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Openness and Wage Inequality in Kenya, 1964–2000

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Summary. — This paper analyzes the evolution of wage inequality in Kenya during 1964–2000. 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–labor ratio, educational attainment, relative labor 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.
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Key words — globalization, Kenya, trade liberalization, trade policy, wage inequality

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 industrialized countries, but there have been a number of studies of less-industrialized Latin American and Asian countries as well (Arbache, Dickerson, & Green, 2004; Goldberg & Pavcnik, 2004; Williamson, 1997). Due to lack of good data there have so far been few studies of African economies, but we have been able to analyze 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.

The rest of the paper is organized as follows. Section 2 provides a brief theoretical review. Section 3 then describes the evolution of the Kenyan labor market and introduces our wage series. Section 4 presents our data on changes 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 discusses the compatibility of our results with those of Manda and Sen (2004), who looked at the effects of trade liberalization on wages and employment in Kenya. Section 8 summarizes and draws conclusions.

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2. THEORY

Most discussion of wage inequality has been based on the standard 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 two-country, two-factor (skilled and unskilled labor), two-good model, barriers to trade drive wedges between the prices of the two goods in the two countries. The price of the unskilled-labor-intensive exportable may be kept low in the less developed country, and *vice versa* for the industrialized country. A move from autarky to free trade, or from high tariffs to low tariffs, can then lead to increased production of unskilled-labor-intensive exportables in the poor country, which in turn leads to increased demand for unskilled labor and reduced demand for skilled labor. 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 Wood, 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 labor 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 labor in non-tradables production could then be larger than the increase in demand for unskilled labor in the labor-intensive tradables sector. Thus it is theoretically possible (though not likely) that there would be a drop in unskilled wages.

Adding another factor of production 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 labor in production, then this would increase the demand for skilled labor. This might lead to an unexpected result in terms of relative wages. In a poor LDC, however, land is probably complementary to unskilled labor, 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 labor. 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 the 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 labor-market rigidities hindering smooth reallocation of labor 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 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 labor 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 a tariff reduction could increase

Table 1. *Kenyan industrial production, export, and formal employment, 1964–2000*

Year	Index of industrial production (constant Kshs)	Index of manufacturing production (constant Kshs)	Index of agricultural production (constant Kshs)	Percentage share of industry in GDP	Percentage share of manufacturing in GDP	Percentage share of agriculture in GDP	Percentage share of labor force in formal employment
1964	100	100	100	16.7	10.4	41.5	12.2
1970	135.8	137.6	118.3	19.8	12.0	33.3	11.6
1975	274.6	285.2	164.8	20.3	12.0	34.2	12.5
1980	372.2	416.0	199.0	20.9	12.8	32.6	12.8
1985	427.6	501.9	231.4	19.1	11.7	32.5	12.5
1990	548.2	663.1	284.9	19.1	11.8	29.1	12.7
1995	593.9	750.8	285.7	16.0	9.9	31.1	11.7
2000	633.7	801.2	304.6	18.6	12.9	19.7	10.9

Sources: Data for industry and manufacturing and manufacturing export from World Bank (2005), and data for labor from Statistical Abstract, various years (Kenya, annual).

the relative wages of skilled labor. Arbache *et al.* (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 (see 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 LABOR MARKET

In this section, we present data on wages for employees in manufacturing and agriculture used in the analysis, and describe the Kenyan labor market. However, first we provide some basic facts about the role of manufacturing and agriculture in the development of the Kenyan economy. Table 1 shows that industry, including manufacturing, has grown sixfold since independence, but at a slow rate in the 1990s, particularly during the period 1995–2000. The share of industry in value added has varied between 17% and 21% with a peak in 1979 and a dip in the mid-1990s, while manufacturing has made up around two-thirds of the industrial sector. The share of agriculture was more or less constant at about 30% during 1970–95 but it fell somewhat in the late 1990s. Hence, overall there was not much structural transformation during the period under study. There also seems to be little structural change within the manufacturing sector; in a recent study of openness and employment, Sen (2004) fails to find significant increases in the efficiency of labor or shifts in industry-specific capital–labor ratios.

