Trade Reform and Wage Inequality in Kenya, 1964-2000*

by

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Abstract

This paper analyses the evolution of wage inequality in Kenya between 1964 and 2000. Our measure of wage inequality is the ratio of wages in manufacturing to wages in agriculture, which can be seen as an indicator of sectoral wage-inequality or as a proxy for skilled to unskilled wages. We find that changes in relative wages have primarily been driven by the degree of openness, while other factors such as the capital-labour ratio, educational attainment, relative labour-productivity, and the ratio between agricultural and manufacturing prices had no significant effect. We conclude that international market integration has reduced wage-inequality in Kenya.

JEL Classification: D13, F13, O55.

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1. Introduction

In recent years there has been extensive research on changes in wage inequality in response to rising global trade. Most studies have dealt with highly industrialised countries, but there have been a number of studies of less-industrialized Latin American and Asian countries as well (Williamson, 1997; Arbache, Dickerson, Green, 2004; Goldberg, Pavcnik, 2004). Due to lack of good data there have so far been few studies of African economies, but we have been able to analyse the impact of greater openness to trade on the evolution of wage-inequality in Kenya.

Since poverty is more prevalent in Africa than in any other continent, it is particularly important to take the distributional impact of economic reforms into account. Not surprisingly, the debate about the impact of structural adjustment and trade-reforms on income-distribution in Africa has been heated. Sahn, Dorosh, and Younger (1996, p. 719) concluded on the basis of extensive empirical work[1] that “in most countries, adjustment policies, when implemented and sustained, improve income distribution and do not adversely affect the poor". However, De Maio, Stewart, and van der Hoeven (1999), and others, have found evidence of increasing inequality in several countries. They argue that the distributional impact of reforms has often been negative, but it is hard to be sure about policy impacts without a counterfactual. We attempt in this paper to isolate the impact of trade-reform in Kenya during 1964-2000 on one aspect of inequality, wage-inequality, while controlling for other important determinants.

Section 2 provides a brief theoretical review. Section 3 then describes the evolution of the Kenyan labour market and introduces our wage-series. Section 4 presents our data on changes

[1] There is a comprehensive presentation of the results in Sahn, Dorosh, and Younger (1997).

in factor endowments in Kenya, while Section 5 provides information about the extent of international economic integration of Kenya during the period. Section 6 presents our econometric analysis of the determinants of wage inequality in Kenya, and Section 7 summarizes and draws conclusions.

2. Theory

Most discussion of wage-inequality has been based on the Heckscher-Ohlin model, which says that a country tends to export goods that use its abundant factor intensively and import goods that use its scarce factor intensively (Leamer, 1995; Wood, 1997). However, in a simple 2-country, 2-factor (skilled and unskilled labour), 2-good model, barriers to trade drive wedges between the prices of the two goods in the two countries. The price of the unskilled labour-intensive exportable may be kept low in the less developed country, and vice versa for the industrialized country. A move from autarchy to free trade, or from high tariffs to low tariffs, can then lead to increased production of unskilled-labour-intensive exportables in the poor country, which in turn leads to increased demand for unskilled labour and reduced demand for skilled labour. In accordance with the Stolper-Samuelson theorem, we would thus expect to see increased unskilled wages in Kenya in response to recent trade reforms. This result would not change fundamentally if we had more tradable goods (see Woods, 1997), although the magnitude of the effect of lower tariffs on wages might be smaller.

The effect of introducing non-traded goods into the model will depend on the pattern of consumption substitution between traded and non-traded goods. If the poor country produces a non-traded good which uses unskilled labour intensively, and demand for this good is a close substitute for the more skill-intensive of the two traded goods, then opening up will reduce the price and increase sales of the skill-intensive tradable. The resulting drop in demand for unskilled labour in non-tradables production could then be larger than the increase in demand for unskilled labour in the labour-intensive tradables sector. Thus it is theoretically possible (though not likely) that there would be a drop in unskilled wages.

Adding another factor to the model might also complicate matters. If our poor country has an abundance of land it might have a comparative advantage in the production of a land-intensive tradable. Greater openness would then increase production of the land-intensive tradable. If land is a complement to skilled labour in production, then this would increase the demand for skilled labour. This might lead to an unexpected result in terms of relative wages. In a poor LDC, however, land is probably complementary to unskilled labour, and thus introducing it into the analysis would not generate such results.

If instead we added the factor capital to the model it would be more reasonable to assume that it is complementary to skilled labour. Assuming the less developed country does not have a comparative advantage in production of skill-intensive goods, this would not change our prediction for the effect on wage-inequality. Just as in our original model, we would expect to see increased unskilled wages in response to greater openness to trade.

There is an abundance of evidence that exporting sectors in LDCs are generally less skill-intensive than import-competing sectors (see survey in Wood, 1997). The evidence for African countries specifically is limited, but seems consistent. In accordance with the model outlined above and the Stolper-Samuelson theorem, unskilled wages should then increase relative to skilled wages when the economy opens up.

The standard Stolper-Samuelson effect is not the only one that can influence wage inequality. An alternative channel through which trade reforms can affect wage inequality is via changes in industry wage premiums. Goldberg and Pavcnik (2004) note that industrial wage-premiums account for a significant portion of wage-inequality in poor countries If there are labour-market rigidities hindering smooth reallocation of labour across sectors, this channel might be important. Sectoral adjustment to tariff-changes might then come via changes in wages rather than changes in employment. A tariff-cut in the Kenya might therefore also in this case translate into a fall in the wage-premium in manufacturing.

