MSSD DISCUSSION PAPER NO. 44

DYNAMICS OF AGRICULTURAL WAGE AND RICE PRICE IN BANGLADESH: A RE-EXAMINATION

by

Shahidur Rashid

Markets and Structural Studies Division

International Food Policy Research Institute
2033 K Street N.W.
Washington, D.C. 20006
Tel. (202) 862-5600 and Fax (202) 467-4439
http://www.cgiar.org/ifpri

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Contact: Candice Cohen
c.cohen@cgiar.org
Tel. (202) 862-8120 or Fax. (202) 467-4439

Shahidur Rashid is a Postdoctoral Fellow at the International Food Policy Research Institute (IFPRI), 2033 K Street N.W., Washington, D.C. 20006 – U.S.A. email: s.rashid@cgiar.org.

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TABLE OF CONTENT

LIST OF TABLES

LIST OF FIGURES

ABSTRACT

Like many other Asian countries, the causal relationship between agricultural productivity and the incidence of rural poverty has been a widely debated subject in Bangladesh. A number of studies argued that the real agricultural wage rate was declining during the period when the country had experienced overall agricultural growth. This paper contributes to this debate in two ways: i) it re-examines the methodological aspects of past studies and presents alternative estimates; and ii) analyzes dynamics of agricultural wage and rice price using the most recent data. Multivariate co-integration techniques are used to examine the long and short-run relationships among agricultural wage rate, rice price, urban wage rate, and other prices. The results show that agricultural wage and rice price maintained strong co-integrating relationships during the periods 1949/50—1979/80; and the elasticities of agricultural wage rate with respect to rice price are substantially higher than what past studies had reported. Analyses of post-famine data (1976/77—1998/99) suggest that rice price, which was strongly co-integrated with agricultural wage rate until the early 80s, is no longer a significant determinant of wage formation in Bangladesh.

1. INTRODUCTION

The dynamics of agricultural wage formation and its implications for poverty and income distribution has been a subject of much debate in Bangladesh. There are three major questions at the center of this debate: i) has an increase in agricultural productivity through green revolution technology translated into a reduction in rural poverty through increased wage income[1]? ii) what has been the trend in the long-term movement of real agricultural wages and what welfare implications does it have for various groups of rural households? and iii) how do agricultural wages respond to changes in prices and other wages in the economy? In the face of low per capita income and increasing dependence of the poor on labor markets for their livelihood, these questions are of significant importance to the policy makers and development practitioners in the country[2]. However, the answers to these questions, largely based on time series analyses, continue to remain elusive.

While there is overwhelming micro-level evidence to suggest that the new technology has had positive impacts in alleviating poverty through direct productivity gains and through indirect linkage effects, macro-econometric analyses present a different set of results[3]. Using time series data from 1949/50 to 1980/81, available studies argue that real agricultural wage has been on a downward trend during the period in which the country experienced positive growth in overall agricultural productivity (Khan 1984, Ravallion 1990, Boyce and Ravallion 1991, Ravallion 1994). Furthermore, on the basis of their finding of low

[1] In the context of green revolution in India, the debate over this question went on for long time. A collection of papers in the book edited by Mellor and Desai (1985) discuss different dimensions of this basic issue.
[2] In fact, this study was initiated in response to a request from the Ministry of Food, Government of Bangladesh, to re-examine the issue.
[3] For micro-level evidence in Bangladesh, see Hossain (1988); for farm non-farm growth linkage, see Haggblade and Hazell (1989).

short-run elasticity of wages to rice price, Boyce and Ravallion (1991) express concerns that an increase in nominal rice price can have detrimental affects on the poor, for whom wage is the main source of income and rice occupies a major share of their expenditure[4].

This study undertakes two tasks: i) it re-examines the methodological aspects of past studies and ii) analyzes the wage-price dynamics using the most recent data. The first task is motivated by the new developments in time series econometrics, namely multivariate co-integration techniques, which have revolutionized the analyses of historical data. Since all of the past studies are based on classical regression method, we consider it to be a worthwhile effort to re-examine the data under the new analytical framework. The rationale for the second task, on the other hand, is two-fold. First, with a war in 1971, famine in 1974, and a couple of military coup d'états, the Bangladesh economy was in chaos during the 70s. Therefore, inclusion of data for this decade into the analyses is likely to affect the parameter estimates. Second, in addition to rebuilding the infrastructure that was destroyed during the war, the country has undertaken a substantial liberalization initiative during the 80s and early 90s, which is likely to affect the adjustment process of agricultural wages and other prices.

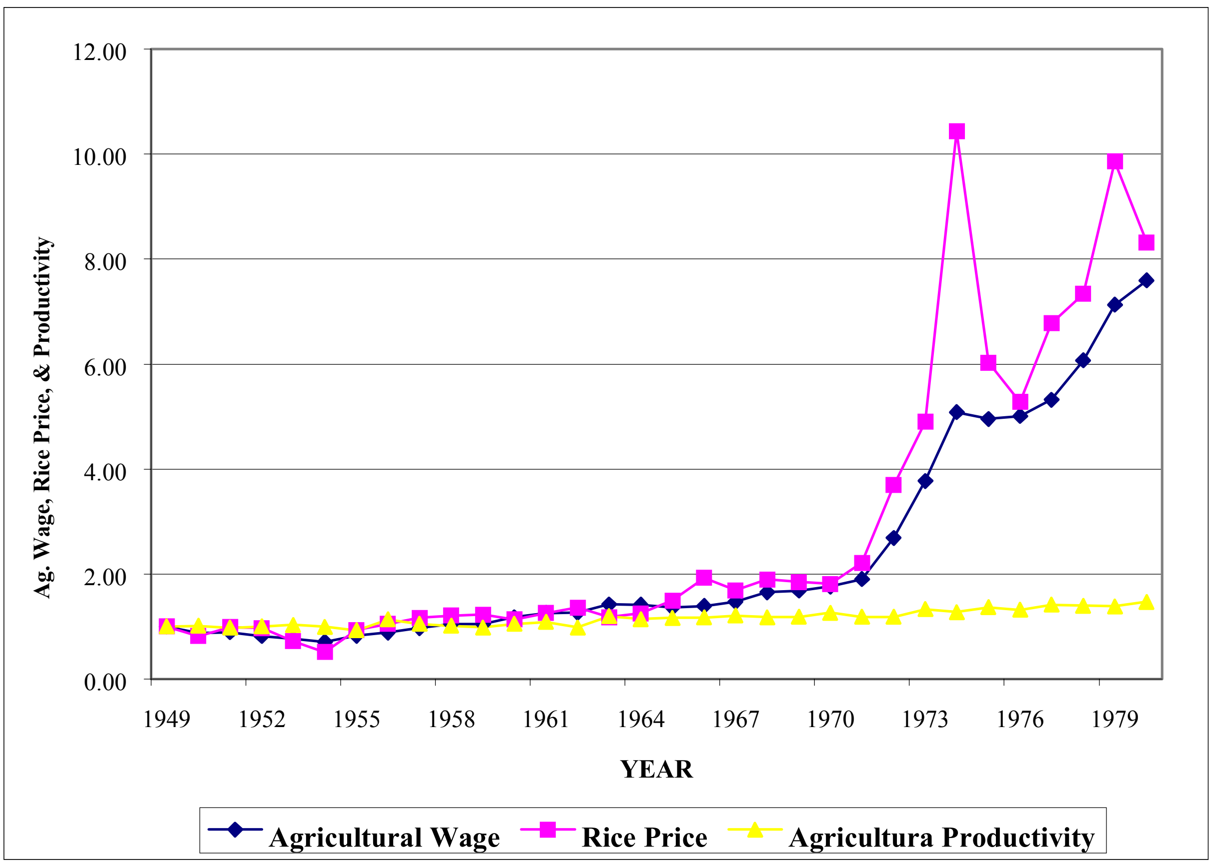
Four important results come out of the re-examination of past studies. First, under a variety of unit root testing, we find that all variables used in the previous studies are non-stationary, which invalidate the use of the classical regression method. Second, within the framework of Johansen's multivariate cointegration technique, our analyses suggest that agricultural wages and rice prices

[4] Available studies suggest that expenditure elasticity of rice for the bottom quartile of the rural population in Bangladesh ranges from 0.68 to 1.01. See Ahmed and Shams (1996), Goletti (1994), and Hossain (1988).

maintained a strong co-integrating relationship throughout the period from 1949/50 to 1979/89. This result is in sharp contrast with past conclusions that the real agricultural wage in Bangladesh was in an alarming downward trend from the mid-sixties to early eighties. Third, we argue that inclusion of the agricultural productivity variable—measured as and index of *per acre production*—into the long run co-integrating relationship is incorrect. This is because variables like nominal wages and prices are inflationary and can be affected by changes in other sectors, as well as through macroeconomic policy levers e.g., exchange rate policies. On the other hand, *per acre production* is non-inflationary and is constrained by the resource endowment and the technology of a country at a given point in time. Therefore, a positive long run relationship between this variable and nominal wages is very unlikely. Figure 1 supports our argument. It clearly shows that while agricultural wage and rice price co-move throughout the period, they diverge from *per acre productivity* from the mid-sixties onward. Finally, a number of tests on the residuals suggest that models used in previous studies suffer from misspecification problems. For instance, non-normality of residuals for three out of six variables in the Palmer-Jones (1993) full model cannot be rejected at conventional level of significance, implying that the model violates the critical assumption of error terms being white noise.

