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Price Escalation of Rice in Bangladesh: A Time Series Approach

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Abstract

To meet its overall objective of ensuring food security for all households, the Government of Bangladesh undertakes several activities: it intervenes in markets to stabilize prices, targets food distribution to poor households and provides emergency relief after natural disasters. This paper measures the fluctuation pattern of rice prices over time. Stabilization of rice prices is a serious concern of Bangladesh. In this paper, we attempt to explore two types of temporal variations in rice prices: inter-year (seasonality) and intra-year. The distinction between the two types is important in the context of policy instruments that are brought to bear on prices in order to contain extreme fluctuations. Therefore, issues and relations underlying price fluctuations need to be properly understood before formulating and implementing any policies. The developed model shows a good fit of the data obtained.

Gel Classification: C12, G32, E17

Introduction

Throughout the world, agriculture is subject to more or less intense government intervention. In less developed countries, agriculture tends to be taxed while in developed countries it is subsidized (e.g., Peterson, 1979; Bale and Lutz, 1981). " Agriculture is not just about putting things in the ground and then harvesting them....... it is increasingly about the social and environmental variables that will in large part determine the future capacity of agriculture to provide for eight or nine billion people in a manner that is sustainable". These were the words of Achim Steiner, Executive Director of the UN Environmental Programme. However, the crop sector is of strategic importance to Bangladesh, as in most other low-income countries. It is the source of staple food for 150 million people and the major means of livelihood of 13 million farm households in the country. Rice is the most important cereal crop, which occupies about 74 percent total cropped area in Bangladesh and 82.56 percent of the total irrigated land is under this crop (BBS, 1996). It is estimated that about 73 percent of the Bangladeshi people's average calorie intake and about 55 percent of protein intake comes from rice (BBS, 1995). Though, rice prices play an important role in the economy of Bangladesh and fluctuation in rice prices has a

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great economic impact on the people of Bangladesh, the dis-equilibrium of demand and supply in agriculture is a universal phenomenon. The nature of the supply, demand and output conditions in agriculture are such that it creates inherent instability. The agricultural prices are notoriously unstable (Chandra, 1985). Consequently, price instability leads to uncertainty in the income of the producers. This uncertainty retards investment in agriculture resulting in slow growth of agricultural output. Moreover natural calamity, instability of international fuel market, fertilizer and electricity crisis are also very important causes in the rise of the price of rice. The price of rice is increasing more rapidly than the income of the poor. Moreover, farmers do not earn huge profit by cultivating rice. This is because the production cost of the cultivation of rice is increasing rapidly. Food prices overall, according to the World Bank, have risen by 83 percent in the past three years (Madely, 2008). Since Bangladesh agriculture traditionally depends heavily upon the natural factors, production varies tremendously which in turn leads to significant fluctuations in prices of agricultural commodities. So the role of prices is crucial and quantitative impact of price changes has to be known. Therefore, the study is an attempt to analyze the nature of price movements of different varieties of Bangladeshi rice between the periods of 1986 to 2006.

Key Research Question

The price of rice is increasing day by day, while the rate of escalation may not be a significant rise in any section of that upward trend. However, the question is still remaining as to what is the actual trend of that price escalation of rice over the years. If that is known then the policy makers could make an appropriate food policy. It is known that in harvesting season the supply of rice is higher than at any other time, which is attributable to the seasonal variation in price of rice. Therefore, this study is an attempt to identify the impact of price hike and how to solve the problem.

Justification of the Study

Rice plays an important role in Bangladesh economy. Rice is not only the staple food of Bangladesh but also the driving force of Bangladesh agriculture. Rice contributes about 50% of the agricultural; GDP; about 78% of the total cultivated area is under rice production, while 55 to 60 percent of the total agricultural labour force is employed in rice production, processing, marketing and distribution.

This paper has discussed about the trend of price escalation of rice and its seasonal variation of price. If the impact of seasonal variation of the price of rice can be reduced, then it will be ensured that the poor farmers are benefitted. For keeping the price level within the tolerable limit, the factor of subsidiary sector of production has to be analyzed properly. If the year-to-year price changes and the impact on seasonal variation are reduced then government can ensure food for every household and the stability of price of rice will be ensured. The study will help in keeping the price of rice stable.

Review of Literature

In Bangladesh, several empirical studies on price behavior, vis-à-vis area, production and yield behavior of different agricultural crops have been undertaken since 1960. Some of the studies may not be entirely relevant to the present study, but their findings, methodology of analysis and suggestions have a great influence on the present study.

Barua and Alam (2000) have studied the key concern of this research which is to assess the growth, fluctuation and price flexibility of Aus, Aman, Boro, Wheat and Jute crops during the last two decades of the twentieth century. Real prices of all the crops have been falling significantly during the study period. The price instability was higher depending on area, production and yield fluctuation of all the crops. It was observed that supply of Aus, Wheat and Jute production played insignificant role to determine their own post-harvest prices, but Boro, Aman and Aus production had significant role on the prices of crops of Aus, Wheat and Jute through significant cross effects. On the other hand, Aman and Boro production had significant influence on post-harvest price determination of these rice varieties as revealed by price flexibility coefficients. So price policy measures cannot be taken on a single crop basis in Bangladesh.

Khalek (2005) examined the temporal trends and fluctuations in the output of some crops in Bangladesh over the period of 25 years. He showed, by cross-commodity comparison by means of output instability index, very high degree of temporal instability in the production of Mango, Pulses, Potato, Wheat and Jute in Bangladesh.

Farouk (1970) in his study analyzed the nature of spatial price relationships of rice in several markets in the country using the field data during the three crop seasons for the year 1967 - 68. He also studied the extent of temporal price variations in the different rice markets for the year 1967 - 68. He showed the extent of temporal price variations in the different rice markets for the same period and compared those with storage costs.

Shahabuddin (1983) estimated yield and price risk of different crops in some selected regions of Bangladesh. The variance - covariance matrices of random disturbances associated with both output and prices were estimated utilizing aggregate time series data in four districts. The residuals represented the estimates of random components from which the relevant variance - covariance matrices were subsequently computed. Ranking of crops in terms of the estimated variance of price disturbance were pulse, which occupied the top position (0.153) followed by oilseeds (0.122) jute (0.099), Aman rice (0.090), IRRI Boro (0.071) and Aus rice (0.070).

Hossain (1983) studied the trend in agricultural prices for the period 1969 - 70 to 1980 - 81. He observed that the price index first accelerated in the early seventies and within a period of three years from 1971 - 72 to 1974 - 75, the index moved from 121 (1969 - 70=100) to 370, increasing at a rate of 37 percent per annum. The trend rate of growth for the entire period was 6.3 percent per annum.

Sarker and Husain (1984) conducted a study to analyze the nature of price movements and the terms of trade between agriculture and non-agriculture sectors for the period 1908 - 81. They revealed that during the study period, agricultural prices were more fluctuating than non-agricultural prices and instability in all prices increased in Bangladesh since 1971, but again the degree was higher for agricultural prices than for non-agricultural prices.

Ahmed and Bernard (1989) computed correlation co-efficient between district prices of Aman and Aus rice during 1976 - 82. They found that in case of Aman price, 63 out of 190 pairs were statistically insignificant and 51 out of 63 insignificant correlations pertained to Barisal, Patuakhali, Dinajpur and Bogra.

Ravallian (1986) developed a market integration model, which can estimate the extent to which local prices are influenced by price in the reference market (e.g. Dhaka market for Bangladesh). He employed his model for rice prices in Bangladesh just prior to and during the 1974 famine. His test rejected the hypothesis that rice markets are segmented i.e. totally lacking of integration. But the test on short run integration of markets was inconclusive.

Sabur and Elahi (1992) examined trend, annual and seasonal rice price fluctuations in Bangladesh. The study revealed that annual rice price fluctuation in the pre-liberation period and in the 80’s was more stable than in the 70's.