If there is imperfect competition in product- and labour-markets, profitable protected industries might have shared their rents with their employees, because of union power, for example. If reduced tariffs now force firms to accept lower mark-ups, manufacturing wages will be forced down as well via these lower industrial rents. There is substantial evidence (e.g. Harrison, 1994) that mark-ups have in fact declined in response to trade-liberalization.

On the other hand, increased openness could lead to changes in the availability of technology or to productivity-improvements in industries exposed to stiffer competition. If the externally induced technological changes are skill-biased, or if productivity-gains are shared with workers, then lower tariffs might lead to higher wage-premiums, and tariff reduction could increase the relative wages of skilled labour. Arbacha, Dickerson, and Green (2004) found that this happened in Brazil. However, Kenya has had very limited inflow of foreign direct investment or foreign technology, so this does not seem likely to have happened there (Bigsten, Kimuyu, 2001).

To summarize, then, if tariffs are reduced in manufacturing in Kenya, we would expect a reduction in wage-inequality either because of the Stolper-Samuelson effect or if the lower tariffs reduce the rents in the previously protected sector. There could be a contrary effect if lower tariffs induced productivity-improvements in the previously protected sector, but we would not expect that effect to dominate in Kenya. Our presumption is thus that lower tariffs have reduced wage-inequality in Kenya.

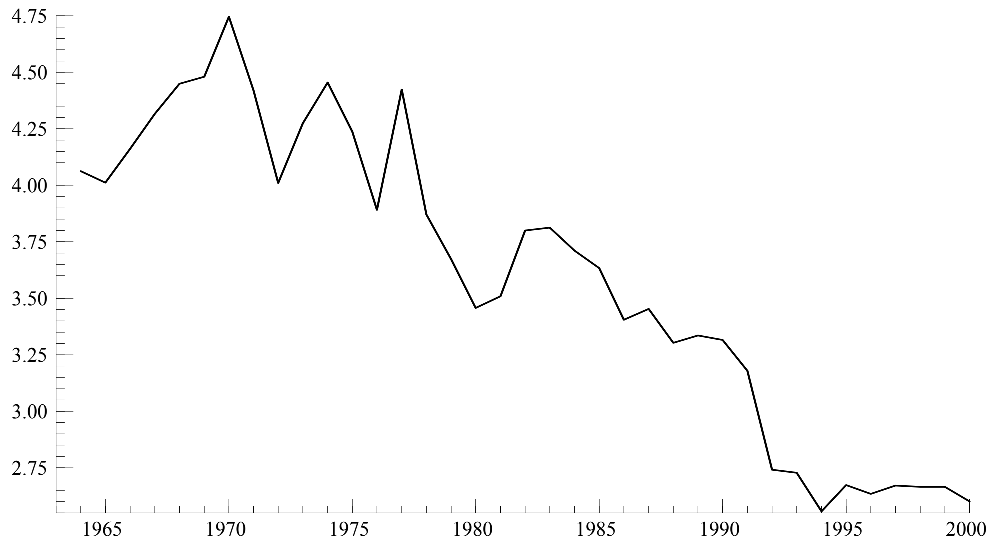
3. Relative Wages and the Labour Market

Time series studies of wage inequality have been plagued by several problems. One is that it has been hard to get time-series data for a well-defined measure of wage-inequality. The Heckscher-Ohlin model discussed above would require data on skilled wages relative to unskilled wages, but such time series data by levels of education are not available in Kenya. We therefore choose to use manufacturing employees and agricultural employees as proxies for skill-categories. These groups are not homogenous, but agricultural workers certainly have lower education on average.

Figure 1 shows our time series of relative wages in Kenya. From 1965, until the late 1970s, manufacturing wages were roughly four times as high as agricultural. During the 1980s and the beginning of the 1990s the ratio declined, and has been about 2.6 times since 1994.

One might wonder whether wages in agriculture also reflect wages in the sprawling informal sector. There is no systematic time-series evidence on informal sector wages in Kenya, but surveys from 1999 show that average formal-sector income then was slightly more than twice that in the micro- and small-enterprise sector (Kenya, 1999). Since average wages in the formal sector are close to those in manufacturing, there is at least some correspondence between agricultural-sector wages and informal-sector wages. Hence, it is not unreasonable to argue that agricultural wages reflect the wider labour-market for less skilled workers.

Figure 1: Manufacturing wages relative to agricultural wages, 1964-2000



Note: The data are from various issues of Statistical Abstract of Kenya.

As Figure 1 shows, the gap between agricultural and non-agricultural wages increased rapidly during the1960s, largely because of a wage-policy aiming to increase urban formal-sector wages. With Independence in 1963, Kenya needed much more qualified manpower in the public sector, which increased its relative wages dramatically. As a consequence of this, private sector real wages also rose. [2] For skilled workers, the late 1960s were a period of rapid increases in earnings, but this trend reversed during the 1970s, when private sector real wages fell. Since then formal sector employment has grown very slowly relative to informal-sector employment. This may reflect wage-rigidities that keep wages above market-clearing levels in

[2] Between 1963 and 1965 public sector real wages increased by 48 per cent, while real wages increased by merely 6 per cent in the private sector (Collier, Lal, 1986, p. 62).

the formal sector, or economic advantages to conducting business in the informal sector (Bigsten, Kimuyu, ánd Lundvall, 2004).[3]

Relative wages can also be affected by changes in labour-market regulation and institutions. Before independence formal-sector minimum wages were important, and increased rapidly until about 1968. Then the new government sought to control the trade union movement, and by the early 1970s many workers' rights conceded by the colonial government had been circumscribed (Collier and Lal, 1986). Trade unions in Kenya are financially weak and their ability to strike is limited. Since the 1960s there have been a series of "tripartite agreements" between workers, employers, and government, in which employers pledged to increase employment by a certain percentage if workers refrained from demanding wage-increments over a given period. An industrial court was also established to help resolve disputes between organised labour and employers, including implementation of the government's wage-guidelines.

