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An Empirical Investigation of Short and Long-run Agricultural Wage Formation in Ghana

by

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ABSTRACT

This paper investigates empirically the factors that influence real agricultural wage rates in Ghana, based on 1957 to 1991 data. The Johansen cointegration framework is used to examine long-run relationships among agricultural and urban wage rates, the domestic terms of trade between agriculture and non-agriculture, urban unemployment, capital stock in agriculture and the size of the rural population. An error correction model is then used to investigate short-run dynamic relationships among the variables. The results show that: (1) there is only one stable equilibrium relationship among agricultural wage rates and their determinants in the long-run; (2) a 1 percent change in the domestic terms of trade between agriculture and non-agriculture leads to a 0.48 percent change in the real agricultural wage rate in the short-run and a 0.83 percent change in the long run; (3) the analysis suggests a one-time and one way upwards structural shift of 3.6 percent in real agricultural wages during the 1980s.

1. INTRODUCTION

This paper addresses a key strategic issue in the adjustment of small, semi-open West African economies to major structural changes in relative output prices of the type that many such countries have experienced over the past thirty to forty years since independence from colonial rule. These structural changes include precipitous drops in export crop prices in the 1980's and rapid rises in the domestic price of importables in the structural adjustment era. The paper adopts a structural approach to the long-run adjustment of real agricultural wage rates, using aggregate national data for Ghana from 1957 to 1991. It takes into account the impact of inter-sectoral mobility of labour in the tradition of Harris and Todaro (1970), and specifically allows for differences in the short-run and long-run responses of real agricultural wage rates to changes in their underlying determinants.

Long-run changes incorporate both the direct effects of shifting relative prices for outputs on real agricultural wage rates, and indirect effects arising from second-order effects of changes in income distribution between town and countryside and within the latter (Boyce and Ravallion, 1991). A key issue is whether changes in real wage rates in response to structural changes in relative prices converge to a long run term equilibrium path or instead fluctuate with the interplay of indirect effects. A major application is to see the scope for policy to be able to have lasting impact on fundamental economic variables, such as the relative price of nontradables such as labour in agriculture (Sarris and Shams, 1992).

Small, relatively poor West African countries, such as Ghana, are primarily price-takers, even if they do face a downward-sloping demand curve in some cases, such as cocoa. Because the role of government and external donors was very high by world standards prior to the early 1980's, and is still large, domestic polices still have considerable latitude to directly affect the domestic terms of trade between agriculture and non-agriculture in these countries, even if they have little direct impact on key shadow prices, such as real agricultural wages. The latter, on the other hand, are key elements in the economic adjustment of these countries to policy initiatives in product markets and in government spending (Delgado and Mellor, 1984).

The main strategy promoted to increase agricultural productivity in the structural adjustment period since 1981 in West Africa has largely focused on trying to induce factor flows into agriculture, or stem factor flows out, through promoting a shift in the internal terms of trade in favour of agriculture. This was designed to be implemented through macroeconomic and sectoral reforms intended to devalue the real exchange rate, through reduction of fiscal outlays, elimination of price taxation through government involvement in the marketing of agricultural tradables, and elimination of explicit and implicit subsidies to primarily urban consumers. The object is to raise the domestic price of agricultural tradables, such as exportable cocoa and importable grain, relative to non-tradables such as labour, thus shifting income to net producers of tradables (Husain, 1994).

Throughout much of Africa, and Ghana is a good example, agricultural marketing is subject to a wide gap between import and export parity prices for products. This implies that a considerable share of agricultural output is consumed within the household or traded locally, falling within the price band in which it neither pays to import nor to export the good in question. Under these conditions, a significant share of the adjustment of rural households to Structural Adjustment policies of the type outlined above can be expected to occur through the labour market, rather than through observed differences in marketed output (Oyejide, 1986).

Despite the key strategic importance of the formation of agricultural wages to countries such as Ghana, there has been relatively little empirical investigation of the determinants of agricultural wage rates in either Ghana or more generally in Africa. Empirical studies on the determinants of real agricultural wages over multiple crop cycles in other developing areas have not specifically considered the impacts of inter-sectoral labour mobility on the evolution of agricultural wage rates (Khan, 1984; Boyce, 1989; Boyce and Ravallion, 1991; Palmer-Jones, 1993).

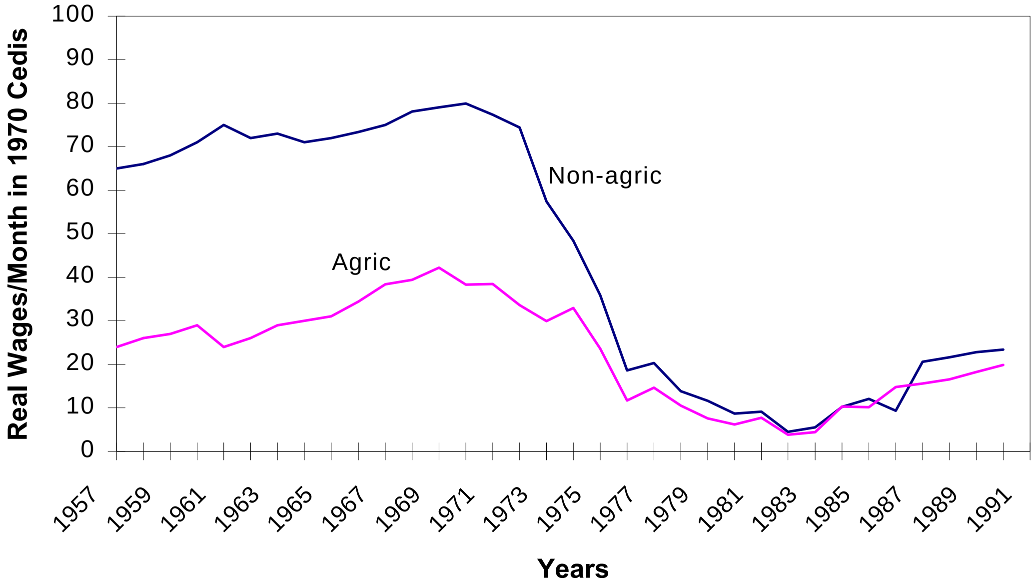
This paper extends previous research in two ways. First, the effects of inter-sectoral mobility of labour is taken into consideration in its theoretical framework. Second, it applies multivariate cointegration techniques pioneered by Johansen (1988) to examine the long-run equilibrium relationships between agricultural wage rates and its determinants in Ghana over the 1957 to 1991 period. Cointegration is then used to estimate a dynamic error correction model that takes account of the dynamics of short-run adjustment towards long-run equilibrium in the agricultural labour market, yielding empirical estimates of the long and short-run response of real agricultural wages to policy variables and exogenous shocks. The remainder of the paper is organised as follows. The next section presents trends in and significance of real agricultural wages in Ghana. Section three provides the theoretical framework. Section four sets out the econometric methodology used. The data and empirical results are described in section five. The final section provides a discussion of the implications of the results, and some conclusions.