The results of the analyses on post-famine data are striking. Our analyses suggest that there has been a major shift in the dynamics of agricultural wage formation during the last twenty-five years in Bangladesh. Using data from 1976/77 to 1998/99, we demonstrate that rice price, which was strongly cointegrated until the late 70s, is no longer a significant determinant of agricultural wages in Bangladesh. This implies that the widespread concern about declining purchasing power due to a sluggish adjustment of wages to rice is no longer valid for the country. Furthermore, our cointegration results indicate that urban wage rate of unskilled workers in major cities is the single most important variable to explain the long-run dynamics of agricultural wage rates during this period. The rest of the paper is organized as follows. The next section provides a brief discussion on past studies. Section three describes data and empirical methodology. Empirical results of the analyses of data sets from past studies are presented in section four. Section five discusses results of post-famine analyses and conclusions, and policy implications are drawn in the final section.

Figure 1. Trends in selected variables (1949/50—1979/80)



2. A BRIEF NOTE ON EXISTING LITERATURE

There are at least a dozen studies that have analyzed various aspects of agricultural wage formation in Bangladesh. Three major issues addressed by these studies are: i) whether or not agricultural wages adjust to food prices and other wages in the economy (Ravallion 1987, Boyace and Ravallion, 1991; and Palmer-Jones 1993); ii) whether productivity gains through green revolution technology trickled down to laborers through increased wage rates (Bose 1974; Khan 1984; and Boyce 1989); and iii) whether market theory of labor demand and labor supply provides a better explanation of wage determination than *subsistence* or *efficiency wage* theories (Ahmed 1981, and Hossain 1990). The first two issues are the focus of this study.

An obvious starting point to reexamine these issues is Boyce and Ravallion (1991), which is by far the most rigorous econometric study on the dynamics of agricultural wages in Bangladesh. Using annual data from 1949/50-1980/81, the study pointed out several shortcomings of past empirical research on the subject and presented a new dynamic model for agricultural wage determination. In particular, the authors demonstrated that Khan's (1984) findings of significant positive relationships between real wage, agricultural productivity, and improved agricultural terms of trade were spurious due to the failure to account for autocorrelation. Thus, following Sargan (1964) and Hendry, Pegan, and Sargan (1984), they reformulated the wage equation in an *error-correction* framework, which included prices of rice, jute, and cloth (p^r, p^j, p^c), the manufacturing wage (w^m), agricultural output per net cropped area Q^b, and a time trend as explanatory variables. Assuming log-linear functional form, the wage equation was represented in 1-year vector autoregresson in level as follows:

$$
\begin{aligned}
w_t^a &= \alpha_0 + \alpha_1 w_{t-1}^a + \beta_0 p_t^r + \beta_1 p_{t-1}^r + \gamma_0 p_t^j + \gamma_1 p_{t-1}^j + \delta_o p_t^c \\
&\quad + \delta_1 p_{t-1}^c + \varepsilon_0 w_t^m + \varepsilon_1 w_{t-1}^m + \phi_0 Q_t^B + \phi_1 Q_{t-1}^B \\
&\quad + \pi_0 t + \pi_1 t^2 + v_t \ \ (t = 1,.....,T) \qquad (1)
\end{aligned}
$$

To give error-correction representation, the above equation was re-written in the following equivalent form:

$$
\begin{aligned}
\Delta w_t^a &= \alpha_0 + (\alpha_1 - 1)(w_{t-1}^a - p_{t-1}^r) + (\gamma_0 - \gamma_1)(p_{t-1}^j - p_{t-1}^r) \\
&\quad + (\delta_0 - \delta_1)(p_{t-1}^c - p_{t-1}^r) + (\varepsilon_0 - \varepsilon_1)(w_{t-1}^m - p_{t-1}^r) \\
&\quad + \beta_0 \Delta p_t^r + \gamma_0 \Delta p_t^j + \delta_0 \Delta p_t^c + \varepsilon_0 \Delta w_t^m \\
&\quad + (1 - \alpha_1 - \beta_0 - \beta_1 - \gamma_0 - \gamma_1 - \delta_0 - \delta_1 - \varepsilon_0 - \varepsilon_1) p_{t-1}^r \\
&\quad + \phi_0 Q_t^B + \phi_1 Q_{t-1}^B + \pi_0 t + \pi_1 t^2 + v_t \qquad (2)
\end{aligned}
$$

By imposing homogeneity restriction, i.e., setting $(1-\alpha_1-\beta_0-\beta_1-\gamma_0-\gamma_1-\delta_0-\delta_1-\varepsilon_0-\varepsilon_1)=0$, and following a *general to specific* modeling technique to drop the insignificant variables, the authors estimated the following parsimonious model:

$$
\begin{aligned}
\Delta w_t^a &= \underset{(0.51)}{0.045} + \underset{(5.7)}{0.22}(p_t^r - p_t^c) + \underset{(8.6)}{0.47}(w_t^m - w_{t-1}^a) - \underset{(9.8)}{0.32}(w_{t-1}^m - p_t^c) \\
&\quad - \underset{(2.9)}{0.00037}\, t^2 + \hat{v}_t \qquad (3)
\end{aligned}
$$

The major conclusions of Boyce-Ravallion (1991) and Ravallion (1990) are based on these results. Pointing to low short-run elasticity of the agricultural wage rate with respect to a change in the rice price, which is estimated to be 0.22, the study expresses concerns that a sluggish adjustment of wage and rice prices can have detrimental affects on the poor, whose major source of income was presumed to be agricultural wages.

A second major conclusion was that increases in agricultural productivity (measured by the value of agricultural output per net cropped area) did not have a significant impact on agricultural wages. This conclusion was based on the insignificant coefficient of the agricultural productivity variable in the initial regressions that led to this variable being dropped in the "parsimonious" model.

Palmer-Jones (1993) criticized the Boyce and Ravallion analysis on the grounds that the "parsimonious" model failed the prediction and stability tests when estimated over the period up to 1989. The modeling framework of Palmer-Jones is essentially the same as Boyce and Ravallion, except the fact that he included a dummy variable for 1972-1974, and a discontinuous time trend (with a value of 1 in 1949 up to 16 in 1964 and 0 afterward) into his analyses. These new time trends and dummy variables are important for Palmer-Jones's specification to support his conclusions that the longer period estimates do not support the earlier Boyce and Ravallion finding that the implicit long run real wage rate has been on an alarming downward trend.

In his rebuttal, Ravallion (1994) acknowledged the prediction failure but forcefully criticized the use of dummies and the time trend, which in his view was an *odd half-time* trend because it suddenly dropped down to zero from 1964 onward. Using a new set of variables and a new parsimonious model, he shows that the longer time series can produce an equation that fits the 1980s data nearly as well as the Palmer-Jones. Moreover, unlike the Palmer-Jones model, his model is homogenous, i.e. the real wage depends only on real, not nominal, variables. There are significant changes relative to the earlier model, however. There is no significant time trend and growth rate in output now has a significant effect on real agricultural wages. In conclusion of the rebuttal, Ravallion stood by the conclusions of Boyce and Ravallion (1991) that the real wages were declining much of the 1960s and 1970s.

There are at least two reasons to be cautious about these studies. First, all of these studies are based on Ordinary Least Square (OLS) estimation with non-stationary explanatory variables, which gives rise to spurious relationships among the variables and inconsistency of parameter estimates (Dicky et. al 1994)[5]. As we demonstrate through a variety of unit root tests in the following section, all variables are indeed found to be nonstationary, suggesting that a co-integration framework is necessary to establish the long run relationship. Second, although there is no *a priori* reason, these studies make an implicit assumption that all right hand side variables of the wage equation are exogenous. Such an assumption has important econometric implications, as endogeneity of any right hand side variable will lead to an *identification* problem. In fact, a weak exogeneity test gives a χ^2 value of 5.86 for rice price, indicating that it cannot be considered to be exogenous to the system and a VAR method is required. This study addresses these issues in analyzing the dynamics of wages and prices.

