

Macro Effects on Agricultural Prices in Different Time Horizons

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Abstract:

Using monthly data covering 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). By controlling factors that determine the long-run trend of relative agricultural prices, the results show that long-term changes in real exchange rates have had a significant negative correlation with the long-term changes in relative agricultural prices. Conversely, changes in the general price significantly affect short-term changes in the relative agricultural price.

Keywords: Relative agricultural price, exchange rates, inflation rates, unit root test, canonical cointegration regression, money neutrality

Relative Agricultural Price Changes in Different Time Horizons

Several studies have focused on the impact of monetary variables on relative agricultural prices. The most common assertion has been that a domestic monetary shock causes a temporal overshooting problem of agricultural prices, which dies out in the long run (Barnett, Bessler, and Thompson 1983; Devados and Meyers 1987; Orden and Fackler 1989; Roberson and Orden 1990; and Karungu, Reed, and Allen 1995). More recently, Saghaian, Reed, and Marchant (2002) reached the same conclusion concerning short-term overshooting of agricultural prices in response to a monetary shock, in the case of the United States.¹

However, these studies have ignored the existence of a long-run relationship between prices in their model specification. Without considering the long-run relationship between prices in their empirical model, researchers could reach misleading results. Tweeten (1989) noted that, unlike insurable short-term price variation, cyclical long-term relative agricultural price movements have a greater impact on the U.S. farm economy. He argued that relatively high agricultural price during the Carter administration (1976-1980) caused a boom in the U.S. agricultural sector, while relatively low agricultural prices during the period of the Regan administration (1981-1988) had a disastrous impact on the U.S. farm economy, particularly small farms. This feature of relative agricultural price variation has been characterized by “long swings” and has deviated from its historical trend for a long time period. In studying the impact of macroeconomic variables on relative agricultural prices it is important to examine the impact of macro variables on relative agricultural prices in different time horizons.

In addition, U.S. dollar movements have played an important role in medium and long-term variability in relative agricultural prices (e.g., Gardner 1981; Batten and Belongia 1986).

U.S. dollar movements can seriously affect agricultural prices relative to other prices for two main reasons: (1) the U.S. agricultural sector is heavily involved in international trade: more than 30 percent of domestic agricultural products are exported (Tweeten *op. cit.*); and (2) the demand for agricultural products is inelastic compared to that for manufacturing products (Kliesen and Poole 2000).²

The idea that U.S. dollar movements influence long-term variation in relative agricultural prices is not generally accepted under the assumption that foreign exchange markets are efficient, mainly because the macroeconomic shocks only cause a temporal overshooting problem of the nominal exchange rate at best (e.g., Dornbusch 1976). However, Rogoff (1996) found that the real exchange rate tends to move toward purchasing power parity (PPP) in the very long run with a slow speed of convergence (the half-life of convergence is 3 or 4 years). Thus, the important feature of the U.S. dollar movements under the floating exchange rate system is that they are characterized by “long swings.”³ Therefore, if exchange rates cause the variation of domestic relative agricultural prices via international trade, their impacts are expected to be relatively long-term rather than short-term.

The objective of this study is to analyze the impact of macroeconomic variables on relative agricultural price changes in different time horizons. This study also discusses a common misspecification problem found in the current literature and develops an empirical model to remedy this problem.

The remainder of this paper is organized as follows. Section 2 discusses the misspecification problem in previous studies and presents our parsimonious empirical models to examine this issue. Econometric procedures and the main empirical findings are presented in the next section, and a summary follows in Section 4.

Empirical Model Derivation

Several studies (e.g., Vining and Elwertowski 1976; Parks 1978; Fisher 1981; Lach and Tsiddon 1992; Bomberger and Makinen 1993; Debelle and Lamont 1997) have examined the effect of nominal monetary shocks (or inflation rate) on different commodity prices within a macroeconomic context. The empirical question is whether 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,

$$RPD_t = \alpha + \beta \cdot \Delta \ln p_t + \gamma \cdot \Delta \ln z_t + \eta_t, \quad (1)$$

where RPD_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 change in other relevant variables. Inter-market price change dispersion is usually measured by a variation (or standard deviation) of changes in relative prices compared to general inflation rates such as⁴:

$$RPD_t = \frac{1}{N} \sum_1^N (\pi_{it} - \pi_t)^2, \quad (2)$$

where $\pi_{it} (= \ln p_{it} - \ln p_{it-1})$ is the rate of change of the i th 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 (1) with the measure (2) is appropriate to examine the effect of the general inflation rate on the relative price dispersion problem at the macroeconomic level, the model is clearly not appropriate to examine the issue of the relative price change of a

specific commodity group compared to other commodity groups.

