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DYNAMICS OF AGRICULTURAL WAGES AND RICE PRICES
IN THE PHILIPPINES

presented by

MARIE-CHRISTINE D. LASCO

has been accepted towards fulfillment
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M.S. degree in AGRICULTURAL ECONOMICS

Richard N. Bernsten
Major Professor's Signature

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DYNAMICS OF AGRICULTURAL WAGES AND RICE PRICES IN THE PHILIPPINES

By

Marie-Christine D. Lasco

A THESIS

Submitted to
Michigan State University
in partial fulfillment of the requirements
for the degree of

MASTER OF SCIENCE

Department of Agricultural Economics

2005

ABSTRACT

DYNAMICS OF AGRICULTURAL WAGES AND RICE PRICES
IN THE PHILIPPINES

By

Marie-Christine D. Lasco

Rice trade liberalization in the Philippines will likely cause domestic rice prices to decrease. This may have a significant impact on agricultural wages and agricultural wage earners. This study examines the long-run (LR) and short-run (SR) relationship of agricultural wages and rice prices in the Philippines using a neoclassical wage determination model. Two frameworks are used in the analysis: A co-integration/error correction framework which assumes nonstationarity of model variables and an OLS framework that assumes that the model variables are stationary. The results show that in the short run, wages adjust partially with a SR elasticity of 0.33-0.41, while in the long run the adjustment is greater. An analysis of welfare implications for various groups of agricultural wage earners suggest that most households will benefit from a decrease in rice prices, although immediate mitigation measures are needed by adversely affected households that are heavily reliant on agricultural wages for income.

ACKNOWLEDGEMENTS

This project would not have been possible without the following people: Dr. Richard Bernsten (Major Professor), Dr. Robert Myers (Thesis Advisor), Dr. Eric Crawford and Dr. John Giles of Michigan State University; Dr. David Dawe and the staff of the Social Sciences Division at the International Rice Research Institute and the very helpful staff of the Bureau of Agricultural Statistics.

I am very grateful for my professors and colleagues at the Department of Agricultural Economics at Michigan State University, and my friends Cathy, Courtney, Cristobal, Esteban, Josh, Kyunghyun, Kirimi, Neslie, Ricardo, and Valeria for all their support and encouragement.

I would also like to thank my family for their constant love and encouragement. Most importantly, I would like to thank God, who gave me gave me strength and enabled to finish this project, and blessed me with friends and experiences that I will always cherish.

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1. Introduction

The relationship between agricultural wages and staple food prices is an important empirical issue. Agricultural wages influence rural welfare, especially the welfare of the poorest groups who rely heavily on these wages for income. In the Philippines, rice is the staple food of 80% of the population, and rice farming is the largest agricultural sector. Thus, it is very likely that changes in rice prices will have a significant impact on agricultural wages. Rice prices in the Philippines have been largely determined by the government, so it is important to inform policymakers about the impacts of rice price movements on agricultural wages. This issue becomes even more relevant because of the country's commitment to the WTO to liberalize rice trade. Under the Agreement on Agriculture, all Quantitative Restrictions (QRs) on rice imports have to be removed and tariffied. Reduction in tariffs is also expected to follow. Because domestic rice prices are significantly higher than world prices, rice trade liberalization is expected to cause a decrease in the price of rice. The removal of QRs was supposed to have taken effect on January 1, 1995. However, the Philippines, along with Japan and South Korea has been granted concessions to extend this deadline for at least 10 years (Cororaton, 2004, David, 1997). As of August 2005, the Philippines has yet to implement the removal of QRs. The government's hesitation to open up rice trade may be caused in part by the perceived negative impact to rice producers and agricultural laborers, although it is clear that urban consumers will benefit from the decrease in rice prices. In order to inform this policy decision, it is important to analyze of the impacts of a decrease in rice prices to various sectors in the economy. This study will focus on the effect of a decrease in rice prices on agricultural wage earners by looking at how changes in rice prices affect the agricultural wage rate.

There is some contention regarding the direction and magnitude of wage response to changes in rice prices in the Philippines. Several studies in Bangladesh have found a positive relationship between agricultural wages and rice prices (Boyce and Ravallion, 1991, Ravallion, 1994, Palmer-Jones, 1993, Palmer-Jones and Parikh, 1998). The general hypothesis of these studies is that the relationship is positive with partial adjustment in the short run and full adjustment in the long run. The reason for this is that an increase in price will generally increase incentives for farmers to engage in rice production. This could lead to a greater demand for factor inputs, including labor, which causes wages to increase. Similarly, if prices decrease then this may lead to a decrease in factor demand and wages. However, there are also some studies which show that agricultural wages and rice prices have no relationship or a negative relationship. A recent study in Bangladesh by Rashid (2002) found that agricultural wages and rice prices have no long run relationship. In the Philippines, Dawe (2003) argued that a decrease in the price of rice will exert upward, rather than downward pressure on agricultural wages. He said that a decrease in rice prices will not decrease demand for labor in the agricultural sector because farmers will diversify to other crops such as vegetables. Since vegetables are more labor intensive than rice, there may even be a higher demand for labor and which may lead to increased wages. Moreover, he contends that the effect of a decrease in rice prices will primarily affect land rents (as opposed to wages and prices of tradable inputs) received by landowners of rice farmlands. This is because farmers have little control over prices of tradable inputs like fertilizer and pesticides which are imported. Thus, the adjustments must be absorbed by either land rents or agricultural wages. Since land is the more inelastic input, it will be affected more than wages.

This study informs this debate by explicitly looking at the relationship between agricultural wages and rice prices, both in the long-run and the short-run. The developing country literature on the subject is limited and no previous empirical study has been conducted to measure agricultural wage response to changes in rice prices in the Philippines.

This study addresses the general research question of how agricultural wages are affected by changes in the price of rice. Specifically, this study 1) tests whether there is a long-run relationship between the agricultural wage rate and the price of rice in the Philippines and 2) estimates the long-run and short-run elasticities of the wage rate with respect to the price of rice. Knowledge of short-run and long-run agricultural wage responses will provide insight into the direction, magnitude and persistence of the wage adjustment. This will aid in the analysis of the impacts of rice trade liberalization on the Philippines' rural sector. In this paper, the empirical findings are used to assess welfare implications of a decrease in rice price to agricultural wage earners in the Philippines who are net suppliers of agricultural labor.

2. Empirical studies on wage determination and agricultural wage response

Although the literature on wage response to changes in rice prices in the Philippines is limited, several studies have sought to measure wage response to changes in rice or other food prices in other developing countries. Boyce and Ravallion (1991) (hereafter BR), conducted a study on wage determination in Bangladesh using annual data from 1949/50 to 1980/81. BR assumed a long-run equilibrium relationship among the nominal agricultural wage and the nominal prices of rice, jute and corn, the manufacturing wage, and a productivity index. From this they estimated an error correction model (ECM) to analyze the short-run (SR) wage adjustment. They found that wages responded significantly to rice price in the SR with an elasticity of 0.22. Based on the SR error correction model, BR derived a model for the long-run (LR) wage adjustment mechanism and reported a LR elasticity of 0.47 for the nominal agricultural wage and 0.53 for wage deflated by the price of rice.

Palmer-Jones (1993), however, argued that the previous BR model failed prediction and stability tests. Extending the BR data set to 1949/50 to 1989/90, he proposed a different model wherein nominal agricultural wage is a function of its own lagged value, nominal prices of rice and (lagged) jute and (lagged) manufacturing wages. Moreover, he added a dummy variable for the period 1972-1974[1] and an ad-hoc time trend which takes on the value of time from 1949-1950 to 1964, and zero onwards. In a rebuttal, Ravallion (1994) criticized Palmer-Jones' modeling technique, claiming that the model is not homogenous, i.e., the implicit long run real wage rate depends on real and nominal variables. Furthermore, he questioned the use of a sub-period dummy and a half-time trend because this imposes a model structure that is not appropriate for the case of Bangladesh. Interestingly, Palmer-Jones arrived at a similar result (0.22) for the short-run elasticity of agricultural wages with respect to rice price. Moreover, the long-run elasticity reported (0.46) was also close to that of BR's.