Sabur and Haque (1993) carried out another study to investigate the trend, seasonal and cyclical variations of retail and wholesale prices of rice in Mymensingh town market and forecast the future prices based on ARIMA model. The compound growth rates of real prices indicate that consumers seem to be better off with respect to rice price since independence, particularly during 80's.

It is evident from the above discussion that since a long time price fluctuation in agricultural crops adversely affected growth in agricultural sector as well as in the whole economy. Reviewing the above studies the researchers felt the need for conducting an analysis of rice price escalation, which would be highly beneficial and derives some significant findings and suggestions, which might help adopt appropriate price policy formulation. Despite some limited improvements, there still exist a number of limitations**.** But the researchers believe that the results of this study are expected to provide useful information**,** which will help in further research and improvements of knowledge.

Objectives of the Study

The main purpose of the study is measuring the seasonal fluctuation of rice price and bringing about its solution. More specifically, the study has the following objectives:

- To interpret and evaluate the changes occurring in the price of rice.
- To examine the time series properties of rice price in Bangladesh.
- To isolate and measure the impact of seasonal pattern in the time series.
- To develop a time series model of Bangladeshi rice price.

Data Sources and Research Methodology

In order to achieve the specific objectives of the study secondary sources of data are used. The majority of the data is collected from Bangladesh Agriculture Marketing Department. We also, collect data from different issues of the Yearbook of Agricultural Statistics of Bangladesh; published by Bangladesh Bureau of Statistics (BBS,2006). Time series data have been used in the analysis. A time series may be defined as a collection of readings belonging to different time periods of some economic variables or composite of variables (Gupta, 1988).

For determining the long-term changes in time series data, the time series properties were checked at first. For measuring the seasonal fluctuation of rice price we used the moving average. The moving average is measuring seasonal variation and also the most widely used method in practical work. This method allows more flexibility when the time series is trend stationary. In this article, we will develop a model of determinants for the rice price policy decisions of Bangladesh governments.

Model and Preliminaries

We consider the following unit root regression model

$$y_t = \alpha y_{t-1} + u_t \; ; \qquad \ldots \qquad (1)$$

for t = 1, . . . , T. We are interested in testing the null of a unit root, $\alpha = 1$, against the stationarity alternative, $|\alpha| < 1$. The initial value y_0 does not affect our subsequent asymptotic analysis as long as they are stochastically bounded, and therefore we set it at zero for expositional brevity. The errors (u_t) in the model (1) are serially correlated and specified as an AR(p) process given by

$$\alpha(L)u_t = \varepsilon_t \qquad \ldots \qquad (2)$$

where L is the usual lag operator and $\alpha(z) = 1 - \sum_{k=1}^{p} \alpha_k z^k$.

As a first step, we carried out Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) tests (see Dickey and Fuller, 1979, and MacKinnon, 1991 for the critical values) on the log our variables. The general regression model is the following:

$$\Delta y_t = \gamma_0 + \gamma_1 y_{t-1} + \sum_{i=1}^{n} \phi_i \Delta y_{t-i} + \varepsilon_t$$

where y_t is the log of the study variables, the γ's and ϕ's are constant parameters and ε_t is a random disturbance term. Rejection of the null of non-stationarity requires γ_1 to be negative and significantly different from zero. A lag length of one was chosen for the countries under investigation on the basis of the Akaike Information and Schwarz Information Criteria (AIC, SIC).

ADF Unit Root Test

If the e_t has autocorrelation more than one period, the unit root test can be modified as

$$\Delta Y_t = \alpha + \beta T + \rho Y_{t-1} + \sum_{i=1}^{k} \lambda_j \Delta Y_{t-i} + e_t$$

The number of lagged difference terms to be included is often determined empirically. The null hypothesis is still same as the DF test (i.e., H_0: ρ=1; and H_0: δ=0). It is so called Augmented Dicky-Fuller (ADF) test.

The test procedure is firstly run the unrestricted model and obtains $RSS_{UR.}$

$$\Delta Y_t = \alpha + \beta T + \rho Y_{t-1} + \sum_{i=1}^{k} \lambda_j \Delta Y_{t-i} + e_t$$

Secondly, run the restricted model and obtains RSS_R

$$\Delta Y_t = \alpha + \beta T + \sum_{i=1}^{k} \lambda_j \Delta Y_{t-i} + e_t$$

Then compute the F-statistic as: $F^* = \dfrac{(RSS_R - RSS_{UR})/2}{RSS_{UR}/(n-k)}$

Compare F to the critical values Φ that are tabulated by Dicky-Fuller (1981).

The null hypothesis:

H_0: β=0, δ =0 (or ρ=1) (unit root, non-stationary)

The KPSS test

Denis Kwiatkowski, Peter C.B. Phillips, Peter Schmidt and Yongcheol Shin (1992) proposed a test of the null hypothesis that an observable series is stationary around a deterministic trend. The series is expressed as the sum of deterministic trend, random walk, and stationary error, and the test is the LM test of the hypothesis that the random walk has zero variance. KPSS type tests are intended to complement unit root tests, such as the Dickey–Fuller tests. By testing both the unit root hypothesis and the stationarity hypothesis, one can distinguish series that appear to be stationary, series that appear to have a unit root, and series for which the data (or the tests) are not sufficiently informative to be sure whether they are stationary or integrated.

Hadri (2000) proposes residual-based Lagrange Multiplier tests for the null hypothesis that all the time series are stationary (either around a level or a deterministic time trend), against the alternative that some of the series are nonstationary. The Hadri tests are panel versions of the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (1992) stationarity tests. Following Hadri (2000), consider the models:

$$y_{it} = r_{it} + \varepsilon_{it} \quad (3)$$

and

$$y_{it} = r_{it} + \beta_i t + \varepsilon_{it} \qquad (4)$$

where r_{it} is a random walk, $r_{it} = r_{i,t\text{-}1} + u_{it}$, and ε_{it} and u_{it} are mutually independent normal distributions. Also, ε_{it} and u_{it} are *i.i.d* across i and over t, with $E[\varepsilon_{it}] = 0$, $E[\varepsilon_{it}^2] = \sigma_{\varepsilon,i}^2 > 0$, $E[u_{it}] = 0$, $E[u_{it}^2] = \sigma_{u,i}^2 \geq 0$, $t = 1,...,T$ and $i = 1,...,N$. The null hypothesis that all the series are stationary is given by $H_0 : \sigma_{u,i}^2 = 0$, $i = 1,...,N$, while the alternative that some of the series are nonstationary is $H_1 : \sigma_{u,i}^2 > 0$, $i = 1,...,N_1$ and $\sigma_{u,i}^2 = 0$, $i = N_1 + 1,...,N$.

Let $\hat{\varepsilon}_{it}$ be the residuals from the regression of y_i on an intercept, for model **Error! Reference source not found.**(or on an intercept and a linear trend term, for model (4)). Then, the individual univariate KPSS stationarity test is given by:

$$\eta_{i,T,k}(\hat{\omega}_i) = \frac{\sum_{t=1}^{T} S_{it}^2}{T^2 \hat{\sigma}_{\varepsilon_i}^2},$$

where S_{it} denotes the partial sum process of the residuals given by $S_{it} = \sum_{j=1}^{t} \hat{\varepsilon}_{ij}$, and $\hat{\sigma}_{\varepsilon_i}^2$ is a consistent estimator of the long-run variance of $\hat{\varepsilon}_{it}$ from the appropriate regression. In their original paper, KPSS propose a nonparametric estimator of $\hat{\sigma}_{\varepsilon_i}^2$ based on a Bartlett window having a truncation lag parameter of $l_q = \text{integer}\left[q(T/100)^{1/4}\right]$, with $q = 4,12$. However, Caner and Kilian (2001) have pointed out that stationarity tests, like the KPSS tests, exhibit very low power after correcting for size distortions. Thus, in our paper we follow recent work by Sul, Phillips and Choi (2005), who propose a new boundary condition rule that improves the size and power properties of the KPSS stationarity tests. In particular, Sul et al. suggest the following procedure. First, an AR model for the residuals is estimated, that is:

