the second step, the equation (3.10) is estimated by replacing the estimated long-run coefficients obtained in the first step.

4. Data

Seasonally adjusted monthly consumer price indices for food items and all items are used as proxy variables of the agricultural price and general price level. Consumer price indices of all commodities less food items, service items, and all items less food items are selected for comparison. We believe the consumer price index of commodities less food items can represent the manufacturing prices, while the consumer price index of service items can represent the price level of non-tradable goods. These data were collected by the Bureau of Labor Statistics (BLS) web site (www.bls.gov). Total trade weighted real exchange rates between the United States and major importing countries are used as a proxy variable for movements in the U.S. real exchange rate. The data are obtained from the Economic Research Service (ERS) of the U.S Department of Agriculture (USDA) web site (www.ers.usda.gov). Because the trade weighted real exchange index represents the U.S. dollar value compared to currencies of importing countries, an increase in the index represents an appreciation of the U.S. dollar. Finally, the sample consists of 348 observations extending from 1974:1 to 2002:12.

5. Empirical Results

As a first step of the analysis, the long-run cointegration vectors in equation (3.3) and (3.4) are estimated by the following procedures. First, we examine the stationarity of each variable with two different unit-root tests: the Said-Dickey (1984) and Philips-Perron (1988) tests. Second, because the test results suggest that all the price indices are difference stationary, we estimate the cointegration vector using Park's (1992) Canonical

Cointegration Regression (CCR) method, which is more efficient than the least squares estimator suggested by Engle and Granger (1987).

5.1. Unit-Root Tests

Preliminary graphical investigation suggested that all the price indices have obvious time trends so that, under the alternative of trend stationarity, the Said- Dickey (SD) (1984) and Phillips-Perron (PP) (1988) tests were applied.

Table 1: Unit-Root Test Results: Sample period 1974:1~2002:12.

	SD(1)	SD(3)	SD(5)	PP(1)	PP(3)	PP(,5)
All	-2.184	-1.904	-1.826	-3.023	-2.564	-2.356
Food	-2.131	-2.104	-2.190	-2.420	-2.290	-2.260
All less food	-2.245	-1.942	-1.935	-2.926	-2.484	-2.287
Commodity less Food	-2.029	-1.941	-1.932	-2.321	-2.175	-2.104
Service	-2.047	-1.844	-1.818	-1.950	-2.099	-1.950

Notes: Critical values for 1, 5, and 10 percent significance levels are –3.99, -3.43, and –3.14 for SD and PP test under the alternative of trend stationarity. The critical values come from MacKinnon (1991).

Because these tests are sensitive to the choice of order of autoregression, we report test results based on different orders of autoregression: one, three, and five. The results presented in Table 1 suggest that all the series are first difference stationary rather than trend stationary. Thus, a cointegration approach is used to obtain long-run relationship between variables[6].

[6] We also used the Park's G(p, q) test (1990) under the null hypothesis of trend stationarity. The test results also suggest the variables are first difference stationary rather than trend stationary.

5.2. Canonical Cointegration Regression

To obtain cointegration vectors in the equations (3.3) and (3.4), we applied Park's Canonical Cointegration Regressions (CCR). Park's nonparametric method may have some advantages as compared to Johansen's (1988) Maximum Likelihood (MLS) approach. The CCR method does not require a normality assumption and any assumption about the lag specification. Park and Ogaki (1991) show that, in Monte Carlo simulations, the CCR procedure consistently outperforms the ML approach in small samples. Asymptotically, the CCR and ML approach will give the same results, if the number of lags in vector autoregression (VAR) representation is true for Johansen's approach. We also applied the Park's H(p, q) test for testing cointegration relationships. Park's H(p, q) test is computed by the CCR residuals. Under the null of the cointegration, H(p, q) tests have asymptotically χ^2 distributions with q-p degrees of freedom, while under the alternative of no cointegration , the test statistic diverges to infinity. Therefore, unlike conventional tests (e.g., Augmented Dickey Fuller test), we can conclude the estimates are cointegration vector when the test statistics fail to reject the null hypothesis. In our model, each variable is treated as the first difference stationary with drift. Because of the drift, each variable can possess a linear deterministic trend as well as a stochastic trend. Therefore, we applied H(1,q) test statistics to the null hypothesis of stochastic cointegration[7].