3. DATA AND METHODOLOGY

We use two different sets of data for this study. To re-examine past studies, we use the data set reported in Palmer-Jones (1993), which contains data from 1949/50 through 1989/90. It is to be noted that the first thirty-year's of data in this data set are identical to the data used in Boyce and Ravallion (1991)[6]. For the purpose of analyzing dynamics of wages and prices during the post-famine period, we constructed four time series variables—agricultural wages, urban wages, rice price, and cloth price index—using various publications of the Bangladesh Bureau of Statistics (BBS). Note that agricultural wage rates in Bangladesh vary depending on whether or not employers provide meals to their

[5] Ravallion (1994) and Palmer-Jones (1994) indicated about the possibility of non-stationarity, but neither of them carried out any testing or attempted alternative modeling.
[6] To check data accuracy, we re-estimated Boyce and Ravallion (1991) model and all parameters were found be identical to the ones they reported in their paper.

workers. We use district-wise wage rates of agricultural workers without meals[7]. Daily average wage rates of unskilled construction workers in seven major cities (Dhaka, Chittagong, Narayangonj, Rajshahi, Khulna, Sylhet, and Rangpur) are taken as proxies for urban wage rates. This is different from the manufacturing wage rate used in the previous studies, which was an average wage rate of the unskilled workers in the manufacturing sectors such as jute and textile industries. Since labor unions are significantly strong in the industrial sector in Bangladesh, wage rates of unskilled construction workers are more likely to better reflect urban wage than the average wages of the unskilled workers in the unionized manufacturing sectors. Rice price is the daily average of the district level prices of coarse rice, commonly consumed by the poor. Finally, cloth price index is the merged cloth price index reported in the monthly statistical bulletin of the BBS. All variables are transformed into natural logarithm and 1977/78 is taken as the base year.

Our empirical analyses are based on cointegration techniques. The concept of cointegration states that if a long run relationship exists among a set of non-stationary variables, then the deviation from the long run equilibrium path should be bounded. In other words, the existence of a long-run relationship implies that cointegrated variables cannot wander too far away from each other. Formally, two non-stationary series x_t and y_t are said to be cointegrated if the following two conditions are satisfied: i) both series are integrated of the same order, and ii) a linear combination of x_t and y_t exists which is $I(0)$ i.e., stationary. Therefore, the first step in cointegration analyses is to examine the order of integration (stationarity) of relevant variables (Banarjee et. al., 1993; Muscatelli and Hurn, 1992).

[7] Two years' of data, namely 1988/89 and 1998/99, were constructed using monthly data published in the Monthly Statistical Bulletin of Bangladesh. We attempted to construct all variables using monthly data, but there is a large number of missing values in wage data. In particular, there were no monthly wage data for entire 1990 and for January 1992 through June 1993. Surprisingly, however, annual data for those years are reported in the Statistical Yearbook.

Once the order of integration is determined, the next step is to determine whether the non-stationary variables form a co-integrating or long run relationship among them. Engle and Granger (1987) demonstrate that for any set of *I(1)* variables, error correction and cointegration are an equivalent representation. In other words, if an error correction model provides sufficient depiction of the variables, then the variables must be cointegrated and vice versa. This provides the basis for using an error correction model in the cointegration framework for analyzing the dynamics of agricultural wage formation. Although Engle and Granger (1987) proposed a simple estimation technique, their methodology was subsequently criticized to have serious econometric shortcomings[8]. The methodology presented in Johansen (1988) and Johansen and Juselius (1990) overcomes those shortcomings and provides ways to test parameter restrictions and other structural hypotheses in a cointegrated system. One of the striking features of this method is that since estimation is carried out in a multivariate VAR framework, it does not make any *a priori* assumption on endogenous or exogenous divisions of variables. It also enables the determination of whether more than one co-integrating relationships exists. Estimation of our model in this framework is described in section 4.2.

[8] See Enders (1995) for a detail discussion.

4. EMPIRICAL RESULTS

Empirical results are presented in three subsections. The first subsection discusses the results of stationarity testing; the second section analyzes the long run relationship; the third section presents results of some hypothesis testing; and the final section analyzes the short-run dynamics.

4.1 Tests for stationarity

Despite availability of other statistical tests, the Dicky Fuller and Augmented Dicky Fuller (ADF) tests continue to be the most popular and commonly used to determine the order of integration. However, a number of recent studies have pointed out that the standard ADF test often fails to reject null hypothesis of non-stationarity, even when null is false (see Banerjee et al. 1992). Therefore, to double check the results of the ADF test, another test developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992)—hereafter KPSS—is used to test for the stationarity of relevant variables. The basic difference between two testing procedure is that, while the ADF test begins by testing a null hypothesis of $I(1)$ against $I(0)$, KPSS does the reverse i.e., it tests $I(0)$ against an $I(1)$ alternative.

Another issue that deserves special attention in doing time series analyses with Bangladeshi data is that the country suffered a number of shocks in the 70s, including a war in 1971 and a famine in 1974. As Perron (1989 and 1990) has shown, such structural shocks to the mean of a stationary variable may bias the ADF test towards non-rejection of the null hypothesis of unit root. In other words, in the presence of such shocks, one may conclude a time series to be $I(1)$, which in reality is $I(0)$. Therefore, in addition to ADF and KPSS, Perron (1989) tests are performed to account for the major shocks in the 70s. Two break years, 1971 and 1974, representing war and famine respectively, were tried in implementing the Perron test for unit root.

Results for the stationarity tests on Boyce and Ravallion (1991) variables for 1949/50 to 1979/80 and 1949/50 to 1988/89 are presented in Table 1. All three tests were performed both on levels (i.e., original series) and on first differenced series. In carrying out these tests, Akaike Information Criterion (AIC) was used to determine appropriate lag length for every series. The results of all three tests clearly reveal that while the variables in level are non-stationary, the first difference of each of the series is stationary. In other words, all variables are integrated of order one i.e., $I(1)$. Graphical results are also in conformity with these test results. For the sake of brevity, we have only plotted natural log of agricultural wage from 1949/50 to 1979/80 in Figure 2. Graphs for other variables are similar. Thus, an error correction model can be used to determine the co-integrating relationship among the variables.

Figure 2. Natural log agricultural wage in *levels* and *first differences* (1949/50—1979/80)

