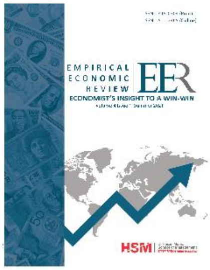

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Article: **Impact of Money Supply and Exchange Rate on Agricultural Prices in Pakistan**

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Impact of Money Supply and Exchange Rate on Agricultural Prices in Pakistan

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Abstract

This study analyzed the impact of the long-run neutrality of money supply and exchange rate on Pakistan's agricultural prices using data from 1975 to 2019. Engle and Granger and Johansen and Jusileius techniques were used to analyze the data. The results showed that the exchange rate's neutrality does not hold in the long-run. Simultaneously, the money supply coefficient was found to be insignificant in the long-run, emphasizing money's neutrality. The study concluded that monetary authorities can control the exchange rate by designing and implementing appropriate policies to overcome the overshoot problem of agricultural prices in Pakistan.

Keywords: agricultural prices, error correction, exchange rate, money supply, Pakistan

JEL codes: E40; E42; E43; E59; Q11

Introduction

Agriculture, since independence, has been a major productive sector of Pakistan’s economy. Even though efforts have been made for decades to reform and shift the economy towards high value industrial and service-centric production, agriculture still significantly impacts the country’s economy. The agriculture sector's significance is paramount because it contributes a sizeable 20% of the Gross Domestic Product (GDP) of Pakistan and employs 43.7% of the country’s total labour force. A significant chunk of this labour force, about 90%, hails from the low-income and fixed-income households of the rural areas of the country, whereas 62% of the entire population depends on this sector for their livelihood (Government of Pakistan, 2019).

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Standard theory dictates that prices are the lubricant that keeps the economic wheel moving. For this study, we focused exclusively on the agriculture sector and were primarily interested in agricultural prices and their dynamics. To put the problem in perspective, we argued that a significant portion of the population attached to this sector is from middle to low-income households (GoP, 2019). Any change (instability) in prices, whether direct or indirect, can significantly impact their living standard. Therefore, it is imperative to determine the major macroeconomic factors that influence agricultural prices and develop measures to reasonably predict these prices in the future.

A fair amount of literature investigates the instability of agricultural prices due to the volatility of the exchange rate and changes in the monetary policy. Pakistan also trades agricultural products internationally, so it can be assumed that domestic agricultural prices are affected by any changes in the exchange rate. This is especially true for an economy operating under the floating ER system because, at times, the nominal exchange rate can overshoot and cause severe distress on prices. Subsequently, the domestic purchasing power of the households is adversely affected.

The overshooting model argues that monetary policy changes have short-run effects on agricultural prices. Moreover, money in the short-run is non-neutral because it can change relative prices. In the absence of government intervention, the prices of agricultural commodities remain flexible and are determined in competitive circumstances. Simultaneously, the prices of manufacturing goods are mostly sticky (Barnet et al., 1983; Krugman, 1986; Betts & Devereux, 1996; Blinder, 2007; Holzer & Bittmann, 2020). Since monetary policy affects the agriculture sector in both the short- and long-run, its study becomes essential from an analytical perspective because farmers' income is susceptible to market price changes. Even if money supply is neutral in the long-run; still, in the short run it has a tremendous impact on farmers' income. Any change in the prices of agricultural commodities is a matter of concern for both the public and the policymakers because price fluctuations affect the agriculture sector's productivity. Before 2007, agricultural prices were comparatively low. However, after 2007, there was a hike in crop prices. Several internal and external factors were responsible for this price hike. In Pakistan, agriculture policy mainly focuses on increasing the farmers' income and providing cheap food items for urban consumers, as well as raw materials for the industrial sector at low prices. (Saghian et al., 2002; Siftain et al., 2016) suggested that an expansionary monetary policy can boost agricultural prices, increase income, uplift the farmers' living standard, and increase their investment capacity.

A relatively tight monetary policy can be drafted to keep agricultural prices in check to support the urban population and decrease domestic inflation. The mechanism of support price is used for controlling the prices of primary commodities. Thus, the agriculture pricing policy plays a pivotal role in boosting crop production and farmers' income. It is also essential to understand the supply response price mechanism (Nerlove & Bachman, 1960).