[7] More detailed discussion about the concepts of deterministic and stochastic cointegration is presented in Park and Ogaki (1998).

Table 2 presents the estimated cointegration vectors and H (1, q) test results[8]. In the case of the food price, the estimated coefficient is 0.8001, which indicate the disproportionate increase of nominal food price compared to general price level during the sample period. In the case of other prices, the estimate coefficients are generally more than one (1.0460, 1.0701, and 1.1391), indicating these are more proportionately increased than general price level. As we discussed before, these results might be due to the different income elasticities and productivity growth rate of each commodity group.

Table 2: CCR Results (Sample: 1974:1~2002:12)

	Constant	*Trend*	$\ln p_t$	$H(1,3)$	$H(1,4)$	$H(1,5)$
Food	0.8597[a]	0.0005[a]	0.8001[a]	4.2357	4.6041	5.4162
	(0.029)	(0.0001)	(0.008)	(0.120)	(0.203)	(0.247)
All less food	-0.1996[a]	-0.0001[a]	1.0460[a]	4.7742[c]	5.5998	5.6622
	(0.014)	(0.00002)	(0.004)	(0.092)	(0.133)	(0.226)
Commodity less Food	-0.1992	-0.0012[a]	1.0701[a]	3.1922	3.8331	5.4030
	(0.239)	(0.0002)	(0.059)	(0.203)	(0.280)	(0.248)
Service	-0.6319[a]	0.00007	1.1391[a]	5.3557[c]	5.3751	7.8510[c]
	(0.027)	(0.00005)	(0.008)	(0.069)	(0.146)	(0.097)

Note: Numbers in parenthesis are the estimated standard errors; **a**, and **c** denote significant at the 1, and 10 percent levels.

The corresponding cointegration tests suggested by Park (1990) cannot reject the null of cointegration at the five percent level; we conclude that the estimates of CCR represent long-run relationship between variables in all cases.

[8] To implement CCR and Park's tests, Gauss routines programmed by Ogaki (1993) is used. In this program, QS kernel and Andrews' (1991) automatic bandwidth selector is used to obtain long-run covariance parameters.

5.3. Relative Price Changes in Different Time Horizons

If macroeconomic variables are important to explain changes in relative prices, the variables should cause more deviation in one price from its long-run equilibrium level than in other prices. Without considering these long run relationships determined by unidentified real factors, the regression results could be biased due to the omitted variable problem. To avoid this possibility, we construct price series as deviations of their long-run equilibrium levels using the estimated cointegrating vectors for each price variables, and then estimate the model (3.10). We present the results showing the changes of relative agricultural price compared to three selected price series (commodity less food items, service items, all less food items) in six different time horizons (1, 12, 24, 36, 48, and 60 months). Because preliminary test results suggest that, in the case of k=1, there are autoregressive conditional heteroskedasticity (ARCH) type errors, we report the results of GARCH (1, 1) model suggested by Bollerslev (1986) in this case. In other cases, however, the serial correlation is more serious problem so that we report the results with a heteroskedasticity and autocorrelation consistent (HAC) standard error suggested by Newey and West (1987).

Food vs. Non-Food Commodity Items

The first case is the relative price movement between food items and commodity less food items, which is expected to represent the relative price movements of the agricultural and other manufacturing goods. The estimation results are presented in Table 3. We find a significant linkage of one-month changes in general price level and one-month changes in relative agricultural prices. The estimated coefficient is negative

(-0.2915) and is significant at the one percent level. However, this linkage is disconnected when the time horizon is lengthened. The changes in general prices do not have any explanatory power for more than one-month changes in relative food and agricultural prices. None of the estimated coefficients except for the one-month changes are statistically significant at the ten percent level.