New developments in econometric time series analysis have shown that many

[1] According to Palmer Jones, this was a period of great disruption and inflation following independence and the associated national and political disturbances in Bangladesh.

macroeconomic data, such as income and prices, appear to be nonstationary. If variables are nonstationary and not co-integrated, classical regression, as used by Palmer-Jones, leads to spurious relationships among variables and inconsistency of parameter estimates (Pindyck and Rubinfeld, 1998). In the case of the BR study, a co-integrating relationship must first be established among the variables before an ECM can be used. Techniques based on a co-integration framework have been developed to analyze the relationships among nonstationary time series variables. More recent studies have adopted these methodologies.

Rashid (2002) re-analyzed the first two studies by BR and Palmer-Jones, using a co-integration framework. Rashid pointed out that tests on the variables used in the previous studies showed that all variables were nonstationary, and therefore the use of classical regression was inappropriate. He also pointed out that the models used in both studies face an identification problem. Although these studies assumed *a priori* that all right-hand side variables were exogenous, a weak exogeneity test showed that rice price cannot be exogenous. Furthermore, Rashid argued that the inclusion of the agricultural productivity variable (measured as an index of per acre production) in the BR model was inappropriate because prices and wages are inflationary while productivity is not. Analyzing the same data used in the two studies using co-integration techniques, Rashid reported higher estimates of SR and LR elasticities for both models. The short-run elasticities of wage to rice price were 0.32 and 0.25 for the BR and Palmer-Jones models respectively. Moreover the corresponding long-run elasticities were 0.72 and 0.69.

In the same study, Rashid also analyzed data from 1976/77-1998/99, which he called 'post-famine' data. He hypothesized a long-run relationship among nominal agricultural wage, rice price, and the urban wage rate. However, he found that in the post-famine period, rice prices and wages were no longer co-integrated (i.e. there is no stable LR relationship between them).

Palmer-Jones and Parikh (1998) also applied co-integration techniques to the data set from the 1993 Palmer-Jones study. From a wage determination model, they estimated the nominal agricultural wage-rice price SR elasticity to be 0.11, while the LR elasticity was 0.48.

Datt and Olmsted (2004) used a Generalized Method of Moments framework to measure short and long-run elasticities of nominal agricultural wages to food prices in Egypt. Using panel data from 18 governorates from 1976-1993, they found that over a 16-month period, the SR elasticity of wages to food prices was only 0.27. Moreover, it took up 5-7 years for wages to adjust to 90-95% of the food price increase.

Although there are differences in the literature about the appropriate model and estimation procedure, almost all studies point to a significant but sluggish response of agricultural wages to changes in rice price in the short-run. In the long-run, most studies reported a greater wage adjustment, although the LR elasticity is still less than one.

3. Methodology

The empirical model is derived based on the neoclassical utility maximizing framework for a representative farm household. In a study of wage determinants and labor supply in rural India, Rosenzweig (1980,1984) argued that the neoclassical competitive framework is appropriate, even for developing countries. Using district and household level data, he found rural wage rates to be endogenously determined and highly responsive to changes in labor supply. A similar study in the Muda River Valley, a farming community in Northwest Malaysia also found evidence for competitive labor markets in rural areas (Barnum and Squire, 1979).

This model assumes that agricultural households derive utility from the consumption of three goods: an agricultural commodity produced by the household (X_a), a market purchased good (X_m), and leisure (X_l). The household chooses a consumption level that maximizes overall utility subject to income, time and production constraints (Equation 1):

$$\max_{X_a, X_m, X_l, L} U = U(X_a, X_m, X_l) \text{ s.t. } P_a X_a + P_m X_m = W(T - X_l) + \Pi$$

where:

P_m – nominal retail price of market goods
P_a – nominal farm price of household-produced agricultural commodity
W – nominal agricultural wage rate
T – total time available to household
$\Pi = P_a Q(L,A,Z) - WL$
Q(L,A,Z) – production function of agricultural commodity
L – total labor input allocated to farm production
A – area devoted to agricultural production
Z – other non-labor, non-land inputs

The household is a price taker in the markets for X_a and X_m. Price-taking behavior is plausible even for X_a because most agricultural commodities are also imported and/or subject to price controls. Family labor and hired labor are assumed to be perfect substitutes and all labor is valued at W, the agricultural wage rate. It is assumed that there is separability of production and consumption decisions. That is, households make production decisions to maximize profit and then choose consumption based on realized income from their production decision. Furthermore, it is assumed that urban and rural markets are separate, i.e. non agricultural wages do not affect the decision to supply agricultural labor. This is reasonable in the case of the Philippines where there is significant unemployment in the urban and rural areas. In 2002, the urban unemployment rate was 13.2% while the rural unemployment rate is 7.3% (National Statistics Coordination Board, 2002). This means that all jobs in the urban areas are taken by urban laborers before they become available to rural laborers. Moreover, in many rural areas in the Philippines, labor is not mobile from rural to urban areas.

To maximize profit, a household chooses a level of labor and inputs such that the marginal revenue product of labor and other inputs are equal to the wage rate and other input prices respectively. In this model only labor is considered as variable. In the Philippines, this stylization is reasonable because the major input to agricultural production is labor. For example, in rice farming, 50% of production cost is for labor. The second major input is fertilizer which is 16% of production cost (2004, International Rice Research Institute- Social Sciences Division).

Taking the derivative of (1) with respect to L, the labor demand from the agricultural household is derived:

$$L^* = L^*(W, P_a, A, Z) \tag{2}$$

Substituting optimal profit Π^* into (1) and taking the derivative with respect to X_l, the demand for leisure is obtained:

$$X_l^* = X_l^*(P_m, P_a, W, T, \Pi^*) \tag{3}$$

In equilibrium under a single representative household, $L = T - X_l$. Imposing this restriction and using (2) and (3) gives:

$$F(W, P_a, A, Z, P_m) = 0 \tag{4}$$

which defines the long-run equilibrium relationship among the variables in the system[2]. Note that $\Pi^* = \Pi(W, P_a, A, Z)$ so (4) completely characterizes the equilibrium relationship between wages and prices.

In the econometric model, the Consumer Price Index is used to reflect P_m. P_a is represented by the nominal prices of rice (P_r) and corn (P_c), which are the primary agricultural crops in the Philippines. Since there are no data to quantify other non-labor, non-land inputs, a time trend (t) is used to capture changes in Z, as well as other time-varying omitted variables, such as labor productivity. Define X as the vector of variables in the equilibrium relationship, that is:

$$X = (W, P_r, P_c, P_m, A, t) \tag{5}$$

The empirical relationship among these variables is examined in both the long run and in the short run.

Annual time series data from 1975-2003 are used in this analysis. Data on agricultural wages, rice and corn prices and agricultural area are from various surveys conducted by the Bureau of Agricultural Statistics (BAS) in the Philippines. Wages are obtained by taking the weighted average of wages received by rice, corn, coconut and sugar laborers who do not receive meals. Weights are determined by the size of production of the abovementioned commodities in the province. Since there are differences in wages across provinces, the provinces which produce more of a given commodity are given more weight. Rice and corn prices are national annual averages computed from daily and monthly prices for different provinces in the Philippines. Area is computed by summing the land areas planted to rice and corn. The data on Consumer

[2] T is assumed to be a constant and was therefore excluded from the relationship.

Price Indices are from various volumes of the Philippine Statistical Yearbook, published by the National Statistics Coordination Board (NSCB).

If the variables in (5) are covariance stationary then the short-run and long-run relationship between wages and prices can easily be estimated using OLS, assuming that the following reduced form relationship exists:

$$W = f(P_r, P_c, P_m, A, t). \tag{6}$$

The assumption of exogeneity of the right-hand side variables (6) is a fairly strong assumption. Due to the lack of available instrumental variables, the endogeneity of each of the explanatory variables was not tested. This limitation should be noted when interpreting the results of this model under the assumption of stationarity.

It is also possible that some of the variables in X may have unit roots (i.e., are integrated of order one or I(1)) and that there exist a long-run equilibrium among these I(1) variables. To evaluate this hypothesis, the Augmented Dickey-Fuller (ADF) and Phillips-Perron tests are used. The ADF test is the most commonly used procedure for testing for unit roots. The Phillips-Perron test is similar to the ADF test, but allows for more relaxed assumptions on the error term. (i.e., it does not assume that errors are uncorrelated with constant variance.) The following equation is used to implement the ADF:

$$\Delta y_t = a_0 + \gamma y_{t-1} + a_1 t + \sum_{i=1}^{m} \delta \Delta y_{t-i} + e_t \tag{7}$$

where y_t is the variable being tested for unit root, Δ is the first difference operator, t is a time trend and e_t is the residual term.