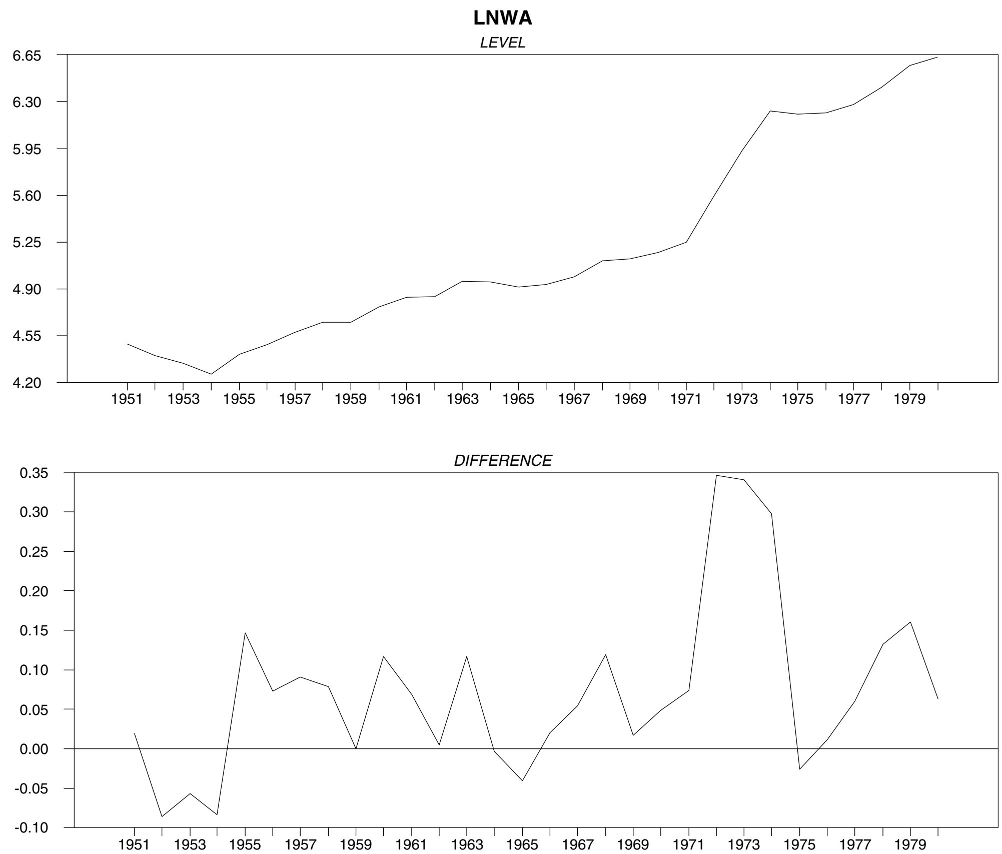


Table 1. Stationarity Tests on Selected Variables

Series	Levels			First-difference		
	ADF	KPSS	PERRON	ADF	KPSS	PERRON
1949/50—1980/81 Data						
p_t^r	-2.69	0.667	-2.697	-5.73	0.229	-5.738
p_t^c	-2.08	0.599	-1.625	-4.31	0.201	-4.309
p_t^j	-3.44	0.646	-3.445	-7.67	0.152	-7.668
w_t^a	-1.84	0.672	-1.997	-3.65	0.412	-3.618
w_t^m	-0.99	0.709	-0.988	-4.04	0.387	-4.352
q_t^b	-1.13	0.707	-5.619	-9.62	0.164	-7.617
1949/50—1980/89 Data						
p_t^r	-3.20	0.870	-3.207	-6.88	0.188	-6.886
p_t^c	-2.50	0.831	-1.986	-4.86	0.156	-4.860
p_t^j	-0.82	0.855	-4.069	-6.52	0.163	-8.628
w_t^a	-2.34	0.864	-2.342	-3.98	0.387	-3.979
w_t^m	-0.74	0.835	-0.740	-5.52	0.432	-5.524
q_t^b	-1.36	0.903	-3.269	-9.01	0.340	-6.383

[a] All variables are in natural logarithm. Note that while ADF tests $H_0 : X_i \sim I(1)$ against $H_1 : X_i \sim I(0)$, KPSS method tests the reverse. The ADF and KPSS critical values at 5% level of significance are -3.07 and 0.463 respectively. Critical value for ADF test is taken from Mackinnon (1991). The break functions for Perron test are calculated as $\lambda = Tb/T,$ where Tb is total number of observations until the break year and *T* is total sample size. In this analysis, 1971 is considered to be the break year. For 1949/50-197980 data, $\lambda = 0.40$ and T=30; the critical value at 5% level of significance for this test (null hypothesis of unit root) is –3.99. For the other data set, $\lambda = 0.50$ and T=39; the corresponding critical value at 5% level of significance is –3.56.

4.2 Analyses of long run relationship

Since all variables are tested to be $I(1)$, Ordinary Least Square (OLS) is no longer a valid method of estimation as it can generate spurious regression estimates. The obvious choice is Johansen's multivariate maximum likelihood cointegration method. Following this methodology, the long run wage function of Boyce and Ravallion (1991) can be formulated as a 6-dimensional vector autoregressive model with Gausian errors as,

$$Z_t = A_1 Z_{t-1} + \cdots + A_k Z_{t-k} + \mu + \delta t + \xi_t , \qquad (t = 1, \cdots, T) \tag{5}$$

where $Z_t = \left(p^r, p^j, p^c, w^a, w^m, Q^b \right)_t$;is a vector of explanatory variables; $Z_{-k+1}, \cdots, Z_0$ are fixed, $\xi_1, \cdots, \xi_T$ are i.i.d. $N_p(0, \Sigma)$; μ is a vector of constants, and δ is a trend coefficient. The above equation can be represented in a VAR in error-correction form as follows:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \cdots + \Gamma_{k-1} \Delta Z_{t-k+1} + \Pi Z_{t-k} + \mu + \delta t + \xi_t \qquad (t = 1, \cdots, T) \tag{6}$$

where $\Gamma_i = -(I - A_1 - \cdots - A_i)$ with $(i = 1, \cdots, k-1)$; $\Pi = -(I - A_1 - \cdots - A_k)$; $\Delta = (1 - L)$, and L is a lag operator. Under this framework, the cointegration hypothesis is formulated as a reduced rank of the $\Pi -$matrix as,

$$H_1(r) : \Pi = \alpha \beta',$$

where α and β are $p \times r$ matrices of full rank; and p and r are number of variables and number of co-integrating vectors respectively. The hypothesis $H_1(r)$ implies that the process ΔZ_t is stationary, Z_t is non-stationary, but $\beta' Z_t$ is stationary (Johansen, 1988 and 1991). There are three possible cases in this testing procedure. First, if $Rank\ (\Pi) = p$, the Π – matrix has full rank implying that the vector process Z_t is stationary i.e., all variables in the VAR are integrated of order zero. Second, if $Rank\ (\Pi) = 0$ (i.e., no independent co-integrating vector), Z_t is a VAR in first differences. Finally, if $(0 < rank\, \Pi = r < p)$, there are r co-integrating vectors.

There are three crucial parts in implementing the Johansen method of cointegration. First, as Boswijk and Franses (1992) point out, results of the VAR model are sensitive to the choice of lag-length. Therefore, before a long-run relationship can be sensibly estimated, it is important to determine the maximum lag-length that makes residuals to be approximately white noise. This paper uses Doornik-Hansen's (1994) modified version of Shenton-Bowman (1977) method for testing residual normality of each of the equation in the VAR system.

The second important part in estimating a long-run relationship with the Johansen method is the determination of cointegration rank. It is important because interpretation of the long-run relationship, as well as hypothesis testing, critically depend on the rank of Π matrix, such as the number of independent cointegrating vectors. For instance, if the cointegration rank is more than one, any linear combination of the cointegration relations can preserve the stationarity, leading to an identification problem (Hansen and Juselius1995).

Finally, selection of deterministic components is important because asymptotic distributions of the test statistics for determining cointegration rank critically depend on whether or not deterministic components (a constant and trend) enter the cointegration space. We use the so-called Pantula principles, presented in Johansen (1992), to resolve these problems by simultaneously determining the rank of Π and appropriate deterministic variables for the model. By examining plots of all variables in levels and first differences, two different models are considered. The first model allows a linear trend in data but no deterministic term in the cointegration space. The second model, on the other hand, includes both a linear trend in the variable and a non-zero intercept in the cointegration space.

We begin by estimating Boyce and Ravallion's full model using their data set (1949/50 to 1979/80) as well as Palmer-Jones's extended data series. To ensure that degrees of freedom remain close, we use k=1 for Boyce Ravallion data set and k=2 for Palmer-Jones data set. Table 2a and 2b presents the results of rank tests that determine number of co-integrating vectors. Now, Johansen's (1992) model selection methodology states that, starting from the hypothesis of zero co-integrating vectors and most restrictive model, model selection testing (based on $\lambda_{\max}$ and *Trace statistics*) should continue until the null is accepted. To see how this method works, read table 2a *row by row* and from *left to right*. For $r = 0$ (i.e., no co-integrating vector), $\lambda_{\max}$ statistics clearly reject the null hypothesis of no co-integrating relationship for both models. In the next row, ($i.e.,\ r = 1$), the hypothesis of one co-integrating vector is accepted for Model-1. Thus, on the basis of $\lambda_{\max}$**,** we conclude that there is one co-integrating vector

Table 2a. Determination of co-integration rank and selection of model for deterministic components (1949/50 to 1979/80, k=1)

Null Hypotheses	Model 1		Model 2	
	$\lambda_{\max}$ *stat*	$\lambda_{\max}(0.95)$	$\lambda_{\max}$ *stat*	$\lambda_{\max}(0.95)$
$\lambda_{\max}$ ***test:***				
$r=o$	41.91→	39.37	48.46	43.97
$r \leq 1$	31.69	33.46**	32.54	37.52
$r \leq 2$	17.42	27.07	20.60	31.46
$r \leq 3$	15.97	20.97	16.21	25.54
$r \leq 4$	10.35	14.07	14.33	18.96
$r \leq 5$	2.04	3.76	2.85	→12.25
Trace test:	Trace Stat	Trace (0.95)	Trace Stat	Trace (0.95)
$r=o$	119.38→	94.15	134.99	114.90
$r \leq 1$	77.47	68.52	86.53	87.31**
$r \leq 2$	45.77	47.21	53.99	62.99
$r \leq 3$	28.35	29.68	33.38	42.44
$r \leq 4$	12.13	15.41	17.18	25.32
$r \leq 5$	2.04	3.76	2.85	→12.25

* 95% Critical values are taken from Osterwald-Lenum (1992).
** Indicates that the null hypothesis of number of co-integrating vectors indicated by first column cannot be rejected for that particular model

Table 2b. Determination of co-integration rank and selection of model for deterministic components (1949/50 to 1979/89, k=2).