2. SIGNIFICANCE OF AND TRENDS IN REAL AGRICULTURAL WAGE RATES IN GHANA

Agricultural wages are important for rural dwellers because they constitute a significant proportion of their income. Available evidence from the 1960 and 1970 population censuses show that about 33 percent and 31 percent of the rural labour force-particularly in the export crop producing parts of southern Ghana where migrant labour constitutes a significant percentage of total hired labour-depended primariy on agricultural wage employment for their livelihood in 1960 and 1970, respectively. In addition, a substantial proportion of the labour force depended on wage employment as a secondary source of income (Bequele, 1983). Beals and Menezes (1970) also show that more than 50 percent of the working age males in the North engaged in seasonal migration in the 1950s. Most of these migrants found employment in agriculture. A study conducted by Honny (1983) also revealed that while about 20 percent of the labour used in the North was hired labour, as much as 41 percent of total labour used in the middle belt was hired suggesting that hired labour constitutes a significant proportion of total labour used in agriculture.

The real agricultural wage series in Ghana reveals a steady rise up to the late 1960s(Figure 1).[1] In the early 1970s real wages declined sharply till 1982 but started rising again in 1983. The main striking features of the real agricultural wage series are the rapid decline in the 1970s and steady rise in the 1980s, and any factors associated with these are likely to exert considerable effect on the overall explanation of real agricultural wage rates. Although wages in the non-agricultural sector have followed a similar trend, agricultural wages lagged behind non-agricultural wages until the early 1980s when they virtually converged. While real wages in agriculture have declined, those of urban workers have declined more. The result is that, relative to their counterparts, agricultural workers have improved their wage conditions. As rightly noted by Bequele (1983), the decline in real agricultural wages in the 1970s and early 1980s, implied a significant rise in rural poverty. Sarris and Shams (1992 show that rural poverty in Ghana increased drastically between 1970 and 1984. According to their estimates the number of rural people living under the poverty line increased from 1.95 million in 1970 to 6.4 million in 1984. This contrasts with the decline in rural poverty between the second half of the 1980s and the beginning of the 1990s (World Bank, 1995).

Figure 1: Real Agricultural and Non-agricultural Wage Rates in Ghana, 1957-91



The wide wage differentials in the 1960s and 1970s contributed to rural-urban migration, which then led to a decline in the acreage under cultivation and consequent reduction in agricultural output and productivity.[2] The narrowing of the wage gap between the two sectors can be partly attributed to an increasing integration of the agricultural and non-agricultural labour markets. The economic recovery programme (ERP) instituted in 1983 appears to have contributed to the upward trend observed since 1983. Specific measures adopted centred around (i) liberalization of the trade regime and curbing the rate of inflation, and (ii) a progressive reappraisal of the country's domestic currency to reflect the scarcity value of foreign exchange. As will be shown later in this study, improvement in the domestic terms of trade for agriculture, and to some extent the lower rates of inflation have helped raise real agricultural wages. The fact that the prevailing agricultural wage rates in 1980s were as high as 3 to 4 times the official minimum wages in the 1980s suggest that other factors than government stated wage rates contributed to the formation of agricultural wage rates in the country (Huq, 1988). Sarris and Shams (1992) argue that the gradual increase in real agricultural wage rates—which they use to represent real rural wages—since the economic reforms is partly due to the increased demand for rural labour by the cocoa sector.

3. THEORETICAL FRAMEWORK

In general, it is assumed that the interplay between demand for and supply of rural labour determines the level of the agricultural wage rate. The supply and demand functions of agricultural households are specified under the assumption of a competitive rural labour market.[3] Although the assumption is made for analytical convenience, structural conditions on the rural labour markets in several developing countries infact justify this assumption. Market forces exert a harsher influence and play a greater role in a poor agrarian economy than in a developed country, as minimum-wage legislation is much less effective, and if at all implemented, unionization of the rural labour force is rare, and the state does not pay unemployment benefits.[4]

On the demand side for agricultural labour, agricultural product prices, labour productivity and the real agricultural wage rate are hypothesised to influence the amount of labour that agricultural producers are willing to engage in agricultural production. This is based on the notion that, for a particular level of productivity, increases in the relative prices of agricultural products stimulate farmers to increase output, resulting in increased demand for labour. On the other hand, an increase in the prices of agricultural products, in particular food, could result in a temporary decline in food purchasing power of the agricultural wage.[5] For a given relative price of agricultural product, the higher the labour productivity the more labour services producers are willing to purchase. The aggregate demand for agricultural labour may therefore be specified as:[6]

$$N^d = N^d(AGW, PAPN, AGP) \qquad (1)$$

where *AGW* is the real agricultural wage rate, *PAPN* is the relative domestic price of agricultural to non-agricultural goods, and *AGP* is capital stock in agriculture. The capital stock and state of technology tend to determine the labour productivity in agriculture. A productivity shock increases the demand for labour but also increases the real wage by reducing the price of output. The net contemporaneous effect on demand depends on whether or not the real wage rises by enough to offset the employment effect of the increase in demand.

