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**The Relative Impact of National Monetary Policies and International  
Exchange Rate on Long-term Variations in Relative Agricultural Prices**

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## **Abstract**

This paper seeks to explain the causes of the long-term variation in food and agricultural prices compared to the overall price level in the United States, over the period of 1974-1996. Using cointegration methods, this study confirms a general consensus of long-run neutrality of national money (money supply) and gives practical evidence of the real impact of international money (exchange rate) on the long-term variation of relative agricultural price in the United States, especially during the 1974-1988 period.

**Keywords:** long-term variation of agricultural prices, exchange rates, money supply

## Highlights

Chronic instability is the most serious problem in the U.S. agricultural sector, and one of the critical factors contributing to this instability is the long-run movement of agricultural price. Although agricultural economists (e.g., Gardner 1981; Tweeten 1989) recognized the exchange rate instability problem of the U.S. agricultural sector after experiencing unexpected U.S. dollar movements during the post-Bretton Woods era, the long-run linkage between exchange rates and relative food and agricultural prices in the United States has been ignored due to the strong influence of the monetary economic view of the flexible exchange rate system. However, in practice, empirical studies have generated substantial evidence which contradicts the monetary economic view: finding that cyclical movements of the U.S. dollar have been persistent and substantial during the flexible exchange rate system.

This study tests the long-run neutrality of the domestic money supply and exchange rates on the movements of relative agricultural prices in the United States over the period of 1974-1996. In this study, we derive a simple and new empirical model to test the long-run neutrality of the money supply and the U.S. dollar. A non-parametric cointegration method (CCR) developed by Park (1992) is implemented for estimation. This method is more efficient than the method by Johansen (1988) for a small sample.

The main findings in this study confirm a general consensus concerning long-run neutrality of national money (money supply) and give practical evidence of the real impact of international money (exchange rate) on the long-term variation of relative agricultural price in the United States. In fact, we found a 1 percent real appreciation of U.S. dollar is associated with a 0.131 percent decrease (increase) of food prices compared to the aggregate price level in the long-run.

# The Relative Impacts of National Monetary Policies and International Exchange Rates on Long-term Variations in Relative Agricultural Prices

Guedae Cho, MinKyoung Kim, and Won W. Koo\*

## INTRODUCTION

Chronic and cyclical instability is the most serious problem in the U.S. agricultural sector. One of the important factors causing instability is the movement of agricultural price, especially its long-term variation. Unlike insurable short-term price volatility, long-term price variability, which exists continuously over several years, can have a detrimental impact on the U.S. farm economy. Tweeten (1989) argues that relatively low agricultural prices during the mid-1980s brought about a major restructuring of the U.S. farm economy. The cyclical and prolonged variation of relative agricultural price is easily observable in Figures 1 and 2. Compared to the historical trend, real food prices surged during the 1974~1980 period, while they plummeted during the 1981~1989 period.

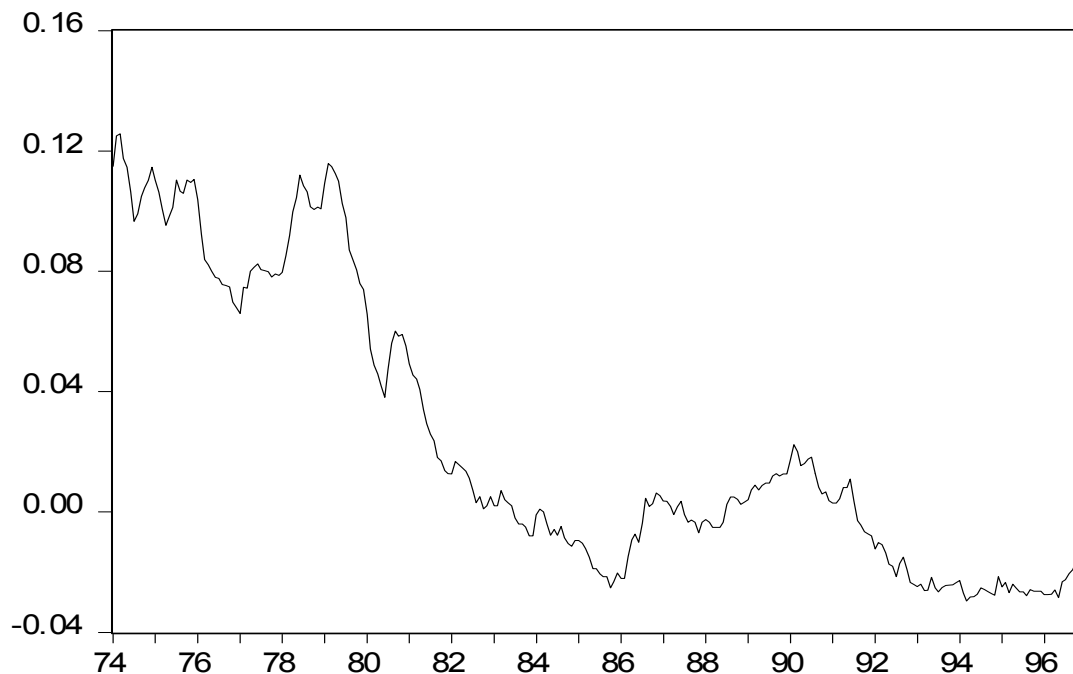
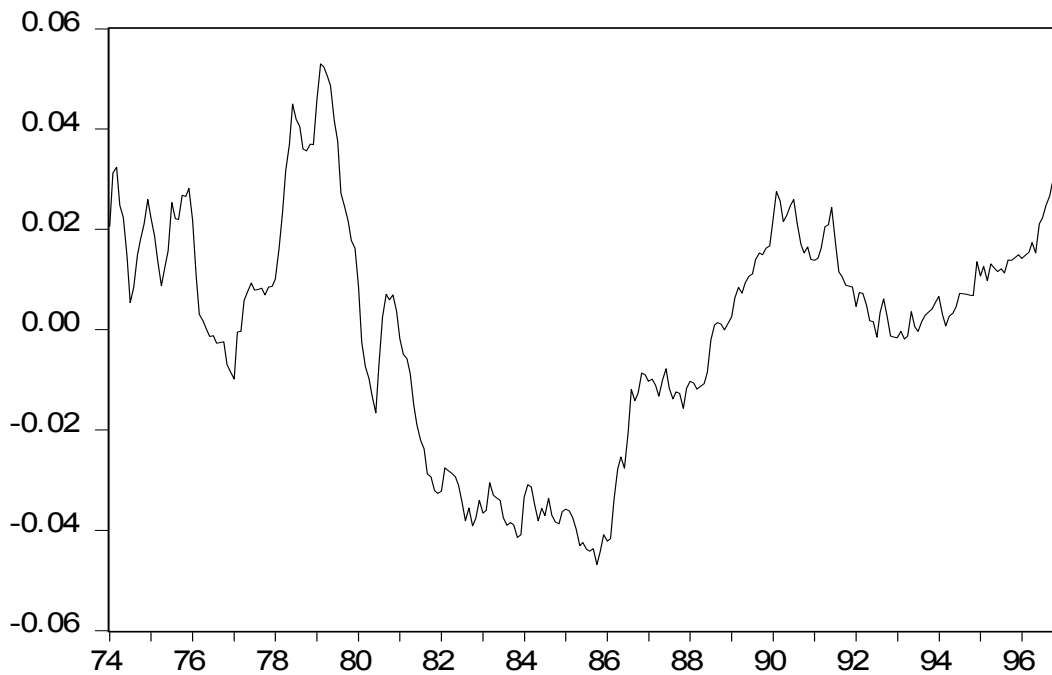