Null hypotheses	Model 1 $\lambda_{\max}$ *stat*	$\lambda_{\max}(0.95)$	Model 2 $\lambda_{\max}$ *stat*	$\lambda_{\max}(0.95)$
$\lambda_{\max}$ ***test:***				
$r=o$	41.07→	39.37	42.79	43.97
$r \leq 1$	35.79	33.46	36.19	37.52**
$r \leq 2$	22.23	27.07	25.72	31.46
$r \leq 3$	10.22	20.97	16.23	25.54
$r \leq 4$	8.13	14.07	9.45	18.96
$r \leq 5$	2.19	3.76	6.17	→12.25
	Trace Stat	*Trace* (0.95)	Trace Stat	*Trace* (0.95)
Trace test:				
$r=o$	119.62→	94.15	136.54	114.90
$r \leq 1$	78.56	68.52	93.75	87.31
$r \leq 2$	42.76	47.21**	57.56	62.99
$r \leq 3$	20.53	29.68	31.85	42.44
$r \leq 4$	10.32	15.41	15.62	25.32
$r \leq 5$	2.19	3.76	6.17	→12.25

* The critical values for $\lambda_{\max}$ and *trace* are taken from Osterwald-Lenum (1992).

** Indicates that the null hypothesis of number of co-integrating vectors indicated by first column cannot be rejected for that particular model

and that a model with a linear trend in data is the preferred model. The bottom panel of the table presents the *trace statistics*. Although this test suggests the same number of co-integrating vectors as $\lambda_{\max}$, it picks up the second model to be the preferred model. Given it has sharper alternative and better power properties (Johansen and Juselius 1990), we decided in favor of $\lambda_{\max}$ i.e., one co-integrating vector with Model 2 as the preferred model. By similar logic, the Palmer-Jones's data set also suggests one co-integrating vector with Model 2 being the preferred model[9]. After imposing appropriate weak exogeneity restrictions on loading parameters $(\alpha 's)$ and normalization by w^a, the following long-run relationships are estimated[10]:

1949/50 to 1979/80 data set

$$w^a = \underset{(2.84)}{0.42}\, p^r + \underset{(5.26)}{0.65}\, w^m + \underset{(1.06)}{0.14}\, p^c + \underset{(3.48)}{0.28}\, p^j - \underset{(7..11)}{2.51} Q^b \tag{7}$$

1949/50 to 1988/89 data set

$$w^a = \underset{(5.20)}{0.76}\, p^r + \underset{(5.00)}{0.0.45}\, w^m + \underset{(0.57)}{0.014}\, p^c + \underset{(0.11)}{0.042}\, p^j - \underset{(4..87)}{2.45} Q^b + \underset{(2.56)}{0.015}5 \tag{8}$$

where the figures in parenthesis are likelihood ratio test statistics, derived by imposing exclusion restriction on each of the variables, which follows a χ^2 distribution with one degree of freedom. There are three observations to make about these cointegration relationships. First, the estimated coefficient for Q^b is significantly negative for both equations, which is an economically nonsensical result. As we have already argued, this might be due to the fact that Q^b is a non-inflationary process and is not likely to form a positive co-integrating relationship with inflationary variables like nominal prices. Second, price of cloth (p^c), which was significant in Boyce-Ravallion's parsimonious preferred model, turns out to be insignificant in both equations. Finally, misspecification tests, presented in Table 3, indicate that neither model convincingly passes ARCH and normality tests on the residuals. Three out of six equations for 1949/50-1088/89 data and two out of six in 1949/50-1988/89 data fail a normality test on their residuals. More importantly, residual of the agricultural wage (w^a) equation, the most important equation in this context, fails a ARCH test for the first data set and normality test on the second data set. This implies that, at least from an

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[9] Note that while $\lambda_{\max}$ and *trace stat* give conflicting results. This is, however, not unusual in Johansen's method of cointegration. For a detail discussion on the subject, see Johansen (1991), Schotman and Dijk (1989) and references therein.
[10] Manufacturing wage rate (W_m), cloth price (P^c), and agricultural productivity (Q^b) were found to be weakly exogenous.

Table 3. Residual misspecification tests for equations 7 and 8. *

Equations	1949/50—1979/80 data		1949/50—1988/89 data	
	ARCH Test	Normality	ARCH Test	Normality
w^a	9.47	4.63	0.51	6.44
w^m	5.70	1.46	2.60	4.26
p^r	0.34	2.18	0.75	2.10
P^c	0.02	16.85	0.27	36.95
p^j	0.91	7.64	0.06	8.10
Q^b	3.14	1.33	3.06	0.62

* Test statistics for normality should be compared with $\chi^2_{(2)}$, the critical value of which is 5.99 at 5% level of significance. For ARCH test, degrees of freedom for χ^2 is equal to the number of lags used in each model. For example, tests statistics for 1949/59—1979/80 and 1949/50—1979/80 are compared with $\chi^2_{(1)}$ and $\chi^2_{(2)}$ respectively.

econometric point of view, earlier models do not provide adequate representation for the dynamic of agricultural wages.

Based on the above analyses, deriving the parsimonious model was not difficult. We first dropped Q^b and p^c from the analyses and carried out a rank test for co-integrating relationship on the rest of the variables. By imposing exclusion restrictions on each of these variables, p^j could be excluded to leave us with a model that consists of agricultural wage rate (w^a), rice price (p^r), and manufacturing wage rate (w^m). Rank test results for this set of variables are presented in Table 4a and 4b. Based on these results, our model selection remains the same i.e., Model 1 for the first data set and Model 2 for the second

Table 4. Rank tests on parsimonious model (1949/50 to 1988/89, k=2)

	Null Hypotheses	Model 1 λ_{max} *stat*	$\lambda_{max}(0.95)$	Model 2 λ_{max} *stat*	$\lambda_{max}(0.95)$
λ_{max} *test:*					
	$r=o$	22.04	20.97	28.58	25.54
	$r \leq 1$	17.33	14.07	17.51	18.96**
	$r \leq 2$	1.50	3.76	6.63	→ 12.25
		Trace Stat	*Trace* (0.95)	Trace Stat	*Trace* (0.95)
Trace test:					
	$r=o$	40.87	29.68	47.73	42.44
	$r \leq 1$	18.83	15.41	24.15	25.32**
	$r \leq 2$	1.50	3.76	6.63	→12.25

* 95% Critical values are taken from Osterwald-Lenum (1992).
** Indicates that the null hypothesis of number of co-integrating vectors indicated by first column cannot be rejected for that particular model

Table 4b. Rank tests on parsimonious model (1949/50 to 1979/80, k=2)

	Null Hypotheses	Model 1 λ_{max} *stat*	$\lambda_{max}(0.95)$	Model 2 λ_{max} *stat*	$\lambda_{max}(0.95)$
λ_{max} *Test:*					
	$r=o$	21.32	20.97	21.33	25.54
	$r \leq 1$	12.99	14.07**	12.99	18.96
	$r \leq 2$	4.37	3.76	6.08	→ 12.25
		Trace Stat	*Trace* (0.95)	Trace Stat	*Trace* (0.95)
Trace test:					
	$r=o$	38.58	29.68	40.40	42.44
	$r \leq 1$	17.26	15.41	19.08	25.32**
	$r \leq 2$	4.27	3.76	6.08	→12.25

* 95% Critical values are taken from Osterwald-Lenum (1992).
** Indicates that the null hypothesis of number of co-integrating vectors indicated by first column cannot be rejected for that particular model

data set. After imposing weak exogeneity restrictions on insignificant loading parameters, the following long-run models were estimated:

1949/50 to 19879/80 Data set

$$w^a = \underset{(8.68)}{0.72}\, p^r + \underset{(3.55)}{0.21} w^m \tag{9}$$

1949/50-1988/89 Data set

$$w^a = \underset{(5.91)}{0.70} p^r + \underset{(2.72)}{0.42} w^m - \underset{(1.74)}{0.13} \tag{10}$$

Notice that in terms of the relationship between agricultural wage and rice price, the two equations are not substantially different. Moreover, when we restrict the constant term in the equation 10 to zero, two equations become even more similar to each other with coefficients of w^a and p^r being 0.69 and 0.31 respectively[11]. That is, with this restriction, equation 10 can now be written as,

$$w^a = \underset{(22.56)}{0.69} p^r + \underset{(7.15)}{0.31} w^m \tag{11}$$

Residual misspecification tests for these equations are presented in Table 5. Note that unlike the full models in equation 7 and 8, these equations pass both normality and ARCH tests on their residuals. Therefore, we consider these to be the preferred models, on which we carry out further analyses. How do these results differ from previous studies? One of the striking differences is that our estimates of long run elasticities of agricultural wage rate with respect to rice price are much larger than those reported in the past studies. In particular, long

[11] The likelihood Ratio test for this exclusion along with two weak exogneity restrictions gives a χ^2_3 value of 4.98, implying that the hypothesis of zero restriction cannot be rejected.