As a first step in developing an empirical model which can examine relative changes of a specific commodity group, assume that there are two goods in an economy: agricultural and non-agricultural products. Following the previous studies in this area (e.g., Grennes and Lapp 1986; Robertson and Orden 1990; Saghalian, Reed, and Marchant 2002), the linear long-run relationship between general price and price levels of each commodity group is specified as

$$\ln p_t^a = \alpha_0 + \beta_0 \cdot \ln p_t + \eta_t \quad (3)$$

$$\ln p_t^{na} = \alpha_1 + \beta_1 \cdot \ln p_t + \mu_t, \quad (4)$$

where p_t is the general price level; p_t^a is the nominal 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 stochastic components including macroeconomic shocks and idiosyncratic shocks for each commodity group.

Subtracting (4) from (3) yields

$$\ln p_t^a - \ln p_t^{na} = (\alpha_0 - \alpha_1) + (\beta_0 - \beta_1) \cdot \ln p_t + (\eta_t - \mu_t). \quad (5)$$

Typically, equations (3) and (4), or equation (5), have been used to test long-run neutrality of money and inflation rate on the relative agricultural prices. If $\beta_0 \neq \beta_1$, then studies typically conclude money (or inflation rate) is not neutral in the long run. For instance, Saghalian, Reed, and Marchant (2002) found that $\beta_0 < \beta_1$ and interpret it as contradictory evidence of long-run money neutrality.

To isolate the relationship between prices that is expected to be determined by an unidentified market structure in *different time horizons*, we rewrite equations (3) and (4) as

$$\ln p_t^a - \beta_0 \cdot \ln p_t = \alpha_0 + \eta_t \quad (6)$$

$$\ln p_t^{na} - \beta_1 \cdot \ln p_t = \alpha_1 + \mu_t. \quad (7)$$

In this case, the coefficients β_0 and β_1 simply represent stochastic trends of prices for each commodity group compared to the general price in the long run. We cannot identify various factors that play an important role in determining these stochastic trends with the given dataset. Therefore, we adapt the idea of rational expectation equilibrium. If market information is perfect, the relative price movements between commodities contain all the information of structural variables (Radner 1979). Even if the market information is not perfect, the market eventually finds a proper price in the long run (Auman, 1969), Therefore, if underlying structural parameters are stable, relative prices between commodities have a long-run relationship.

Subtracting equation (7) from equation (6), we have

$$(\ln p_t^a - \beta_0 \cdot \ln p_t) - (\ln p_t^{na} - \beta_1 \cdot \ln p_t) = (\alpha_0 - \alpha_1) + (\eta_t - \mu_t). \quad (8)$$

If we decompose the stochastic terms 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

$$(\ln p_t^a - \beta_0 \cdot \ln p_t) - (\ln p_t^{na} - \beta_1 \cdot \ln p_t) = \kappa + \gamma \cdot \ln p_t + \delta \cdot \ln r_t + \zeta_t, \quad (9)$$

where $\kappa = \alpha_0 - \alpha_1 + \theta_0 - \theta_1$; $\gamma = \lambda_0 - \lambda_1$; $\delta = \pi_0 - \pi_1$; and $\zeta_t = \omega_t - \varepsilon_t$.

With equation (9), we can address the question of which macroeconomic factors cause *deviation* in the agricultural price in the long-run equilibrium path relative to the non-agricultural price, rather than the question of which macroeconomic factors determine the variation in relative agricultural price in the long run.

If agricultural prices are more (less) sensitive than non-agricultural prices to the changes in general price level, the estimated coefficient γ is expected to be positive (negative). If the U.S. dollar movement is important, the estimated coefficient δ is expected to be significant.

Furthermore, by differencing equation (10) with lag length k , our final empirical model is

$$\Delta_k(\ln p_t^a - \beta_0 \cdot \ln p_t) - \Delta_k(\ln p_t^{na} - \beta_1 \cdot \ln p_t) = \kappa + \gamma \cdot \Delta_k \ln p_t + \delta \cdot \Delta_k \ln r_t + \zeta_t. \quad (10)$$

Different choices of lag lengths are important in examining the main hypothesis. If the inflation rate causes only short-term effects for the changes in relative agricultural price, then the significance of the estimated coefficients should die out when 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. The deviation of real exchange rate from PPP 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. This can better explain relative price changes than the inflation rate in a longer time period.

In practice, we use a two-step estimation procedure. In the first stage, we estimate the cointegration vector, which explains the long-run relationship between the general price level and the price level of each commodity group. In the second step, equation (10) is estimated by replacing the estimated long-run coefficients obtained in the first step.