Figure 1. Historical Movement of U.S. Real Food Prices: 1974:1~1996:12

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**Figure 2. Historical Movement of U.S. Real Food Prices (Detrended)**

Studies have long discussed macroeconomic factors as causes of the long-term variation in relative agricultural prices, and exchange rates have often been mentioned as an important factor. Since the U.S. agricultural sector has generally been more heavily involved in international trade than other industries, U.S. dollar movement can disproportionately affect agricultural trade, which causes over-(under-) supply of domestic agricultural products. This results in variation of relative agricultural prices (e.g., Schuh 1974; Gardner 1981; Batten and Belongia 1986; Tweeten 1989).

However, there have been few empirical studies examining the relationship between exchange rate and long-term relative agricultural price since the mid-1980s because of the strong influence of the monetary economic view about exchange rates under the floating exchange rate system. The flexible and sticky price monetary models (e.g., Dornbusch 1976) assume that purchasing power parity (PPP) holds even in the short-run, so that the nominal exchange rate cannot deviate from the inflation ratio (or monetary fundamentals) between countries for a substantial period. Consequently, the U.S. dollar movement is not expected to distort relative purchasing power between countries, which implies the U.S. dollar movement has an insignificant real impact on agricultural exports, as well as on domestic relative agricultural prices. Therefore, even if we accept the hypothesis that exchange rate movements specially affect the U.S. agricultural sector, the existence of large and persistent real over- (under-) valuation of the U.S. dollar is not easily supported by these standard macroeconomic models.

Under the assumption that PPP holds even in the short-run, the observable long-term variations of relative agricultural price could be explained by the variation in the domestic monetary policy (or money supply).<sup>1</sup> Many studies have analyzed the effect of the money supply (or inflation rate) on the relative agricultural price,<sup>2</sup> and empirical evidence finds overshooting of relative agricultural price in response to money supply (e.g., Devadoss and Meyers 1987; Robertson and Orden 1990; Zanas 1998; Saghaian *et al.* 2002).<sup>3</sup> However, these studies have concentrated on the issue of short-term (monthly, quarterly, and yearly) variation of relative agricultural price, due to the prevailing belief of long-run money neutrality. Thus, the observable *long-term* cyclical fluctuation of relative agricultural price during the post-Bretton Woods period has been ignored.

Recent studies in international macroeconomics and finance provide clues in explaining the factors causing long-term cyclical variation of relative agricultural prices. Empirical evidence indicates that there is some degree of inefficiency in the foreign exchange market (e.g., Frankel and Froot 1987, 1990; Ito 1990; Frankel and Rose 1995), resulting in large and persistent deviation of the nominal exchange rate from its PPP (or monetary fundamentals) (Dornbusch 1987; Rogoff 1996). The speed of convergence to PPP is very slow: deviations appear to dampen out at a rate of roughly 15 percent per year (Rogoff 1996). Therefore, it may be hypothesized that the U.S. dollar movement has a long-run relationship with the movement of relative agricultural price, even when the domestic money supply is neutral in the long-run.

The primary objective of the study is, therefore, to test the *long-run* neutrality of the money supply and exchange rates on the movements of relative agricultural prices in the United States over the 1974-1996 period. Empirically, Park's (1992) Canonical Cointegration Regression (CCR), which is known as a non-parametric counterpart of the widely used Johansen (1988) method, is implemented to examine the hypothesis. This study develops an improved empirical model to generate a proper explanation for the existence of exchange rate impacts on the relative agricultural price in the long-run. The main finding of the paper is simple but interesting, namely, national money (money supply) is neutral, but international money (exchange rate) is not neutral in the long-term variation of relative agricultural price in the United States.

The paper is organized as follows. The next section presents a discussion about the exchange rate movement under the floating system. The third derives an empirical model designed to test

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<sup>1</sup> The debate between Grennes *et al.* (1980) and Chambers and Just (1980) show how difficult it is to justify the effect of the exchange rate on relative agricultural price under the assumption of holding PPP even in the short-run.