Labour-market reforms started as part of the structural-adjustment efforts in the 1980s and became more extensive in the 1990s. In 1994 it became much easier for employees to negotiate wage increases, as trade unions were allowed to seek full compensation for price increases without being hindered by wage-guidelines (Manda and Sen, 2004). Redundancy laws were amended in 1994, which made it possible for employers to lay off employees.

[3] Real wages in the formal sector have recently been increasing rapidly, which suggests that the market is not reacting to the supply pressure. During 1998-2000 real earnings in the formal sector increased by 22.2%, 14.2% and 5.5% respectively (Kenya, 2001, p. 56). It thus seems as if workers in the informal sector cannot underbid those in the formal sector.

4. **The International Economic Integration of Kenya**

We are exploring how changes in international economic integration have affected factor-rewards in Kenya. International factor mobility has been limited, so the link to the world market has essentially come via goods prices.

Kenya inherited a policy of import substitution from the colonial government, but from about 1967 on it was pursued more vigorously. This was particularly the case after the foreign-exchange crisis of 1971 when the government chose to introduce strict import-controls rather than to devalue and undertake macroeconomic adjustment. This policy was temporarily relaxed in 1976, when the coffee boom led to a massive inflow of foreign exchange. The boom lasted just a few years, but as coffee prices came down and oil prices shot up in 1979 the government was forced to embark on a policy of structural adjustment, including trade-liberalization.[4] In November 1981 the government stopped requiring "no-objection certificates" from domestic producers[5], gradually replacing quantitative restrictions with equivalent tariffs, followed by tariff-reductions and rationalizations. Macroeconomic problems made the government halt some of the reforms in 1982, but the reform process started slowly again in 1983. The reforms gathered pace in 1987-88, and this time they were more successful. Between then and 1997-98 maximum tariffs were reduced from 170% to 25%, and the number of tariff bands reduced from 24 to 4; the average tariff fell from 49% to 17% (O'Brien and Ryan, 2001). Import-licensing schedules were abandoned in 1993, and in 1993-94 virtually all current and capital account transactions were fully liberalized. This is also the year in which Kenya was classified as "open" according to Sachs and Warner (1995).

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[4] Glenday and Ryan (2003) discuss the stages economic liberalization in Kenya and the links between trade liberalization and economic growth.
[5] Domestic producers of specific products could object to import of competing goods.

There was another policy-reversal in 1997, however, in conjunction with loose macroeconomic policy in the run-up to the election that year.

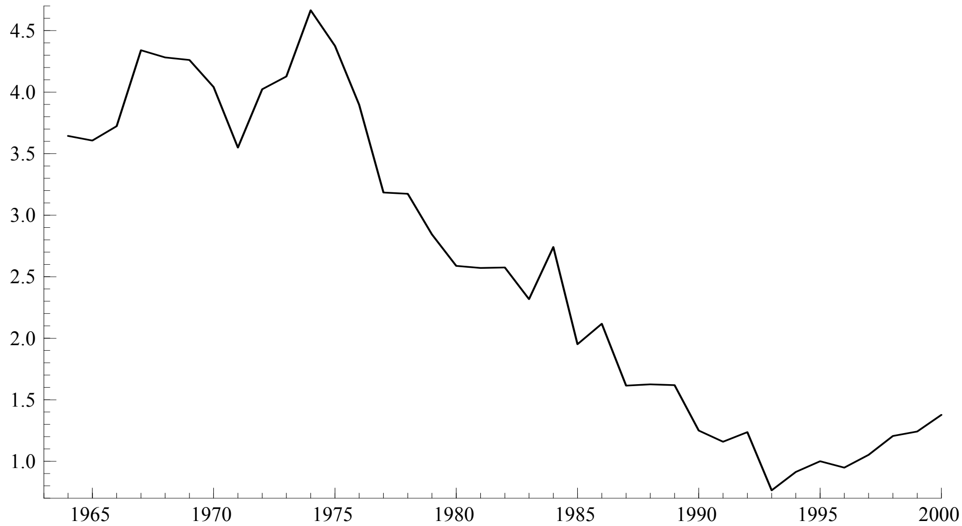
For our empirical analysis we had to come up with an estimate of the extent of international economic integration or liberalisation. Standard measures of openness to trade are the average tariff-rate and the coverage-ratio for non-tariff barriers. To use this measure we would also need a consistent time-series of import tariffs, which we do not have, although we have data from the mid-1970s on tariffs collected as a share of imports (World Bank, 2002).[6] Besides only covering part of the study-period, this is a very imperfect indicator of the level of protection. The tariff-average tends to under-estimate the impact of the high tariff-rates, because the corresponding import-levels are low. And the actual tariff rates on the books were higher than what this measure shows, but some importers could get their import-taxes waived. Moreover, there were major quantitative restrictions on imports, which obviously had an effect on domestic prices, though no tariff revenue was collected.

Since we did not have tariff-data for the whole period, nor adequate information on quantitative restrictions, we instead looked for price-data reflecting changes in the degree of protection. A possible indicator would be relative prices between manufacturing and agricultural goods in Kenya, but commodity-booms and weather conditions have had strong effects on agricultural prices, obscuring the impact of relative price changes due to long-term international market-integration. So we chose to measure the impact of trade-policy reform by comparing the change of Kenyan market-prices of manufactured goods with the change of world market-prices of manufactured goods (converted to Kenya Shillings using the official

[6] Export taxes have not been very important in the case of Kenya.

exchange-rate).[7] Presumably opening up reduces domestic prices relative to world prices. We used the GDP deflator for manufacturing in Kenya, and the price-index for industrial production in the UK, the latter being a major trading-partner.[8] The resulting time-series is shown in Figure 2.