In formulating the specification for aggregate labour supply, the ingredients of the Harris and Todaro (1970) model are incorporated. They postulated that migration will occur from rural to urban areas if the expected urban real wage rate is greater than *AGW*.[7] Migration is considered in the present analysis as a response to the short-run disequilibrium in inter-sectoral labour markets. In their study of rural-urban migration in Ghana, Rouke and Sakyi-Gyinae (1972) used agricultural wage rates for rural incomes, to show how rural-urban income differentials affect migration in the country.[8] Although other factors such as uncertainty have been mentioned in the literature as additional factors affecting inter-sectoral labour mobility, Larson and Mundlak (1997) have shown in a recent empirical work that the income differential between sectors alone seems to have a large impact on intersectoral migration. This lead them to the conclusion that uncertainty or market imperfections do not seem to have important quantitative impacts on labour mobility.

Assuming that the migration cost is not prohibitive, a change in the expected wage rate in favour of the urban sector affects rural-urban migration in the long-run, which in turn affects the agricultural labour supply in the following manner:[9]

$$\mathrm{D}\, N^{s} = N^{s}\,(\mathrm{D}W_{ru}) \qquad (2)$$

where $\mathrm{D}N^{s}$ indicates a change in the rural labour force through out-migration in response to short-run rural-urban wage differential $\mathrm{D}W_{ru}$ increases in favour of the urban sector. The number of migrants is expected to depend not only on the prevailing wage rate in the urban sector, but also on the rate of unemployment in the sector. The later reflects the probability of obtaining employment in the urban sector.[10] Assuming the real agricultural wage rate is taken as the expected rural wage rate, the expected urban wage can be defined as the prevailing real urban wage *URW* corrected by the urban unemployment rate *UREM*. In the long-run, the increase in labour force in the urban sector and its decrease in the rural sector reduces the sectoral wage gap observed in the short-run. Migration then stops when wages are roughly equal across sectors (Larson and Mundlak, 1997).

The aggregate supply of agricultural labour is then related to the urban wage rate, the rate of urban unemployment and the agricultural wage rate. This may be specified as:

$$N^s = N^s(AGW, URW, UREM) \qquad (3)$$

where *AGW* is as originally defined, *URW* is the real urban wage rate and *UREM* is the rate of urban unemployment. Assuming there is competitive equilibrium in the aggregate agricultural labour market, the reduced-form wage function can be obtained from equations 1 and 3:

$$AGW_t = (AGP_t, URW_t, UREM_t, PAPN_t, Z_t, e_t) \qquad (4)$$

where Z_t represents exogenous shifters not included in equation 4 and e_t is an error term summarizing the influence of all other omitted variables.

While the direct effects of an increase in food prices might reduce real wages in the short-run, the long-run effects of higher agricultural prices may be advantageous to agricultural workers. A shift in the inter-sectoral terms of trade in favour of agriculture will stimulate farmers to increase farm output. This requires an increase in labour inputs, which, in turn, requires an increase in the real wage rate if wages are flexible and the labour supply curve is positively sloped. A number of authors have argued in favour of the positive impacts of increased agricultural prices on real agricultural wage rates (Khan, 1984; Sah and Stiglitz, 1987). Khan (1984) demonstrated in an empirical study that improved terms of trade for agriculture result in higher real wages, while Sah and Stiglitz (1987) suggested, without empirical evidence, that the agricultural wage rate in a poor rural economy responds proportionally to the price of the staple foodgrain.

Capital accumulation and the state of technology are associated with increasing labour productivity that tends to increase real wages. If, however, capital accumulation and technological change involves a shift from labour intensive to labour saving technologies this could as well lead to a decline in agricultural wages. Evidence from other developing countries, however, suggest that on the whole technological change resulted in a net increase of labour demand and real wage rates (Khan, 1984; Boyce, 1989). Thus, a positive productivity shock is expected to have a positive contemporaneous and lagged effects on real wages.

A component of the vector of exogenous variables (*Z*) is the size of the rural population (*RUP*). There is an ambiguous effect of population on the real wage rate. First is the negative effect of population growth primarily attributed to increased labour supply, which *ceteris* paribus, causes the real wage to decline, unless either the demand for labour is perfectly elastic or the supply of labour is unlimited at the ruling real wage. It has, however, been argued that capital accumulation and technological and institutional changes can result in situations in which population growth is accompanied by rising real wages (e.g., Boyce, 1989). Thus, whether induced changes in labour demand could outweigh the effect of increased labour supply upon wages is an open question, the answer to which may vary in the course of time and from place to place. Empirical investigation is essential to assess the relationship between population growth and the real wage rate.

The issue of causality between the right hand variables and the agricultural wage rate will be considered later in this paper, since the exogeneity of some of the variables is questionable. For example, the domestic terms of trade might be thought of as endogenous, with higher agricultural wage rates driving up agricultural prices. Similarly, higher wages could result in higher population densities. This could arise from Malthusian effects (higher wages leading to higher fertility and/or lower mortality), or from inter-sectoral migration in response to wage differentials, or both. Furthermore, as explained in the efficiency wage hypothesis, wages could be a cause of differences in productivity across workers. All these issues indicate that the question of causality remains an empirical matter.

In the following analysis, Johansen's system approach is used to estimate and test the long-run relationship between the real agricultural wage rate and the explanatory variables, the so-called cointegrating relationship. Granger Causality is then explicitly tested to confirm the assumption of exogeneity of the RHS variables. An error correction model (ECM) that incorporates the long-run equilibrium relationship as well as the short-run dynamics of the agricultural labour market in Ghana is then developed.