<sup>2</sup> In fact, several authors have suggested different possibilities concerning the link between money supply (or inflation rate) and relative price variation in an economy (e.g., Lucas 1973; Barro 1976; Shehinski and Weiss 1977; Bordo 1980; Frankel 1986; Ball and Romer 1989; Ball and Mankiw 1995). Although these theoretical models can explain the short-term variation in relative price, they usually assume that money is neutral in the long-run.

<sup>3</sup> Saghaian *et al.* (2002) find that monetary policy is related to the relative price-overshooting problem of agricultural product in the case of the United States. They conclude that money is not neutral even in the long-run, reviving the issue of the long-run money neutrality.

for the existence of impacts (or long-run neutrality) of both money supply and exchange rates. The subsequent section presents data and preliminary test results of the variables, and the econometric method and the main empirical findings of the paper are presented in sections 5 and 6, respectively. The final section consists of a summary and concluding remarks.

## EXCHANGE RATE MOVEMENT UNDER THE FLOATING SYSTEM

It is important to discuss the question of whether real exchange rate movements can cause a prolonged overshooting problem for agricultural prices in a theoretical ground. As mentioned earlier, this long-run relationship between the exchange rate and the relative agricultural price is not supported by standard macroeconomic models. Consider the theory of purchasing power parity (PPP), which is one of the fundamental assumptions in a flexible price monetary model. Under the assumption of a fully integrated world goods market, this theory assumes that the following PPP condition holds internationally:

$$\frac{P_t}{P_t^*} = \theta \cdot S_t \quad \text{or} \quad \frac{P_t}{P_t^* S_t} = R_t = \theta, \quad (1)$$

where  $P_t$  and  $P_t^*$  are the aggregate price levels of the home and foreign countries;  $S_t$  and  $R_t$  are the nominal and real exchange rates between the home country and a foreign country (i.e., units of home currency required to buy one unit of foreign currency).  $\theta$  represents potential factors that cause a deviation of the nominal exchange rate from PPP, such as transportation costs and trade barriers, which are assumed to be constant in the long-run. It is important to note that the real exchange rate,  $R_t$ , is assumed to be a constant  $\theta$  in both the short-run and the long-run.

Therefore,  $\ln R_t$  is a fixed parameter in the long-run, so we do not expect that  $\ln R_t$  has a long-run relationship with any cyclical variation of real agriculture prices. Simply put, a constant factor cannot be related to the cyclical deviation of food and agricultural prices from their long-run equilibrium path.

However, there is an argument about possible deviations of nominal exchange rate from PPP. Under the sticky-price monetary model (Dornbusch 1976), a short-run deviation of nominal exchange rates from PPP may be possible due to the stickiness of nominal wages and prices. In other words, the adjustment speed of the financial sector in response to a nominal shock is assumed to be faster than that of real sectors, which could possibly cause temporary overshooting of nominal exchange rates. However, according to the model, it is also true that the nominal shock should cause only temporal overshooting of nominal exchange rates, which might cause short-term volatility rather than long-term cyclical variation of real exchange rates.

Recent empirical evidence in international macroeconomics and finance indicates the strong possibility of a persistent deviation of nominal exchange rates from PPP, suggesting the long-run fluctuation of real exchange rates. Empirical evidence also indicates the possibility of some degree of inefficiency in the foreign exchange market. Monetary fundamentals, at best, only explain the long-run movement of nominal exchange rates (Mark 1995); PPP, therefore, might hold only in the long-run (Rogoff 1996). A rational speculative model is proper to explain the

long-term deviations.<sup>4</sup> If the nominal exchange rate moves in one direction, it will drift in the same direction for a long time unless there is an important economic event which can change the direction of expectations held by foreign exchange market participants (e.g., Frankel and Rose 1995; Frankel 1996).

Under both the sticky-price model and the rational speculative bubble model, we can define the real exchange rate as:

$$\frac{P_t}{P_t^* S_t} = R_t = \theta + f_t, \quad (2)$$

where  $f_t$  relates to unobservable stochastic factors that cause fluctuation of real exchange rates. Under the sticky-price monetary model,  $f_t$  might be serially correlated, but the coefficient of autoregression is far less than one, so the deviation of real exchange rates from an arbitrary constant  $\theta$  would die out within a short time period. By contrast, under the assumption of a rational speculative bubble model,  $f_t$  could possibly have a unit root or near unit root (the coefficient of autoregression is almost one). Therefore, only by assuming the existence of a rational speculative bubble in the foreign exchange market can real exchange rate movements have any explanatory power for the prolonged overshooting of food prices in the United States.

### EMPIRICAL MODEL SPECIFICATION

Following Grennes and Lapp (1986), Robertson and Orden (1990), Zanas (1998), and Saghalian *et al.* (2002), in order to test the long-run neutrality of money and exchange rate, *long-run* relationships among the stock of money, real exchange rates, nominal food and agricultural prices, and the aggregate price level can be specified as

$$\ln P_t^A = \alpha_0 + \alpha_1 \ln M_t + \alpha_2 \ln R_t + \varepsilon_t, \quad (3)$$

$$\ln P_t = \beta_0 + \beta_1 \ln M_t + \beta_2 \ln R_t + v_t, \quad (4)$$

where  $\ln M_t$  is the log of the stock of money at time  $t$ ;  $\ln R_t$  is the log of real exchange rates between the United States and its trading partners;  $\ln P_t^A$  and  $\ln P_t$  are the log of food and

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<sup>4</sup> 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, we do not have any universally accepted reason about what initiates the 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).

agricultural price and aggregate price level, respectively, in the United States; and  $\varepsilon_t$  and  $v_t$  are the stochastic terms in respective equations which include unidentified macroeconomic shocks and commodity group specific idiosyncratic shocks.