Figure 2: Ratio of Kenyan to UK manufacturing prices, 1964-2000



Note: This is the ratio of the deflator for the Kenyan manufacturing sector to the UK price-index of industrial production converted to Kenya Shillings using the official exchange-rate; !995 = 1.

Our indicator of trade-reform coincides with the general description above of the Kenyan experience. Protection increased from Independence until 1975; the coffee-boom in 1976 led to opening up; there was a reversal in the beginning of the 1980s and then further reform; and by 1993 Kenya was an open economy. During the latter half of the 1990s there were some setbacks after macroeconomic prudence was thrown aside in the run-up to the 1997 election.

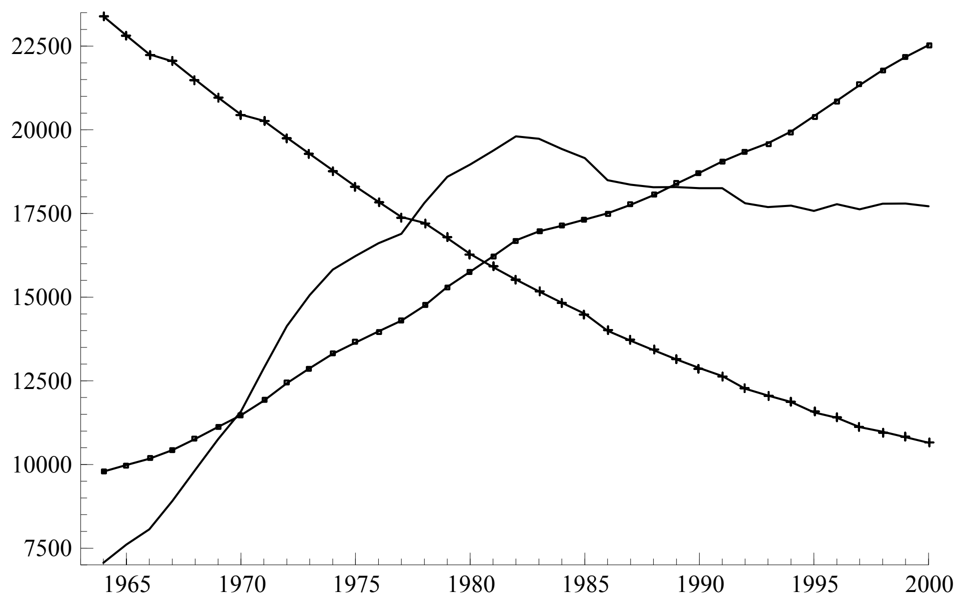
[7] A similar measure was used by Athukorala and Rajapatirana (2000) for an analysis of the Sri Lankan experience.
[8] We also calculated the index with US prices but the overall pattern was the same. Since imports from Europe were 4.5 times those from the US; and exports to Europe about 10 times those to the US, we preferred to use the UK rather than the US price-deflator (Kenya, 2003).

5. **Changes in Factor Endowments**

Since factor abundance is a basic determinant of the pattern of specialisation and factor-prices in an economy, we included endowment-ratios in the analysis. Figure 3 shows changes in the capital-labour ratio, land-labour ratio, and capital-land ratio for the period 1964-2000. The estimate of capital stock relies on assumptions about depreciation rates, but we would argue that the estimates of factor-availability are reasonably reliable.

Since Kenya has seen a rapid increase in labour throughout the period while the amount of arable land has changed very little, there has been a continuous decrease in the land-labour ratio. The capital-labour ratio increased until 1982, but then the trend was reversed.

Figure 3: Relative factor-endowments in Kenya, 1964-2000.



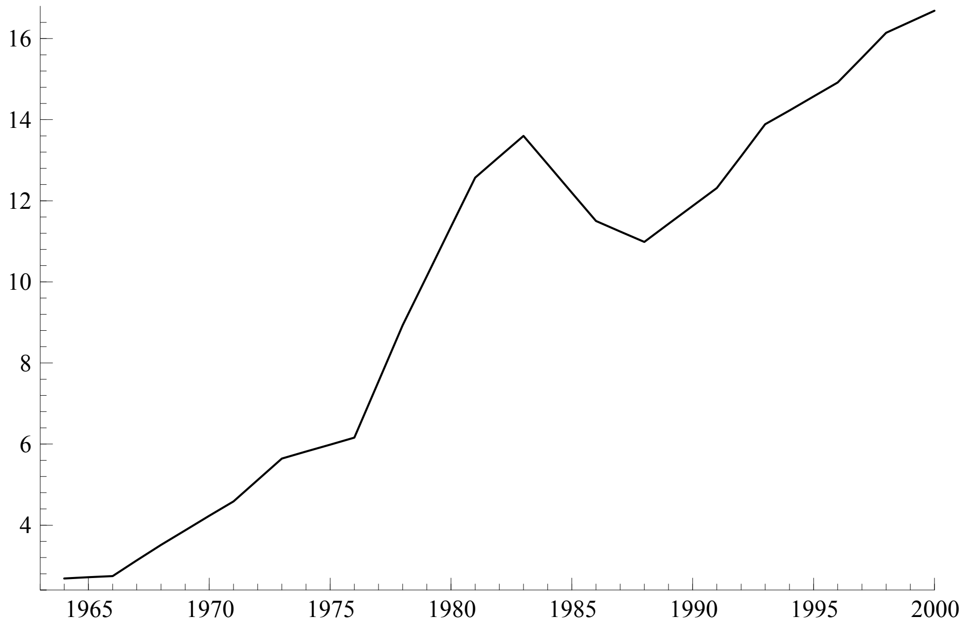
Capital-labour ratio ——— **Land-labour ratio** —+—+—+ **Capital-land ratio** =—□—□—□.
Note: The variables have been mean- and variance-adjusted to increase readability of the graph.
Source: The capital-stock series is from Ryan (2002), while labour and arable land are from World Bank (2002).