Ejaz (2007, 2009) conducted several studies about the agriculture sector's monetary impacts in Pakistan. In these studies, exchange rate was not incorporated. As Pakistan is a small open economy, it would be better to include the exchange rate. Siftain et al. (2016) incorporated the monetary variables with exchange rate in their study to investigate the impacts of monetary policy on food prices in the both long- and short-run using the (Saghaian et al., 2002) model.

However, Siftain et al. (2016) did not focus on the impact of the long-run neutrality of the exchange rate on the movement of relative agricultural prices in Pakistan and showed the long-run relationship only. In this study, we strived to determine the impact of monetary policy and exchange rate on the relative prices of the agriculture sector in Pakistan by determining the impacts of their long-run neutrality on the movement of comparable agricultural prices.

There is also an additional long-run relationship of agricultural prices or food prices with overall prices. Friedman (1975) noted that an expansionary monetary policy affects the economy's overall price strucrture, while the demand and supply of commodities determine its relative prices. This shows that agricultural prices move differently in the long-run as compared to the general price level, even if money supply does not change.

The factors that influence agricultural prices are essential to study for a developing country, such as Pakistan. Historically, relative prices have been determined mainly by the actual demand and supply factors. Nominal money factors have had a lesser role in determining relative prices, affecting only the general price level. Money supply and demand only determine the general price level and minimally affect the relative prices. Schuh (1974) suggested for the first time that exchange rate significantly affects agricultural prices. Later, researchers attempted to determine the implication of several other nominal variables, such as money supply and discount rate, along with agricultural prices. Studies show ambiguity about the relationship between agricultural prices and monetary variables. Lapp (1990) showed that money supply does not significantly affect food prices. However, later studies such as (Saghain et al., 2002; Asfaha & Jooste, 2007; Ejaz et al., 2007; Bekkers et al., 2017; Islam et al., 2017; Taghizadeh et al., 2019; De & Kakar, 2021) found that monetary policy has a significant impact and substantial implications for the agriculture sector. Several macroeconomic variables along with money significantly impact agricultural and food prices. Policies and changes in relative prices impact the farmers' investment decisions, farm productivity, and income. There is a need to understand which factors affect agricultural prices because it is essential to sustain productivity in this sector and in the whole economy.

The current study aimed to achieve the following objectives:

i. To ascertain whether there is a short-run impact of money supply and exchange rate on agricultural prices.
ii. To determine whether, in the long-run, money supply and exchange rate remain neutral in determining the relative agricultural prices.
iii. To find out the relationship of overall prices with agricultural prices.

Data and Methodology

This section discusses the theoretical foundation of the proposed empirical model. It also includes the econometric specification of the model as well as the relevant information regarding data sources and variables.

Theoretical Framework

Schuh (1974) is the seminal work on the issues faced by the agriculture sector and its relationship with monetary and other macroeconomic variables. This is vital because monetary policy directly affects agricultural prices and influences the living standard. Our main problem is to check whether agricultural and non-agricultural prices respond to monetary changes in the long-run or not. Furthermore, we want to check the hypothesis of money neutrality in the short-run.

Observational data suggested that agricultural prices are more competitive than any other sector, so these prices are less sticky. Consequently, an expansionary monetary policy favours the agriculture sector, while a contractionary monetary policy has an adverse effect (Devadoss & Meyers, 1986; Bakucs & Ferto, 2005; Frankel, 2008; Begum, 2021). Many studies conducted in this regard showed that agricultural prices adjust faster to changes in the monetary policy in the short-run than prices in the non-agriculture sectors, although money neutrality does not hold in the long-run (Saghaian et al., 2002; Asfaha & Jooste, 2007).

Dornbusch's (1976) model explained the link between exchange rate, money supply, and commodity prices. According to the model presented by Saghaian et al. (2002), which is an extended version of Dornbusch's model and incorporates international trade, a short-run deviation from the nominal exchange rate may be possible when prices are sticky. So, this overshooting may cause a short-run variation in the real exchange rate.