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 1964, until the late 1970s, manufacturing wages were roughly four times as high as agricultural wages. During the 1980s and the beginning of the 1990s the ratio declined, and since 1994 it has been about 2.6 times.

The wages series used are from the formal sector, which during the period covered 11–13% of the labor force. Most of the remainder of the labor force were smallholders, but an

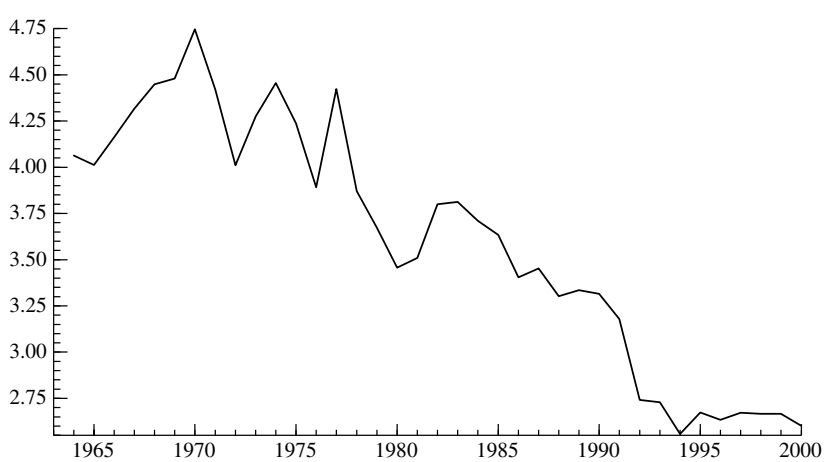

Figure 1. *Manufacturing wages relative to agricultural wages, 1964–2000. Source: Kenya (annual).*

increasing share was engaged in the non-agricultural informal sector. An interesting question is thus whether wages in agriculture also reflect wages in the 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 or the informal 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 assume that agricultural wages reflect the wider labor market for less skilled workers.

As Figure 1 shows, the gap between agricultural and non-agricultural wages increased rapidly during the 1960s, 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 was 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 the formal-sector, or economic advantages to conducting business in the informal sector (Bigsten, Kimuyu, & Lundvall, 2004). [3]

Relative wages can also be affected by changes in labor-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 & Lal, 1986). Trade unions in Kenya were financially weak and their ability to strike was 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 organized labor and employers, including implementation of the government's wage guidelines.

Labor-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 & Sen, 2004). At the same time, redundancy laws were amended, which made it easier for employers to lay off employees. These reforms meant that market forces generated rapid wage increases during the 1995–2000 period (see IMF, 2003, for details).

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, [4] 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 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. [5] In November 1981, the government stopped requiring "no-objection certificates" from domestic producers, [6] 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 & 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). Although Kenya remained an open economy during the latter half of the 1990s, several factors such as loose macroeconomic policy, macroeconomic instability, and the collapse of the IMF program in 1997, set in motion a process which led to an appreciating real exchange rate, that is, a process that resembles reduced openness.

For our empirical analysis we had to come up with an estimate of the extent of international economic integration or liberalization. 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, 2005). [7] 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, or 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. Consequently, 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 exchange rate). [8] Presumably opening up reduces domestic manufacturing prices relative to world prices, since the whole point of the import-substitution policy was to keep manufacturing prices above world market prices to make it possible for domestic producers to enter the market. We used the GDP deflator for manufacturing in Kenya, and the price index for industrial production in the United Kingdom, the latter being a major trading-partner. [9] The resulting time series is shown in Figure 2.

Our indicator of trade reform coincides with the general description given 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.