We hypothesize that capital and skilled labour are complements so that an increase in the capital-labour ratio will lead to higher demand for skilled workers and thus increase wage-inequality. Similarly we hypothesize that land and unskilled labour are complements, which means that increased demand for land would reduce wage-inequality. Increased supply of educated workers would tend to hold skilled wages down.

Educational attainment has also changed during the period of analysis, which could be a problem since the skill-composition of employees in both agriculture and manufacturing is likely to have changed, as more and more of the labour-force have more education. For example, in manufacturing the share of employees with at least secondary education went from 39% in 1978 to 65% in 2000 (Manda and Sen, 2004, p. 38).

Since it is possible that the average educational level has increased more in manufacturing than in agriculture, we added a variable in our regressions to at least partially control for this: the ratio of those over 15 who have at least have initiated secondary school to all those over 15 (Figure 4). The data on educational attainment was taken from the database of Barro and Lee (2000). There are only observations for every fifth year (1965, 1970, and so on) so the series was interpolated to a smooth trend. The trend-increase is only broken by a decline in the 1980s.

Figure 4: Changes in educational attainment, 1964-2000



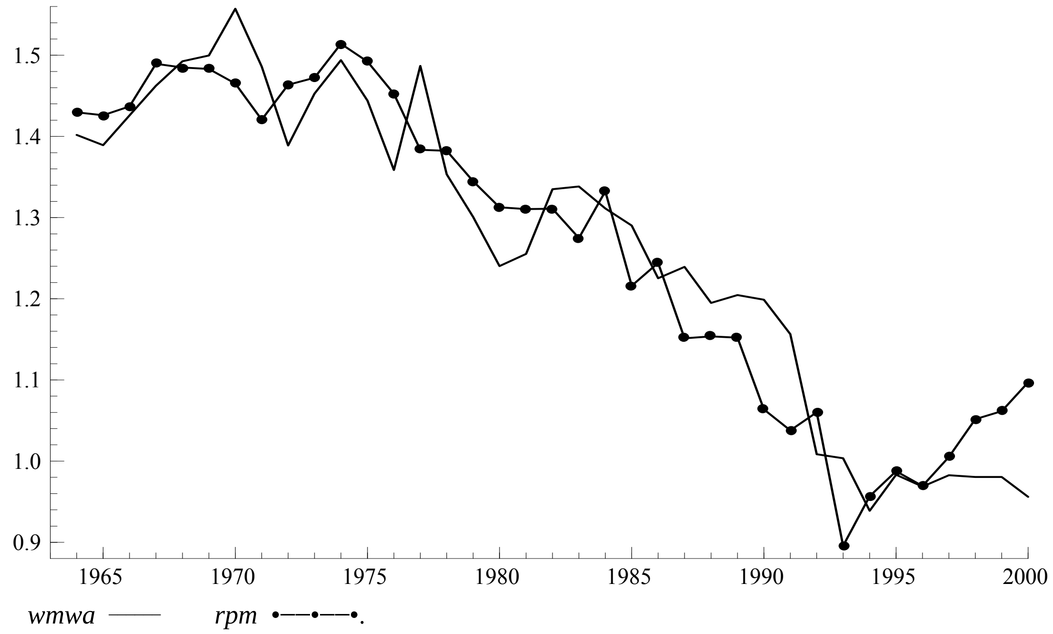
Note: This is the percentage of those over 15 who had at least started some secondary education.

6. Regression Results

We are now ready to look at how globalisation, that is, trade-policy reform, has affected relative wages in Kenya. Our focus will be mainly on the explanatory power of the changes in the ratio between Kenyan and UK manufacturing prices. However, several other variables might also have affected relative wages, so we analysed the roles of the Kenyan capital-labour and labour-land ratios, levels of education, and relative labour-productivity in agriculture and manufacturing. We also included the alternative to our preferred measure of economic integration, the relative price of agricultural and manufacturing goods in Kenya; and to capture short-run dynamics, we tested changes in terms of trade, the price of coffee beans and the exchange rate.

Before reviewing the results from the empirical analysis it may be helpful to see a plot of the two variables of major interest, the logs of relative wages and of Kenya/UK relative manufacturing prices, denoted *wmwa* and *rpm*, respectively (see Figure 5). The ratio of Kenyan manufacturing to agricultural wages fell substantially in the late 1970s, probably due to a rise in agricultural wages as a result of the 1975-78 coffee boom. The ratio then fell further in response to the liberalisation process started in the early 1980s. Kenya/UK relative manufacturing prices followed a similar pattern, although they rose quickly during the latter half of the 1990s.

Figure 5: The logs of the ratios between Kenyan manufacturing and agricultural wages (*wmwa*), and between Kenya and UK manufacturing prices (*rpm*), 1964-2000



Note: *rpm* has been mean- and variance-adjusted.

Sources: *wmwa* was taken from various issues of Statistical Abstract of Kenya; *wm* is wages for employees in the manufacturing sector, *wa* is wages for farm workers in the private sector; *rpm* was calculated from the GDP deflator for manufacturing in Kenya, taken from Ryan (2002), and industrial prices for the UK obtained from the IFS database.