Table 3: Estimation Results: Food vs. Non-Food Commodity

	$k = 1$	$k = 12$	$k = 24$	$k = 36$	$k = 48$	$k = 60$
Constant	0.0009**b**	0.0080	0.0090	0.0069	0.0075	0.0115
	(2.095)	(1.489)	(1.374)	(1.023)	(0.972)	(1.244)
$\Delta \ln p_t$	-0.2915**a**	-0.1632	-0.0930	-0.0479	-0.0367	-0.0491
	(-3.752)	(-1.588)	(-1.497)	(-1.267)	(-1.102)	(-1.464)
$\Delta \ln r_t$	0.0206	-0.0878**b**	-0.1071**a**	-0.1159**a**	-0.1142**a**	-0.1005**a**
	(1.260)	(-2.045)	(-3.663)	(-5.897)	(-6.847)	(-5.341)
DW-statistics	1.408	0.130	0.093	0.095	0.082	0.059
Adj-R^2	0.023	0.126	0.208	0.324	0.357	0.325

Notes: *z*-ratios are in parenthesis in the case of the *k*=1: In other cases, Newey-West HAC standard errors are used to calculate the *t*-ratios; **a**, and **b** denote significant at the 1, and 5 percent levels.

In the case of the real exchange rates, however, the explanatory power increases when the time horizon is lengthened. The sign of the coefficients are negative for the twelve-month changes , and the absolute sizes of the coefficients and significance levels are increased from -0.0878 at *k*=12 to 0.1142 at *k*=48. At *k*=60, the significance levels and sizes of the coefficients become smaller than those at *k*=48. The adjusted R^2 also increases from 0.023 at *k*=1 to 0.357 at *k*=48. The negative signs imply that real appreciation of the U.S. dollar causes a decrease in food and agricultural price compared to other manufacturing commodity prices, which is consistent with our expectation.

Food vs. Service Items

Table 4 presents the relative food price movement compared to that of service items, which is expected to represent the prices of non-tradable goods. In the case of the one-month changes, inflation rate has a statistically significant explanatory power. The estimated coefficient is positive (0.2783) and significant at the ten percent level. However, the general price changes do not have an important role in explaining the long-term changes of the relative food and agricultural price. In any case except for the one- month changes, the estimated coefficients are not statistically significant.

Table 4: Estimation Results: Food vs. Service

	$k = 1$	$k = 12$	$k = 24$	$k = 36$	$k = 48$	$k = 60$
Constant	-0.0012[a] (-4.899)	-0.0087[b] (-2.143)	-0.0156[a] (-2.664)	-0.0223[a] (-2.720)	-0.0283[a] (-3.239)	-0.0322[a] (-3.672)
$\Delta \ln p_t$	0.2703[a] (4.092)	0.0952 (0.876)	0.0739 (0.987)	0.0732 (1.105)	0.0692 (1.487)	0.0510 (1.485)
$\Delta \ln r_t$	0.0097 (1.032)	-0.0511 (-1.541)	-0.0623[b] (-2.013)	-0.0708[b] (-2.404)	-0.0844[a] (-3.169)	-0.0807[a] (-3.468)
DW-statistics	1.331	0.124	0.068	0.046	0.044	0.036
Adj-R^2	0.021	0.045	0.074	0.113	0.187	0.195

Notes: *z*-ratios are in parenthesis in the case of the *k*=1: In other cases, Newey-West HAC standard errors are used to calculate the *t*-ratios; **a**, and **b** denote significant at the 1, and 5 percent levels.

In the case of the real exchange rate, it does not have explanatory power in explaining relatively short-term, one-month and twelve months, changes in relative agricultural prices. However it has statistically significant explanatory power in the cases of the 24, 36, 48, and 60 months changes. The significance of the estimated coefficient increases when the time horizons are increased.

<u>Food vs. Non-Food Items</u>

Finally, we present the estimation results of relative price movements between food items and non-food items in Table 5. As whole, the economic implication of the results is similar to the previous cases. In the case of the general price changes, we do not find any significant results in any of the time horizons except the one-month changes.