Previous studies considered  $\alpha_1 = \beta_1$  as a condition of long-run money neutrality. That is, one percent increase in the money supply should generate the same percentage increase in agricultural and overall prices. However, as Friedman (1975) notes, an increase in the money supply causes an increasing *average* price level in an economy, and long-run relative prices between commodities should be determined by the movements of relative underlying supply and demand conditions of their products. This indicates that it is possible for food and agricultural prices to move disproportionately to aggregate prices regardless of changes in money supply in the long-run. Thus, if nominal food and agricultural prices were observed to increase disproportionately compared to aggregate prices, the impact of the money supply ( $\ln M_t$ ) on food and agricultural prices ( $\ln P_t^A$ ) would be seemingly different from the impact on aggregate prices ( $\ln P_t$ ). In this case,  $\alpha_1$  should be smaller than  $\beta_1$  ( $\alpha_1 < \beta_1$ ). Although this result is interpreted as evidence against the long-run money neutrality hypothesis (e.g., Saghayan *et al.* 2002), it is a simple long-run relationship; the empirical model is not proper to test the long-run money neutrality hypothesis.

Therefore, there is an additional, important long-run relationship that should not be ignored: between food and agricultural prices and aggregate price. As Kliesen and Poole (2000) discussed, the long-run relationship could be explained by unobservable relative movements of factors, such as the different degrees of income and demand elasticities. In practice, however, it is not easy to include all the structural variables into the empirical model. In our analysis, we include an additional long-run equilibrium relationship between nominal food and agricultural prices and general price by adapting the idea of the rational expectation model, which suggests that variation of relative price realizes all the variations of underlying movement of relative supply and demand over time, especially in the long-run.

Let us assume there is the following long-run relationship between food and agricultural price and general price, which is determined by unobservable real factors:

$$\ln P_t^A = \gamma_0 + \gamma_1 \ln P_t + \eta_t . \quad (5)$$

Pre-multiply Equation (4) by  $-\gamma_1$ , and add (3) to (4) to obtain the following long-run relationship:

$$\ln P_t^A - \gamma_1 \ln P_t = \alpha_0 - \gamma_1 \beta_0 + (\alpha_1 - \gamma_1 \beta_1) \ln M_t + (\alpha_2 - \gamma_1 \beta_2) \ln R_t + (\varepsilon_t - \gamma_1 v_t) . \quad (6)$$

Or, equivalently,

$$\ln P_t^A = \delta_0 + \gamma_1 \ln P_t + \delta_1 \ln M_t + \delta_2 \ln R_t + \xi_t , \quad (6')$$

where  $\delta_0 = \alpha_0 - \gamma_1 \beta_0$ ;  $\delta_1 = \alpha_1 - \gamma_1 \beta_1$ ;  $\delta_2 = \alpha_2 - \gamma_1 \beta_2$ ; and  $\xi_t = \varepsilon_t - \gamma_1 v_t$ .

If food and agricultural prices have reacted more sensitively than the aggregate price level in response to changes in the money supply,  $\delta_1 > 0$  and  $\alpha_1 > \gamma_1 \beta_1$ ;  $\delta_1 < 0$  and  $\alpha_1 < \gamma_1 \beta_1$ , otherwise. If food and agricultural prices have reacted more sensitively in response to real exchange rates,  $\delta_2 < 0$  and  $\alpha_2 < \gamma_1 \beta_2$ ;  $\delta_2 > 0$  and  $\alpha_2 > \gamma_1 \beta_2$ , otherwise. Finally, under the assumption of money and real exchange neutrality,  $\delta_1$  and  $\delta_2$  are expected to be zero, meaning that  $\alpha_1 = \gamma_1 \beta_1$  and  $\alpha_2 = \gamma_1 \beta_2$ .

Formally, there are three possible cases for the empirical model (6').

**(Case 1):** if the relationships (3), (4), and (5) are the true long-run relationships among variables, and if long-run money and real exchange rate neutrality holds, the estimated coefficients  $\delta_1$  and  $\delta_2$  in Equation (6') should be equal to zero. The estimated coefficient  $\gamma_1$  should be a cointegration vector because  $\xi_t$  should be a stationary process by assumption of the underlying innovations.

**(Case 2):** if relationships (3), (4), and (5) are true long-run relationships, but either long-run money or real exchange rate neutrality does not hold, the estimated coefficients  $\delta_1$  and  $\delta_2$  are equal to zero. However, the estimated coefficients in Equation (6)' are a long-run cointegration vector by assumption of the innovations.

**(Case 3):** if the estimated coefficients of (6') are not a cointegration vector, so that the estimated residual  $\xi_t$  is a non-stationary process, Equations (3), (4), and (5) should not be the true long-run equilibrium conditions. Or, with given variables, we cannot identify the true long-run relationship among variables because there might be unobservable factors which have caused cyclical deviation of food and agricultural prices from their long-run equilibrium path.

Note that the estimated coefficient  $\delta_2$  tests the sensitivity of food and agricultural prices in response to movements of real exchange rates relative to aggregate prices. In other words, even when estimated  $\delta_2$  equals zero, it does not indicate that real exchange rates have not caused any real effect on food and agricultural exports and domestic food and agricultural prices. Rather, it means that real exchange rate variations have almost the same long-run effect on food and agricultural prices and the aggregate price level.

## DATA AND PRELIMINARY TESTS

### *Data*

We used seasonally adjusted monthly data (1974:1~1996:12) of the consumer price index of food as a proxy variable for the price of food and agricultural commodities ( $P_t^A$ ) and the index of all items as a proxy variable for the aggregate price level or inflation rate ( $P_t$ ). These data are collected from the Bureau of Labor Statistics (BLS). The money stock M1 from the Federal Reserve Bank in St. Louis is used as a proxy variable of the money supply ( $M_t$ ). Finally, agricultural trade-weighted real exchange rates between the United States and major trading countries are used as a proxy variable of U.S. real exchange rate movements ( $R_t$ ). The exchange rate data are obtained from the U.S. Department of Agriculture (USDA). Because the trade-weighted real exchange index represents the dollar value compared to currencies of all trading countries, an increase in the exchange rate index represents an appreciation of the U.S. dollar. Finally, the number of observations is 276 for each variable.