Since both series are clearly non-stationary and our main interest was in the long-run relationships between the variables, we first used the Johansen (1988, 1995) approach to test for integration and cointegration between *wawm* and all the other variables. After that, to highlight the adjustment to long-run equilibrium as well as short-run dynamics, a model was estimated with stationary variables. Because of the small number of observations and the intractability of testing for cointegration with many variables at the same time, we did the cointegration analysis for groups of variables. There did not appear to be cointegration between *wawm* and any of the other variables, or combinations of them, except *rpm*.[9] Since the cointegration-tests produced a lot of output we only report these latter results; the other results can be obtained from the author upon request. However, we report results from tests of the short-run impact of all the variables in Table 4.

To test for cointegration, we first estimated a VAR for the period 1966-2000, with two lags on *wmwa* and *rpm* and an impulse dummy for 1993. The dummy captures a sharp drop in *rpm* that year due to a major devaluation. The F-test on model-reduction, reported in Table 1, indicates that the longest lag could be removed, one lag seemed sufficient to capture the dynamics. This result is supported by the decline in the value of the Schwartz criteria. Table 1 also reports multivariate misspecification tests for the model with one lag. The tests for serial correlation (AR 1-2 test), heteroscedasticity and normality are all insignificant.

Table 1 here

Table 2 reports the trace-tests for cointegration, the eigenvalues of the long-run matrix, the characteristic roots of the companion matrix, the cointegrating vectors (β), and feedback coefficients (α). Since the trace-test relies on asymptotic distributions, and we have very few observations, it is only indicative of the number of cointegrating relations; other sources of information should also be used to determine the number of stationary relations (see Hendry and Juselius, 2001).

Table 2 here

[9] The land-labour ratio comes out significant in some specifications. It has a downward trend coinciding with that of *wmwa* over the period 1977-1993 but not before or after. Assuming that *lanlab* and *wmwa* cointegrate, and adding a linear combination of the two variables to the error correction model reported in Equation (1), does not alter the interpretation our results, though the absolute value of some coefficients becomes lower.

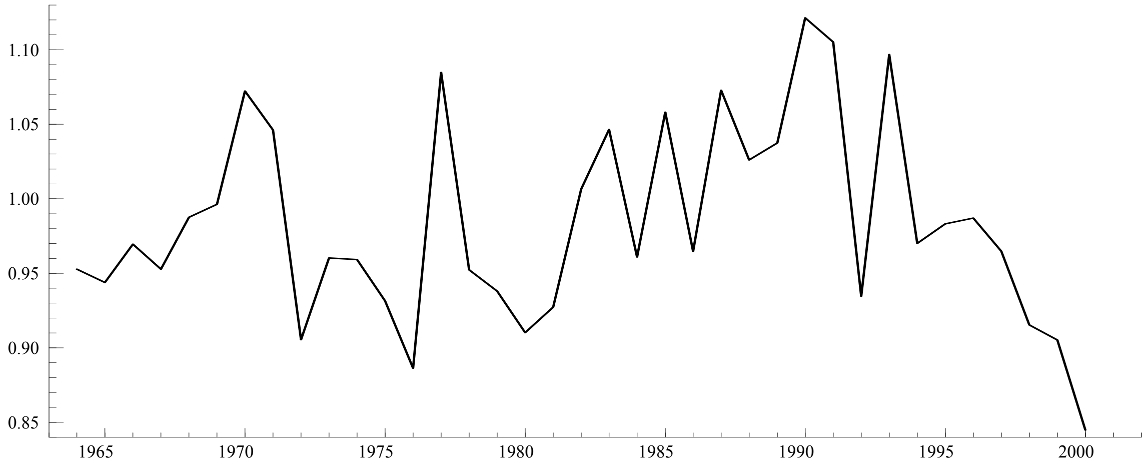
According to the trace tests there is one cointegrating vector; the test-statistic for the null hypothesis of a rank of zero, r=0, is significant at the 99% level, while the null of a rank of one, r=1, is insignificant. The eigenvalues support this result: The first one is clearly larger than zero (0.40), while the second is close to zero (0.06). The presence of one cointegrating vector is also evident from the roots of the companion matrix of the long-run matrix; one is close to unity (0.94), while the other is 0.35. This is also the case with the feedback coefficients, which are high in the first vector but very low in the second.

The next step was to identify the cointegrating vector. We started by testing whether either of the two variables was stationary by itself, that is, whether one variable could be excluded from the cointegrating vector. Then we tested for weak exogeneity. Since we expected *rpm* to determine *wmwa* in the long run, our model would not be valid if *wmwa* were weakly exogenous. Table 3 shows that both variables were nonstationary and neither could be excluded, and that *rpm* was weakly exogenous. This means that *wmwa* adjusts to changes in *rpm* in the long run.

Table 3 here

The cointegrating vector, $CI = wmwa - 0.35rpm$, is plotted in Figure 6. It shows the deviations from the long-run equilibrium relation between *wmwa* and *rpm*.

Figure 6: The cointegrating vector, 1964-2000



Next we estimated a general error correction model (ECM) with one lag of each variable, plus the rate of change of world coffee prices measured in constant Kenyan Shilling (*Δpcoff*); initially we also included several other variables but they proved to be insignificant.[10] Table A1 in the Appendix reports the general ECM and diagnostic tests, which are all insignificant. After removing variables with insignificant coefficients we obtained the preferred model,

$$\underset{(SE)}{\Delta wmwa} = \underset{(0.12)}{0.56} - \underset{(0.12)}{0.58}\left[wmwa - 0.35rpm\right]_{t-1} + \underset{(0.024)}{0.046}\Delta pcoff_{t-1}$$

$$R^2 = 0.45; \quad \hat{\sigma} = 0.045; \quad T = 1966\text{-}2000; \quad F_{AR}(2,30) = 0.54\ [0.59];$$
$$F_{ARCH}(1,30) = 0.004\ [0.95]; \quad F_{Het}(4,27) = 0.402\ [0.80]; \quad \chi^2_{Norm}(2) = 2.26\ [0.32];$$
$$F_{RESET}(1,31) = 0.860\ [0.36]; \quad \text{Test of reduction to preferred model: } F(4,28) = 0.296\ [0.88]. \tag{1}$$