Agricultural prices and exchange rate are assumed to be flexible as they have their own unqiue adjustment paths and adjust quickly to the shocks in the monetary policy. In contrast, prices in the non-agriculture sector are assumed to be sticky. The study asserted that as a result of monetary shocks, agricultural and services sector prices move away from their long-run equilibrium. The study concluded that when monetary shocks occur, the burden of adjustment in the sector where prices are sticky is shared by the sector where prices remain flexible. An economy with a floating exchange rate system is less prone to agricultural price hikes due to monetary shocks.

Model Specification

The goal was to test for money neutrality in the long-run. For this purpose, we followed in the footsteps of (Lapp & Grennes, 1990; Robertson & Orden, 1990; Zanias, 1998; Saghaian et al., 2002). We set up the equations for the nominal prices of food and agriculture, money stock, real exchange rate, and aggregate price level as follows:

$$\ln P_t^A = \alpha_0 + \alpha_1 lnM_t + \alpha_2 lnR_t + \varepsilon_t \quad (2.1)$$
$$lnP_t = \beta_0 + \beta_1 lnM_t + \beta_2 lnR_t + v_t \quad (2.2)$$

where

$\ln P_t^A$ denotes the log of agricultural food / product prices
lnM_t denotes the log of money supply
lnR_t denotes the log of real exchange rate
lnP_t denotes the log of manufacturing products prices

If a one percent increase in the money supply generates the same percentage increase in the general price level and agricultural prices, this would indicate the long-run neutrality of money. In previous studies, $\alpha_1 = \beta_1$ was taken as a condition to test this hypothesis. However, suppose a percentage increase in money supply translates into a higher average price level, as Friedman (1975) argued. In this case, it becomes imperative that the relative prices of commodities, in the long-run, are determined by the changes in the existing supply-demand conditions. Therefore, complying with Friedman's argument, it is possible that agricultural prices do not always move in conjunction with the general prices, regardless of how the stock of money changes. Conversely, if the stock of money changes, where agricultural and general prices are moving disproportionately, its impact on both would be different. As per our hypothesis, the impact of money supply on agriculture would be different as compared to the overall prices. In this case, α_1 should be smaller than β_1. This empirical model is not suitable to test money neutrality.

Another significant relationship prevails among food and agricultural prices and prices in general. There is a relative movement of factors to explain the long-run relationships, as Kliesen and Poole (2000) noted the demand and income elasticity. However, it is impossible to include all such structural variables in the analysis. We incorporated the relationship between agricultural and food prices and prices in general, using the rational expectation approach. This approach suggests that the relative movements of demand and supply over time are realized in the variation of relative prices in the long-run.

It was assumed that real factors determine the long-run relationship between agricultural prices and general prices.

$$lnP_t^A = \gamma_0 + \gamma_1 lnP_t + \eta_t \tag{2.3}$$

Multiply equation (2.2) by -γ_1 and add equation (2.1) and (2.2) for the following long-run relationship:

$$lnP_t^A - \gamma_1 lnP_t = \alpha_0 - \gamma_1\beta_0 + (\alpha_1 - \gamma_1\beta_1)lnM_t + (\alpha_2 - \gamma_1\beta_2)lnR_t + (\varepsilon_t - \gamma_1 v_t) \tag{2.4}$$

Or, equivalently

$$lnP_t^A = \delta_0 + \gamma_1 lnP_t + \delta_1 lnM_t + \delta_2 lnR_t + \xi_t \tag{2.5}$$

If agricultural prices react more than overall prices in reaction to any change in the money supply, then $\delta_1 > 0$ and $\alpha_1 > \gamma_1\beta_1$; $\delta_1 < 0$ and $\alpha_1 < \gamma_1\beta_1, otherwise$. If agricultural prices respond more sensitively in response to the variations in the real exchange rate $\delta_2 < 0$ and $\alpha_2 < \gamma_1\beta_2$; $\delta_2 > 0$ and $\alpha_2 > \gamma_1\beta_2$, otherwise. If we take money and exchange rate to be neutral, that is, δ_1 and δ_2 equal to zero, then $\alpha_1 = \gamma_1\beta_1$ and $\alpha_2 = \gamma_1\beta_2$.

Model (2.5) contains three possible cases. Formally, they are expressed below:

Case 1: If the long-run relationship signified in equations (2.1), (2.2), and (2.5) holds and if exchange rate and money remain neutral in the long-run, then the values of the estimated coefficients δ_1 and δ_2 in the model (2.5) should be zero. Under given innovations, ξ_t should be a stationary process, which implies that the coefficient γ_1 should be a cointegration vector.