### *Unit-Root Tests*

The stationarity of each variable in (6') is examined with four different unit-root tests. The usual Said-Dickey (1984) and Philips-Perron (1988) tests are used with the null hypothesis of non-stationarity. To confirm the test results, Park and Choi's (1988) J(p, q) test and G(p, q) test are implemented with the null hypothesis of the first difference stationarity and stationarity, respectively. Preliminary graphical investigation suggests that the log of food prices ( $\ln P_t^A$ ), the log of the overall price level ( $\ln P_t$ ), and the log of the stock of money ( $\ln M_t$ ) have time trends, while the log of the trade-weighted real exchange rate ( $\ln R_t$ ) does not. Therefore, the unit root tests are implemented with the alternative hypothesis of trend stationarity in the cases of  $\ln P_t^A$ ,  $\ln P_t$ , and  $\ln M_t$ , while with the alternative of stationarity without trend for  $\ln R_t$ .

Table 1 reports the results of the Said- Dickey (SD) and Phillips-Perron (PP) tests. Because these tests are sensitive to the choice of autoregression order, they are conducted based on different autoregression orders, 1, 3, and 5. In the case of  $\ln P_t^A$ ,  $\ln P_t$ , and  $\ln M_t$ , both the SD and PP tests do not reject the null hypothesis of the stationarity for the first difference and reject the alternative hypothesis of trend stationarity. In the case of the  $\ln R_t$  series, both tests also do not reject the null hypothesis of difference stationarity and reject the alternative of stationarity without trend. Thus, the tests conclude that all the variables in Equation (6') are difference stationary processes.

**Table 1. Unit-Root Test Results: Sample Period 1974:1~1996:12**

	SD(1)	SD(3)	SD(5)	PP(1)	PP(3)	PP(5)
$\ln(P_t)$	-1.392	-1.392	-1.355	-1.694	-1.510	-1.445
$\ln(P_t^A)$	-1.670	-1.699	-1.721	-1.826	-1.779	-1.770
$\ln(M_t)$	0.682	-0.118	-0.556	2.323	1.496	1.031
$\ln(R_t)$	-1.725	-1.703	-1.5431	-2.231	-1.516	-1.606

Notes: Critical values for 1, 5, 10 percent significance levels are -3.99, -3.43, and -3.14 for SD and PP tests under the alternative of trend stationarity, and -3.46, -2.87, and -2.57 under the alternative of stationarity without trend. The critical values come from MacKinnon (1991).

In addition, Park's (1990)  $J(p, q)$  and  $G(p, q)$  tests are conducted to confirm the results of SD and PP tests. When a series has a deterministic time trend,  $J(1, q)$  and  $G(1, q)$  tests are recommended. On the other hand, if a series does *not* have a deterministic trend,  $J(0, q)$  and  $G(0, q)$  tests should be utilized.

The test results are reported in Table 2. In the cases of  $\ln P_t^A$ ,  $\ln P_t$ , and  $\ln M_t$ , both tests consistently suggest nonstationarity of variables, which is consistent with results of the previous two tests. However, in the case of  $\ln R_t$ , the test gives mixed results. For instance,  $J(0, 3)$  and  $G(0, 3)$  test results suggested stationarity of the variable, while the  $G(0, 4)$  test results suggested nonstationarity. This contradictory evidence is quite interesting with respect to recent debates about the stationarity of real exchange rates under the floating exchange rate system. Many empirical studies based on panel datasets have suggested stationarity of real exchange rates (e.g., Frankel and Rose 1996; Taylor and Sarno 1998). However, it is also true that economists do not confirm the stationarity of real exchange rates when univariate time-series data are used. Without the existence of a universally most-powerful test, it is difficult to confirm the stationarity of real exchange rates using any particular test method. As Engel (2000) indicated, real exchange rates are viewed to have a near unit root, which has caused the quite persistent departure of nominal exchange rates from their PPP level. In this paper, real exchange rates will be treated as a non-stationary variable based on the results of the two popular tests.

**Table 2.  $J(p, q)$  and  $G(p, q)$  Test Results**

	$J(0, 3)$	$J(1, 5)$	$G(0, 3)$	$G(0, 4)$	$G(1, 2)$	$G(1, 3)$
$\ln(P_t)$	--	13.93	--	--	11.82 (0.001)	12.29 (0.002)
$\ln(P_t^A)$	--	7.099	--	--	11.63 (0.001)	12.13 (0.002)
$\ln(M_t)$		1.007			2.935 (0.087)	8.268 (0.016)
$\ln(R_t)$	0.323	--	3.498 (0.321)	7.928 (0.094)	--	--

Notes: Critical values for 1, 5, 10 percent significance levels are 0.123, 0.295, and 0.452 for  $J(1, 5)$ , and 0.112, 0.338, and 0.577 for  $J(0, 3)$  test. The critical values come from Park and Choi (1988).  $G(p, q)$  has an asymptotic chi-square distribution with  $q-p$  degree of freedom. Numbers in parenthesis are p-values.



## ESTIMATION METHOD<sup>5</sup>

To analyze the long-run relationship among variables, the least squares estimator (LSE) suggested by Engle and Granger (1987) has been successful. It is known for being super-consistent in estimating long-run cointegration vectors under some regularity conditions. However, it is not asymptotically efficient because it does not incorporate short-run dynamic adjustment of variables into a model. There is another relatively new method, Park's (1992) Canonical Cointegration Regression (CCR). This method is developed in a non-parametric framework, so that it does not rely on rigid assumptions such as normality of error terms, and is known as a more efficient estimator than the LSE. Both methods are utilized and the results are compared in this study.