$$\hat{\varepsilon}_{it} = \rho_{i,1}\hat{\varepsilon}_{i,t-1} + ... + \rho_{i,p_i}\hat{\varepsilon}_{i,t-p_i} + \upsilon_{it} \qquad (5)$$

where the lag length of the autoregression can be determined for example using the general-to-specific algorithm proposed by Campbell and Perron (1991). Second, the long-run variance estimate of $\hat{\sigma}_{\varepsilon_i}^2$ is obtained with the boundary condition rule:

$$\hat{\sigma}_{\varepsilon_i}^2 = \min\left\{ T\hat{\sigma}_{\upsilon_i}^2, \frac{\hat{\sigma}_{\upsilon_i}^2}{\left(1 - \hat{\rho}_i(1)\right)^2} \right\},$$

where $\hat{\rho}_i(1) = \hat{\rho}_{i,1}(1) + ... + \hat{\rho}_{i,p_i}(1)$ denotes the autoregressive polynomial evaluated at $L = 1$. In turn, $\hat{\sigma}_{\upsilon_i}^2$ is the long-run variance estimate of the residuals in equation (5) that is obtained using a quadratic spectral window Heteroscedastic and Autocorrelation Consistent (HAC) estimator.

The Hadri (2000) panel stationarity test statistic is given by the simple average of individual univariate KPSS stationarity tests:

$$\overline{LM}_{T,N} = \frac{1}{N}\sum_{i=1}^{N}\eta_{i,T},$$

which after a suitable standardisation, using appropriate moments, follows a standard normal limiting distribution. That is:

$$Z = \frac{\sqrt{N}\left(\overline{LM}_{T,N} - \overline{\xi}\right)}{\overline{\zeta}} \Rightarrow N(0,1)$$

where $\overline{\xi} = \frac{1}{N}\sum_{i=1}^{N}\xi_i$ and $\overline{\zeta}^2 = \frac{1}{N}\sum_{i=1}^{N}\zeta_i^2$.

The Monte Carlo experiments of Hadri (2000) illustrate that these tests have good size properties for T and N sufficiently large. However, Giulietti *et al.* (2008) show that even for relatively large T and N the Hadri (2000) tests suffer from severe size distortions in the presence of cross-sectional dependence, the magnitude of which increases as the strength of the cross-sectional dependence increases. This finding is in line with the results obtained by Pesaran (2007) on both the Im, Pesaran and Shin and the Maddala and Wu panel unit root tests. In order to correct for the size distortion caused by cross-sectional dependence, Giulietti et al. (2008) apply the bootstrap method and find that the bootstrap Hadri tests are approximately correctly sized.

To implement the bootstrap method in the context of the Hadri tests, we start off by correcting for serial correlation using equation (2) and obtain $\hat{\upsilon}_{it}$, which are centred around zero. Next, as suggested in Maddala and Wu (1999), the residuals $\hat{\upsilon}_{it}$ are resampled with replacement with the cross-section index fixed, so that their cross-correlation structure is preserved; the resulting bootstrap innovation $\hat{\upsilon}_{it}$ is denoted $\hat{\upsilon}_{it}^*$. Then, $\hat{\varepsilon}_{it}^*$ is generated recursively as:

$$\hat{\varepsilon}_{it}^* = \hat{\rho}_{i,1}\hat{\varepsilon}_{i,t-1}^* + \ldots + \hat{\rho}_{i,p_i}\hat{\varepsilon}_{i,t-p_i}^* + \upsilon_{it}^*,$$

where, in order to ensure that initialisation of $\hat{\varepsilon}_{it}^*$, i.e. the bootstrap samples of $\hat{\varepsilon}_{it}$, becomes unimportant, we follow Chang (2004) who advocates generating a large number of $\hat{\varepsilon}_{it}^*$, say $T+Q$ values and discard the first Q values of $\hat{\varepsilon}_{it}^*$ (for our purposes we choose $Q=30$). Lastly, the bootstrap samples of y_{it}^* are calculated by adding $\hat{\varepsilon}_{it}^*$ to the deterministic component of the corresponding model, and the Hadri LM statistic is calculated for each y_{it}^*. The results shown in Table 1 are based on 1,000 bootstrap replications used to derive the empirical distribution of the LM statistic.

The DF–GLS test

Conventional unit root tests are known to lose power dramatically against stationary alternatives with a low order MA process: a characterization that fits well to a number of macroeconomic time series. Consequently, these original tests have been largely supplanted in many researchers' toolkits by improved alternatives. Along the lines of the ADF test, a more powerful variant is the DFGLS test proposed by Elliott, Rothenberg and Stock (ERS, 1996), described in Baum (2000, 2001), and implemented in Stata as command performs the ERS efficient test for an autoregressive unit root.

This test is similar to an (augmented) Dickey-Fuller test, as performed by, but has the best overall performance in terms of small-sample size and power, dominating the ordinary Dickey-Fuller test. The DF–GLS test "has substantially improved power when an unknown mean or trend is present" (ERS, p.813). The DF–GLS test applies a generalized least squares (GLS) detrending (demeaning) step to the variable name:

$$y_t^d = y_t - \hat{\beta}' z_t$$

For detrending, $z_t = (1, t)'$ and $\hat{\beta}_0$, $\hat{\beta}_1$ are calculated by regressing $[y_1, (1-\bar{\alpha} L)y_2, \ldots, (1-\bar{\alpha} L)y_T]$ onto $[z_1, (1-\bar{\alpha} L)z_2, \ldots, (1-\bar{\alpha} L)z_T]$ where $\bar{\alpha} = 1 + \frac{\bar{c}}{T}$ with $\bar{c}$ = -13.5, and L is the lag operator. For demeaning, $z_t = (1)'$ and the same regression is run with $\bar{c}$ = -7.0. The values of $\bar{c}$ are chosen so that "the test achieves the power envelope against stationary alternatives (is asymptotically MPI (*most powerful invariant*) at 50 percent power" (Stock, 1994, p.2769; emphasis added). The augmented Dickey-Fuller regression is then computed using the y_t^d series:

$$\Delta y_t^d = \alpha + \gamma t + \rho\, y_{t-1}^d + \sum_{i=1}^{m} \delta_i \Delta y_{t-i}^d + \varepsilon_t$$

where m = *maxlag*. The *notrend* option suppresses the time trend in this regression.
Approximate critical values for the GLS detrended test are taken from ERS, Table 1 (p.825). Approximate critical values for the GLS demeaned test are identical to those applicable to the no–constant, no–trend Dickey–Fuller test, and are computed using the *dickey-fuller* code.

Findings of the Study

Time Series Properties of the Variable

A time series data heavily depends on time series properties. In checking the time series properties, we focus on the presence or absence of unit roots or stochastic trends in the variable used in this article. In order to form a statistically adequate model, the variable should first be checked as to whether they could be considered as stationary or non-stationary.

Unfortunately, it is well known that unit-root tests have low power and that results can vary with the types of test used and on the number of lags included in the test equations. For this reason, it becomes a strategy among the researchers to examine the results of several test procedures in order to draw conclusions regarding variable integration. With this in mind, three unit root tests procedures are performed: (i) most widely used Augmented Dicky-Fuller (ADF) test of Dicky and Fuller (1979, 1981) (ii) the asymptotically most powerful DF-GLS (Generalized Least Square) test of Elliott *et al*. (1996) and (iii) the Kwiatkowski *et al*. (1992) LM test (KPSS). The null hypothesis of ADF and DF-GLS tests is that a time series variable has a unit root while that of KPSS test is that a variable is stationary. A common strategy is to present results of both ADF/DF-GLS and KPSS tests, and show that the results are consistent (e.g., that the former reject the null while the later fail to do so and vice-versa). The lag length is selected by using the Akaike Information Criteria (AIC).