Case 2: If an authentic long-run relationship holds in (2.1), (2.2), and (2.3), but without holding long-run neutrality either in money or in the real exchange rate, the values of the coefficients δ_1 and δ_2 should be zero. However, the coefficients in the model (2.5) represent a long-run cointegration vector under the assumption of given innovations.

Case 3: If residual ξ_t is a non-stationary process, it means that the model's (2.5) coefficients do not comprise a cointegration vector. Hence, there should be no authentic long-run relationships in the estimated equations, that is, equations (2.1), (2.2), and (2.3). Alternatively, it might be that we are unable to identify the long-run relationship with the variables because there could be unobservable factors causing cyclical variations in the long-run equilibrium path of food and agriculture prices.

Note that the δ_2 coefficient indicates the extent to which food and agricultural prices are sensitive and respond to movements in actual exchange rates and aggregate prices. Even when δ_2 is zero, it does not imply the exclusion of any actual effect of the real exchange rate, either on food and agricultural exports or on domestic food and agricultural prices. Instead, real exchange rate variability affects food and agricultural prices and the aggregate price level in the long-run.

We followed a proper econometric procedure to calculate our estimates. We started with a fundamental model OLS which gave us insight into econometric problems, such as endogeneity and autocorrelation, as expected. Then, we checked whether the data was stationary or not. Augmented Dicky Fuller (ADF) test was used for unit root analysis. Secondly, we used cointegration based on unit root analysis for estimating long-run relationships. We used Least Square Estimator (LSE) Engle and Granger (1987) to check the long-run relationships. LSE is known for its consistency in estimating such relationships. We also used another technique known as Johansen and Juselius' (JJ) (1990) cointegration technique for comparison.

Data and Variables

The variables used in this study are money supply, real exchange rate, agricultural prices, and general prices. Consumer Price Index (CPI) of food was used as proxy for agricultural prices and an index of commodities as proxy for general prices. Money stock (M1) data was used for the money supply variable and the Real Effective Exchange Rate (REER) was measured as nominal effective exchange rate divided by a price deflator. Annual time series data was used for the period 1975 to 2019. Data on money supply was gathered from the data source of the State Bank of Pakistan (SBP). International Financial Statistics (IFS) database was used to collect the exchange rate data. The index of food prices and overall price index was collected from the Pakistan Bureau of Statistics (PBS) data source.

Estimation Method

For our analysis of the relationships among variables in the long-run, we used the most successful technique, that is, the Engle and Granger Two-Step Estimation Method, as put forth by (Engle & Granger, 1987). Error Correction Model (ECM) was used to check the short-run dynamics of the variables. However, this method is not asymptotically efficient because of the non-existent dynamic short-run adjustments. It is only consistent under a few regular conditions for estimating long-run cointegrating vectors. JJ technique also allowed us to test the hypothesis on cointegrating relationships, which Engle and Granger does not (Brooks, 2008). The latter cannot find the cointegrating vectors if there is more than one cointegrating vector, so the JJ method was applied to find out more than one cointegrating vectors.

We start the discussion with an explanation of the Least Square Method (LSM). Let Z_t be a $n \times 1$ vector of a variable that is both random and stationary at the first difference (ΔZ_t denotes stationarity). Under the condition where there is a non-zero vector of real number a, such that $a'Z_t$ is stationary, then it is associated with a cointegrating vector a. Assuming that the first element of a is zero, then seperating Z_t by $Z_t = (y_t\ , X_t')$ and normalize a by $a = (1, -c)$. Here, y_t is a difference stationary process, X_t is a vector difference stationary process, and $\boldsymbol{c}$ is a normalized associating vector.

The cointegration system (2.5) can be written as follows:

$$y_t = X_t'c + \varepsilon_t \tag{2.6}$$

$$\Delta X_t' = v_t \tag{2.7}$$

Here, $y_t = lnP_t^A, X_t' = [1,\ lnP_t, lnM_t, lnR_t]\ and\ c' = [\delta_0, \gamma_1, \delta_1, \delta_2]$. Then, $y_t\ and\ X_t$ are stationary at first difference. While, $\varepsilon_t\ and\ v_t$ are stationary and their mean is zero.