The first method, the least squares estimator (LSE), is briefly explained as follows. Suppose  $\mathbf{Z}_t$  is a  $n \times 1$  vector of difference stationary random variables, with  $\Delta \mathbf{Z}_t$  being stationary. If there exists a nonzero vector of real number  $\mathbf{a}$  such that  $\mathbf{a}'\mathbf{Z}_t$  is stationary, then  $\mathbf{Z}_t$  is said to be cointegrated with a cointegrating vector  $\mathbf{a}$ . It is often convenient to normalize one element by one. Suppose that the first element of  $\mathbf{a}$  is zero, then partition  $\mathbf{Z}_t$  by  $\mathbf{Z}_t = (y_t, \mathbf{X}_t')$  and normalize  $\mathbf{a}$  by  $\mathbf{a} = (1, -\mathbf{c})$ . Here,  $y_t$  is a difference stationary process,  $\mathbf{X}_t$  is a vector difference stationary process, and  $\mathbf{c}$  is a normalized cointegrating vector.

The cointegration system (6') can be expressed in a convenient form,

$$y_t = \mathbf{X}_t' \mathbf{c} + \varepsilon_t \quad (7)$$

$$\Delta \mathbf{X}_t' = \mathbf{v}_t \quad (8)$$

where  $y_t = \ln P_t^A$ ,  $\mathbf{X}_t' = [1, \ln P_t, \ln M_t, \ln R_t]$ , and  $\mathbf{c}' = [\delta_0, \gamma_1, \delta_1, \delta_2]$  in our case. The  $y_t$  and  $\mathbf{X}_t$  are first difference stationary, and  $\varepsilon_t$  and  $\mathbf{v}_t$  are stationary with zero means. Let

$$\mathbf{w}_t = (\varepsilon_t, \mathbf{v}_t)'. \quad (9)$$

Define  $\Phi(i) = E(\mathbf{w}_t \mathbf{w}_{t-i}')$ ,  $\Sigma = \Phi(0)$ ,  $\Gamma = \sum_{i=0}^{\infty} \Phi(i)$ , and  $\Omega = \sum_{i=-\infty}^{\infty} \Phi(i)$ . In fact, the  $\Omega$  is the long-run variance matrix of  $\mathbf{w}_t$ . Partition  $\Omega$  as

$$\Omega = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{bmatrix} \quad (10)$$

where  $\Omega_{11}$  is a scalar, and  $\Omega_{22}$  is  $(n-1) \times (n-1)$  matrix, and partition  $\Gamma$  similarly.

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<sup>5</sup> Most of the explanation in this section is based on Ogaki and Jang (2001).

Define

$$\Omega_{11.2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21} \quad (11)$$

and  $\Gamma_2 = (\Gamma_{12}, \Gamma_{22})'$ .

The LSE is a parametric correction for long-run correlation of short-run movements of the variables and error term, for instance Johansen's (1988) maximum likelihood estimation. Because our interest is not short-run dynamics but a long-run relationship among variables, Park's CCR is more convenient to estimate the long-run cointegrating vector, which is considered as a second type of efficient estimation methods. To briefly explain the idea, consider the following transformations:

$$y_t^* = y_t + \Pi_y' \mathbf{w}_t \quad (12)$$

$$\mathbf{X}_t^* = \mathbf{X}_t + \Pi_x' \mathbf{w}_t. \quad (13)$$

Because  $\mathbf{w}_t$  is stationary,  $y_t^*$  and  $\mathbf{X}_t^*$  are cointegrated with the same cointegrating vector  $(1, -\mathbf{c})$  as  $y_t$  and  $\mathbf{X}_t$  for any  $\Pi_y$  and  $\Pi_x$ . The idea of the CCR is to choose  $\Pi_y$  and  $\Pi_x$ , so that the OLS estimator is asymptotically efficient when  $y_t^*$  is regressed on  $\mathbf{X}_t^*$ . The suggested matrices are

$$\Pi_y = \Sigma^{-1}\Gamma_2\mathbf{c} + (0, \Omega_{12}\Omega_{22}^{-1})' \quad (14)$$

$$\Pi_x = \Sigma^{-1}\Gamma_2. \quad (15)$$

In practice, long-run covariance parameters in these formulas should be estimated, and estimated  $\Pi_y$  and  $\Pi_x$  are used to transform  $y_t$  and  $\mathbf{X}_t$ . As long as these parameters are estimated consistently, the resultant CCR estimator is asymptotically efficient. In order to implement the CCR practically, the long-run covariance parameters  $\Omega$  and  $\Gamma$  are estimated using the Gauss module programmed by Ogaki (1993). In the program, the VAR pre-whitening method introduced by Andrews and Monahan (1992) is used in the first step of the estimation. Andrews (1991) automatic bandwidth parameter with QS kernel is implemented to estimate the sample counterpart of the  $\Omega$  and  $\Gamma$ .<sup>6</sup> Park's nonparametric method may have some advantages compared to Johansen's maximum likelihood (ML) estimator. The CCR method does not require any assumption about the lag specification, as the estimation is done directly on the cointegrating regression. Park and Ogaki (1991) show in Monte Carlo simulations that the CCR procedure consistently outperforms the ML approach in small samples. Asymptotically, the CCR and ML approaches will give the same results, if the number of lags in VAR representation is true for Johansen's approach.

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<sup>6</sup> See Ogaki (1993) for a more detailed discussion about the program.

## RESULTS

Table 3 summarizes the results of the LSE and the CCR estimations. The CCR estimation is conducted twice, with and without the money supply ( $\ln M_t$ ), to examine the robustness of our argument.