4. COINTEGRATION ANALYSIS

The concept of cointegration states that if there exists a long-run relationship between two variables, then the deviations from the long-run equilibrium path should be bounded, and if this is the case, then the variables are said to be cointegrated. Cointegration requires that the variables concerned are integrated of the same order. Cointegration analysis therefore begins with an examination of the order of integration of the variables (Banerjee et al., 1993; Muscatelli and Hurn, 1992). Order of integration (stationarity) is commonly established using the Augmented Dicky Fuller (ADF) test and a test recently developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992)--hereafter KPSS test--are used in this study to examine stationarity of the individual series (Silvapulle and Jayasuriya, 1994). While the ADF procedure tests the null of non-stationarity (i.e., $I(1)$) against stationarity (i.e., $I(0)$), the KPSS procedure tests the null hypothesis of a stationary against a non-stationary alternative.

Engle and Granger (1987) show that if two series are cointegrated, then there must exist an error correction representation; and, conversely, if an error correction model (ECM) provides an adequate representation of the variables, then they must be cointegrated. The variables are stationary in the error correction formulation. This provides the basis for using cointegration approach and the ECM model in examining agricultural wage formation, and to seperate the long-run relationships from the short relationships. A representative ECM would reflect the change in one variable to the change in the other variable, to past equilibrium errors, and to past changes in both variables.

Johansen (1988) and Johansen and Juselius (1990) present a methodology for testing parameter restrictions in cointegrated systems. This method enables estimation of the cointegration vectors by maximum likelihood and to determine the number of cointegrating relationships in the data. The procedure also permits one to test linear parameter restrictions on the cointegrating vector(s).

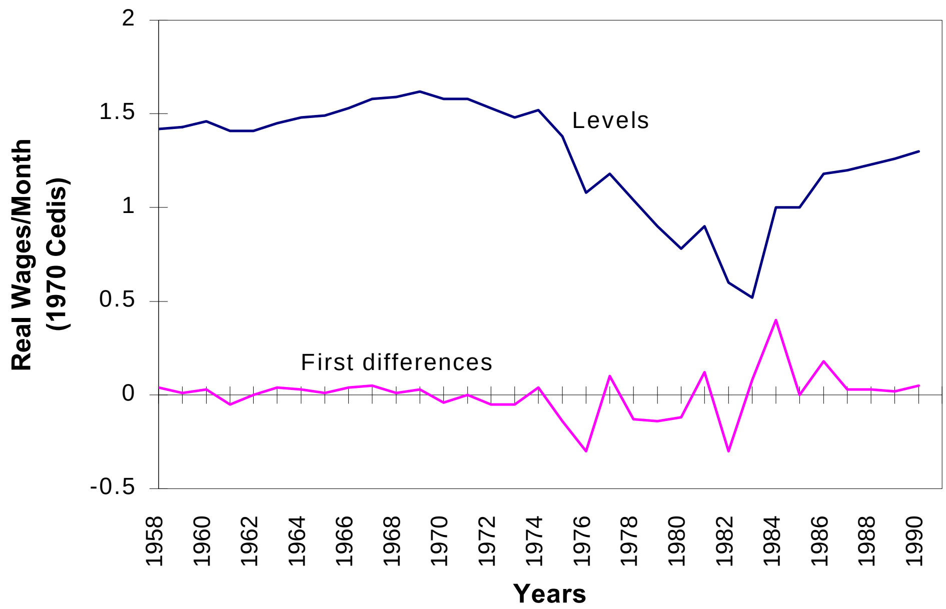
5. EMPIRICAL RESULTS AND DISCUSSION

DATA AND INTEGRATION TESTS

The model is estimated using secondary annual data for the period 1957-91. The real agricultural wage rate is obtained by deflating the nominal annual agricultural wage rate by the Consumer Price Index (CPI) for rural areas, with 1970 = 100. The wage rate in the construction sector is used in this study as a proxy for urban wage rate. This is deflated by the urban CPI to obtain the real urban wage rate.[11] The source of data for nominal wages is the International Labour Organisation's (ILO) *Yearbook of Labour Statistics* and the *Ghana Quarterly Digest of Statistics* (QDSS). The agricultural wage and construction wage rates used in the analysis corresponds to that of basic wage rates paid for normal hours of work. The former includes wages of workers in agriculture, forestry, hunting, and fishing. The wage rates used includes payments in kind, as well as the value of meals and lodging furnished by the employer. The data which is collected through a questionnaire normally takes place during the period from December of a given year to 15 February of the following year. Details of the concepts and sampling techniques can be found in ILO, *Labour Statistics*, Sources and Methods, Volume 2, Employment, Wages, Hours of work and Labour cost.

The agricultural terms of trade indices represent the domestic price index of agricultural commodities divided by the domestic price index of non-agricultural goods (obtained from QDSS). As in Garcia and Llamas (1988), the rate of urban unemployment is defined as the ratio of the difference between total employment and agricultural employment to the difference between the economically active population and agricultural employment. This assumes that the economically active population in agriculture is equal to the number of people employed in the sector. The source of employment data is the *ILO's Yearbook of Labour Statistics* and the *FAO Production Yearbook*. Capital stock in agriculture is obtained by dividing the real value added in agriculture by total employment in the sector. Source of data for the agriculture value added is the *United Nations National Accounts*. All the series are considered in logarithmic form.