Now

$$Wt\ =\ (\varepsilon t, V\mathrm{t})' \tag{2.8}$$

Let $\Phi(i) = E(w_t w_{t-i}'), \sum = \Phi(0),\ \Gamma = \sum_{i=0}^{\infty} \Phi(i)$, and $\Omega = \sum_{0=-\infty}^{\infty} \Phi(i)$. In detail, Ω is the long-run variance matrix of w_t. Further, Ω is explained in matrix form as follows:

$$\Omega = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{1221} & \Omega_{22} \end{bmatrix} \tag{2.9}$$

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

$$\Omega_{11.2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21} \tag{2.10}$$

$$and\ \Gamma_2 = (\Gamma'_{12}, \Gamma'_{22})'$$

LSE was used to correct the short-run movements and error terms in the model. An example of this correction technique is the Maximum Likelihood (ML) estimation presented by (Johnson, 1988). As we are more interested in the long-run association of variables in the model rather than short-run estimates, therefore, Johansen and Juselius (1990) cointegration technique was also used in this study. The Johansen and Juselious (1990) method follows the ML method and finds the cointegrating equation in a non-stationary time series called Vector Autoregressive (VAR). With restrictions imposed, it is known as Vector Error Correction Model (VECM). For more brief understanding, take into account the following equations:

$$y_t^* = y_t + \Pi'_y w_t \tag{2.11}$$

$$X_t^* = X_t + \Pi'_x w_t \tag{2.12}$$

As w_t is stationary, y_t^* and X_t^* are cointegrated in the same order. When y_t^* is regressed on X_t^*, the matrices for the purpose are as follows:

$$\Pi_y = \Sigma^{-1}\Gamma_2 c + (0, \Omega_{12}\Omega_{22}^{-1})' \tag{2.13}$$

$$\Pi_x = \Sigma^{-1}\Gamma_2 \tag{2.14}$$

Practically, through these equations, long-run covariance parameters can be estimated and Π_y and Π_x are transformed into y_t and X_t.

Results and Discussion

This section presents the estimation results and a detailed discussion of these results. Section 3.1 provides unit root test results and also incorporates the long-run analysis discussed in the subsequent sections.

Unit Root Test

Augmented Dicky Fuller (ADF) and Dicky Fuller (DF) tests are widely used to check unit roots in the data set. DF captures only the AR (1) process, whereas ADF test captures the higher order process as well. ADF is an improved version of DF and three different forms of DF tests were used to amend the ADF test. Null hypothesis $\delta = 0$ was used in ADF test against the alternative hypothesis $\delta < 0$. Alternatively, alternative hypotheses was accepted, that is, $\delta < 0$, whereas the null hypothesis was rejected, that is, $\delta = 0$. So, it was determined that the series remained stationary and unit root did not occur.

We also applied the ADF test to check the stationarity of the series. The results of all unit root tests are presented below in tables.

Table 1

Unit Root Test of Variables at Level

Variable	1% critical	5% critical	10% critical	t-Statistic	Prob.*
ln Food Prices	-3.606	-2.934	-2.607	-0.082	0.945
ln General Prices	-3.606	-2.937	-2.607	-0.210	0.929
ln M1	-3.600	-2.935	-2.606	0.343	0.978
ln Real Effective ER	-3.601	-2.935	-2.606	-1.946	0.309

Table 1 shows the result of the unit root at level. Based on the probability value of all variables, we cannot reject the null hypothesis. Henec, it was concluded that all variables have a unit root at level. Therefore, we rechecked the unit root after taking the first difference and the results are reported in Table 2.

Table 2

Unit Root Test of Variables at 1st Difference

Variable	1% critical	5% critical	10% critical	t-Statistic	Prob.*
ln Food Prices	-4.212	-3.531	-3.196	-4.281	0.0084
ln General Prices	-3.610	-2.934	-2.608	-3.669	0.0080

Variable	1% critical	5% critical	10% critical	t-Statistic	Prob.*
ln M1	-3.610	-2.934	-2.608	-5.624	0.0000
ln Real Effective ER	-3.610	-2.934	-2.608	-4.741	0.0004

Graphical representation of the series suggested that the log of food prices has a time trend. So, we applied the ADF test, accordingly. Table 2 reports the results of the ADF test. These results reject the null hypothesis, that is, the food price series has a unit root. The probability value indicates that the series is stationary after taking the first difference.