International trade seems to have responded to these policy changes as expected. During the first half of the 1980s, there was a decline

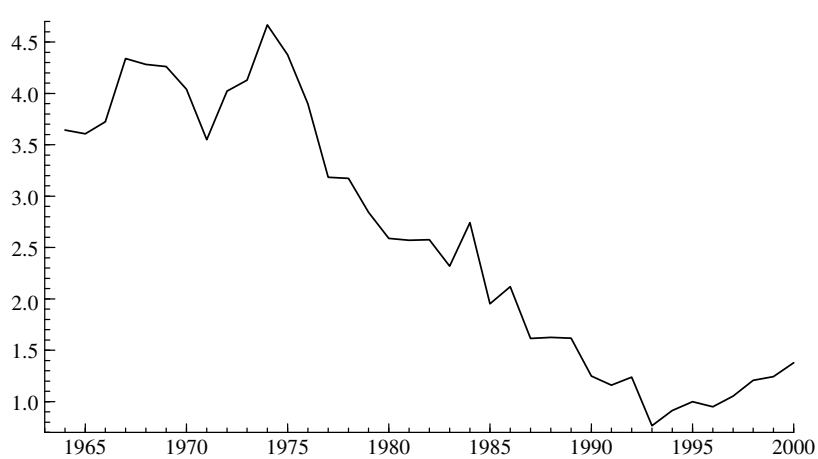

Figure 2. *Ratio of Kenyan to UK manufacturing price indexes, 1964–2000 (1995 = 1).*
Source: Ryan (2002) and IMF (2004).

in manufactured exports, partly due to global recession. However, during 1985–90 the real US$ export value grew by 54%, and during 1990–95 it grew by 114%.[10] The second half of the 1990s saw a decline in manufactured exports by 27%, probably as a result of the impact of the Asian crisis and exchange rate appreciation; the real effective exchange rate increased by about 40% from 1993/94 to the late 1990s.[11]

5. CHANGES IN FACTOR ENDOWMENTS

Since factor abundance is a basic determinant of the pattern of specialization and factor prices in an economy, we included endowment ratios in the analysis. Figure 3 shows changes in the capital–labor ratio, land–labor 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 labor throughout the period, while the amount of arable land has changed very little, there has been a continuous decrease in the land–labor ratio. The capital–labor ratio increased until 1982, but then the trend was reversed.

We hypothesize that capital and skilled labor are complements so that an increase in the capital–labor ratio will lead to higher demand for skilled workers and thus increase wage inequality. Similarly, we hypothesize that land and unskilled labor 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 labor force have more education. For example, in manufacturing the share of employees with at least secondary education went from 36% in 1978 to 61% in 2000 (Manda & 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. However, there are no time series available on the educational attainments specifically in agriculture and industry so we used the ratio of those over 15 who at least have initiated secondary school to all those over 15 (Figure 4). The data on educational attainment were taken