The null hypothesis of the test is $\gamma=0$ (nonstationarity). OLS is used to estimate (7) and the τ statistic associated with the γ parameter is compared to the critical value in the Dickey-Fuller tables to determine if the null hypothesis can be rejected. The lag length m is chosen by adding additional lags until no more autocorrelation is found in the residuals. The procedure outlined by Enders (2004) for determining whether to include a constant and a time trend in (7) is used. This method involves estimating the most general form which includes a constant and a time trend and sequentially imposing restrictions that test whether a regressor can be excluded from the equation.

The classical approach to dealing with nonstationarity is to difference the data and then use OLS estimation. However, differencing removes important information about the long-run relationship among the levels of the variables. In order to capture this information, a co-integration framework can be used to analyze the relationship among these variables. If the variables are indeed co-integrated then, consistent with the Granger Representation Theorem, an error correction model (ECM) can be used to analyze the short-run adjustment.

Co-integration implies a long-run equilibrium relationship among I(1) variables, although in the short-run there may be substantial deviations from this long-run equilibrium. Engle and Granger (1987) give a formal definition of co-integration: "The components of the vector $\boldsymbol{x}_t$ are said to be co-integrated of order d, b, denoted $\boldsymbol{x}_t \sim CI(d,b)$ if (i) all components of $\boldsymbol{x}_t$ are $I(d)$; (ii) there exists a vector $\boldsymbol{\alpha}(\neq 0)$ so that $z_t = \boldsymbol{\alpha}'\boldsymbol{x}_t \sim I(d-b)$, $b>0$. The vector is called the co-integrating vector." The co-integrating vector (CV) is not unique. A scalar combination of the CV is also a co-integrating vector. Moreover there may be other equilibrium relationships among a subset of the variables in the co-integrating regression.

Engle and Granger have developed a simple two-step procedure for estimating the co-integrating vector and formulating the corresponding ECM. This involves estimation of the co-integrating regression by OLS and incorporating lagged errors of this regression into a differenced equation. However, there are some limitations to this method. First, the results of the co-integration test may be sensitive to the choice of dependent variable, although asymptotically the results should be consistent. Secondly, the method does not allow for determining if there is more than one co-integrating vector. Lastly, because a two-step procedure is used, it is possible that errors in the first step will be carried over to the next step. For example, an estimation bias in the parameters of the co-integrating regression could be transmitted to the regression that tests for co-integration (Davidson and MacKinnon, 1993, Enders). Johansen (1988) developed a method based on the estimation of a multivariate Vector Autoregression (VAR) by maximum likelihood which overcomes these limitations. However, an undesirable characteristic of the VAR is that it requires a lot of parameters to be estimated. If the VAR were to be used is this analysis, the very low power of the test will not yield useful results. Due to the fact that only a limited number of observations are available, the two-step procedure proposed by Engle and Granger is used in this study. Some of the limitations mentioned above are addressed by testing for co-integration using each of the model variables as the dependent variable. This evaluates whether the results are robust to the choice of normalization variable.

The first step in the Engle-Granger methodology is to use OLS to estimate a model containing the variables that are hypothesized to be co-integrated. If the variables are not co-integrated, then OLS leads to 'spurious regressions', where the regression results are characterized by a high R-squared and significant t-values, but the relationship has no economic meaning. However, an interesting result of OLS estimation is that if the variables are co-integrated, then the parameters are *super consistent* estimators of the long-run equilibrium relationship. This means that the parameters converge to their true value faster than OLS estimates involving stationary regressors. However, R-squared and standard errors are still misleading, and drawing any statistical inference about the parameters from conventionally estimated standard errors is inappropriate.

A log-log functional form is used in the model. Since wage response is the focus of this paper, wage is assigned to be the dependent or normalizing variable. The equation to be estimated is:

$$w_t = \beta_0 + \beta_1 p_{rt} + \beta_2 p_{ct} + \beta_3 p_{mt} + \beta_4 a_t + \beta_5 \mathrm{t} + \varepsilon_t \qquad (8)$$

where lower case italicized letters represent natural logarithms of the model variables. If the variables in (8) are co-integrated, then the βs can be interpreted as long-run elasticites.

To determine if the variables in (8) have a co-integrating relationship, the residuals from the regression are tested for stationarity using the Engle-Granger test for co-integration. The following equation is used for this test:

$$\Delta\hat{\varepsilon}_t = b_0 \hat{\varepsilon}_{t-1} + \sum_{i=1}^{k} b_i \Delta\hat{\varepsilon}_{t-i} + v_t \qquad (9)$$

where $\hat{\varepsilon}_t$ is the estimated residual from estimation of (8).

The null hypothesis is $b_0 = 0$ (nonstationarity). If autocorrelation is found in v_t, then successive lags of $\Delta\hat{\varepsilon}_{t-1}$ are included in the equation. The test above is similar to the ADF test, but in this case, standard Dickey-Fuller tables cannot be used for hypothesis testing. This is because the test is based on estimated residuals, and this makes the procedure biased towards stationarity. Several tables of critical values, which correct for this bias, have been developed. Here, the values presented by Davidson and MacKinnon are used. The critical values are determined by the number of I(1) variables in the co-integrating relationship and the nature (trend, constant) of the nonstochastic regressors.

Once it is established that there is a co-integrating relationship among the variables in the model, the co-integrating vector can be used to formulate the ECM. The intuition behind the ECM is that deviations from the long-run equilibrium in the previous period are corrected in the current period. Thus, it is necessary to establish a LR equilibrium first before using an ECM. The ECM is formulated as follows:

$$\Delta w_t = \alpha_0 + \alpha_1 \Delta p_{rt} + \alpha_2 \Delta p_{ct} + \alpha_3 \Delta p_{m_t} + \alpha_4 \Delta a_t + \lambda[w_{t-1} - \boldsymbol{\beta}' \mathbf{x}_{t-1}] + \upsilon_t \quad (10)$$

where **x** is a vector consisting of the variables p_r, p_c, p_m, a and t and $\boldsymbol{\beta}$ is composed of the parameters (not including the coefficient of the normalizing variable) from the co-integrating regression. The α parameters can be interpreted as short run elasticities. The term $[w_{t-1} - \boldsymbol{\beta}' \mathbf{x}_{t-1}]$ is the error correction (*EC*) term and λ is the *speed of adjustment* parameter which measures how past period deviations of the wage rate from its long-run equilibrium are 'corrected' or adjusted in the current period. The larger the value of λ the greater the adjustment to deviations from the previous period's disequilibrium. The EC term can be conveniently replaced by the lagged residuals from the co-integrating regression because the magnitude of the residuals reflects the deviations from the long-run equilibrium in the past period.

4. Empirical Results

A. Unit Root Tests

The results of the ADF and Phillips-Perron tests are presented in Table 1. Both levels and differences were tested. The second column shows the form of the equation used in the test (Random walk with drift (RWD), Random Walk (RW)). The inclusion of the drift was determined by testing for its significance in the regression equation for each of the variables. The levels of all the variables were found to be nonstationary using a 5% significance level, while the differences of all the variables were found to be stationary at the same significance level. These results suggest that all variables are I(1).

Table 1. Unit root tests

VARIABLE	Null Hypothesis	ADF τ Stat	Phillips-Perron τ Stat	τ Critical Value (5%)	τ Critical Value (10%)	Order of Integration
LEVELS						
w	RWD	-1.696	-1.696	-2.99	-2.62	I(1)
p_r	RWD	-1.218	-0.982	-2.99	-2.62	I(1)
p_c	RWD	-1.287	-1.129	-2.99	-2.62	I(1)
p_m	RWD	-1.938	-2.296	-2.99	-2.62	I(1)
a	RW	-0.331	-0.265	-1.95	-1.60	I(1)
DIFFERENCES						
$d.w$	RW	-2.358	3.800	-1.95	-1.60	I(0)
$d.p_r$	RW	-2.974	1.551	-1.95	-1.60	I(0)
$d.p_c$	RW	-4.697	1.306	-1.95	-1.60	I(0)
$d.p_m$	RW	-2.009	4.080	-1.95	-1.60	I(0)
$d.a$	RW	-7.389	-0.265	-1.95	-1.60	I(0)

All variables are in natural logarithms.