Table 5. Residual misspecification tests on equations 9 and 11

Equations	1949/50—1979/80 data		1949/50—11988/89 data	
	ARCH Test	Normality	ARCH Test	Normality
w^a	0.176	5.92	0.32	5.34
w^m	0.835	0.75	1.87	1.75
p^r	0.540	4.50	0.62	5.58

All test statistics follow χ^2 distribution with two degrees of freedom, the 5% critical value of which is 5.99.

run elasticity of 0.72 in equation 9 compares to 0.47 in Boyce-Ravallion (1991) and 0.69 in equation 11 compares to 0.46 of Palmer-Jones (1993).

4.3 Hypothesis Testing

To further explore the co-integrating relationship between agricultural wage and rice price, we test two hypotheses on equation 9 and 10. The first hypothesis is motivated by the theoretical finding of Sah and Stiglitz (1987) that agricultural wage rate in a poor rural economy responds proportionally to the price of the staple food. In our analytical framework, this hypothesis means that variables w^a, p^r, and w^m enter the co-integrating relation as $(a, -a, *, *)$; where asterisks mean that coefficients on w^m and the constant are left unrestricted. That is, with weak exogeneity, parameter restrictions can be written as: $(\alpha_{wa} = \alpha_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$. Following Johansen and Juselius (1992), this hypothesis can be formulated as follow:

$$H_1 = \begin{bmatrix} 1 & 0 & 0 \\ -1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix}$$

The above hypothesis is tested using a Likelihood Ratio test, where eigenvalues of the full model are compared with the eigenvalues of the restricted model. Test results are presented in Table 6. According to these results, the

hypothesis cannot be rejected at a conventional level of significance, implying that agricultural wage rate and rice price maintained a proportional relationship during the time period considered. For the time period 1949/50—1979/80 and 1949/50-1988/89, the Likelihood Ratio test yields χ^2_3 values of 4.04 and 4.65 respectively, which are compared with a critical value of 7.81 at 5% level of significance[12].

Table 6. Test of proportionality and stationarity of agricultural wage rate and rice price

Parameter Restrictions	LR test statistics	P-Value
1949/50—1979/80 data		
$\alpha_{wa} = \alpha_{wm} = 0$	χ^2_2 = 0.51	0.78
$(\alpha_{wa} = \alpha_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$	χ^2_3 = 4.04	0.26
$(\alpha_{wa} = \alpha_{wm} = \beta_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$	χ^2_4 = 5.46	0.24
1949/50—1988/89		
$\alpha_{wa} = \alpha_{wm} = 0$	χ^2_2 = 1.74	0.42
$(\alpha_{wa} = \alpha_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$	χ^2_3 = 4.65	0.20
$(\alpha_{wa} = \alpha_{wm} = \beta_0 = \beta_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$	χ^2_5 = 9.29	0.10

[12] Note that degree of freedom is 3 because in addition to proportionality restriction, there are two weak exogeneity restrictions on the loading parameters in equations 9 and 10.

The second hypothesis we test on the relationship of agricultural wage and rice price is more restrictive. In addition to proportionality, it restricts all other parameters to zero. In other words, this hypothesis tests whether agricultural wage and rice price maintained a stationary relationship. This hypothesis can be formulated as $H_2 = \begin{bmatrix}1 & -1 & 0 & 0\end{bmatrix}'$, with $(\alpha_{wa} = \alpha_{wm} = \beta_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$ as parameter restrictions[13]. Results of the Likelihood Ratio test, reported in Table 6, suggest that while the hypothesis can be rejected for the 1949/50-1988/89 data at a 10% level of significance, it holds true for the 1949/50-1979/80 data. Given agricultural wages increased faster than rice price in the 1980s, these results are not surprising. In fact, available studies suggest that real agricultural wage relative to real rice price in Bangladesh increased throughout the 80s and 90s (Dorosh 2000). Two major policy events contributed to this process. First, the government of Bangladesh implemented several input market liberalization policies, which boosted agricultural production through higher rate of HYV adoption in this decade (Ahmed 1998). Second, by encouraging foreign direct investment, the country experienced substantial growth in Garment Industries during the 80s. Since the garment factories are labor intensive, this triggered massive migration of rural workers to major cities. This result is also consistent with previous finding of Palmer-Jones (1993) that the Boyce and Ravallion (1991) model failed a prediction test for the extended time series.

4.4 Short-run dynamics of agricultural wages

The co-integration analysis explains the existence and nature of long-run relationships among the economic variables. It does not provide any explanation as to how these variables adjust in the short run. This section analyzes the short-

[13] Note that this hypothesis is similar to testing whether real agricultural wage in terms of rice price followed a random walk process. In other words, we are testing whether $w_t^a = p_t^r + \varepsilon_t$ holds true. I would like to thank professor Salim Rashid of the University of Illinois for suggesting to carry out this test.

run adjustment behavior of the variables used in equations 9 and 11. Following Engle and Granger (1987), the general structure of the short-run dynamics in error correction framework can be represented as follows:

$$\Delta w_t^a = \alpha_0 + \sum_{i=1}^{k} \alpha_{1i} \Delta w_{t-i}^a + \sum_{i=1}^{k} \alpha_{2i} \Delta w_{t-i}^m + \sum_{i=1}^{k} \alpha_{3i} \Delta p_{t-i}^r - \lambda \left(w_{t-1}^a - \hat{\beta}_1 w_{t-1}^m - \hat{\beta}_2 p_{t-1}^r \right) + \varepsilon_t \quad (12)$$

where ε_t is white noise, k is lag length, and the α_i's are *Speed of Adjustment*. The expression within the bracket represents the error correction mechanism through which deviations of wage rate from its long-run equilibrium level gets adjusted in the subsequent periods[14]. For example, if wage rate falls below its long-run level at time *t-1*, the term within the bracket will become negative. However, since λ is negative, the resulting consequence of this deviation will be an increase in the observed wage rate at time *t*. Similarly, if agricultural wage rate increases beyond the long-run equilibrium level, the error-correction mechanism will pull it down in the subsequent period.

Given unspecified lag length, and hence a large number of right hand side variables, the immediate task in estimating the equation (12) is to derive a parsimonious model that adequately explains short-run dynamics in a given equation. We adopted Hendry's *general to specific modeling technique*, suggested in Hendry, Pagan, and Sargan (1984), to achieve such parsimony and to derive the final estimating model. Under this technique, estimation begins with a set of economically plausible explanatory variables, and then simplifications are made by excluding the variable through model selection tests. Since all variables in equation (12) are *I(0),* Ordinary Least Square method is efficient and statistical inferences based on t- and F- test are valid. Therefore, the model was estimated by OLS and the parsimonious model was derived by imposing sequential

[14] Error Correction term is computed using the parameter estimates of equation 9 and 11.

restrictions on estimated parameters. Each restriction was tested by F-test against a slightly less restrictive model preceding it in the sequence. Finally, a series of diagnostic tests are performed to ensure the adequacy of the final model. We started out with a general model that included a constant, two lags of each of the explanatory variables and a time trend. However, on the basis of model selection tests, second lags of all explanatory variables, except w^m for 1949/50-1988/89 data set could be convincingly excluded from the models. Through further exclusion of insignificant variables, the following parsimonious models were derived for the two data sets used in the analyses (*t-statistics* in parenthesis).