**Table 3. Results of Cointegration Approaches (LSE vs. CCR)**

	LSE	CCR	CCR without $M_t$
Constant	1.411 <sup>a</sup> (33.4)	1.129 <sup>a</sup> (9.10)	1.097 <sup>a</sup> (11.6)
$\ln(P_t)$	0.911 <sup>a</sup> (96.5)	0.874 <sup>a</sup> (29.8)	0.898 <sup>a</sup> (163.3)
$\ln(M_t)$	-0.029 <sup>a</sup> (-4.11)	0.019 (0.87)	
$\ln(R_t)$	-0.152 <sup>a</sup> (-20.1)	-0.139 <sup>a</sup> (5.47)	-0.131 <sup>a</sup> (6.36)
$R^2$	0.99	---	---
H(1,2)	---	1.596 (0.207)	1.323 (0.250)
H(1,3)	---	6.575 <sup>b</sup> (0.037)	5.134 <sup>c</sup> (0.077)
H(1,4)	---	13.52 <sup>a</sup> (0.004)	6.193 (0.103)
H(1,5)	---	13.97 <sup>a</sup> (0.007)	6.269 (0.180)

\* The corresponding ADF and PP test statistics for LSE are -5.074, -4.582, and -3.556 for ADF(1), ADF(3), and ADF(5); -3.907, -4.145, and -4.108 for PP(1), PP(3), and PP(5), which all reject the null hypothesis of no cointegration at the 10 percent level; **a**, **b**, and **c** denote the estimated coefficients are significant at the 1, 5, and 10 percent levels. The t-statistics are reported in parentheses in the cases of the estimated coefficients, and p-values are in parentheses in the cases of H (p, q) test results.

In the case of the LSE, the estimated coefficient of  $\ln P_t$  is 0.911 and highly significant. This result implies that a 1 percent increase in the nominal aggregate price level has been associated with a 0.911 percent increase in nominal food prices, which describes the *disproportionate* increases of the nominal food prices (or downward trend of real food and agricultural prices) compared to other goods during the sample period. In fact, Kliesen and Poole (2000) extensively discuss the reasons why real agricultural price has had a downward stochastic trend. The suggested factors are relatively lower income elasticity (Engel's law) and inelastic demand and supply functions of agricultural products. Income growth rates in the United States have been generally positive due to the country's technological progress and positive growth rates of productivity. Engel's law implies that the consumption of food and agricultural products are increasing less proportionately than income. The inelastic demand for food consumption combined with low income elasticity causes a disproportionate increase in relative nominal food and agricultural prices.

With respect to the estimated coefficients of  $\ln M_t$  and  $\ln R_t$ , if estimated coefficients are statistically insignificant, both the money supply and the real exchange rate movements will have a similar effect on food and aggregate price levels, which is interpreted as long-run neutrality of the variables. Hence, the estimated coefficient of  $\ln R_t$  (-0.152) indicates that a 1 percent real appreciation (depreciation) of the U.S. dollar has been associated with a 0.152 percent decrease (increase) in nominal food prices compared to the aggregate price level. It is clear that real exchange rate movements are not neutral in terms of explaining long-term overshooting of food and agricultural prices. On the other hand, the estimated coefficient of money supply is also statistically significant. The estimated coefficient is -0.029, which means a 1 percent increase in money supply has been related to a 0.029 percent decrease in nominal food prices compared to the aggregate price level. This result is not consistent with a general consensus of money neutrality. Considering the fact that annual growth rates of the U.S. money supply have been positive during most of the sample period, this finding implies that the downward trend of real agricultural price is connected with the positive growth of money even when we separate out the long-run relationship between agricultural and aggregate prices. In fact, Saghaian *et al.* (2002) reached a similar conclusion of long-run money non-neutrality. However, the estimated coefficient in our study is far less than their finding. The corresponding Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) tests all reject the null hypothesis of no cointegration at the 10 percent level.

Although the LSE is super-consistent, it is not asymptotically efficient and is biased in a small sample. To overcome the potential problem of the LSE and confirm the results, we re-estimate Equation (6') using the CCR methods, which is more efficient than LSE. The second column of Table 3 presents the estimated coefficients in Equation (6') with the corresponding H (p, q) test results. In general, the estimation results of the CCR method are similar to those of the LSE approach. The estimated coefficients of  $\ln P_t$  and  $\ln R_t$  are 0.874 and -0.139, respectively, which are statistically significant and similar to the LSE results. However, the estimated coefficient of  $\ln M_t$  differs between the LSE and the CCR estimation. The estimated coefficient of  $\ln M_t$  is *not* statistically significant in the CCR estimate, which confirms the long-run money neutrality hypothesis. Another important result is that the corresponding cointegration tests suggested by Park (1990, 1991) are mixed.<sup>7</sup> The null of stochastic cointegration is rejected for the H (1, 3), H (1, 4), and H (1, 5) test, while the null hypothesis is not rejected with the H (1, 2) test. This indicates that test results do not strongly support the cointegration relationship among variables. One way to interpret the mixed results is that  $\ln M_t$  might not be an important variable in the cointegration system, as the *t*-statistic (0.87) indicates. Because of this, the