Figure 2: Logarithm of Real Agricultural Wage Rates in Ghana



To obtain further insight into the data, the statistical properties (order of integration or stationarity) of the variables are examined. As stated in section 4, the order of integration of the variables are determined empirically by the ADF test and the KPSS test. The logarithm of real agricultural wage rate and its first-differenced series are presented in Figure 2. Visual inspection of Figure 2 suggests that *AGW* series are non-stationary but the first-differenced series are stationary. Logarithmic series of the other variables and their first-differences also exhibit similar patterns.[12]

The empirical results are presented in Table 1. Akaike's Information Criterion is used to determine the appropriate lag-length truncation in each case.[13] The results of the ADF and KPSS tests reveal that while the levels of each of the series are non-stationary, the first differences are stationary. The results are therefore consistent with the hypothesis that non-stationarity characterises each of the series in this study. To ensure that structural breaks in the series themselves are not underlying the non-stationarity in the series, Perron (1989) test for unit roots with structural break was employed to examine the individual series. The test statistics shown in the Table under PERRON—with 1973 as break point—suggest that the null hypothesis of unit root could not be rejected in all the series at the 5 percent level, thus validating the ADF and KPSS tests.[14]

Table 1: Results of stationarity tests for the series †

		Levels			First-differences		
Series	**lag-length**	**ADF**	**KPSS**	**PERRON**	**ADF**	**KPSS**	**PERRON**
AGW	1	-2.19	0.659	-0.91	-6.72	0.098	-4.13
URW	2	-1.47	0.568	-1.46	-7.51	0.109	-3.86
RUP	2	-2.24	0.703	-2.01	-5.86	0.124	-5.89
UREM	3	-2.92	0.624	-1.16	-6.53	0.143	-4.27
PAPN	2	-2.35	0.856	-0.84	-5.62	0.131	-3.98
AGP	2	-1.81	0.510	-1.63	-4.28	0.189	-4.52

† All variables are in logarithms. The ADF technique tests H_0: $X_i \sim I(1)$ against H_1: $X_i \sim I(0)$, while the KPSS procedure tests the reverse. The 5 percent ADF and KPSS critical values are -3.07 and 0.463, respectively. Critical values for the ADF test are obtained from Mackinnon (1991). The critical value for the Perron test (null hypothesis of unit root) calculated for λ=0.5, at the 5 percent level is -3.76. The corresponding break fractions are calculated as $\lambda = Tb/T$, where Tb and T represent the number of observations until the breakyear and the whole sample size, respectively.

LONG-RUN RELATIONSHIPS

Having found that all variables are *I*(1), the Johansen procedure is employed to establish the presence of cointegration among the variables, to estimate the associated cointegrating vectors. Boswijk and Franses (1992) emphasise that the results from VAR models are sensitive to lag-length choice. They therefore suggest the application of the Johansen procedure for the determination of the different lag-lengths and to base the final choice of parameters on the absence of serial correlation in the residuals and the significance of parameters at higher lags using LR-tests. This procedure was adopted in the presented analysis. A VAR model incorporating two lags of each variable is selected from the test applied.

Table 2: Johansen-Juselius test statistics model with two lags and 1983 step dummy variable

Testing for the number of cointegrating vectors				
Null hypothesis	**λ_{max} statistic**	**λ_{max} (0.95)**	**Trace statistic**	**Trace (0.95)**
$r = 0$	50.26	40.30	123.33	102.14
$r \leq 1$	31.48	34.40	66.74	76.07
$r \leq 2$	22.39	28.14	49.01	53.12
$r \leq 3$	17.06	22.00	28.24	34.91
$r \leq 4$	9.78	15.67	12.16	19.96
$r \leq 5$	0.96	9.24	0.28	9.24

† The critical values, λ_{max} (0.95) and Trace (0.95) are taken from Osterwald-Lenum 1992.

Table 2 presents the results of the cointegration analysis.[15] The calculated λ_{max} and trace statistics are compared with the critical values of these test statistics. Both λ_{max} and trace test suggest the presence of only one cointegrating vector. This vector can be interpreted as an estimate of the long-run cointegrating relationship between the variables concerned.

The sequential Chow test was employed to test the stability of the cointegrating vector (Gregory et al., 1996; Johansen and Strom, 1997). The test yielded $F(6, 23) = 4.65$ which is above the critical value, suggesting a shift in the long-run relationship in 1983. A step dummy and a trend were tried for the post-1983 period, and the step dummy was the preferred choice. The cointegration relationship was re-estimated in the multivariate cointegration system including the step dummy (Table 2).

The long-run relationship normalised by the agricultural wage rate[16] can be represented as:

$$\begin{aligned} AGW = {} & 0.48\, AGP + 0.83\, PAPN + 0.56\, URW - 0.29\, UREM \\ & - 0.42\, RUP + 0.036\, ERP + 0.015 \end{aligned} \tag{5}$$

This is a stable equilibrium relationship to which the variables have a tendency to return in the long-run (Engle and Granger, 1987). Real agricultural wage rates increase by 0.48 percent if the capital stock in agriculture increases by 1 percent; they increase by 0.83 percent if the domestic terms of trade improve by 1 percent in favour of agriculture; they increase by 0.56 percent if the urban wage rate rises by 1 percent; they decline by 0.29 percent if the rate of urban unemployment rises by 1 percent; and they decrease by 0.42 percent if the rural population increases by 1 percent. The positive coefficient of the step dummy suggests a *permanent* increase in the real agricultural wage rate by 3.6 percent as a result of the economic reforms. As indicated earlier, the economic reforms have created a macroeconomic environment that is conducive for sectoral and general economic growth. Per capita agricultural production which declined by an annual average of about 3 percent between 1970-81, increased by an annual average of almost 5 percent between 1982-1991, while domestic terms of trade have improved in favour of agricultural commodities. In the four years after 1983, the price received by farmers for a tonne of cocoa—which constitutes a large proportion of the agricultural price index—rose in real terms by a cumulative 174 percent before falling in real terms by a cumulative 34 percent during the next three years as price increases failed to compensate for rises in the consumer price index and the overall impact of cocoa taxation (World Bank, 1991). Compared to the enormous decline in agricultural wages in the 1970s, this positive impact of economic reforms is quite modest. It, however, represents a reversal of the declining real wage rates that was observed prior to the adjustment programme.

Since the coefficient of the domestic terms of trade in the estimated long-run relationship is 0.83, a linear hypothesis (H_0^*; $b_{1,2} = b_{1,6}$)[17] was formulated to investigate whether this coefficient is equal to unity. This was to verify if improvements in domestic terms of trade resulting from better macro-economic and sector-specific policies are completely passed on to agricultural workers in the form of higher wages in the long-run. A likelihood ratio of 5.06 was obtained, which exceeds the critical value of $\chi_{0.05}^2 = 3.84$. The null hypothesis of a complete transmission of improvement in domestic terms of trade effects on the agricultural wage rate was therefore rejected.