Before beginning the formal tests for unit roots, the variables should be plotted against time to visually determine if a trend exists in the time series. The necessity of this step is simply due to the fact that the critical values of the tests depend on the sample size and the inclusion of deterministic components, i.e., the inclusion of a constant and a time trend. Price of rice variable in level has been graphed against time in Figure 1(a) to 1(c) over the period 1986q3-2006q4. By referring figure, it is visually evident that price series presents upward trend. This upward trend indicates that the series of rice prices are stationary, but it is difficult to guess whether the trend is deterministic or stochastic.

Sample correlograms of rice price in level (d=D=0) for sample autocorrelations (SAC) and sample partial autocorrelations (SPAC) are depicted in Figures 2. The key feature of this SAC is that the spikes decay rapidly to zero in Figure 2(a). Most of the autocorrelation functions lie in the 95-percent confidence interval. This type of pattern is generally an indication that the time series is stationary. But, in case of Figure 2(b) and 2(c) some sample partial autocorrelations (SPAC) gives a little bit different scenario. That is, the feature of this sample correlogram is that a few autocorrelation functions lie outside the 95-percent confidence interval. Overall conclusion about our time series, type of pattern is generally an indication that the time series is stationary. Since rice price shows upward trend, both constant, and constant trend are used in the models to test for unit roots. Table 1 contains the results for three test procedures mentioned above. For the level series, both ADF and DF-GLS tests reject the null hypothesis of unit root, while KPSS test accept the null hypothesis of stationarity.

The tests carried out are the asymptotically most powerful DF-GLS test for the null of unit root of Elliott *et al.* (1996), the Kwiatkowski *et al.* (1992) LM test for the null of stationarity (KPSS) as well as the PP test of Philips and Perron (1988) for the null of unit root. A common strategy is to present results of both ADF/PP and KPSS tests, and show that the results are consistent (e.g., that the former reject the null while the latter fails to do so, or vice -versa). The lag length is selected by the Akaike Information Criteria (AIC). The results are shown in Table 1.

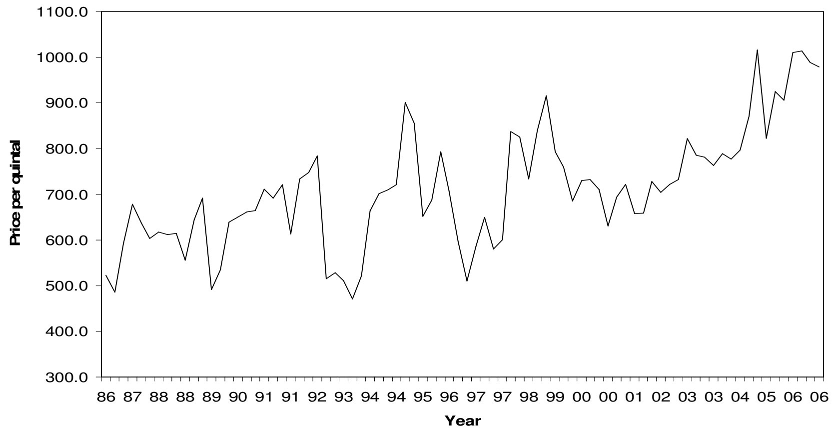


Figure 1(a). The pattern of Aman rice price in level over the period 1986q3-2006q4

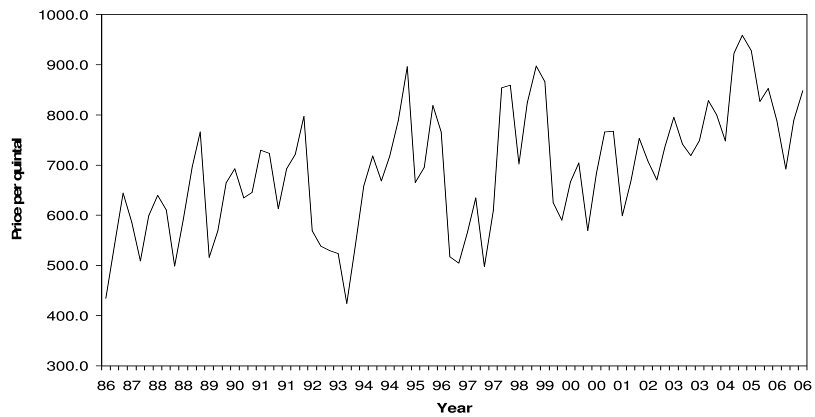


Figure 1(b). The pattern of Aus rice price in level over the period 1986q3-2006q4

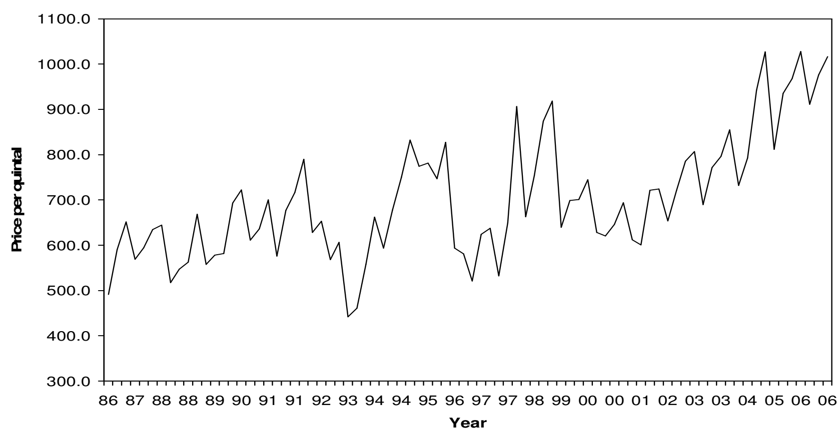
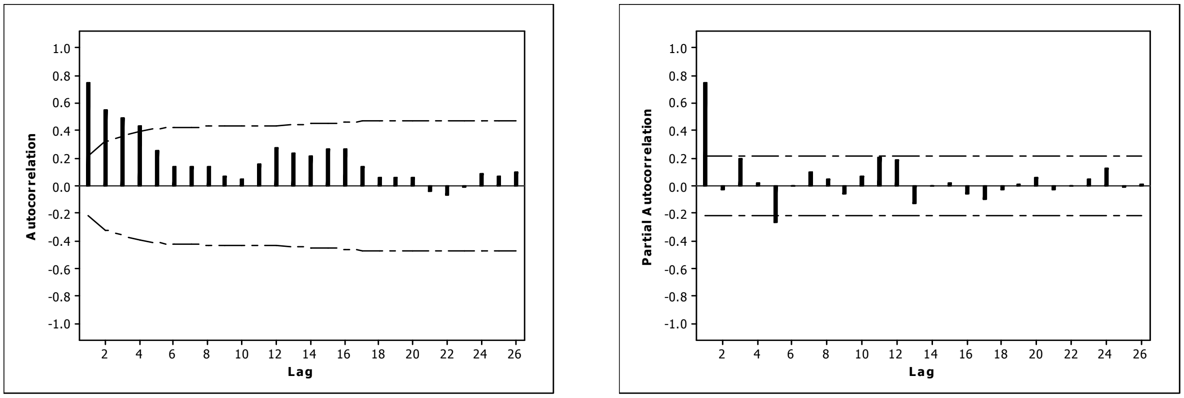
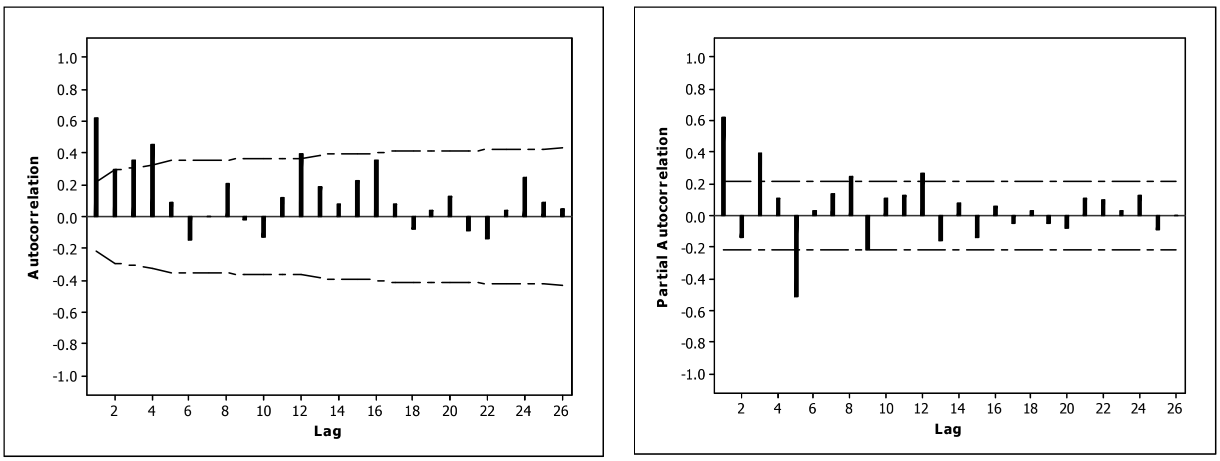