1949/50-1979/80 Data

$$\Delta w_t^a = \underset{(3.54)}{0.08} + \underset{(3.38)}{0.53}\Delta w_t^m + \underset{(7.52)}{0.32}\Delta p_t^r - \underset{(4.96)}{0.48}\Delta EC_{t-1} \tag{13}$$

$$\overline{R}^2 = 0.76;\ SSE = 0.082;\ DW = 1.73;$$
$$\text{LM Test}: AR(1) = 1.09;\ AR(2) = 2.43$$

1949/50-1988/89 DATA

$$\Delta w_t^a = \underset{(2.54)}{0.07} + \underset{(2.42)}{0.31}\Delta w_{t-1}^a + \underset{(2.50)}{0.39}\Delta w_t^m - \underset{(1.67)}{0.31}\Delta w_{t-1}^m + \underset{(5.55)}{0.25}\Delta p^r - \underset{(1.79)}{0.11}\Delta p^r - \underset{(3.9)}{0.44}\Delta EC_{t-1} \tag{14}$$

$$\overline{R}^2 = 0.69;\ SSE = 0.062;\ DW = 1.80;$$
$$\text{LM Test}: AR(1) = 1.84;\ AR(2) = 2.35$$

Two results of short-run analyses are worth noting. First, with estimated coefficients of 0.48 and 0.44 respectively, error correction terms are significant for both data sets. This implies that a substantial fraction of the deviation from long-run equilibrium at a given period gets adjusted during the next period. Second, although larger than previously reported, short-run elasticity of agricultural wage rate with respect of rice price is still low[15]. These figures are 0.32 and 0.25 for 1949/50-1979/80 and 1949/50-1988/89 data respectively, which compare to 0.22 for both Boyce Ravallion (1991) and Palmer-Jones (1993).

[15] Note that our estimates are not strictly comparable to short-run elasticities of previous studies. For example, short-run elasticity of Boyce and Ravallion (1991) can be interpreted either as $\partial w_t^a / \partial p_t^r$ or $\partial \Delta w_t^a / \partial p_t^r$, none of which are identical to our estimates i.e., a measure of the speed of adjustment.

5. ANALYSES OF POST-FAMINE DATA

The main motivation for undertaking an analysis of the post-famine data (1976/77-1998/99) has been to assess how policy reforms of the last two decades affected dynamics of the macroeconomic relationships discussed in the previous section. Following the analyses of the previous section, a long run relationship is assumed to exist among agricultural wage rate, rice price, and urban wage rate. Stationarity test results, presented in Table 7, suggest that all variables are *I(1).* Therefore, Johansen's model selection principle was used to determine the number of co-integrating vectors and deterministic terms in the model. These results are presented in Table 8. While $\lambda_{\max}$ selected a model with linear trend in data, trace *statistics* suggests the model with a trend in the cointegrating space. Both tests, however, suggest that there is only one co-integrating relationship and misspecification test results, presented in Table 9, indicate that residuals are normal. For comparison, we present both models below:

$$w^a = \underset{(11.6)}{0.75}w^u - \underset{(3.19)}{0.86}p^r \qquad (15)$$

$$w^a = \underset{(11.58)}{1.51}w^u - \underset{(2.72)}{0.73}p^r + \underset{(0.74)}{0.012}t \qquad (16)$$

These results are in sharp contrast with the previous section. Rice price, which formed such a strong co-integrating relationship in the past, now turns out to be insignificant, implying that there has been a reversal in the dynamics of agricultural wages in Bangladesh during the last twenty-five years. The most important variables in the co-integration space for this period are agricultural wage rate and urban wage rate.

Table 7. Stationarity tests on 1977/78—1998/99 data

	Levels		First difference	
Series	ADF	KPSS	ADF	KPSS
w^a	-0.13	0.76	-4.45	0.42
w^u	-1.55	0.77	-3.74	0.44
p^r	-2.22	0.60	-6.61	0.22
Q^b	-0.79	1.08	-4.67	0.29

[a] All variables are in natural logarithm. Note that while ADF tests $H_0 : X_i \sim I(1)$ against $H_1 : X_i \sim I(0)$, KPSS method tests the reverse. The ADF and KPSS critical values at 5% level of significance are -3.07 and 0.463 respectively. Critical value for ADF test is taken from Mackinnon (1991).

Table 8. Rank tests on 1977/8—1998/99 data

		Model 1		Model 2	
	Null Hypotheses	$\lambda_{\max}$ *stat*	$\lambda_{\max}(0.95)$	$\lambda_{\max}$ *stat*	$\lambda_{\max}(0.95)$
$\lambda_{\max}$ ***test:***					
	$r=o$	26.30→	20.97	26.77	25.54
	$r \leq 1$	11.17	14.07**	11.25	18.96
	$r \leq 2$	5.73	3.76	5.89	→ 12.25
Trace test:		Trace Stat	*Trace* (0.95)	Trace Stat	*Trace* (0.95)
	$r=o$	43.20	29.68	43.91	42.44
	$r \leq 1$	16.90	15.41	17.14	25.32**
	$r \leq 2$	5.73	3.76	5.89	→12.25

* 95% Critical values are taken from Osterwald-Lenum (1992).

** Indicates that the null hypothesis of number of co-integrating vectors indicated by first column cannot be rejected for that particular model

To further explore the relationship between agricultural wage and rice price, we carried out proportionality and stationarity testing on equation 16. Results of these tests are presented in Table 10. Note that, with χ_2^2 and χ_3^2 values of 6.49 and 9.75 respectively, both hypotheses are strongly rejected for the data set. These results are consistent with recent trends in labor markets of Bangladesh, particularly in the major cities. Due to higher wage rate and other employment opportunities, the country has experienced unprecedented growth in its urban population during the last two decades. For example, according to 1991 census, the population of Dhaka grew by almost 39 percent and Chittagong by more than 26 percent.

What policy implications do these results have for the country? There are at least two clear implications for public policies. First, since labor outlays are a significant proportion of total cost of rice production, an increase in agricultural wage relative to rice price implies a decline in farmers' profitability. Therefore, there should either be a shift from rice production to other high crops that farmers can profitably grow. However, such a shift in production behavior will depend on a number of different factors, including infrastructure, access to market information, and institutions for production and marketing risk management. Currently, these institutions are either weak or do not exist in Bangladesh. There is an ongoing project trying to promote agribusiness and agro-based industries in the country, but it does not provide a clear policy or an institutional framework, portraying that such a project can be successful. For example, agro-industries and agribusiness development requires some basic infrastructure—such as, processing, grading and safety standards—which are generally considered to be public goods in the developing countries. Therefore, there is a clear role for the government in order to make agriculture profitable, and to generate employment in rural areas.

Table 9. Residual misspecification tests on equation 16

Equations	1977/8—1998/99 data	
	ARCH Test	Normality
w^a	0.109	4.62
w^m	0.376	3.48
p^r	0.231	1.86

*Test statistics for ARCH and normality follow χ_1^2 and χ_2^2, the 5% critical values of which are 3.84 and 5.99 respectively.

Table 10. Test of proportionality and stationarity agricultural wage rate and urban wage rate.

Parameter Restrictions	LR test statistics	P-Value
1977/78—1998/99 data		
$\alpha_{rice} = 0$	χ_1^2 = 0.00	0.98
$(\alpha_{rice} = 0 \cap (\beta_{agwage} = \beta_{uwage}))$	χ_2^2 = 6.49	0.04
$(\alpha_{rice} = \beta_{wm} = 0) \cap (\beta_{wa} = \beta_{pr})$	χ_3^2 = 5.46	0.02

Second, our result of a non-stationary relationship between agricultural and urban wage rates indicate that two wages are diverging, suggesting that major cities will continue to attract rural workers. Following the work of Fei and Ranis (1960), it is anticipated that rural-urban migration will take place during the process of development. However, urbanization with high pollution posing a threat to human health is not socially desirable. In our view, the government has a clear role in halting this unbalanced growth of urbanization. In particular, by building the necessary infrastructure in smaller towns, some industries can be relocated from major cities. As a simple example, consider how a typical garment factory in Dhaka operates. After signing a contract, a foreign company ships specified merchandise to Chittagong seaport, which is then transported to Dhaka for stitching and packaging. After stitching and packaging, garment factories in Dhaka ship the finished product back to Chittagong to be exported. There are number of small towns between these two cities, where there are no garment factories due to the lack of necessary infrastructure that would make investment in a garment factory profitable. If such infrastructure is considered a public good, public expenditure is clearly justifiable to build these infrastructures, which have the potential to reverse the current trend of urbanization. African experience suggests that such reversal is possible. For example, Jaeger (1992) showed that narrowing rural-urban wage gap resulted in a significant reverse migration in Ghana.

6. CONCLUSIONS

The main objective of this study has been to examine the dynamics of agricultural wage formation in Bangladesh, paying close attention to its relationship with rice prices. Two specific tasks have been carried out. First, given the non-stationarity of all variables, data from past studies are re-examined in multivariate co-integration framework. Second, to assess whether policy initiatives of the 80s and 90s reflected through changes in the dynamics of wages, similar analyses have also been carried out on the post-famine (1977/78-1998/99) data. Our results are striking on both counts.