Economic Procedures and Empirical Findings

Seasonally adjusted monthly consumer price indices for food items and all items are used as proxy variables for the agricultural price and general price level. Consumer price indices of

all commodities less food items and service 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 goods⁵. These data were collected from the Bureau of Labor Statistics (BLS) web site (www.bls.gov). The 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. The sample consists of 348 observations from 1974:1 to 2002:12.

As a first step of the analysis, the long-run cointegration vectors among the variables are estimated using 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).

Unit-Root Tests

Preliminary graphical investigation suggested that all the price indices have obvious time trends. Thus, under the alternative of trend stationarity, the Said- Dickey (SD) (1984) and Phillips-Perron (PP) (1988) tests were applied. Because these tests are sensitive to the choice of order of autoregression, we report test results based on different orders: 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 the long-run relationship among variables.⁶

Canonical Cointegration Regression

To obtain cointegration vectors among the variables, we applied Park's Canonical Cointegration Regressions (CCR). Park's CCR method may have some advantages compared to Johansen's (1988) Maximum Likelihood (ML) approach. The CCR method does not require a normality assumption or assumption about the lag specification. Park and Ogaki (1991) show that, in Monte Carlo simulations, the CCR method consistently outperforms the ML approach in small samples. Asymptotically, the CCR and ML approaches give the same results, if the number of lags in vector autoregression (VAR) representation is the same in both approaches. We also applied the Park's $H(p, q)$ test for cointegration relationships. Park's $H(p, q)$ test is computed with the CCR residuals. Under the null of cointegration, $H(p, q)$ tests have asymptotically χ^2 distributions with $q-p$ degrees of freedom. Therefore, unlike conventional tests (e.g., Augmented Dickey Fuller test), we can conclude there is a cointegration vector when the test statistics fail to reject the null hypothesis. In our model, each variable is treated as 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

The estimated cointegration vectors and $H(1, q)$ test results are presented in Table 2.⁸ In the case of the food price, the estimated coefficient is 0.8001, which indicates less proportionate 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 larger than one (1.0701 and 1.1391),

indicating these increases are more proportionate compared to the general price level. These results might be due to the different income elasticities and productivity growth rates of each commodity group as indicated in Klisen and Poole (2000).

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.

Relative Price Changes in Different Time Horizons

If macroeconomic variables are important to explain changes in relative prices, the variables should cause deviations in prices from their long-run equilibrium paths. 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, price series are first constructed as deviations of their long-run equilibrium paths using the estimated cointegrating vectors for each price variable, and then Equation (10) is estimated. We present the results showing the changes of relative agricultural price in six different time horizons (1, 12, 24, 36, 48, and 60 months). Because preliminary test results suggest that there are autoregressive conditional heteroskedasticity (ARCH) type errors in the case of $k=1$, we used the GARCH (1, 1) procedure suggested by Bollerslev (1986). Since the serial correlation is a more serious problem in the cases of $k>1$, the heteroskedasticity and autocorrelation consistent (HAC) covariance matrix procedure suggested by Newey and West (1987) is used to estimate equation (10).

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 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. None of the estimated coefficients except for the one-month changes are statistically significant at the ten percent level, indicating that changes in general prices for more than one month do not have any explanatory power in relative agricultural prices. These results are consistent with the consensus of the long-run money neutrality but contradict Saghaian, Reed, and Marchant (2002).

In the case of the real exchange rates, however, the explanatory power increases when the time horizon is lengthened. The signs 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 magnitudes of the coefficients become smaller than those at $k=48$. The adjusted R^2 also increases from 0.023 at $k=12$ to 0.357 at $k=48$. The negative signs imply that real appreciation of the U.S. dollar causes a decrease in agricultural prices more than other manufacturing commodity prices.

Although the heavy involvement of the U.S. agricultural sector in international trade is an important explanation for these results, there is an alternative rationale. According to studies focusing on the price decision behavior of traders (e.g., Krugman 1987; Knetter 1989, 1993; Feenstra, Gagnon, and Knetter 1996), in the case of many manufacturing goods such as automobiles and electronics, international arbitrage is difficult due to differing national standards and warranty service. In this case, producers can exercise price discrimination across the different international markets (Rogoff 1996). Unlike homogenous goods, therefore, prices of these products are sticky in a destination market currency, resulting in relatively small

exchange rate impact on the domestic price. Considering the fact that agricultural products are usually homogeneous, these studies predict that more severe domestic supply shocks can occur in the agricultural sector, compared to manufacturing sectors, in response to an exchange rate shock, resulting in much more relative price variation of agricultural products compared to manufacturing products.