6. GRANGER CAUSALITY

The existence of a cointegrating vector implies Granger causality between the variables.[18] Causality is examined in this study with a test of zero restrictions in the Π matrix. This is achieved through testing for zero restrictions in α matrix using standard tests within the Johansen framework (Hall and Milne, 1994). If the α matrix has a complete column of zeros, then no cointegrating vector will appear in a particular block of the model, thus indicating no causal relationship. The restrictions are tested by direct Wald tests on the loading parameters. The results, shown in Table 3, indicate that the main direction of long run causality flow from *URW*, *RUP*, *UREM*, *PAPN*, and *AGP* into *AGW* and imply that the disequilibrium in the *AGW* does not feed back into the righthand side variables.

Table 3: Loading Weights for the Model

Loading Weights					
AGW	**URW**	**RUP**	**UREM**	**PAPN**	**AGP**
0.098	-0.019	0.006	0.032	-0.04	-0.026
Wald test of the significance of the loading weight (χ^2_1)					
8.62	1.24	0.85	1.53	2.02	1.97

Note: Critical value at 5 percent level of significance is $\chi^2_1 = 3.84$

These findings suggest that the righthand side variables can be treated as weakly exogenous. In particular, the ability of governments to effectively set prices by macroeconomic and sectoral policy reforms suggests that domestic terms of trade are mainly influenced by government policies. *PAPN* thus becomes an exogenous "policy lever" in determining real agricultural wage rates.

7. SHORT-RUN WAGE ADJUSTMENTS

In this section, an analysis of the short-run effects, with the purpose of imposing more economic structure and at the same time obtaining parsimony is undertaken. In line with the error correction framework used here, the general structure of the short-run dynamics may be represented as (Engel and Granger, 1987):

$$\begin{aligned}\Delta AGW = {} & \alpha_0 + \sum_{i=1}^{k}\alpha_1 \Delta AGW_{t-i} + \sum_{i=1}^{k}\alpha_2 \Delta PAPN_{t-i} + \sum_{i=1}^{k}\alpha_3 \Delta AGP_{t-1} \\ & + \sum_{i=1}^{k}\alpha_4 \Delta AGP_{t-i} + \sum_{i=1}^{k}\alpha_5 \Delta RUP_{t-i} + \sum_{i-1}^{k}\alpha_6 \Delta URW_{t-i} \\ & - \lambda\,(AGW_{t-1} - \hat{\gamma}_1 PAPN_{t-1} - \hat{\gamma}_2 AGP_{t-1} - \hat{\gamma}_3 UREM_{t-1} \\ & - \hat{\gamma}_4 RUP_{t-1} - \hat{\gamma}_5 URW_{t-1}) + \varepsilon_t \end{aligned} \qquad (6)$$

where ε_t is a white noise disturbance term and *k* is lag length. The term inside the brackets provide the error correction mechanism, with λ representing the error correction term. If the real wage rate falls below its long-run equilibrium level at time *t*-1, the term in the brackets become negative. However, because λ is negative, its effect at time *t* is to increase the growth rate of the observed wage towards its steady state path. The short-run effects of the changes in the right hand variables on *AGW* are captured by the coefficients (a). Given the existence and importance of wage contracts in agriculture, both current and lagged effects will have impacts on wage rates. This stems from the fact that contractual nominal wages are set at the rate that is expected to clear the labour market, given information available at the time the contract is negotiated. So, for some firms, the wage rate will deviate from its full-information, market clearing level if shocks occur in period *t* that were unanticipated at the end of period *t*-1. For others, with contracts drawn up at the end of *t*-2, unanticipated shocks in period *t*-1 as well as those in period *t* will affect period *t* wages. To this effect, Sargent (1978) argues that failure to appropriately account for lagged relationships in real wage equations can lead to incorrect conclusions

The estimated model is derived using Hendry's general to specific modelling, which starts from a general model, in which the variables are suggested by economic theory, and then narrows down this general model by looking for simplications that are acceptable, given the data (Hendry et al., 1984). The general model was estimated with lagged terms of the variables by Ordinary Least Squares, and model selection test performed to remove irrelevant variables that the model contains. This was done by sequentially imposing restrictions, testing each restriction by the *F* test for significance against the slightly less restricted model preceding it in the sequence. The final preferred equation was subjected to diagnostic checks, to ascertain the adequacy of the specification chosen. The selected agricultural wage equation turned out to be (*t* statistics in parentheses):

$$\begin{aligned}\Delta AGW_t = {} & \underset{(4.82)}{0.57}\Delta AGW_{t-1} + \underset{(5.03)}{0.48}\Delta PNPA_t + \underset{(3.21)}{0.29}\Delta PNPA_{t-1} \\ & + \underset{(2.58)}{0.28}\Delta URW_t - \underset{(2.13)}{0.19}\Delta URW_{t-1} - \underset{(1.89)}{0.12}\Delta UREM_t \\ & - \underset{(2.17)}{0.16}\Delta RUP_t + \underset{(3.82)}{0.35}\Delta AGP_t + \underset{(1.78)}{0.24}\Delta AGP_t \\ & - \underset{(2.64)}{0.40}\Delta U_{t-1} \qquad (7)\end{aligned}$$

Goodness of fit is acceptable with an R^2 of 0.749. The Durbin-Watson test gives $DW = 1.92$.[19]

The cofficient estimates are consistent with the theoretical framework and provide interesting dynamic properties. The short-run elasticities are smaller than the long-run elasticities implied by the model in levels. Growth in real wage in the previous year has a positive effect on current wage rate. The direct impact of a 1 percent increase in the relative prices of agricultural goods is to increase real agricultural wage rate by 0.48 percent. The short-run effect is reduced after one year. The current short-run effect is smaller than the long-run effect from the cointegrating regression. This result is consistent with previous work done by Khan (1984) who reported significantly positive response of agricultural wage rates to favourable agricultural terms of trade. It contrasts however with the findings of Boyce and Ravallion (1991), who obtained significantly negative short-run responses of agricultural wage rate to relative prices of agricultural commodities.[20] The significantly positive coefficients of current and lagged domestic terms of trade variables lend support to the widespread notion that policy manipulation that depresses agricultural prices work against the economic welfare of the rural population, by reducing the real wage rate. The finding also suggests that the indirect effect of increases in agricultural prices tend to outweigh direct effects that might lead to a decline in real agricultural wage rates.