In case of general prices and M1, the ADF test results reject both null hypotheses, that is, $\delta = 0$. It means that both series are stationary at first difference. The last unit root test was used to check the stationarity of the real effective exchange rate. The results reject the null hypothesis that indicates that the series is stationary at first difference. Thus, the ADF test results reported in Table 3 show that all series are stationary at first difference.

Results of Engle-Granger

The two-step Engle-Granger cointegration approach was used to analyze the long-run relationship among variables suggested by (Engle & Granger, 1987). The results of Engle and Granger are given below in Table 3.

The table shows that the value of the coefficient of general prices is 0.870, which is also statistically significant. A one percent increase in general prices raises food prices by 0.870%, which is close to but still lower than a one-to-one increase. However, the difference is significant enough to prove disproportionate movement in general prices and food prices. The reasons behind it were explained extensively by Kliesen and Poole (2000), regarding why food prices have a downward trend. The proposed reasons included a comparatively low income elasticity (Engel's Law) and inelastic demand and supply functions of food products. Engel's law points out that food and agricultural consumption increases less proportionately than income. Low income elasticity and an inelastic demand for food consumption are the reasons underlying the disproportionately increasing movement in food prices.

Table 3

Least Square Estimation Results

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.995	0.176	-5.657	0.000
Ln General Price	0.870	0.033	26.049	0.000
Ln M1	0.009	0.0136	0.674	0.504
Ln Real Effective Exchange Rate	0.103	0.038	2.694	0.010
R^2	0.993	Adjusted R^2	0.992	

The above table shows that the estimated value of the coefficient Ln M1 is 0.009. It is statistically insignificant, which explains the long-run money neutrality of the said variable. The money supply growth rate was positive during the sample period.

The real effective exchange rate coefficient is 0.103 and it indicates that one percent appreciation of the currency leads to 0.103% increase in food prices. The variable explains that the real effective exchange rate movements are not neutral in presenting the overshooting of food prices in the long-run.

Adjusted R^2 has a value of 0.992, which shows the goodness of fit. It indicates that regression explains 99% variation in explanatory variables.

Table 4

Augmented Dicky Fuller Test of Cointegration

Variable	1% critical	5% critical	10% critical	t-Statistic	Prob.*
Residuals	-3.616	-2.941	-2.609	-3.013	0.0426

The current study utilized the Augmented Dicky Fuller (ADF) test to check cointegration. The results reject the null hypothesis of no cointegration at a 5% level.

Error Correction Model (ECM)

The benefit of using ECM to find a long-run relationship is that it takes care of spurious regression. Table 5 offers sufficient evidence regarding the long-run relationship among the said variables. The probability value is 0.043, which implies the rejection of the null hypothesis. This condition became the basis to regress ECM, the results of which are reported in Table 5.

Table 5

Error Correction Model Results

Variable	Coefficient	Std. Error	t-Statistic	Prob.
DLGP	1.451	0.074	19.702	0.000
DM1	0.004	0.013	0.330	0.744
DRER	0.089	0.037	2.356	0.024
U(-1)	-0.365	0.084	-4.351	0.000
C	-0.030	0.007	-4.617	0.000
R^2	0.929	Durbin-Watson stat		1.825

According to Durbin-Watson and R-squared values, ECM regression was not spurious. The coefficient DLGP is counted among short-run coefficients and has a positive sign, indicating a positive relationship between general prices and food prices. This coefficient is also statistically significant at 1% level. Both DM1 and DLRER are also short-run coefficients and positively correlate with food prices. However, our model's money supply variable is statistically insignificant, whereas the real effective exchange rate significantly impacts the model. The coefficient U (-1) is the error correction coefficient. It is a long-run coefficient and has a negative sign as required. The U (-1) coefficient value is -0.36, which explains that the previous period's shock adjusts in the next period by 36%. This variable has a probability value of 0.0001, which confirms the significance and the long-run relationship.