Food vs. Service Items

Table 4 presents the relative food price movement compared to that of service items, which represent the prices of non-tradable goods. In the case of the one-month changes, the inflation rate has statistically significant explanatory power for relative agricultural price. 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 in the relative agricultural price. In any case, except for the one-month changes, the estimated coefficients are not statistically significant. Again, these results support the long-run money neutrality hypothesis. Monetary shock can cause only a short-term overshooting problem on agricultural prices compared to prices of manufacturing and non-tradable goods. In the case of the real exchange rate, it does not have explanatory power in the relatively short-term changes in relative agricultural prices. However, it has statistically significant explanatory power for time horizons longer than 24 months. The significance of the estimated coefficient increases when the time horizon is increased. These results are also consistent with our prior expectation. The supply schedules of non-tradable goods are usually not affected by the flow of international trade. Therefore, even if a real fluctuation in the U.S. dollar is large and persistent, the domestic prices of non-tradable goods should be more stable than agricultural prices, resulting in more variation of relative agricultural prices than prices of non-tradable goods.

Conclusion

Several studies have examined the potential effect of macroeconomic shocks on changes in relative agricultural prices. However, these studies are limited to only short-term changes of relative agricultural prices, and ignore the issue of long-term variation of relative agricultural price.

Using monthly data 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 six different time horizons (1, 12, 24, 36, 48, and 60 months). By controlling factors that determine the long-run trend of relative agricultural prices, the results show that long-term change in the real exchange rate has a significant negative correlation with long-term change in relative agricultural prices. By contrast, change in the general price explains only short-term changes in relative agricultural price.

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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
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 tests under the alternative of trend stationarity. The critical values come from MacKinnon (1991).

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

	<i>Constant</i>	<i>Trend</i>	$\ln p_t$	<i>H(1,3)</i>	<i>H(1,4)</i>	<i>H(1,5)</i>
Food	0.8597 ^a (0.029)	0.0005 ^a (0.0001)	0.8001 ^a (0.008)	4.2357 (0.120)	4.6041 (0.203)	5.4162 (0.247)
Commodity less Food	-0.1992 (0.239)	-0.0012 ^a (0.0002)	1.0701 ^a (0.059)	3.1922 (0.203)	3.8331 (0.280)	5.4030 (0.248)
Service	-0.6319 ^a (0.027)	0.00007 (0.00005)	1.1391 ^a (0.008)	5.3557 ^c (0.069)	5.3751 (0.146)	7.8510 ^c (0.097)

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

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 (2.095)	0.0080 (1.489)	0.0090 (1.374)	0.0069 (1.023)	0.0075 (0.972)	0.0115 (1.244)
$\Delta \ln p_t$	-0.2915 ^a (-3.752)	-0.1632 (-1.588)	-0.0930 (-1.497)	-0.0479 (-1.267)	-0.0367 (-1.102)	-0.0491 (-1.464)
$\Delta \ln r_t$	0.0206 (1.260)	-0.0878 ^b (-2.045)	-0.1071 ^a (-3.663)	-0.1159 ^a (-5.897)	-0.1142 ^a (-6.847)	-0.1005 ^a (-5.341)
<i>DW-statistics</i>	1.408	0.130	0.093	0.095	0.082	0.059
<i>Adj-R²</i>	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 significance at the 1 and 5 percent levels.

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)
<i>DW-statistics</i>	1.331	0.124	0.068	0.046	0.044	0.036
<i>Adj-R²</i>	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 significance at the 1 and 5 percent levels.

Footnotes

1. Other empirical studies include Grennes and Labb (1986) and Zanas (1999).
2. Recently, Kim and Koo (2002) found 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.
3. An important empirical study by Mark (1995) found evidence that monetary fundamentals such as money stock and real income are predictable components only in long-horizon (16 quarter horizon) changes in exchange rate. It implies that nominal exchange rate deviates from its long-run equilibrium level, which implies the long-swinging pattern of real exchange rate.
4. This definition is used in Parks (1978) and Fisher (1981).
5. Initially, we considered the prices of more detailed commodity groups. However, we could not determine the existence of a long-run relationship. It means that commodity specific idiosyncratic shocks dominate the movement of relative prices in detailed commodities.
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.
7. $H(p, q)$ tests are based on the following regression.

$$\hat{e}_t = \sum_{\tau=0}^p \mu_\tau t^\tau + \sum_{\tau=p+1}^q \mu_\tau t^\tau + \eta_t$$

where p is the order of removed deterministic terms of time polynomial and q is the maximum order of included time polynomials for the test. A more detailed discussion about the concepts of deterministic and stochastic cointegration is presented in Park and Ogaki (1998).

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