Figures 1 to 10 show the levels and differences of the natural logs of the variables in the model. While it is apparent from Figures 5-10 that the differences are stationary, the graphs of the levels suggest that the variables could be trend stationary, as opposed to having unit roots. It should be noted that due to the small sample size, the ADF and Phillips-Perron tests have very low power (Enders). Thus it is possible that the variables are stationary, but the null hypothesis in the ADF and Phillips-Perron tests cannot be rejected due to the low power of the test.

In order to allow for the possibility that the variables are stationary and the possibility that the variables have unit roots, both cases are analyzed. In the first part of the next section, it is assumed that the variables are nonstationary and a co-integration framework is used to derive the long run and short run elasticities. In the second part, it is assumed that the variables are stationary and OLS estimation is used to obtain long run and short run elasticities. These alternative specifications allow the results to be evaluated under alternative assumptions regarding the stationary properties of the variables.

Figure 1. Natural log of agricultural wages in levels (1975-2003)

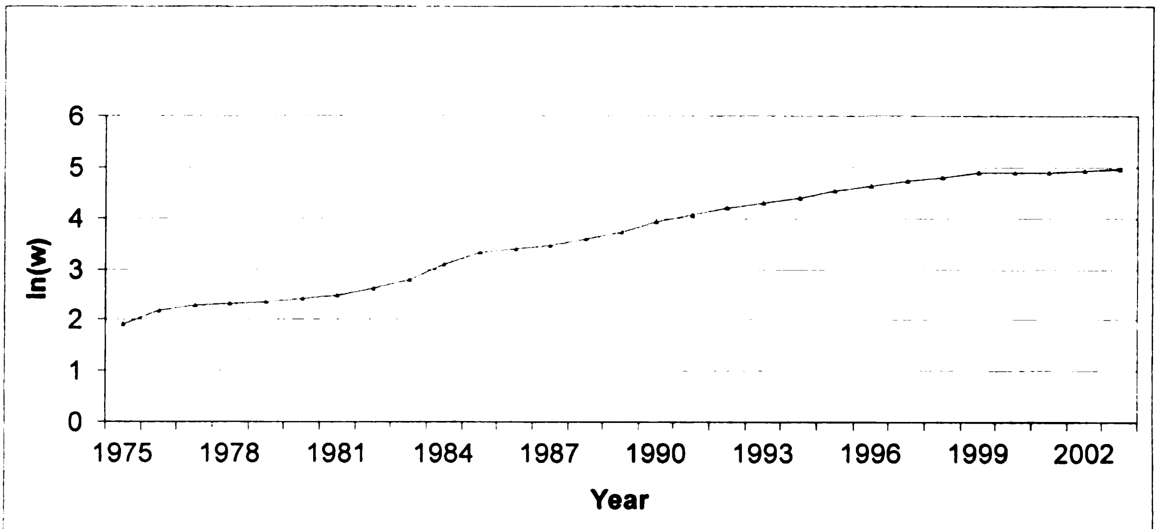


Figure 2. Natural log of nominal rice prices in levels (1975-2003)

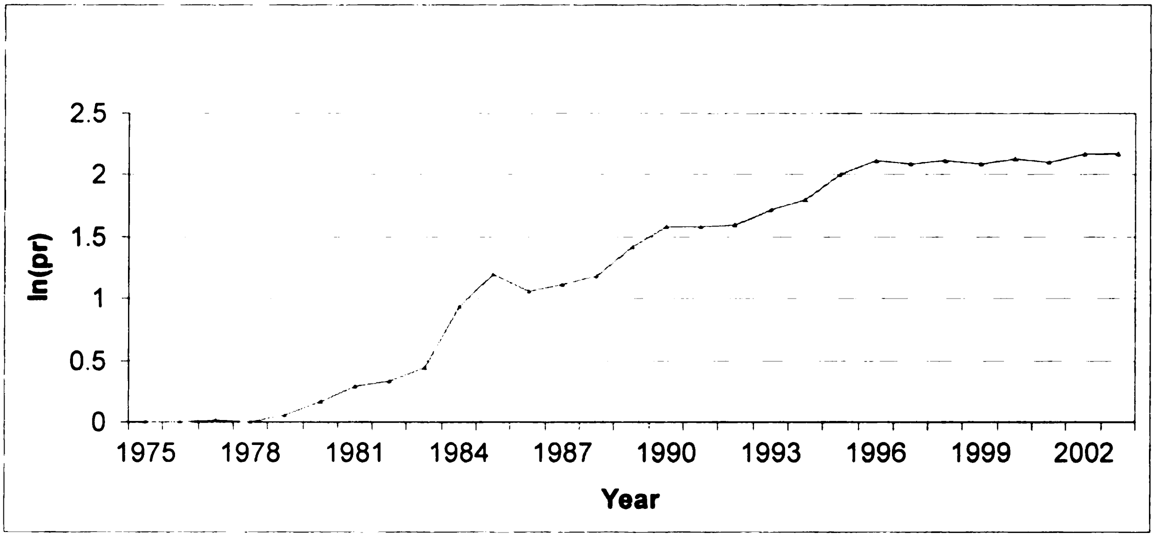


Figure 3. Natural log of nominal corn prices in levels (1975-2003)

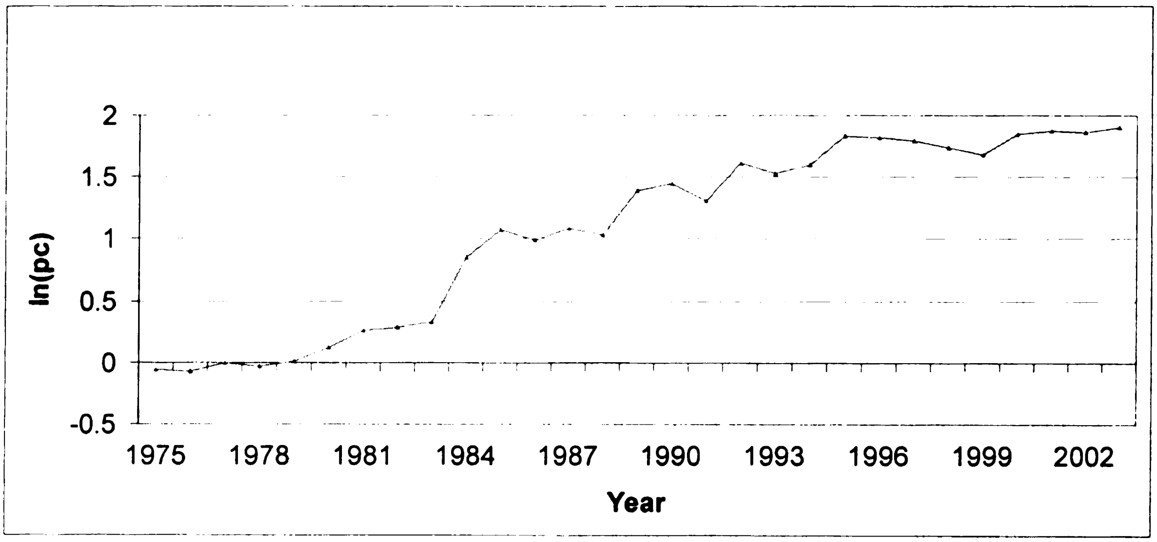


Figure 4. Natural log of Consumer Price Index (1994=100) in levels (1975-2003)