Diagnostic Test

To diagnose the above regression, we applied the tests for autocorrelation and normality. The results of both tests are given below in Table 6.

Table 6

Serial Correlation LM Test

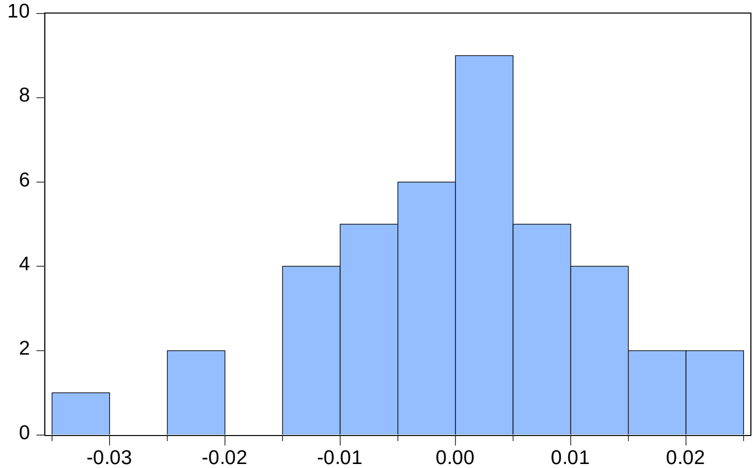
F-statistic	1.064	Prob. F(2,33)	0.357
Obs*R-squared	2.424	Prob. Chi-Square(2)	0.298

We applied the Breusch-Godfrey serial correlation LM test to detect autocorrelation in our model. The results are reported in Table 6 and they highlight the absence of autocorrelation. The probability value of Chi-Square is 0.298. Thus, the null hypothesis is rejected.

Further, we applied the Histogram Normality test to check the distribution of errors. The graph and statistics of the normality test are provided below in Figure 1.

Figure 1

Histogram Normality Test



The above graph shows a normal distribution of error terms. Further, we can also check the statistics provided in the above diagram. The probability value of Jarque-Bera is 0.466. So, the null hypothesis was not rejected and it was assumed that errors are normally distributed.

Johansen and Juselius Cointegration

There are some testing prerequisites to use the Johansen's cointegration technique, such as the time series data must be I(1). ADF test was utilized to check the stationarity of data. We found that all series are stationary at first difference. The results of ADF are given above in tables 1 and 2.

The JJ method was used to estimate the cointegrating equations, following the application of the ML method using a non-stationary time series called the Vector Autoregressive (VAR) model. With restrictions imposed, it is known as Vector Error Correction Model (VECM).

To obtain the optimal lag length for the JJ procedure, we preferred the Akaike Information Criteria (AIC) over the Schwarz Bayesian Information Criteria (SBIC) because of its efficiency (Brooks, 2008). The results of lag length criteria are shown below in Table 7.

Table 7

Lag Length Criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	38.472	NA	1.91e-06	-1.814	-1.642	-1.753
1	287.087	431.804*	9.29e-12*	-14.057*	-13.195*	-13.751*
2	299.352	18.721	1.17e-11	-13.861	-12.309	-13.309
3	316.763	22.909	1.17e-11	-13.935	-11.693	-13.138
4	329.057	13.588	1.66e-11	-13.740	-10.809	-12.697

Johansen (1990) proposed two tests to check cointegration including the maximum eigenvalue test and the trace test. The latter was used to test the alternative hypothesis of no cointegration and the null hypothesis of cointegration. The maximum eigenvalue test was used to check the hypothesis that whether the number of cointegrating vectors is $r + 1$ or equals to r (Brooks, 2008).

After checking the unit root, we applied the JJ cointegration method to study the long-run relationship among the variables. As per the trace test result, we rejected the null hypothesis of no cointegration because the probability value is less than 0.05. The next hypothesis stated one cointegrating equation and according to the probability value, we cannot reject the null hypothesis. Both trace and maximum eigenvalue tests gave the same results. We concluded that in VECM, one cointegrated vector (long-run equilibria) is added with one lag. The results are shown below in tables 8 and 9.