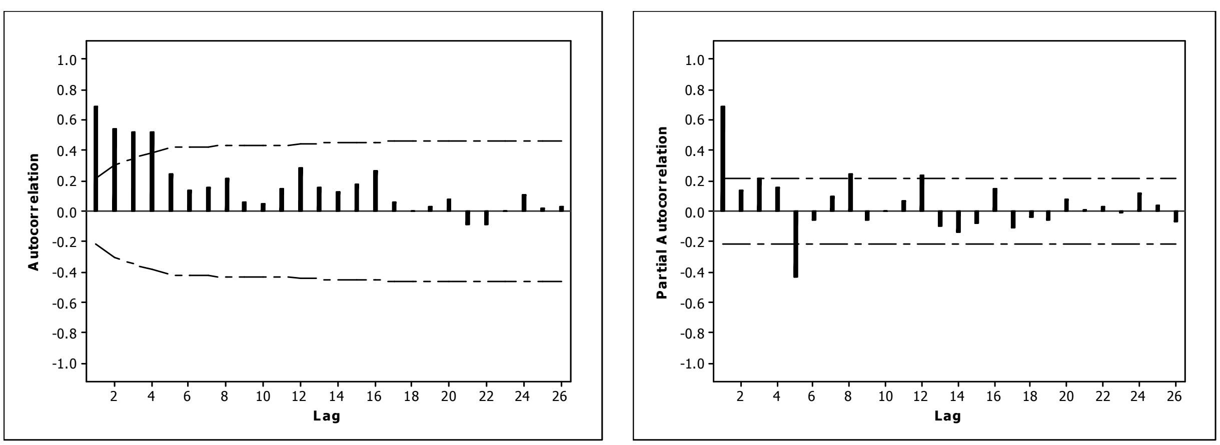
Figure 1(c). The pattern of Boro rice price in level over the period 1986q3-2006q4



(a) Aman price



(b) Aus price



(c) Boro price

Figure 2. SACF and SPACF for rice price in level over the period 1986q3-2006q4

Table 1 : Results of different unit root tests for levels data

Variables	ADF		DF-GLS		KPSS	
	Const.	Const. & Trend	Const.	Const. & Trend	Const.	Const. & Trend
Aman	-2.909^{*} (0.048) LL = 0	-4.307^{**} (0.005) LL = 0	-1.946^{*} (0.046) LL = 0	-4.307^{**} (0.000) LL = 0	0.884^{**} (0.002) BW= 6	0.093 (0.572) BW= 5
Aus	-3.886^{**} (0.003) LL = 4	-6.071^{***} (0.000) LL = 4	1.414 (0.173) LL = 11	-4.828^{**} (0.000) LL = 4	0.838^{**} (0.008) BW= 5	0.044 (0.342) BW= 4
Boro	-0.144 (0.940) LL = 7	-3.772^{**} (0.023) LL = 4	1.004 (0.267) LL =7	-3.740^{**} (0.003) LL = 4	0.843^{**} (0.005) BW = 6	0.119 (0.132) BW= 5

Note: *** and * indicates the significant at 1% and 5% level respectively. Figures in parenthesis (.) indicate p-value. Also LL = Lag length and BW = Bandwidth*

From the results of Table 1, revealed that all methods suggest that price of Aman, Aus and Boro time series are stationary. Since our study data series are trend stationary, so we can easily apply classical methods to analyze and making inference about our time series data set.

Trend Analysis of Rice Price in Bangladesh

Since Independence, the Government of Bangladesh has attempted to reduce variability of rice prices, and especially to prevent sharp increases in price. Here we analyze historical price trends and variability in Bangladesh. First, inter-year (annual) and intra-year (seasonal) prices in Bangladesh are analyzed. In order to separate out price trends from seasonal or random elements, price fluctuations are measured as deviations from the moving average of prices and from a linear trend. Price changes relative to the price in the preceding period are also discussed.

Annual price fluctuations in Bangladesh arise mostly from fluctuation in production, which again can be attributed to the random effect of floods and drought. Prior to 1994, public imports, and to a lesser extent drawdown of stocks, were the main policy instruments to achieve year-to-year stability in prices. As will be discussed below, since the trade liberalization of 1994, the private sector import (rice) trade has been the dominant factor in keeping price rises within acceptable limits in case of a domestic production shortfall. Year-to-year fluctuations in nominal prices of rice in Bangladesh were very high. Prices were especially unstable due to severe rice shortages caused by drought-related production shortfalls and shortage of foreign exchange for government rice imports. In Appendix Table 1 discussed about the trend value and year-to-year change of price of Aman, Aus and Boro respectively. In Aman year-to-year fluctuations greater than 10 percent occurred in 3 out of 7 years during the 1994-2000 as compared with 1 out of 6 years during the 2000-2006.