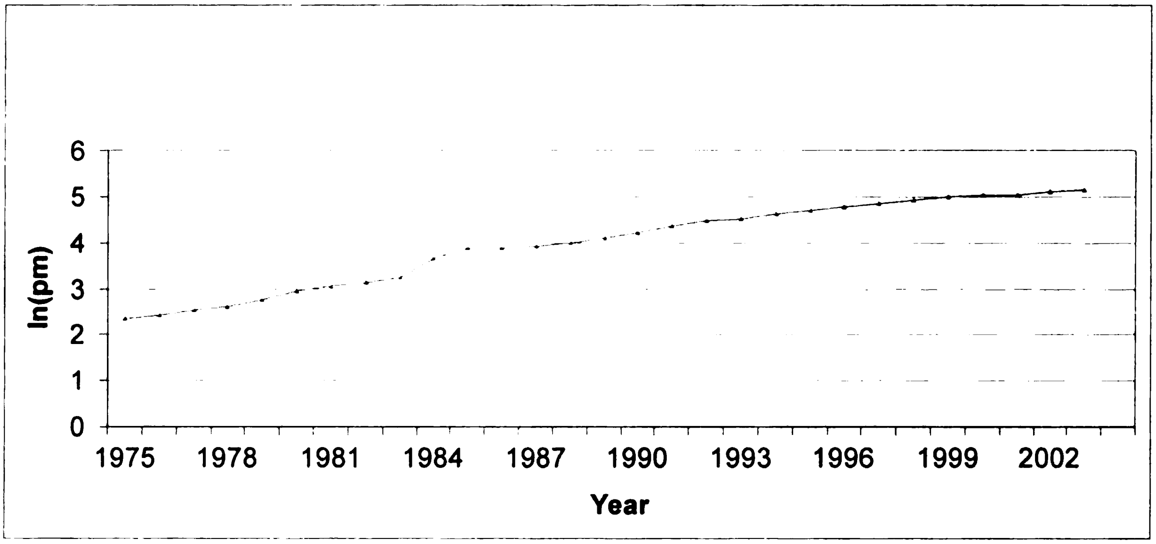


Figure 5. Natural log of agricultural area in levels (1975-2003)

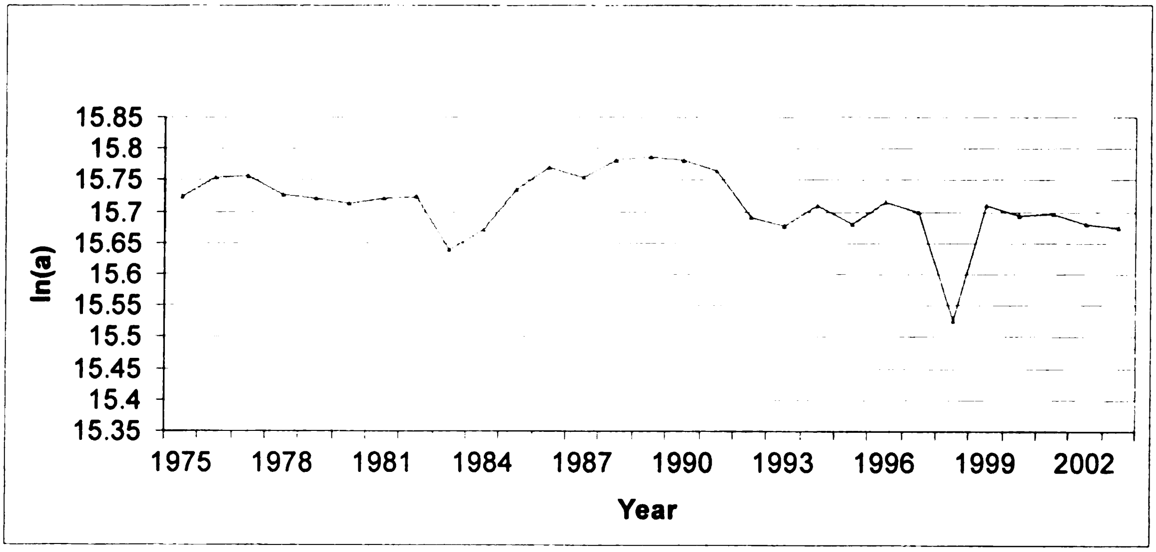


Figure 6. Natural log of agricultural wages in differences (1975-2003)

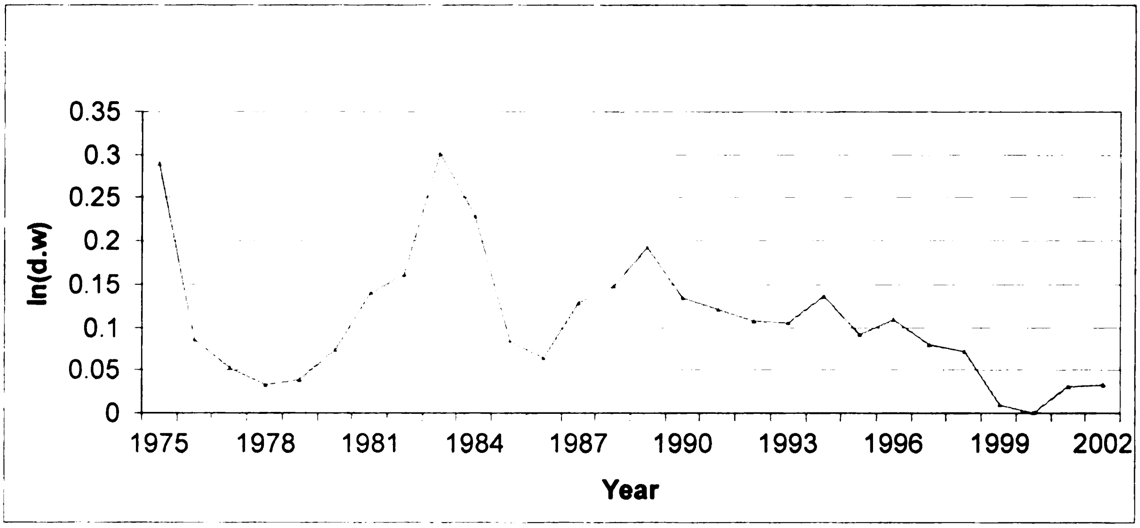


Figure 7. Natural log of nominal rice prices in differences (1975-2003)

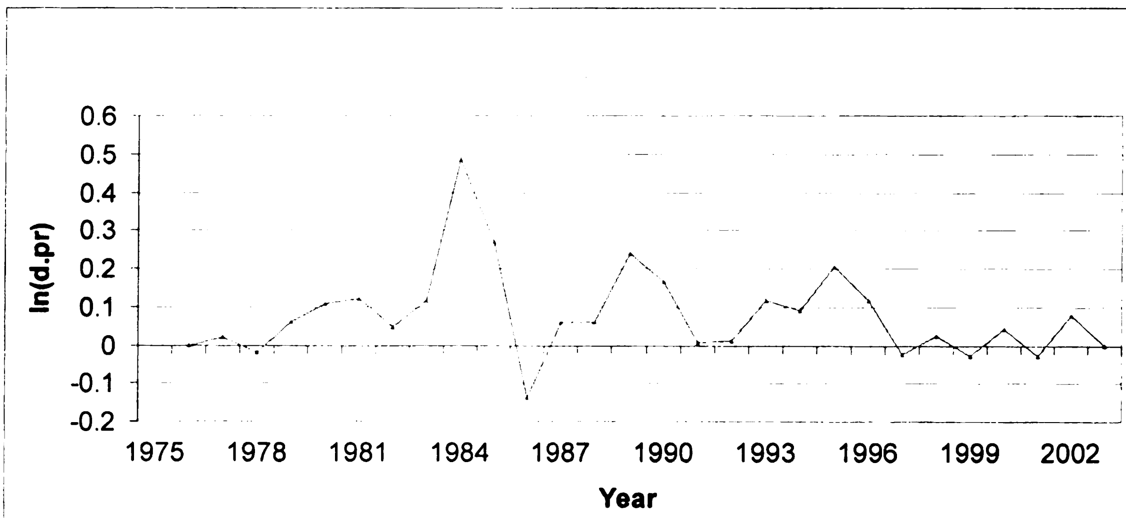


Figure 8. Natural log of nominal corn prices in differences (1975-2003)