Table 8

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.692	74.038	47.856	0.000
At most 1	0.360	29.254	29.797	0.057
At most 2	0.221	12.266	15.495	0.144
At most 3	0.069	2.732	3.841	0.098

Table 9

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.692	44.784	27.584	0.000
At most 1	0.360	16.988	21.132	0.173
At most 2	0.221	9.534	14.264	0.244
At most 3	0.069	2.732	3.841	0.098

VECM Results

After detecting the cointegrating equation, we proceeded with VECM. Table 10 shows the results of VECM. It shows that the long-run speed of adjustment back to its equilibrium is denoted by c(1), which is recognized as the adjustment factor. The VECM coefficient is -0.464. It is also statistically significant, which implies that the system comes back to its equilibrium by 46% in the long-run.

Table 10

Results of VECM

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C(1)	-0.464	0.216	-2.149	0.038
C(2)	0.441	0.605	0.728	0.471
C(3)	1.924	0.860	2.237	0.031
C(4)	0.086	0.057	1.491	0.145
C(5)	-0.272	0.228	-1.192	0.241
C(6)	-0.146	0.041	-3.498	0.001
Adjusted R-squared	0.4793	F-statistic	2.636	
Durbin-Watson stat	1.973	Prob (F-statistic)	0.040	

The results were obtained after estimating VECM. We also applied different diagnostic tests to figure out how fit our model is. If the estimated model clears all the diagnostic tests, then we may conclude that the obtained results are efficient.

Wald Test

We conducted the Wald test to check the joint influence of the variables. The results showed that the variables jointly influenced the dependent variables. The null hypothesis of the Wald test stated that the selected variables were equal to zero. However, we rejected the null hypothesis because the probability value was calculated as 0.0126, which is less than 0.05. The following table shows the result of the Wald test.

Table 11

Wald Test

Test Statistic	Value	Df	Probability
F-statistic	3.187056	(4, 34)	0.0251
Chi-square	12.74823	4	0.0126

Serial Correlation LM Test

The Lagrange Multiplier test, commonly known as the LM test, captures the autocorrelation present in a model. Table 12 shows no autocorrelation in the model.

Table 12

Breusch-Godfrey Serial Correlation LM Test

F-statistic	0.024	Prob. F(2,32)	0.976
Obs*R-squared	0.060	Prob. Chi-Square(2)	0.970

Conclusion and Policy Implications

Conclusion

Instability is a severe issue in the agriculture sector of Paksitan's economy and the long-term volatility of prices is vital in causing this instability. Agriculture economists uncovered severe instability issues caused by unexpected variations in the price of US dollar. Simultaneously, in the Bretton Woods era, the long-run relationship between food, exchange rate, and agricultural prices was ignored due to the stringent influence of the flexible exchange rate system.

The current study tested the impact of the long-term neutrality of local money and exchange rate on the long-run variations in the relative agriculture prices in Pakistan. A new empirical model was derived to test the long-term neutrality of the supply of money and exchange rate. We used the JJ method and LSE to check our results.

We examined the relationship between food prices and other independent variables described above using the annual data for the period 1975-2019. We estimated the short-run coefficients and also identified the long-run equilibrium relation. We also found evidence that an increase in the general price level causes a rise in food prices and the results are consistent with the findings of (Frankel, 2008). Moreover, Begum (2021) also reported the same results in their study. Furthermore, we found in our study that money neutrality holds, as shown by (Choe & Koo, 1993; Cho et al., 2004; Holzer & Bittmann, 2020). The real effective exchange rate also causes an increase in food prices.

On the other side, it is argued by certain economists that a stable monetary policy is not enough to prevent the problem of instability in the future, as money supply plays a neutral role since it is insignificant in the long-run.

Policy Recommendations

Instability in any economic sector is a significant problem for any country. Money supply is not the main factor in increasing agricultural prices. Indeed, it plays a neutral role in the long-run because the coefficient of money supply was found to be insignificant in the long-run. However, it was found that the exchange rate overshoot altered the agricultural prices. Therefore, monetary authorities need to control the exchange rate through suitable policies to overcome the overshoot problem. Some other unobservable factors also exist which cause this problem, such as the demand and supply problems. These problems can be overcome through crop support prices. The government only stipulates wheat support prices to control the illegal export of wheat. However, support prices ensure the stability of prices in the agriculture sector.

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