The coefficient standard errors are shown in parentheses; Δ denotes the first difference, $\hat{\sigma}$ is the residual standard deviation and *T* is the sample period. The diagnostic tests are against serial correlation of order 2, F_{AR}, autoregressive conditional heteroscedasticity of order 1, F_{ARCH}, heteroscedasticity, F_{Het}, non-normality, χ^2_{Norm}, and non-linearity, F_{RESET} (see Hendry and Doornik, 2001, for details).

[10] We report the attempts to add other variables to the error correction model in the form omitted-variables tests, see Table 4 below.

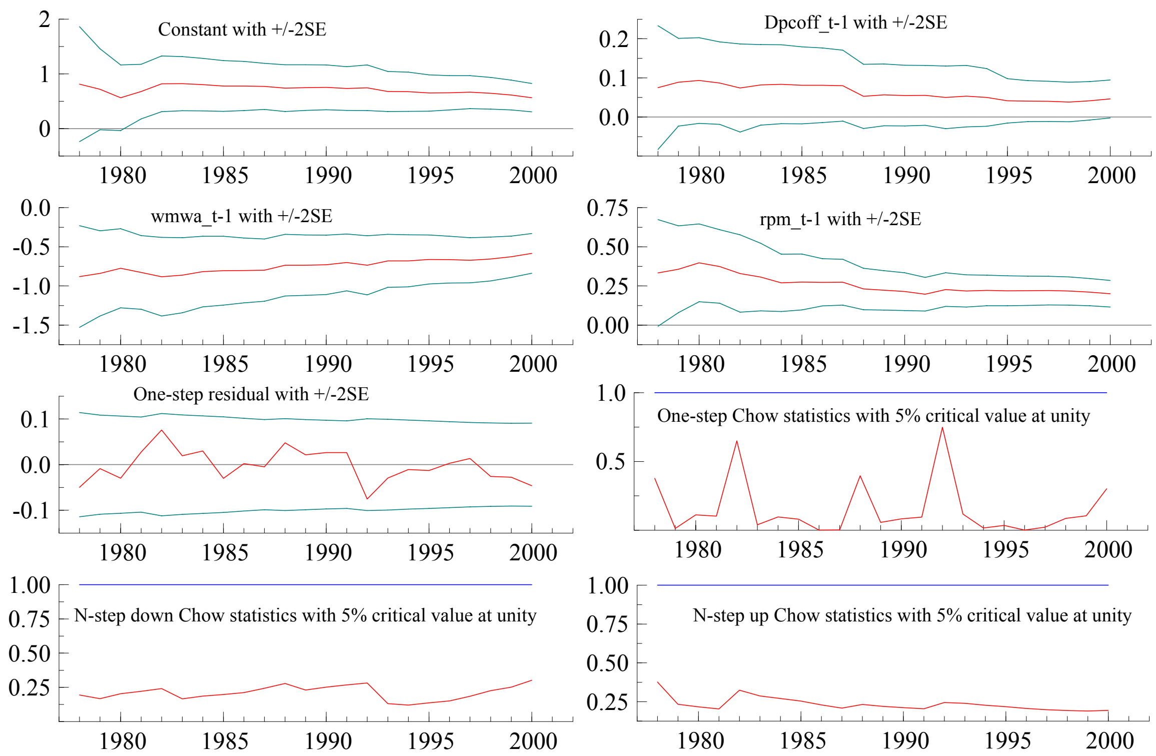
According to this model, our measure of openness (*rpm*) determines relative wages in the long run; an increase in openness reduces *wmwa,* thus reducing wage-inequality. The t-value for the error-correction term is -4.8, and the speed of adjustment towards equilibrium is fairly rapid: 0.58 of a deviation from long-run equilibrium is corrected within a year. The rate of change in relative wages is also affected by coffee prices, though its t-value is only close to 2. An increase in coffee prices increases manufacturing wages relative to agricultural the following year. The positive coefficient is probably due to indirect effects of increased export incomes. Changes in terms of trade and coffee prices have a correlation coefficient of 0.7, catching the same effect, but only the coffee-price variable has a t-value close to 2.

Table 4 shows results from diagnostic tests on omitted variables. It reports F-statistics from Lagrange multiplier tests for adding a variable to Equation (1), showing that none of the omitted variables was significant. The variables are the log-differences of the relative price of Kenyan agricultural and manufacturing goods, *Δpmpa*, the exchange rate, *Δe*, terms of trade, *Δtot,* coffee prices in US dollars, *Δ*p*coffUS,* a measure of educational attainment, *Δeduc,* relative labour-productivity in agriculture and manufacturing, *Δrelprod,* and the capital-labour ratio in first and second difference, *Δcaplab* and *ΔΔcaplab,* and the land-labour ratio, *Δlanlab*.

Table 4 here

The model's empirical constancy was assessed for the period 1978 to 2000 using recursive estimation. The output from this exercise is summarised in the graphs in Figure 7 below, showing recursive estimates of the four coefficients with ± 2 standard errors; one-step residuals with ± 2 estimated standard errors; and sequences of 1-step, break-point (N-step down) and forecast-Chow (N-step up) test-statistics. The latter are scaled such that the straight line at unity matches the 5% significance level. The coefficients are quite stable, there are no outliers, and none of the test-statistics of the three Chow tests are significant at the 5% level. Hence, the model appears to be stable.