By re-analyzing the data used in previous studies, this paper has demonstrated that previously drawn conclusions do not hold when the same models are estimated in cointegration framework. In particular, contrary to previously drawn conclusions of alarming downward trends in agricultural wages relative to rice price, the results suggest that agricultural wage and rice price maintained a strong co-integrating relationship during the 1949/50-1988/89 period. The long run elasticities of agricultural wage rate with respect to rice price are estimated to be 0.72 and 0.69 for Boyce-Ravallion (1991) and Palmer-Jones (1993) data sets, which are substantially higher than their estimates of 0.46 and 0.47 respectively. To further explore the degree of co-integration between agricultural wage rate and rice price, we carried out two structural hypotheses tests—one on the proportionality and the other on the stationary relationship between the two variables. Neither hypothesis could be rejected at a conventional level of significance for the 1949/50-1979/80 data. This suggests that, under our analytical framework, previous conclusions of a decline in real agricultural wage rate with respect to rice price cannot be supported by the data. For the 1949/50-1988/89 data, however, stationarity hypothesis was rejected, which can perhaps be explained by a higher growth of agricultural wages relative to rice prices during the 80s.

Analyses of the post-famine data suggest that there has been a dramatic change in the dynamics of agricultural wage and rice price. Rice price, a variable that maintained strong relationship with agricultural wage until the late 80s, turns out to be insignificant in the wage equation estimated for the 1977/78 to 1998/99-time period. Urban wage rate, measured as wage rate of unskilled construction workers in major cities, is found to be the most significant variable to explain the dynamics of agricultural wage. This is consistent with the fact that real rice price in Bangladesh has been declining during the last two decades. The results of this study also suggest that relative to agricultural wage rate, urban wage rate has been increasing, implying that unless deliberate policies for rural employment are undertaken, the unprecedented growth of urban population is likely to continue.

REFERENCES

Ahmed, A. and Y. Shams. (1996). 'Demand Elasticities in Rural Bangladesh: An Application of AIDS Model,' *Bangladesh Development Studies,* 22(1), pp. 1-25.

Ahmed, I. (1981). 'Wage Determination in Bangladesh Agriculture', *Oxford Economic Papers,* Vol. 33, pp. 298—323.

Ahmed, R. (1995). 'Liberalization of Agricultural input markets in Bangladesh: Process, Impacts, and Lessons.' *Agricultural Economics*, 12: 115-128.

Ahmed, R., Haggblade, S., and Chowdhury, T.E. (eds.) (2000). ' *Out of the Shadow of Famine: Evolving Food Markets and Food Policy in Bangladesh'*, John Hopkins University, Baltimore.

Banarjee, A.D., Rayner, A.J., and D. Greenway (1993). *Cointegration, Error Correction, and Econometric Analysis of Non-Stationary Data*. Oxford University Press, New York.

Bangladesh Bureau of Statistics (Various Years). *Statistical Year Book of Bangladesh*, Bangladesh Bureau of Statistics, Dhaka.

Bose, S.R. (1974). 'Movement of Agricultural Wages in Bangladesh 1949-73', in Mitra, A. (ed.), *Economic Theory and Planning*, Oxford University Press, Calcutta.

Boswijk, H.P., and P.H. Frances (1992). 'Dynamic Specification and Cointegration', *Oxford Bulletin of Economics and Statistics* 54:369-381.

Boyce, J. (1989). 'Population Growth and Real Wages of Agricultural Labourers in Bangladesh', *Journal of Development Studies*, Vol. 25, pp. 467—89.

Boyce, J. and Ravallion, M. (1991). 'A Dynamic Econometric Model of Agricultural Wage Determination in Bangladesh', *Oxford Bulletin of Economics and Statistics,* Vol. 53, pp. 361—376.

Dickey, D.A., Jansen, D.W., and Thorton, D.L. (1994). 'A Primer on Cointegration with an Application to Money and Income' in Rao, B. (ed.) *Cointegration for Applied Economists,* St. Martin Press, New York.

Dorosh, P. (2000). 'Foodgrain Production and Imports: 'Towards Self-sufficiency in Rice?' in Ahmed et al (eds.) ' *Out of the Shadow of Famine: Evolving Food Markets and Food Policy in Bangladesh'*, John Hopkins University, Baltimore.

Doornik, J.A. and Hansen, H. (1994). "An Omnibus Test for Univariate and Multivariate Normality", Working Paper, Nuffield College, Oxford.

Enders, W. (1995). *Applied Econometric Time Series,* John Wiley & Sons, Inc., New York.

Engle, R.F. and Granger, C.W.J. (1984). 'Co-integration and Error Correction: Representation, Estimation, and Testing', *Econometrica,* 55: 251-276.

Goletti, F. (1994). *The Changing Public Role in a Rice Economy Approaching Self-Sufficiency: The Case of Bangladesh,* IFPRI Research Report No. 98. International Food Policy Research Institute, Washington DC.

Granger, C.W.J. (1986). 'Development in the Study of Cointegrated Economic Variables,' *Oxford Bulletin of Economics and Statistics,* 48(3): 213-228.

Granger, C.W.J. and Newbold, P. (1974). 'Spurious Regression in Econometrics', *Journal of Econometrics,* 2:11-20

Haggblade, S., Hazell P. (1989). ' Agricultural Technology and Farm Non-Farm Growth Linkages.' *Agricultural Economics*, 3:345-364.

Hansen, H. and Juselius, K. (1995). *Cats in Rats: Cointegration Analysis of Time Series*, Estima Inc., Evanston, Illinois.

Hendry, D.F., Pagan, A.R. and Sargan, J.D. (1984). 'Dynamic Specification', in Griliches, Z. and Intriligator, M.D. (eds.), *Handbook of Econometrics*, Vol. 2, North Holland, Amsterdam.

Hossain, A. (1990). 'Real Wage Determination in Bangladesh Agriculture: An Econometric Investigation', *Applied Economics*, 22: 1549-1565.

Hossain, M. (1988). *Natures and Impact of Green Revolution in Bangladsh*. IFPRI Research Report No. 67. International Food Policy Research Institute. Washington, D.C.

Johansen, S. (1988). 'Statistical Analysis of Cointegration Vectors', *Journal of Economic Dynamic and Control*, Vol. 12, pp. 231—254.

Johansen, S. (1992). ‘Determination of Cointegration Rank in the Presence of a Linear Trend’, *Oxford Bulletin of Economics and Statistics,* Vol. 54, No. 3, pp.383—397.

Johansen, S. and Juselius, K. (1990). ‘Maximum Likelihood Estimation and Inference on Cointegration—With Applications to The Demand for Money’, *Oxford Bulletin of Economics and Statistics,* Vol. 52, No. 2, pp. 169—210.

Johansen, S. and Juselius, K. (1992). ‘Testing Structural Hypothesis in a Multivariate Cointegration Analysis of the PPP and UIP for UK’, *Journal of Econometrics*, Vol. 53, pp. 211—244.

Kwiatkowski, D., Phillips, C.B. Schmidt, P. and Shin, Y. (1992). “Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure are We that Economic Time Series Have a Unit Root?” *Journal of Econometrics*, 54:159-178. Mellor, J.W. and Desai, G.M. (eds.) (1985). *Agricultural Change and Rural Poverty: Variations on a Theme by Dharm Narain,* John Hopkins University Press, Baltimore.

Osterwald-Lenum, M. (1992). “A Note on the Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test.” *Oxford Bulletin of Economics and Statistics* 54:461-472,

Palmer—Jones, R.W. (1993). ‘ Agricultural Wages in Bangladesh: What the Figures Really Show?’ *Journal of Development Studies*, Vol. 29, pp. 277—300.

Palmer—Jones, R.W. (1994). ‘ An Error Corrected? And What the Figures Really Show’, *Journal of Development Studies*, Vol. 31, No.2, pp. 346—351.

Perron, P. (1989). ‘The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis’. *Econometrica*, 102:109-134.

Ravallion, M. (1985). *Markets and Famines*, Oxford University Press, Oxford.

Ravallion, M. (1990). ‘Rural Welfare Effects of Food Price Changes Under Induced Wage Responses: Theory and Evidence from Bangladesh’, *Oxford Economic Papers*, Vol. 42, pp.574—85.

Ravallion, M. (1994). 'Agricultural Wages in Bangladesh: What do the Figures Really Show? *Journal of Development Studies,* Vol. 31, No. 2, pp. 334—345.

Sah, R. and Stiglitz, J.E (1987). 'Price Scissors and Structure of the Economy'*, Quarterly Journal of Economics*, Vol. 102, pp. 109—34.

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