Table 5: Estimation Results: Food vs. Non-Food Prices

	$k = 1$	$k = 12$	$k = 24$	$k = 36$	$k = 48$	k=60
Constant	0.0003 (0.204)	-0.0009 (-0.241)	-0.0036 (-0.749)	-0.0075 (-1.178)	-0.0106 (-1.456)	-0.0112 (-1.466)
$\Delta \ln p_t$	-0.2089[a] (-2.807)	-0.0285 (-0.292)	-0.0102 (-0.169)	0.0093 (0.192)	0.0148 (0.396)	0.0013 (0.042)
$\Delta \ln r_t$	0.0084 (0.655)	-0.0675[b] (-2.642)	-0.0832[a] (-3.864)	-0.0918[a] (-4.492)	-0.0977[a] (-5.409)	-0.0901[a] (-5.122)
DW-statistics	1.344	0.132	0.087	0.068	0.061	0.048
Adj-R^2	-0.043	0.087	0.164	0.238	0.305	0.318

Notes: *z*-ratios are in parenthesis in the case of the *k*=1: In other cases, Newey-West HAC standard errors are used to calculate the *t*-ratios; **a**, and **b** denote significant at the 1, and 5 percent levels.

However, in the case of the real exchange rate, the significance of the variables increases when the time-horizon is lengthened, similar to the previous cases. The estimated coefficient is 0.0084, which is not statistically significant at *k*=1. However, the sign of the coefficients are changed to negative from *k*=12, and the absolute sizes of the coefficients and significance levels increase from -0.0675 at k=12 to 0.0977 at k=48. The adjusted R^2 are also increased from -0.005 at k=1 to 0.318 at k=60. The signs of the coefficients are all negative and significant.

5. Conclusion

Since the influential paper by Schuh (1974), agricultural economists have long recognized the potential effect of exchange rate movements on the agricultural sector in the United States. When exchange rates became floating, most agricultural economists believed that the new market-based system could substantially mitigate the misalignment problem. Although some agricultural economists (e.g., Gardner 1981; Tweeten, 1989) have recognized the potential instability problem of the U.S. agricultural sector after experiencing unexpected U.S. dollar movements during the post-Bretton Woods era, researches have been concentrated on the linkage between relative agricultural prices and inflation rates, while the potential linkage between exchange rates and relative food and agricultural prices has been largely ignored. This lack of study might be due to influence of the traditional macroeconomic view of the flexible exchange rate system. If the foreign exchange market has been working properly, nominal exchange rates properly align inflation rates between countries, which might prevent permanent over-valuation (or under-valuation) of the U.S. dollar under a fixed exchange rate system. However, in reality, empirical studies have suggested much evidence against the monetary economic view of the floating exchange rate system. Cyclical misalignments of the U.S. dollar have been persistent and substantial during the flexible exchange rate system. In fact, controversial proposals relating to international monetary reform (e.g., Williamson, 1989; Krugman, 1989; Mundell, 1992; McKinnon, 1995)[9] show how seriously these problems are considered in this area.

The economic implication of the present study is simple. The main source of the variation in relative agricultural price is the variation in real exchange rate movements, especially long-term variation. Conversely, variation of the general price has a role in explaining short-term changes in relative agricultural price at best[10]. Considering the fact that the misalignment problem of the U.S. dollar has been cyclical and prolonged, the long-term linkage between the variations in real exchange rate and relative agricultural prices implies that the U.S. agricultural sector has faced a prolonged instability problem due to the U.S. dollar movements. The results imply that the U.S. monetary policy alone might not be enough to prevent the possibility of an instability problem in the U.S. agricultural sector caused by U.S. dollar movements in the future. If an important source of the misalignment problem is some degree of inefficiency in the foreign exchange market, internationally coordinated monetary policy must be important.

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[9] More detailed discussion concerning different proposals on international monetary reform is summarized in Frankel (1996).

[10] However, we do not completely rule out indirect effect of the U.S. monetary policy on movements of the U.S. dollar, and, hence, on relative agricultural prices although we do not find strong linkage between inflation rate and relative agricultural prices. Eichenbaum and Evans (1995) find that around 17 percent variations in the U.S. real exchange rates can be explained by the U.S. monetary policy variation during the flexible exchange rate system. Rogers (1999) also found US monetary policy has been responsible for a minimum 20 percent variation of real exchange rates between the dollar/pound during the period of 1889-1992.

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