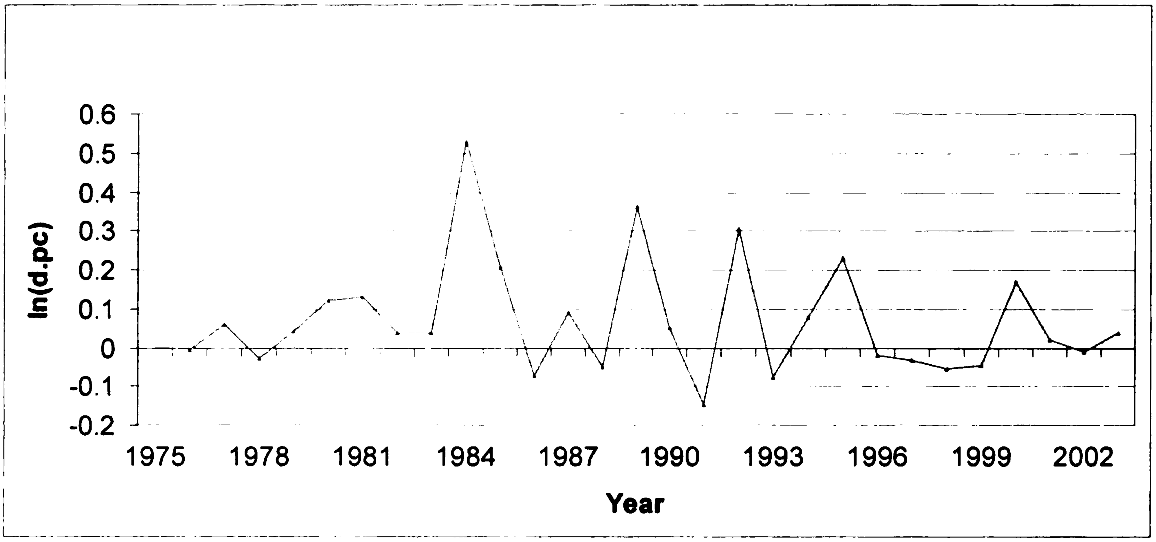


Figure 9. Natural log of Consumer Price Index (1994=100) in differences (1975-2003)

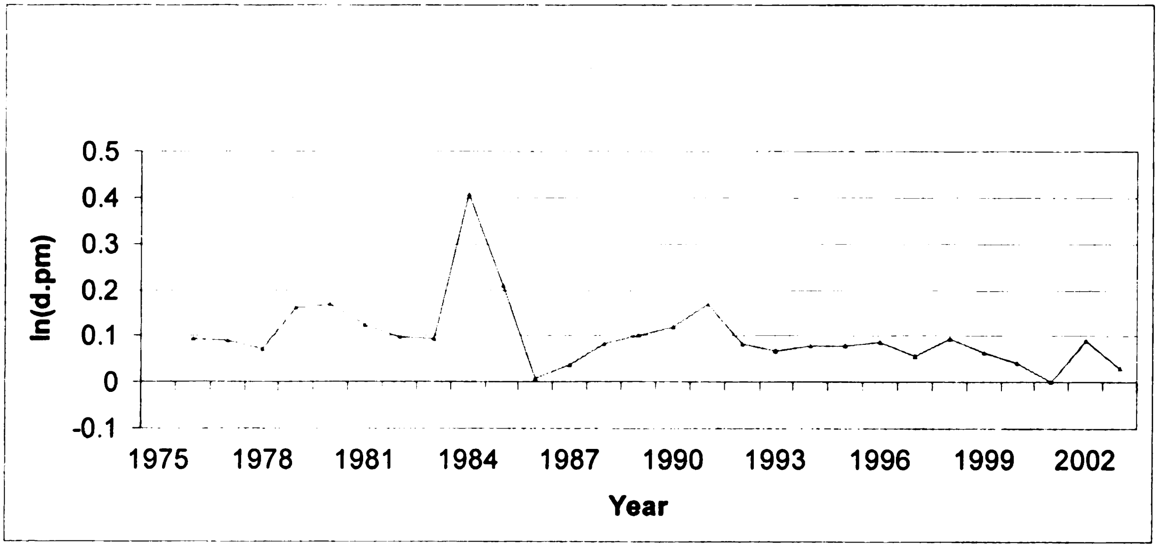
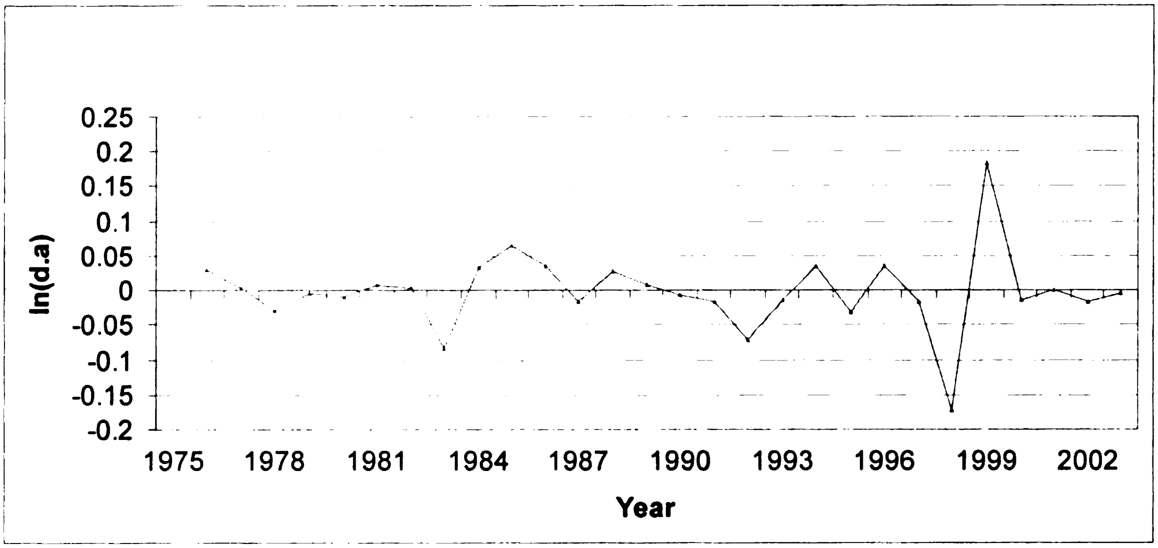


Figure 10. Natural log of agricultural area in differences (1975-2003)



B. Dynamics Under Nonstationarity

Long-run Wage Response

Following the methodology developed by Engle and Granger, OLS was used to estimate the following equation which is hypothesized to represent a co-integrating regression:

$$w_t = 1.2553 + 0.7807 p_{rt} - 0.1234 p_{ct} + 0.0547 p_{mt} + 0.0385 a_t + 0.0501\text{t} + \varepsilon_t \qquad (11)$$

S.E.	(5.5685)	(.2396)	(.2399)	(.1966)	(.3504)	(.0120)
t-stat	(0.23)	(3.26)	(-0.51)	(0.28)	(0.11)	(4.16).

F(5, 22) = 862.96 R-squared = 0.9947
Prob > F = 0.0000 Adj R-squared = 0.9935

If the variables in (11) are co-integrated, then the coefficients can be interpreted as long run elasticities so the long run elasticity of wage with respect to the price of rice is 0.78.

Using error terms from equation (11) to estimate (9) yields:

$$\Delta\hat{\varepsilon}_t = -.4835\,\hat{\varepsilon}_{t-1} + .2929\Delta\hat{\varepsilon}_{t-1} + v_t \qquad (12)$$

S.E.	(.1628)	(.1528)
τ -Stat	(-2.97)	(1.92)

where a lag of $\Delta\hat{\varepsilon}_t$ was added to correct for autocorrelation.

The critical value for five variables in the co-integrating regression (with a trend and a constant) is -4.72 at the 5% significance level, and -4.43 at the 10% significance level. Since the τ-Stat of the lagged error term (-2.97) is not greater (in absolute terms) than the critical values, the null hypothesis of nonstationarity of the residuals cannot be rejected. This implies that the variables in the equation (11) are not co-integrated, and no stable long-run relationship exists among them.

Tests for co-integration are also done under alternative normalizations by varying which variable in (11) is specified as the dependent variable. The results of the Engle-Granger test under these alternative normalizations are presented in Table 2.

Table 2. Engle-Granger test for co-integration with constant

NORMALIZING VARIABLE	Lags	τ Stat	τ Critical Value (5%)
w	1	-2.97	-4.72
p_r	0	-5.14*	-4.72
p_c	0	-6.63*	-4.72
p_m	0	-3.14	-4.72
a	0	-4.91*	-4.72

The null of no co-integration ($\hat{\varepsilon}_t$ *s* are nonstationary) can be rejected in three of the five cases. Since the sample size is limited, it is expected that the Engle-Granger test has low power. Thus, it is possible that the regressions having w and p_m as dependent variables have stationary residuals, and a co-integrating relationship does exist among the variables. The analysis of the short-run dynamics is dependent on the assumption about the long-run relationship among the variables. Since there is evidence for both co-integration and no co-integration, we present two models of the short-run dynamics in the following sections. The error correction model assumes that the variables in the model are co-integrated while the first difference model assumes that the variables are not co-integrated.