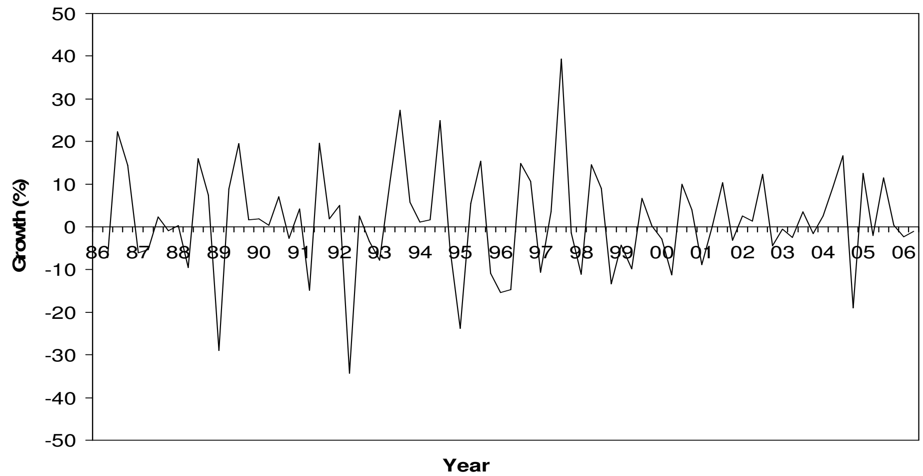


Figure 3(a). Percentage of growth Aman rice price over the period 1986q3-2006q4

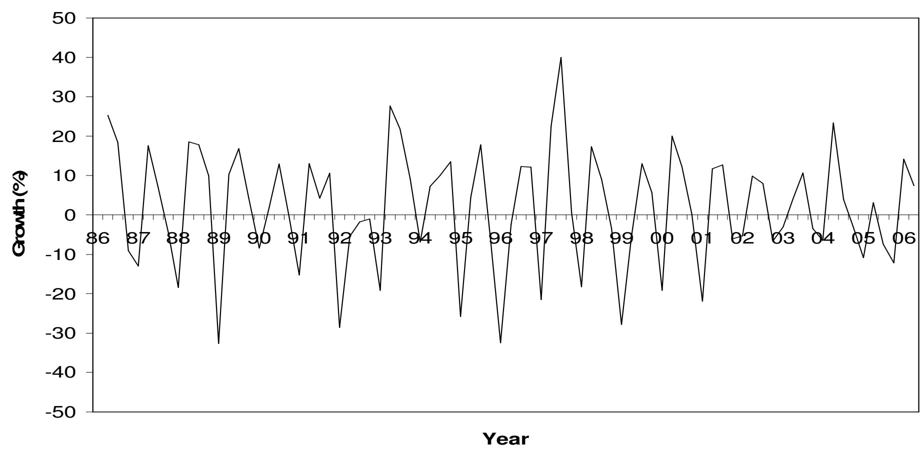


Figure 3(b). Percentage of growth Aus rice price over the period 1986q3-2006q4

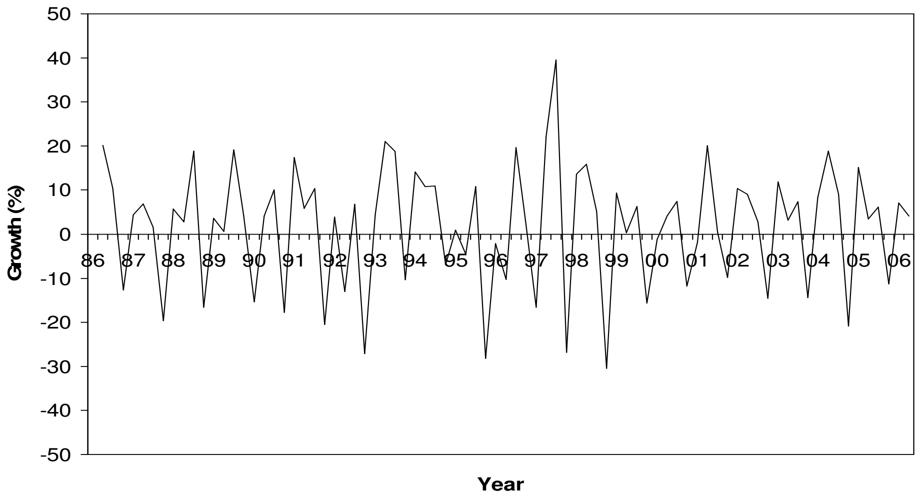


Figure 3(c). Percentage of growth Boro rice price over the period 1986q3-2006q4

In case of Aus price also the changes from preceding year in 2000-2006 is more stable than it was in 1994-2000. Year-to-year fluctuations greater than 10 percent occurred in out of 7 years during the 1994-2000 as compared with 2 out of 6 years during the 2000-2006. Similarly in Boro year-to-year fluctuations greater than 10 percent occurred in 4 out of 7 years during the 1994-2000 as compared with 2 out of 6 years during the 2000-2006. Comparing among Aus, Aman and Boro, the price of Boro was the most instable during 1994-2000. In 1994 the changes occurred in the price of rice with a very high rate because of flood in 1993.

Though the prices were relatively stable of all three kinds of rice in 2000-2006 but during that period the trend of prices were increasing slightly. But in the early of 2000 the prices of rice were more instable but in that period the prices were fluctuating occasionally. The percentage changes of price during 2000-2006 leads the overall change of price to the upward trend. Figure 3(a), 3(b) and 3(c) represent the trend of price escalation during the period 1986q3 to 2006q4 for Aman, Aus and Boro respectively.

In order to distinguish between trend and random elements of fluctuation in prices, trends are calculated using least square method (see, Appendix Table 1). From the above figure 3(a) to 3(b) we study year by year the change in percentage. Observing above figures we can primarily suspect our time series data contains seasonality and its fluctuations caused by seasonality. For concrete decision about seasonality we need further investigation.

Seasonal Fluctuation Analysis of Rice Price in Bangladesh

Seasonal price variations are generated by seasonality in production. The policy instruments that are used to keep seasonal price spreads within acceptable limits are domestic procurement, which attempts to raise average prices (and farmer incomes), and Open Market Sales (OMS) and other sales channels, designed to moderate prices to consumers when there are severe upward pressure on prices.

The seasonal component is defined as the intra-year pattern of variation that is repeated from year to year. To identify seasonality in our data set calculated 12-month (4-quarter) moving average and it is represent in Figure 4. This figure shows the seasonality pattern in our data set. It is clear that our time series data set contain seasonality pattern from this figure.

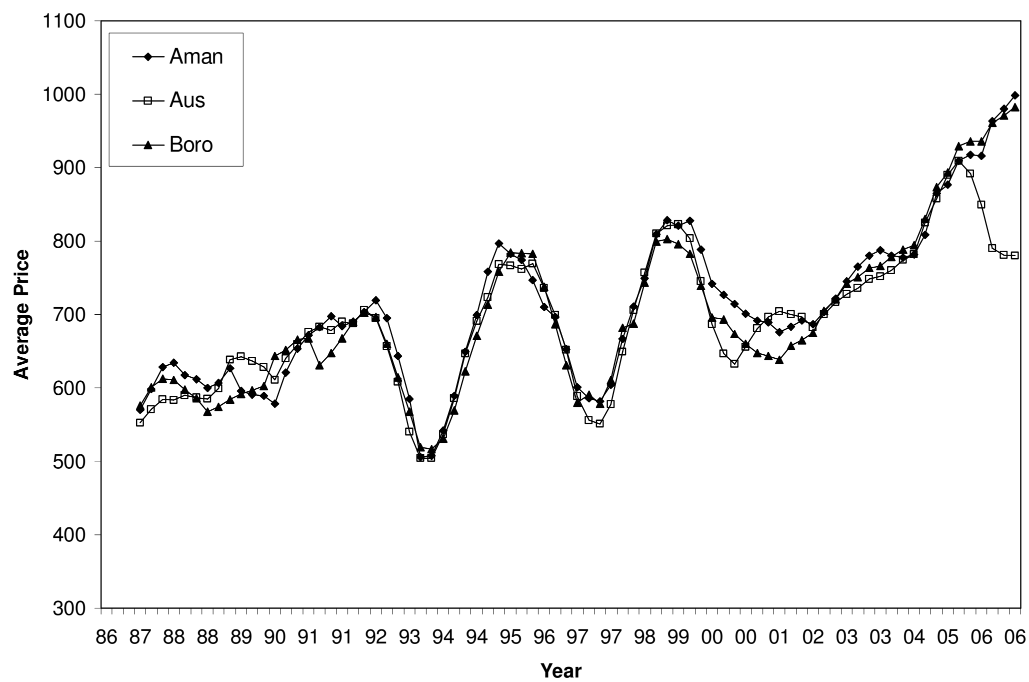


Figure 4. Seasonal pattern of rice price by 4-quarter point moving average over the period 1986q3-2006q4

For the seasonal fluctuation of price of rice we computed 4-quarter moving averages, since the seasonal pattern repeated year after year. In this process, the seasonal and irregular variations were ignored from the original price. On the basis of the 4-quarter moving averages above figure showed the de-seasonalised pattern of rice price in 1994, 1997 and 2004-2006. This de-seasonalised data give the idea about the other components of the time series such as trend, cyclic fluctuation, and irregular variation.

Alternatively, the seasonal level (d=0) of rice price plotted in Figure 2(a-c) moves through a constant value, which indicate that the rice price is stationary or integrated of order one, I(0). But the data in Figure 2(b) and 2(c) shows a strong seasonal variation. SACF and SPACF for the level of rice price series are shown in Figures 2(b) and 2(c) respectively. Significant seasonal pattern is obviously seen in SACF in Figure 2(b) for leveled series. These figures showed that seasonality series returns quickly to a constant overall mean, and the seasonal strong pattern is given. Now, we investigated the intra-seasonal pattern of rice price by taking monthly average of the period 1986 to 2006.