Figure 7: Recursive estimation of coefficients, one-step residuals, and Chow tests



Note: See Hendry and Doornik (2001) for detailed description of the tests.

7. Summary and Conclusions

We have analysed how relative wages in Kenya have been affected by changes in the the degree of international economic integration, as well as by a range of other factors. In spite of having few observations, the results are strong. We find evidence that openness, measured by the ratio between manufacturing prices in Kenya and industrial prices in the UK (Kenya's main trading partner) played a major role in determining relative wages during the period

1964-2000. Of those included in the analysis, the only other variables that influenced relative wages was the world coffee price measured in constant Kenyan Shilling, but it had only a short-run effect. Other variables such as educational attainment, the capital-labour ratio, relative labour-productivity, changes in the nominal exchange-rate, and the ratio between Kenyan agricultural and manufacturing prices, had no significant effects. Our results thus indicate that liberalisation has reduced wage-inequality in Kenya. Whether this is so in other African countries awaits further research.

Table 1: Misspecification tests, test of reduction from VAR(2) to VAR(1)

Multivariate tests:		
AR 1-2 test	$F(8,54) = 0.573$ [0.79]	
Normality test	$\chi^2(4) = 6.170$ [0.19]	
Hetero test	$F(12,66) = 0.818$ [0.63]	
Hetero-X test	$F(15,66) = 1.23$ [0.27]	
Schwartz criteria	two lags	one lag
	-3.568	-3.794
Test of model reduction, 2 lags to 1: $F(4,58) = 1.36$ [0.26]		

Note: *p*-values are reported in brackets. The estimation period is 1966 – 2000. The vector autoregression includes an impulse dummy set to unity in 1993. A description of the tests can be found in Henry and Doornik (2001).

Table 2: Cointegration Analysis, 1965 - 2000

Eigenvalue of Π-matrix	0.40	0.06
Null hypothesis	r = 0	r = 1
Trace test	20.64	2.20
p-value for Trace test	0.007	0.138
Roots of process	0.95	0.35

Variable	*wmwa*	*rpm*
β'_1	1.000	-0.347
β'_2	106.4	1.00
α_1	-0.5763	0.3157
α_2	-0.0001	-0.0014

Note: The vector autoregression includes one lag on *wmwa* and *rpm*, and an impulse dummy set to unity in 1993. Critical values for the trace tests are based on the asymptotic distributions for an unrestricted constant.

Table 3: Cointegrated VAR and tests of significance and weak exogeneity

	Standardised eigenvector	
Variable	*wmwa*	*rpm*
β'	1.00	-0.35
Standard error	-	0.024
	Standardised adjustment coefficients	
α'	-0.58	0.31
Standard error	-0.13	0.34
	Statistics for test of significance and stationarity	
	wmwa	*rpm*
$\chi^2(1)$	14.90**	16.25**
	Weak exogeneity test statistics	
	wmwa	*rpm*
$\chi^2(1)$	15.65**	0.83

Table 4: Diagnostic tests for omitted variables, 1966 -2000

Variables		F-tests
$\Delta pmpa_t$	$\Delta pmpa_{t-1}$	F(2,29) = 0.24 [0.79]
Δe_t	Δe_{t-1}	F(2,29) = 0.17 [0.84]
Δtot_t	Δtot_{t-1}	F(2,29) = 1.23 [0.31]
$\Delta pcoffUS_t$	$\Delta coffpUS_{t-1}$	F(2,29) = 0.45 [0.64]
$\Delta educ_t$	$\Delta educ_{t-1}$	F(2,29) = 0.51 [0.61]
$\Delta relprod_t$	$\Delta relprod_{t-1}$	F(2,29) = 0.09 [0.90]
$\Delta caplab_t$	$\Delta caplab_{t-1}$	F(2,29) = 0.28 [0.76]
$\Delta\Delta caplab_t$	$\Delta\Delta caplab_{t-1}$	F(2,28) = 0.10 [0.90]] [a]
$\Delta lanlab_t$	$\Delta lanlab_{t-1}$	F(2,29) = 0.13 [0.87]

a) The sample is 1967-2000.

Note: *p*-values are given in brackets. A Δ denotes first difference and ΔΔ second difference. All variables are in logs. *papm* is the ratio of prices of agricultural and manufacturing goods; *e* is the nominal exchange-rate; *tot* is terms of trade; *pcoffUS* is the world price of coffee beans; e*duc* is a educational attainment; *relprod* is the relative labour productivity between agriculture and manufacturing; *caplab* is the capital-labour ratio; and *lanlab* is the land-labour ratio.

Appendix: General Error Correction Model

Table A1: General error correction model, 1966 - 2000

	Coefficient	SE	t-value
Constant	0.533	0.182	2.930
$\Delta wmwg_{t-1}$	0.102	0.158	0.648
Δrpm_{t}	0.008	0.060	0.128
Δrpm_{t-1}	0.050	0.089	0.564
$\Delta pcoff_{t}$	0.025	0.031	0.804
$\Delta pcoffp_{t-1}$	0.048	0.025	1.880
$[wmwa\text{-}0.35rpm]_{t-1}$	-0.549	0.185	-2.970

$R^2 = 0.47;\quad \hat{\sigma} = 0.047;\quad T = 1966\text{ - }2000;\quad F_{AR}(2,26) = 0.704\ [0.50];$
$F_{ARCH}(1,26) = 0.752\ [0.39];\quad F_{Het}(12,15) = 1.29\ [0.32];\quad \chi^2_{Norm}(2) = 2.47\ [0.29];$
$F_{RESET}(1,27) = 1.32\ [0.26].$

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