Error Correction Model

If the variables in (11) are co-integrated, then an ECM is appropriate in analyzing short run wage adjustments. The ECM assumes that there is a significant error correction (*EC*) term $[w_{t-1} - \boldsymbol{\beta}'\mathbf{x}_{t-1}]$, which corrects past-period deviations from the long run equilibrium in the current period. Estimation of (10) yields the following (Equation 13):

$$\Delta w_t = 0.0290 + 0.3345\Delta p_{rt} - 0.0387\Delta p_{ct} + 0.1356\Delta p_{m_t} - 0.0364\Delta a_t + 0.3389\Delta w_{t-1} - 0.2617EC + \upsilon_t$$

S.E.	(.0141)	(.1131)	(.0666)	(.1418)	(.1138)	(.0933)	(.0991)
t-Stat	(2.06)	(2.96)	(-0.58)	(0.96)	(-0.32)	(3.63)	(-2.64)

F(6, 20) = 14.21 R-squared = 0.8100
Prob > F = 0.0000 Adj R-squared = 0.7531

where a lag of Δw_t was added to correct for autocorrelation. The residuals of (13) passed tests for white noise using Portmanteau's Q test (Pval> $\chi^2_{(1)}$ = 0.6508) and Durbin H (Pval> χ^2_{stat} = 0.5649) test for first order serial autocorrelation. Because each term in the equation above is stationary, standard asymptotic theories apply, and inferences can be made about the model based on t and F statistics. The individual coefficients of Δp_{rt}, Δw_{t-1} and the error correction term are significant at the 5% level. The coefficient of the error correction term is negative (-0.26). This means that if the *EC* term is positive, or the wage in period t-1 is above the long run equilibrium, then the error correction mechanism will drive the current wage lower. Under the assumptions that the variables in (11) are co-integrated and have an ECM representation, the SR contemporaneous elasticity of wages with respect to rice prices is 0.33. This means that in the very short run, a 1% decrease in nominal rice prices will cause a .33% decrease in the agricultural wage.

First Difference Model

The ECM assumes that there is a co-integrating relationship among the variables in (11). However, the results of the Engle-Granger tests for co-integration also gave some evidence that there is no co-integrating relationship among the variables in the model. To allow for this possibility, the short run relationship among the variables is analyzed using a FDM. This model assumes that the variables in (11) are not co-integrated. Since it is also assumed that the variables have unit roots, first differencing is needed to make the variables stationary. Note that since the data have been differenced, inferences about the LR relationship among the levels of the variables cannot be made. However, short-run elasticities can still be derived. OLS is used to estimate the following model:

$$\Delta w_t = \alpha_0 + \alpha_1 \Delta p_{rt} + \alpha_2 \Delta p_{ct} + \alpha_3 \Delta p_{m_t} + \alpha_4 \Delta a_t + \alpha_5 \mathrm{t} + \omega_t \qquad (14)$$

This model is similar to the ECM model in (10) but the error correction term is excluded because there are no adjustments from a long equilibrium that need to be captured. A time trend is included because there is no underlying long run relationship that would have captured the effects of time varying omitted variables. Estimation of (14) yields the following results (Equation 15):

$$\Delta w_t = 0.0110 + 0.3092\Delta p_{rt} - 0.0339\Delta p_{ct} + 0.1764\Delta p_{m_t} - 0.0108\Delta a_t + 0.0004t + 0.3957\Delta w_{t\text{-}1} + \omega_t$$

S.E.	(.0308)	(.1341)	(.0774)	(.1777)	(.1314)	(.0011)	(.1104)
t-Stat	(0.36)	(2.31)	(-0.44)	(0.99)	(-0.08)	(0.38)	(3.59)

F(6, 20) = 9.77 R-squared = 0.7457
Prob > F = 0.0000 Adj R-squared = 0.6694

where a lag of Δw_t was added to correct for autocorrelation.

Portmanteau's Q (Pval> $\chi^2_{(1)}$ = 0.6603) and Durbin H (Pval> $\chi^2_{(1)}$ = 0.6112) tests showed that the residuals of (15) are uncorrelated of order one and follow a white noise process. The coefficients of $\Delta w_{t\text{-}1}$ and Δp_{rt} are significant at the 5% significance level with t-statistics of 3.59 and 2.31, respectively. These results are similar to the results of the ECM. Moreover, the elasticities of wages with respect to rice prices are close, at 0.33 for the ECM and 0.31 for the FDM.

A F-test was performed on the non-significant variables to see if they are jointly zero. The null that $\alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = 0$ was not rejected at 5% significance level (Pval > F = 0.8604). By removing the variables $\Delta p_{ct}, \Delta p_{m_t}, \Delta a_t$ and t, a more parsimonious model is derived:

$$\Delta w_t = 0.0279 + 0.3608\Delta p_{rt} + 0.4048\Delta w_{t\text{-}1} + \omega_t \qquad (16)$$

S.E.	(.0131)	(.0611)	(.0975)
t-Stat	(2.14)	(5.91)	(4.15)

F(2, 25) = 32.34 R-squared = 0.7294
Prob > F = 0.0000 Adj R-squared = 0.7068

The residuals of this model were found to be white noise using Portmanteau's Q (Pval> $\chi^2_{(1)}$ = 0.125) and Durbin H (Pval> $\chi^2_{(1)}$ = 0.7238) tests. Under this model, which is the preferred model under the nonstationary and not co-integrated assumptions, the short run elasticity of wages with respect to rice prices is 0.36. This means that for a 1% change in rice prices, agricultural wages will change by .36% in the same year.

C. Dynamics Under Stationarity

If the variables in (6) are I(0) then the short run elasticities can be obtained by using OLS estimation. To make the stationary model (SM) fully consistent (that is, terms in both equations are similar) with (16), the model estimated is:

$$w_t = \beta_0 + \beta_1 p_{rt} + \beta_2 p_{rt-1} + \beta_3 w_{t-1} + \beta_4 w_{t-2} + z_t \tag{17}$$

The corresponding long run elasticities can be derived by using the formula

$$\text{LR Elasticity} = \frac{\beta_1}{1-\beta_3} \tag{18}$$

where β_1 is the coefficient of p_{rt} while β_3 is the coefficient of $w_{t\text{-}1}$.

Estimation of (17) yields:

$$w_t = 0.5829 + 0.4128 p_{rt} - 0.1208 p_{rt-1} + 0.8839 w_{t-1} - 0.1261 w_{t-2} + z_t \tag{19}$$

S.E.:	(.1579)	(.0536)	(.0847)	(.1689)	(.1143)
t-Stat:	(3.69)	(7.70)	(-1.43)	(5.23)	(-1.10)

$F(4, 22) = 6435.17$ R-squared = 0.9991
Prob > F = 0.0000 Adj. R-squared = 0.9990

The coefficients of p_{rt-1} and w_{t-2} are individually and jointly non-significant (Pval>F = 0.3397). Removing p_{rt-1} and w_{t-2} from (19) leads to a more parsimonious model:

$$w_t = 0.8498 + 0.4126 p_{rt} + 0.6494 w_{t-1} + z_t \tag{20}$$

S.E.	(.0828)	(.0506)	(.0402)
t-Stat:	(10.26)	(8.14)	(16.16)

F(2, 25) = 862.96 R-squared = 0.9986
Prob > F = 0.0000 Adj R-squared = 0.9985

The residuals of (20) passed Durbin H (Pval> $\chi^2_{(1)}$ = 0.2426) and Portmanteau's Q (Pval> $\chi^2_{(1)}$ = 0.1969) tests for first order autocorrelation and white noise.

To verify if the results of the full model estimation are robust to underlying stationary assumptions, the full model is also estimated assuming that all variables are stationary and all right hand side variables are exogenous:

$$w_t = 3.7348 + 0.5722 p_{rt} - 0.0669 p_{ct} - 0.0948 p_{m_t} - 0.1617 a_t + 0.0066 t + 0.5961 w_{t-1} + z_t \tag{21}$$

S.E.	(2.5923)	(.1268)	(.1106)	(.1021)	(.1627)	(.0082)	(.0742)
t-Stat:	(1.44)	(4.51)	(-0.61)	(-0.93)	(-0.99)	(0.80)	(8.03)

F(6, 21) = 3042.70 R-squared = 0.9989
Prob > F = 0.0000 Adj R-squared = 0.9985

A lag of w_t is added to remove autocorrelation. The residuals of (21) passed Durbin H test (Pval> $\chi^2_{(1)}$ = 0.5157) and Portmanteau's Q test (Pval> $\chi^2_{(1)}$ = 0.4783) for first order autocorrelation and white noise.