The data used in the analysis show that relative to the price of non-agricultural goods, agricultural prices have fluctuated markedly and have shown a tendency to increase over time. There was a gradual increase in the prices of agricultural goods relative to non-agricultural goods throughout the 1960s. The terms of trade, however, turned against the agricultural sector during the early 1970s as real cocoa producer prices declined. An improvement in favour of the agricultural sector took place in the early 1980s, as food prices increased more than the increase in the prices of non-agricultural items. The initial imapct of the adjustment programme was, however, to shift the terms of trade in favour of the non-agricultural sector, as the realignment of the domestic currency resulted in substantial increase in prices of imported non-agricultural goods. The second half of the 1980s again saw relative prices increasing in favour of agricultural goods. The terms of trade improved by about 22 percent between 1985-88, and almost 9 percent between 1988-1990, as cocoa producer prices experienced increases in real terms. This development suggests that incentives created for the agricultural sector, particularly improved cocoa prices, helped increase real agricultural wages.

The positive and significant coefficient of the urban wage variable indicates that an increase in the urban sector wage rate attracts workers out of agriculture, thus bidding up the wage rate in that sector. This finding lends support to Dautey's (1979) hypothesis--although not empirically tested--that higher wages in the urban sector tend to exert an upward pressure on wages in the agricultural sector of the country. It is also in line with the positive relationship between manufacturing wage rate and agricultural wage rate reported by Boyce and Ravallion (1991). Using data collected by the World Bank, Jaeger (1992) demonstrated that the narrowing of the rural-urban wage gap in the country since 1983 resulted in a significant reverse migration from urban to rural areas.

Contemponaneous increases in the rate of urban unemployment decrease real wage growth by 0.19 percent. This effect is moderated in a year later, but the long-run effect is much greater. This finding suggests that increased urban unemployment, *ceteris paribus*, lowers a prospective migrant's perception of the probability of obtaining a job in town, encourages people to stay in rural areas, and thus increases agricultural labour supply over what it would have been otherwise, exerting a downwards influence on agricultural wages.

The error correction coefficient is statistically significant, indicating that last period's equilibrium error has significant impact on subsequent wage changes in the agricultural labour market. The coefficient of -0.40 shows that less than half of the last-period's equilibrium error is eliminated within a year, suggesting that full adjustment to long-run equilibrium is never completed within one year. This is consistent with the high transaction costs--from poor road and transport infrastructure and communication--in rural labour markets of under-developed economies such as Ghana (Oyejide, 1986; de Janvry et al., 1991).

The capital stock variable also appears to have a positive current and lagged effects on wages, indicating that changes in labour demand arising from exogenous shocks on the agricultural supply side had perceptible effects upon agricultural wages. The coefficient suggests that a 1 percent change in the capital stock will result in a 0.35 percent change in the real agricultural wage rate. This finding is in line with the thesis that investment in agriculture that improves productivity contributes to the labour demand effect in increasing wages of farm workers. Public investments on feeder roads, waterways, agricultural research, and irrigation schemes tend to favour productivity growth in the sector. Under the economic reforms, the government has increased investments in agriculture to boost productivity growth through a more effective tax collection system, and a reduction in the expenditure bias against agriculture. While resources were previous transfered from the sector through taxes to support urban development, expenditures for feeder roads and agricultural research has increased in real terms since the implementation of the adjustment programme (Word Bank, 1994).

The coefficient for the size of the rural population variable has a negative sign, and is statistically significant at the 5 percent level, lending support to the view that as the rural population grows, the supply of agricultural labour rises, tending to exert a downward pressure on agricultural wages in the absence of alternative employment in the rural area, or presence of high urban unemployment. This suggest that the labour supply effect of rural population growth upon agricultural wages outweighs the labour demand effect. This finding contrasts with that of Boyce (1989), who reported a positive relationship between population growth and real agricultural wage rates for Bangladesh, but is consistent with that of Palmer-Jones (1993) who found a negative relationship between the variables.

As evident in Figure 2, real agricultural wage rates have been on an upward trend since 1983, in spite of the rising size of the rural population. The effects on aggregate labour demand for agriculture of improved domestic terms of trade for agriculture, gains in agricultural productivity and increases in non-agricultural wages, have presumably outweighed the effects of any increases in labour supply due to rapid population growth over the period (Sarris and Shams, 1992). The direct effects of the increase in real agricultural wages and the indirect general equilibrium effects such as rising non-farm employment and income have contributed to poverty reduction in the country. According to recent estimates by the World Bank (1995), Ghana's poor fell from 37 percent to 32 percent of the population between 1987 and 1992, with rural poverty falling from 42 percent to 34 percent within the same period.

8. CONCLUSIONS AND POLICY IMPLICATIONS

The main contribution of the paper is an examination of the process and time period of adjustments of real agricultural wage rates in Ghana over the 1957-1991 period. Johansen's maximum likelihood cointegration approach is used to determine the causal relationship between the real agricultural wage rate, the agricultural terms of trade, labour productivity in agriculture, the rate of urban unemployment, the urban wage rate and the size of the rural population. Once the presence of cointegration was recognised, the relevant error correction model that incorporates the long-run equilibrium relationship and the short-run dynamics was estimated.