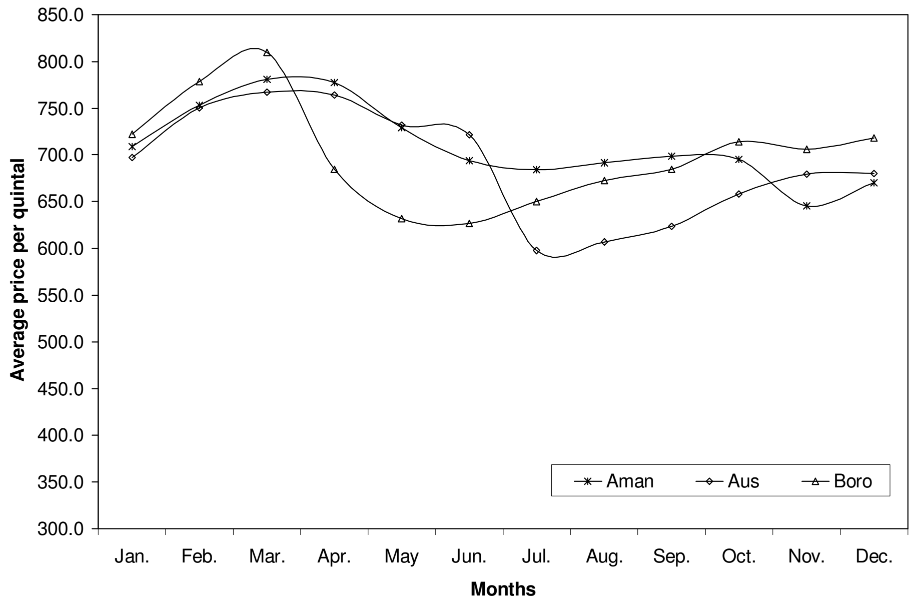


Figure 4. Intra seasonal fluctuations of rice price

Tests for Seasonality

Seasonality can be deterministic and/or stochastic. For quarterly data, deterministic seasonality assumes that the data generating process for the variable y_t is

$$y_t = \delta_1 s_{1t} + \delta_2 s_{2t} + \delta_3 s_{3t} + \delta_4 s_{4t} + \varepsilon_t$$

where s_{st} (=1 in season s, 0 elsewhere, for s = 1, 2, 3, 4) is a seasonal dummy variable. Including seasonal dummy variables in a regression model is appropriate for variables with deterministic seasonality. The absence of these dummy variables will lead to the standard problem of bias associated with the exclusion of relevant explanatory variables. Stochastic seasonality extends the unit root hypothesis to seasonal time series. An integrated seasonal process is a process that contains unit roots at the seasonal frequencies, and appropriate differencing filters are required for seasonally integrated processes.

Table 2 Estimates of auxiliary regression

Variable	Coefficient	Std. Error	t-Statistic	Prob.
S1	747.640	28.691	26.058	0.000
S2	733.532	28.691	25.566	0.000
S3	691.352	28.000	24.691	0.000
S4	670.298	28.000	23.939	0.000

If any variable y_t is taken as the dependent variable, the equation involves the regression of the growth rate of the variable on a set of seasonal dummy variables. Estimates of the δ_s (s = 1, 2, 3, 4) coefficients can be used to observe the pattern of seasonality which are shown in Table 2. As can be seen from the table, except last one all coefficients are highly significant. Thus it can be concluded that rice price exhibit seasonal patterns.

As a formal test for seasonality, we used the auxiliary regression in equation (6) proposed by Miron (1994).

$$y_t = \beta t + \gamma y_{t-1} + \sum_{i=1}^{4} \delta_i s_{it} + \varepsilon_t \qquad (6)$$

where s_{st} (=1 in season s, 0 elsewhere, for s = 1, 2, 3, 4) is a seasonal dummy variable and ε_t is assumed to be a stationary.

Table 3 The coefficient of trend, lag and seasonality.

Variable	R^2	β	γ	δ_1	δ_2	δ_3	δ_4	J-B
Aman	0.760	1.122	0.724	228.417	-82.431	-107.091	-104.894	4.75
		(0.011)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.093)
Aus	0.745	0.849	0.701	238.835	-52.468	-174.798	-28.565	1.048
		(0.027)	(0.000)	(0.000)	(0.012)	(0.000)	(0.156)	(0.592)
Boro	0.801	0.989	0.760	200.762	-178.927	-56.253	-37.153	0.964
		(0.018)	(0.000)	(0.000)	(0.000)	(0.006)	(0.058)	(0.618)

Note: *Figures in parentheses shows p-values*

Small values of R^2 indicate that the model does not fit the data well. The sample R^2 tends to optimistically estimate how well the models fit the population. For all model R^2 is very high, so our estimated model is consistent. From the above table, at a glance we may take decision about the coefficient of trend, lag price and seasonal dummy effects are almost statistically significant. Except in case of Aus price 4^{th} quarter gives insignificant result. Also Jarque-Bera (J-B) suggested that prices of rice were strictly followed normality assumptions, except Aman.

Conclusion

Stabilization of rice prices is of utmost importance for Bangladesh. Though, the prices of rice play an important role in the economy of Bangladesh and its fluctuation has a great economic impact on the people of Bangladesh, the dis-equilibrium of demand and supply in agriculture is a universal phenomenon. The agricultural prices are disgracefully unsteady. The price of rice is increasing more rapidly than the income of the poor. This study finds out two types of temporal variations in rice prices: inter-year (seasonality) and intra-year. Rice price is increasing day by day. This paper examined trend of annual and seasonal rice price fluctuations in Bangladesh. The study reveals that the trend of rice price showed an upward trend used by moving average method. The study also reveals that the stochastic seasonality in rice prices presented positive sign, but in case of 4^{th} quarter of Aus price, it gives insignificant result. Since rice price shows an upward trend, both constant, and constant trend are used in the models to test for unit roots. From the results of the study, researchers revealed that prices of Aman, Aus and Boro time series were stationary. Year-to-year fluctuations in nominal prices of rice in Bangladesh were very high. Prices were especially unstable due to severe rice shortages caused by drought-related production shortfalls and shortage of foreign exchange of government for rice imports. The present study reflected that the prices of all three varieties of rice in 2000-2006 were relatively stable but during that period prices were increasing insignificantly. But in the early of 2000, it was observed that the price of rice was unstable but in that period the variations were at random but with a minor abruption. The changes in percentages of price during 2000-2006 lead the overall change in price to the upward trend. The study also shows that SACF and SPACF for the level of rice price series at the intra-seasonal pattern of rice price by taking monthly average of the period 1986 to 2006.

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Appendix

Table 1: Average annual rates of percentage change and estimated price of Aman, Aus and Boro price during the period, 1986q3 to 2006q4. q= Quarter of a year.

Year	Price per quintal			Changes from previous year (%)			Estimated Trend Values		
	Aman	Aus	Boro	Aman	Aus	Boro	Aman	Aus	Boro
1986q3	522.90	434.70	491.40				558.95	570.88	547.90
1986q4	485.10	544.50	590.40	-7.23	25.26	20.15	562.68	573.79	551.64
1987q1	593.10	644.40	651.60	22.26	18.35	10.37	566.41	576.70	555.39
1987q2	678.60	585.90	568.80	14.42	-9.08	-12.71	570.13	579.61	559.14
1987q3	637.20	509.40	594.00	-6.10	-13.06	4.43	573.86	582.52	562.89
1987q4	603.00	598.50	634.50	-5.37	17.49	6.82	577.59	585.43	566.64
1988q1	617.40	639.90	644.40	2.39	6.92	1.56	581.32	588.34	570.38
1988q2	612.00	611.10	517.50	-0.87	-4.50	-19.69	585.05	591.25	574.13
1988q3	613.80	498.60	547.20	0.29	-18.41	5.74	588.78	594.16	577.88
1988q4	555.30	591.30	562.50	-9.53	18.59	2.80	592.51	597.07	581.63