In Equation 21, the signs of $\Delta p_{ct}, \Delta p_{m_t}$ and Δa_t are not expected. The t-statistics also show that these variables and t are non-significant. A joint F-test indicated that the coefficients of $\Delta p_{ct}, \Delta p_{m_t}, \Delta a_t$ and t are jointly equal to zero at the 5% significance level

(F_{stat} = 0.97). Removing the non-significant variables in (21) we derive a model that is identical to (20) which is the preferred model under this specification.

Under the assumption that the variables are stationary, the short-run elasticity is 0.41 while the long run elasticity is 1.17. This implies that in the SR, a 1% change in rice price will lead to a 0.41% change in the agricultural wage. Under this model, the LR elasticity is greater than one. As mentioned in the previous chapter, results from this model should be interpreted with caution because of the potential endogeneity problem associated with this specification.

Table 3 summarizes the short run and long run elasticities obtained from the error correction model and the stationary model. The first difference model is excluded because long run elasticities cannot be obtained from this model. It is of interest to know

Table 3. Wage elasticities and adjustment under different model specifications

	MODEL	
	ECM	**SM**
Short-Run Elasticity	0.33	0.41
Long-Run Elasticity	0.78	1.18
YEAR	**% Adjustment to the LR Elasticity**	
Year 1	0.57	0.58
Year 2	0.62	0.73
Year 3	0.64	0.82
Year 4	0.64	0.88
Year 5	0.65	0.92
Year 10	0.65	0.99
Year 15	0.65	1.00

how long it takes before the long run adjustment is reached. By summing the annual marginal effect of a percentage change in the price of rice in time t to the agricultural wage, we can obtain the annual percentage adjustment to the long run elasticity.

The ECM model suggests that even after 15 years, long run adjustment has not been reached. In the case of a price decrease, this means that agricultural wages will not decrease as much as rice prices, even in the long run. The SM suggests that 90% of the adjustment to the LR elasticity takes place after 5 years.

D. Welfare Implications

So far the effect of changes in rice price on the agricultural wage has been discussed, but nothing has been said about its effect on agricultural wage earners. Note that agricultural wage earners refer to net demanders of rice and net suppliers of labor (i.e. those households that supply more labor than they hire). The short-run and long-run elasticities provide information about how quickly wages respond to changes in rice prices, and enable the identification of groups who will be adversely affected by the decrease in rice prices. Using a household model similar to Equation 1 to analyze the effect of changes in wages to food prices, Ravallion (1990) showed that the necessary and sufficient condition for households that are net demanders of rice and net suppliers of agricultural labor to benefit from a small increase in food price is that the elasticity of agricultural wages with respect to food prices should be greater than the share of food expenditure in the household's labor earnings. Based on the same reasoning, the condition for households to benefit from a decrease in rice price is if the elasticity of wages to rice prices (η) is less that the share of rice expenditures in labor earnings (η^*). Conversely, households that are net demanders of rice and net suppliers of labor will incur a loss if the elasticity of wages with respect to rice price is greater than the share of rice expenditures in labor income, i.e., $\eta > \eta^*$. To illustrate this concept, Table 4 presents the different welfare implications of decreasing the price of rice for different types of agricultural wage earners, given the SR and LR elasticity. Since rice is a staple food in the Philippines, a substantial proportion of household expenditure is devoted to purchasing rice. The monetary amount is more or less fixed for a household of a given size, but the share in the total budget changes depending on the income. The poorest households in the Philippines typically spend 30% of their income on rice expenditures (Cororaton). Case 1 represents the poorest agricultural household that relies solely on agricultural wages where, $\eta^* = 0.30$. Since the SR and LR η values in the ECM, FDM and SM models are not greater than η^*, these households are expected to incur welfare losses both in the SR and in the LR if rice prices were to decrease.

Table 4. Welfare effects on agricultural wage earners of a reducton in rice price

				ECM		FDM	SM	
Case	% of rice exp. in total HH budget	% of income from ag. labor	% of rice exp. in ag. income (η^*)	SR gain or loss (η_{SR} = 0.33)	LR gain or loss (η_{LR} = 0.78)	SR gain or loss (η_{SR} = 0.36)	SR gain or loss (η_{SR} = .041)	LR gain or loss (η_{LR} = 1.18)
1	30	100	0.30	Loss	Loss	Loss	Loss	Loss
2	30	50	0.60	Gain	Loss	Gain	Gain	Loss
3	20	50	0.40	Gain	Loss	Gain	Gain	Loss
4	10	20	0.50	Gain	Loss	Gain	Gain	Loss
5	10	10	1	Gain	Gain	Gain	Gain	Loss

The value of η^* increases as households rely less heavily on agricultural wages, or incomes increase such that rice expenditures become less relative to the household budget. In general, as in Cases 2-4, households benefit from the decrease in rice price in the SR but eventually lose in the LR, as wages are further driven down in response to the decrease in rice prices. Case 5 shows that a household that is a net supplier of labor can only unambiguously gain from the decrease in rice price if the income is high enough and if the income from agricultural wages is only small portion of total income.

The analysis on Table 4 applies to a specific type of household, which is a net demander of rice and net supplier of agricultural labor. Although other types of households are not the focus of this paper, it is beneficial to briefly extend the welfare application to other household types. Households that are net demanders of food and net demanders of labor will unambiguously benefit from a decrease in the price of rice for all $\eta>0$, because they will have both the benefit of decreased rice expenditures as well as decrease in the cost of hiring labor. Urban consumers, who are net demanders of food, but have no agricultural income can also be categorized under this group. Urban consumers will clearly gain from a decrease in rice price because their food expenditures will decrease while their income will remain the same.

On the other hand, households that are net suppliers of rice and net suppliers of labor will be adversely affected by the price decrease for any $\eta>0$ regardless of how much their own consumption of rice is. This is because their income from selling rice will decrease and at the same time, their wage from agricultural labor will also decrease. For households that are net suppliers of rice but are net demanders of labor, the welfare effect is less predictable. It will depend on the loss of income due to the decrease in price at which rice is sold, relative to the decrease in the cost of hiring agricultural laborers.

5. Summary and Conclusions

The objective of this paper is to determine whether there is a long-run relationship between agricultural wages and rice prices, and to measure the short-run and long-run wage rate elasticities with respect to rice price. This study was motivated by the expected decline of rice prices in the Philippines, as a result of the government's efforts to liberalize rice trade in the country. The impact on wages is important because agricultural wages are the primary source of income of some of the poorest households in the Philippines' rural sector.

The relationship of agricultural wages and rice prices is analyzed using a neoclassical wage determination model where the wage rate is determined by supply and demand of labor in the agricultural sector. The ADF and Phillips-Perron tests for nonstationarity showed that variables in the model are I(1). However, it is recognized that the tests for nonstationarity may have low power. Thus, the LR relationship was analyzed under two assumptions: stationarity and nonstationarity. In the first part, it was assumed that the variables are nonstationary and a co-integration framework was used. In the second part, it was assumed that the variables are stationary and OLS was used to derive SR and LR elasticities. Because the test for co-integration revealed mixed results about the existence of a long run equilibrium relationship among the variables in the model, two models under the nonstationarity assumption are also specified: an error correction model and a first difference model.

The results confirm the general hypothesis that rice prices are an important influence on agricultural wages. Depending on the model specification, the short-run elasticity of wages with respect to rice prices ranges from 0.33-0.41, while the long-run elasticity ranges from 0.61-1.18. In all model specifications, wages adjusted to the LR elasticity by more that 50% after the first year. The SM suggests that at least 90% of the adjustment to the LR elasticity will occur in less than 5 years while the ECM model suggests that the adjustment is slower, with only 64% adjustment even after 15 years.

An analysis of the welfare implications of a decrease in rice prices for agricultural households that are net demanders of rice and net suppliers of labor showed that most households benefit in the SR but lose in the LR. Moreover, households who rely primarily on agricultural wages and spend a large percentage of their wage income on rice will be adversely affected by the decrease in rice prices in both the short-run and the long-run.

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