The results suggest that there is only one cointegrating relationship among these variables. This implies the presence of a stable equilibrium relationship to which these variables have a tendency to return in the long-run. The results permit several conclusions. The most important policy conclusion is that governments can have a strong influence on labour earnings in agriculture in particular, and rural poverty in general, in economies such as Ghana. Almost half of an increase in domestic prices in favour of agriculture is passed on in the agricultural wage within the current year. This rises to over four-fifth in the long-run (more than one year), without complete transmission. This evidence is consistent with the notion that policies that tend to suppress domestic agricultural prices relative to the non-agricultural sector, not only discourage agricultural production, but also depress real wages in the sector and contribute to rural poverty. Conversely, policies that favour agricultural prices relative to non-agriculture still raise agricultural wages, even if they succeed in depressing the price of non-tradables overall.

The second conclusion is that promoting investments that increase average labour productivity in agriculture by 1 percent will increase the real agricultural wage rate by 0.35 percent in the short-run and 0.48 percent in the long-run. This result seems plausible but has rarely been quantified in West Africa. It should be noted that this very positive argument for increased attention to investment in agricultural productivity actually understates the full benefit of investing in agriculture, since it relates to average wages, and does not include the extra benefits of more people staying in rural areas than would otherwise have been the case. It also does not assess the spin-off effects of net increases in rural incomes from spending on non-tradables, which can be expected to create additional employment and income in rural areas.

Third, it is also evident from the empirical results that appropriate macroeconomic and sector-specific policies can provide substantial latitude for improving earnings in the agricultural sector. The effect of the economic reform programme has been to shift real agricultural wages upwards by 3.6 percent. Low agricultural productivity and unfavourable domestic terms of trade for agriculture that obtained in the 1970s and early 1980s was partly responsible for the massive decline in real agricultural wage rates within this period.

Finally, the overall performance of the model suggests that secondary data of the type available in government offices in Ghana can provide useful information on strategic matters such as real agricultural wage determination. The extra benefit of having a long time series is worth the increased measurement error, provided that an adequate time series estimation procedure is used here. Going beyond the main conclusions above, however, will require micro surveys of rural employment and labour markets, social and demographic characteristics of labourers, as well as the terms and conditions of contracts in the labour and related markets.

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10. ENDNOTES

[1] The agricultural wage includes wages of workers in agriculture, forestry, hunting, and fishing. The wage rates used includes payments in kind, as well as the value of meals and lodging furnished by the employer. The real wage is obtained by deflating the nominal annual wage rate by the Consumer Price Index for rural areas, with 1970 = 100. See information in data section for sources and details.

[2] For a detailed discussion on the impacts of rural-urban migration on agricultural production in Ghana, see, for example, Abdulai (1999).

[3] This is satisfactory for rural labour markets in Ghana, although it is less satisfactory for urban labour markets over the period studied. Because of a general excess supply of urban labour in Ghana, the urban wage rate can be assumed to be exogenous to the rural labour market, whereas rural wage rate are endogenous to rural-urban migration.

[4] Note that the assumption of a competitive labour market *per se* does not necessarily guarantee that labour is efficiently allocated, since absence of credit or land markets may create a situation as in Kenya, where labour and other

factor markets are linked (Collier and Lal, 1989). The assumption in our model is that effective supply and demand for labour match in equilibrium.

[5] Thus, while the direct effect of a an increase in food prices is a reduction in thereal wage in the short-run, the indirect effect of increased demand for agricultural labou will offset this with time.

[6] The demand for labour is obtained by equating the real wage and the marginal product of labour.

[7] A number of studies have shown rural-urban income differentials to be the driving force behind observed rural-urban migration in Ghana (e.g., Dautey, 1979; Kasanga and Avis, 1988).

[8] This was done in the context of the Harris and Todaro (1970) model, where real agricultural wage rate is used as a proxy for the rural wage rate. Sarris and Shams (1992) also use the real agricultural wage rate to show how economic reforms have affected rural wage rates and poverty.

[9] Kasanga and Avis (1988) demonstrated how rural-urban migration has affected agricultural labour supply in Ghana.

[10] This implies that when the wage differential is high, it pays to migrate even when the probability of getting a job is less than unity.

[11] It is, however, significant to note that even the expectation of public sector employment may be the key engine of migration. In this case, the formal manufacturing sector wage rate, although difficult to be considered as a market wage could be employed. As pointed out by an anonymous reviewer, for the purposes of migration it does not matter whether the reference wage is market-determined or not.

[12] These figures are not included in this paper, but are available from the authors on request.

[13] Other tests such as Schwarz-criterion and Hannan-Quinn Criterion produced similar results.

[14] The results obtained by using 1983 as break point also indicated the series contained unit roots.

[15] The residuals passed the Jarque-Bera normality test and are uncorrelated according to the Box-Pierce test.

[16] This is the vector obtained from the largest eigenvalue. For space reasons, the eigenvalues are not presented. The eigenvectors are presented in the next section examining causality.

[17] Referring to Table 2, $\beta_{1,2}$ refers to *PAPN* and $\beta_{1,6}$ to *AGW*.

[18] As pointed out by Hall and Milne (1994), in its broadest sense causality is a term which can have no precise empirical implementation. The term causality in this paper refers to a variant of Granger causality, that is *X* Granger causes *Y* if a change in *X* generally predates a change in *Y*.

[19] The Lagrange Multiplier tests for first and second order serial correlation yield χ^2_1 = 0.102 and χ^2_2 = 0.975, well below the relevant critical values. The *RESET* test for functional misspecification yields 2.31, showing no evidence of misspecification at conventional significance level. The Jarque-Bera test for normality yielded a value of 1.86, accepting the hypothesis of normally distributed residuals. The test for *ARCH* residuals gave 1.24, indicating that *ARCH* is not a problem in this model.

[20] The indexing of the agricultural wage rate in Boyce and Ravallion (1991), with the price of the agricultural good is a contributory factor to the negative sign

obtained by them, as they point out. When wages are measured by the purchasing power of the agricultural good, then the short-run effects of higher relative prices for food on real wages would be unambiguously adverse.