Cont. Table

Year	Price per quintal			Changes from previous year (%)			Estimated Trend Values		
	Aman	Aus	Boro	Aman	Aus	Boro	Aman	Aus	Boro
1989q1	644.40	696.60	668.70	16.05	17.81	18.88	596.24	599.98	585.38
1989q2	692.10	765.90	558.00	7.40	9.95	-16.55	599.97	602.89	589.12
1989q3	491.40	515.70	577.80	-29.00	-32.67	3.55	603.70	605.80	592.87
1989q4	534.60	568.80	581.40	8.79	10.30	0.62	607.43	608.71	596.62
1990q1	639.00	664.20	693.00	19.53	16.77	19.20	611.16	611.62	600.37
1990q2	649.80	693.00	721.80	1.69	4.34	4.16	614.88	614.53	604.12
1990q3	661.50	634.50	611.10	1.80	-8.44	-15.34	618.61	617.44	607.86
1990q4	664.20	646.20	636.30	0.41	1.84	4.12	622.34	620.35	611.61
1991q1	711.33	729.90	700.20	7.10	12.95	10.04	626.07	623.26	615.36
1991q2	692.00	723.60	576.33	-2.72	-0.86	-17.69	629.80	626.17	619.11
1991q3	721.00	613.00	676.67	4.19	-15.28	17.41	633.53	629.08	622.86
1991q4	613.33	692.73	716.20	-14.93	13.01	5.84	637.26	631.99	626.60
1992q1	733.33	721.57	789.97	19.57	4.16	10.30	640.99	634.90	630.35
1992q2	747.00	797.40	628.33	1.86	10.51	-20.46	644.72	637.81	634.10
1992q3	784.33	569.30	652.67	5.00	-28.61	3.87	648.45	640.72	637.85
1992q4	515.03	538.50	568.00	-34.33	-5.41	-12.97	652.18	643.63	641.59
1993q1	528.33	529.20	606.60	2.58	-1.73	6.80	655.90	646.54	645.34
1993q2	510.67	523.80	441.67	-3.34	-1.02	-27.19	659.63	649.45	649.09
1993q3	470.67	423.67	461.00	-7.83	-19.12	4.38	663.36	652.36	652.84
1993q4	521.33	541.00	558.00	10.76	27.69	21.04	667.09	655.27	656.59
1994q1	663.67	658.80	662.40	27.30	21.77	18.71	670.82	658.18	660.33
1994q2	701.67	718.20	593.67	5.73	9.02	-10.38	674.55	661.09	664.08
1994q3	709.67	668.67	677.33	1.14	-6.90	14.09	678.28	664.00	667.83
1994q4	721.00	717.40	750.17	1.60	7.29	10.75	682.01	666.91	671.58
1995q1	900.90	789.30	832.50	24.95	10.02	10.98	685.74	669.82	675.33
1995q2	855.80	896.40	774.30	-5.01	13.57	-6.99	689.47	672.73	679.07
1995q3	651.60	665.00	781.33	-23.86	-25.81	0.91	693.20	675.64	682.82
1995q4	687.37	694.93	746.50	5.49	4.50	-4.46	696.93	678.55	686.57
1996q1	793.33	819.00	827.10	15.42	17.85	10.80	700.65	681.46	690.32
1996q2	707.00	765.90	593.67	-10.88	-6.48	-28.22	704.38	684.37	694.07
1996q3	598.33	517.17	580.67	-15.37	-32.48	-2.19	708.11	687.28	697.81
1996q4	510.33	504.33	521.00	-14.71	-2.48	-10.28	711.84	690.19	701.56
1997q1	586.33	566.10	623.70	14.89	12.25	19.71	715.57	693.10	705.31
1997q2	649.00	634.50	637.50	10.69	12.08	2.21	719.30	696.01	709.06
1997q3	580.00	498.00	532.00	-10.63	-21.51	-16.55	723.03	698.92	712.81
1997q4	600.67	610.17	649.57	3.56	22.52	22.10	726.76	701.83	716.55
1998q1	837.00	854.10	906.30	39.35	39.98	39.52	730.49	704.74	720.30

Cont. Table

Year	Price per quintal			Changes from previous year (%)			Estimated Trend Values		
	Aman	Aus	Boro	Aman	Aus	Boro	Aman	Aus	Boro
1998q2	825.00	859.50	663.00	-1.43	0.63	-26.85	734.22	707.65	724.05
1998q3	733.33	702.67	753.67	-11.11	-18.25	13.68	737.95	710.56	727.80
1998q4	840.00	824.67	873.67	14.55	17.36	15.92	741.68	713.47	731.55
1999q1	916.00	898.00	918.33	9.05	8.89	5.11	745.40	716.38	735.29
1999q2	793.67	866.33	639.33	-13.36	-3.53	-30.38	749.13	719.29	739.04
1999q3	759.33	625.00	698.67	-4.33	-27.86	9.28	752.86	722.20	742.79
1999q4	684.67	590.00	700.67	-9.83	-5.60	0.29	756.59	725.11	746.54
2000q1	730.33	666.67	744.67	6.67	12.99	6.28	760.32	728.02	750.29
2000q2	731.67	704.00	628.00	0.18	5.60	-15.67	764.05	730.93	754.03
2000q3	710.67	569.33	620.33	-2.87	-19.13	-1.22	767.78	733.84	757.78
2000q4	630.67	683.33	646.00	-11.26	20.02	4.14	771.51	736.75	761.53
2001q1	693.67	766.33	694.00	9.99	12.15	7.43	775.24	739.66	765.28
2001q2	721.67	767.33	612.00	4.04	0.13	-11.82	778.97	742.57	769.03
2001q3	657.67	599.00	601.00	-8.87	-21.94	-1.80	782.70	745.48	772.77
2001q4	659.00	668.67	721.67	0.20	11.63	20.08	786.42	748.39	776.52
2002q1	727.33	753.33	724.33	10.37	12.66	0.37	790.15	751.30	780.27
2002q2	704.00	708.00	653.33	-3.21	-6.02	-9.80	793.88	754.21	784.02
2002q3	722.00	670.33	720.67	2.56	-5.32	10.31	797.61	757.12	787.77
2002q4	732.00	736.33	785.67	1.39	9.85	9.02	801.34	760.03	791.51
2003q1	822.00	795.00	807.00	12.30	7.97	2.72	805.07	762.94	795.26
2003q2	785.67	742.00	689.67	-4.42	-6.67	-14.54	808.80	765.85	799.01
2003q3	781.67	719.33	771.67	-0.51	-3.05	11.89	812.53	768.76	802.76
2003q4	762.33	749.00	796.00	-2.47	4.12	3.15	816.26	771.67	806.51
2004q1	789.33	828.67	855.00	3.54	10.64	7.41	819.99	774.58	810.25
2004q2	777.00	800.14	731.67	-1.56	-3.44	-14.42	823.72	777.49	814.00
2004q3	796.67	748.33	792.67	2.53	-6.47	8.34	827.45	780.40	817.75
2004q4	871.33	923.33	942.33	9.37	23.39	18.88	831.17	783.31	821.50
2005q1	1016.00	959.00	1027.00	16.60	3.86	8.98	834.90	786.22	825.24
2005q2	822.67	928.00	812.00	-19.03	-3.23	-20.93	838.63	789.13	828.99
2005q3	925.33	827.00	935.33	12.48	-10.88	15.19	842.36	792.04	832.74
2005q4	906.00	852.67	968.00	-2.09	3.10	3.49	846.09	794.95	836.49
2006q1	1010.00	788.85	1027.67	11.48	-7.48	6.16	849.82	797.86	840.24
2006q2	1013.67	692.55	911.33	0.36	-12.21	-11.32	853.55	800.76	843.98
2006q3	989.33	790.18	976.00	-2.40	14.10	7.10	857.28	803.67	847.73
2006q4	979.00	848.33	1016.33	-1.04	7.36	4.13	861.01	806.58	851.48

Source: *Bangladesh Agriculture Marketing Department*