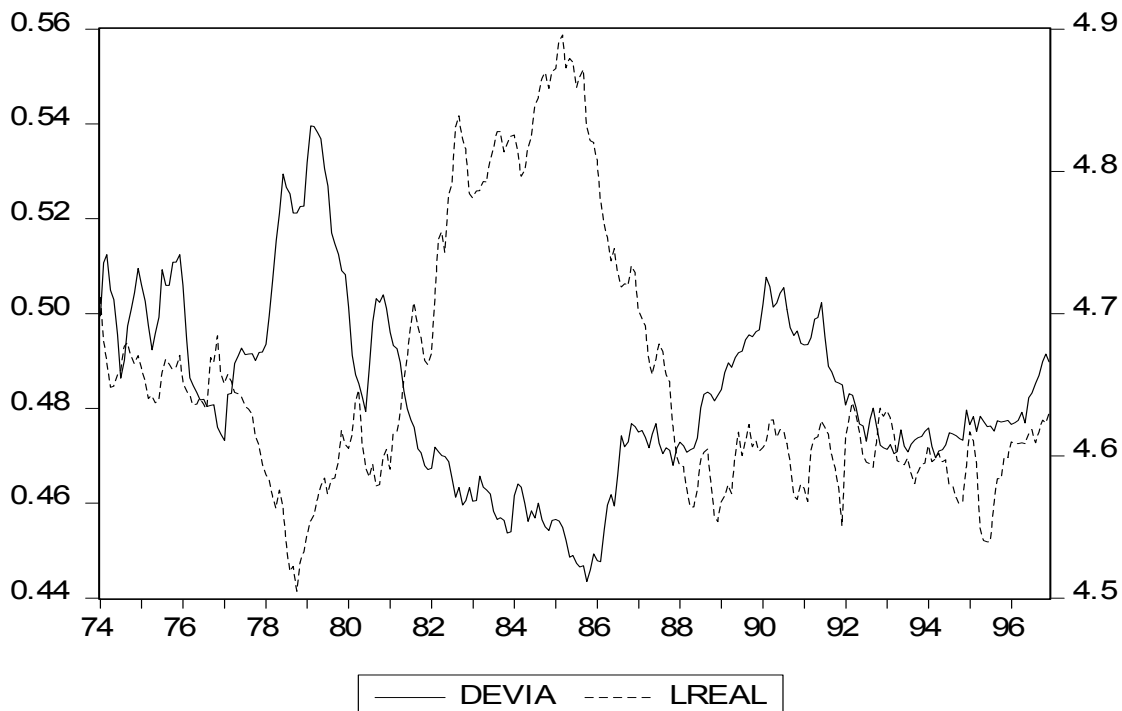
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<sup>7</sup> Park's H (p, q) test is computed by the CCR estimation and essentially applies the G (p, q) test to CCR residuals. Under the null of the cointegration, H (p, q) tests have asymptotically chi-square distributions with q-p degrees of freedom. Meanwhile, under the alternative of no cointegration, the test statistic diverges to infinity. In our model, each variable is treated as the first difference stationary with drift. Because of the drift, some variables 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 (see Park and Ogaki (1998) for more detailed explanation about the concept of deterministic and stochastic cointegration).

cointegration test results show that the estimated cointegration vector is not proper due to the presence of an irrelevant variable.

To examine this possibility, we eliminate  $\ln M_t$  from the empirical model and then re-estimate the cointegration vector. The last column of Table 3 presents the results. The test results are different. The null hypothesis of stochastic cointegration is *not* rejected in the cases of the H (1, 2), H (1, 4), and H (1, 5) tests, while the null hypothesis is rejected in the case of the H (1, 3) test, which is more favorable evidence of cointegration. In other words, the test results suggest that the long-run relationship among the three remaining variables becomes stronger when the money supply is eliminated from the model. The estimated coefficient of real exchange rates is -0.131 and statistically significant, meaning that the real exchange rate has an effect on the food and aggregate price level in the long-run. A 1 percent real appreciation (depreciation) of the U.S. dollar is related to a 0.131 percent decrease (increase) of food prices compared to the aggregate price level in the long-run.

Figure 3 confirms the finding of exchange rate impacts. It shows both the deviation of the food price (DEVIA) from its long-run equilibrium path, calculated from the estimated cointegration vector, and the log of real exchange rate (LREAL) movements during the sample period. A high value of the deviation means that food prices are higher than the long-run equilibrium level, while a low value of deviation means food prices are lower. This deviation is negatively correlated with real exchange rates, especially during the period between 1974 and 1989.



**Figure 3. Deviation of the U.S. Food Prices from Their Long-Run Equilibrium Path and Real Exchange Rate Movements**

As a whole, we find a strong relationship between the real exchange rate and the cyclical deviation of the food and agricultural price from its long-run equilibrium path. However, we do not find strong evidence that the money supply is related to this deviation in the long-run. In fact, these results are quite different from the recent empirical finding of Saghayan *et al.* (2002). The long-run relationship expressed in Equation (5) is considered in our paper, which explains natural divergence of agricultural prices from aggregate prices. By considering the long-run relationship between prices in the analysis, we can achieve the general consensus of long-run money neutrality.

### CONCLUDING REMARKS

Chronic instability is the most serious problem in the U.S. agricultural sector, and one of the important factors contributing to this instability is the long-term variation of agricultural price. Although agricultural economists (e.g., Gardner 1981; Tweeten 1989) recognized the potential instability problem of the U.S. agricultural sector after experiencing unexpected U.S. dollar movements during the post-Bretton Woods era, the potential long-run linkage between exchange rates and relative food and agricultural prices has been ignored due to the strong influence of the monetary economic view of the flexible exchange rate system<sup>8</sup>.

This study tests the long-run neutrality of the domestic money supply and exchange rates on the long-term movements of relative agricultural prices in the United States over the period 1974-1996. In this study, we derive a simple and new empirical model to test the long-run neutrality of the money supply and the U.S. dollar. A non-parametric cointegration method (CCR) developed by Park (1992) is implemented for estimation, which is more efficient than the method by Johansen (1988) for a small sample. The main findings in this study confirm a general consensus concerning long-run neutrality of national money (money supply) and give practical evidence of the real impact of international money (exchange rate) on the long-term variation of relative agricultural price in the United States.

The U.S. trade-weighted real exchange rate appreciated by 25 percent between 1995 and 2000. Moreover, the U.S. dollar appreciated by 42 percent relative to the currencies of its trade competitors during the same period. The study results indicate that this U.S. dollar appreciation caused 3.275 percent additional reduction of the food and agricultural prices in the United States.

Contrary to the argument of Kliesen and Poole (2000), the results of the present study also suggest that the stable U.S. monetary policy alone is not sufficient to prevent the possible instability problem in the U.S. farm economy in the future. Since one of the main sources of exchange rate misalignment is some degree of inefficiency in the foreign exchange market, an internationally coordinated monetary policy could be desirable.

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<sup>8</sup> In fact, recent controversial proposals relating to international monetary reform show how seriously these problems are considered by the economics profession (e.g., Williamson 1985, 1989; Krugman 1989; Mundell 1992; McKinnon 1996).

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