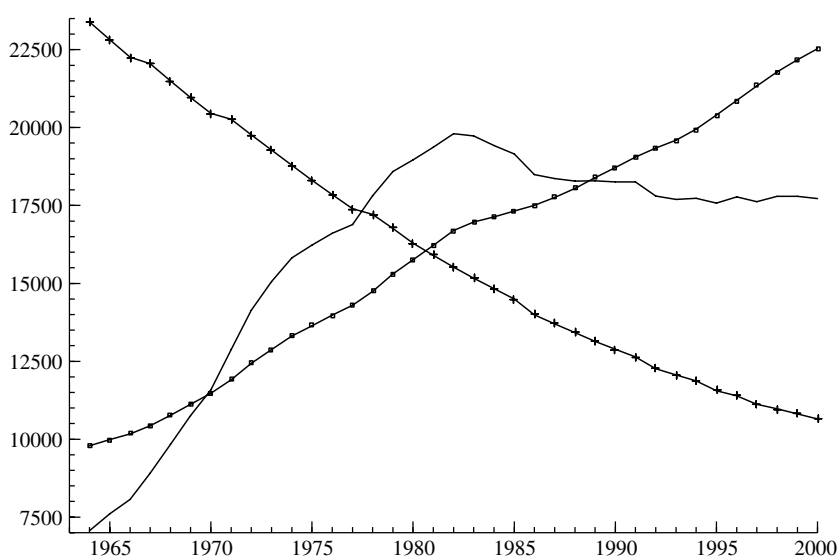

Figure 3. *Relative factor endowments in Kenya, 1964–2000. The variables have been mean- and variance-adjusted to increase readability of the graph. Sources: The capital-stock series is from Ryan (2002), while labor and arable land are from World Bank (2005). Capital–labor ratio (—); land–labor ratio (+–+–+) and capital–land ratio (■–■–■).*

Figure 4. *Changes in educational attainment, 1964–2000. The percentage of those over 15 who had at least started secondary education. Source: Barro and Lee (2000).*

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.

6. REGRESSION RESULTS

We are now ready to look at how globalization, 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 analyzed the roles of the Kenyan capital–labor and labor–land ratios, levels of education, and relative labor 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 liberalization process that started in the early 1980s. [12] Kenya/UK relative manufacturing prices followed a similar pattern, although they rose quickly during the latter half of the 1990s.

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*. [13] Since the cointegration tests produced a lot of output we only report these latter results; the other results can be obtained from the authors upon request. However, we report results from tests of the short-run impact of all the variables in Table 5.

To test for cointegration, we first estimated a vector autoregression (VAR) for the period 1966–2000, with two lags on *wmwa* and *rpm*

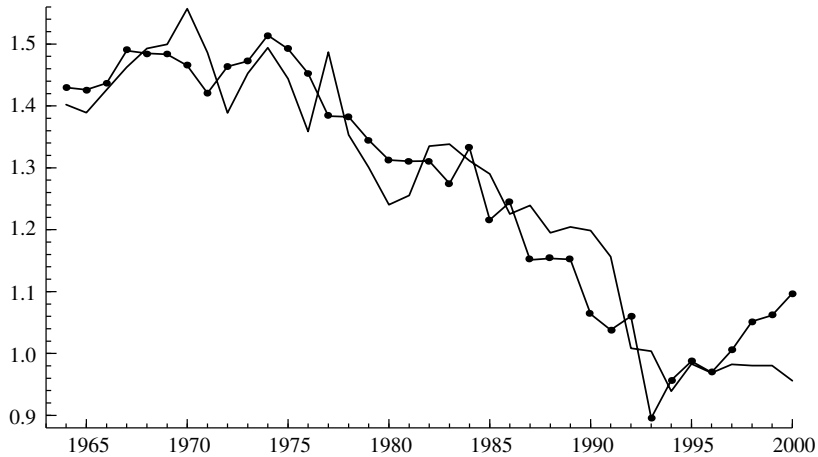

Figure 5. *The logs of the ratios between Kenyan manufacturing and agricultural wages (wmwa), and between Kenya and UK manufacturing prices (rpm), 1964–2000. rpm has been mean- and variance-adjusted. Sources: wmwa is from Kenya (annual); wm is wages for employees in the manufacturing sector and wa is 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 United Kingdom obtained from IMF (2004). 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 2, 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 2 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 3 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 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 & Juselius, 2001).

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.

Table 2. *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]
	Two lags	One lag
Schwartz criteria	−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 Hendry and Doornik (2001).

Table 3. *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.

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 4 shows that both variables were non-stationary and neither could be excluded, and that *rpm* was weakly exogenous. This means that *wmwa* adjusts to changes in *rpm* in the long run.

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

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 ($\Delta pcoff$); initially we also included several other variables but they proved to be insignificant. [14] Table 7 in Appendix A reports the general ECM and diagnostic tests, which are all insignificant. After removing variables with insignificant coefficients we obtained the preferred model,

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

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

Table 4. *Cointegrated VAR and tests of significance and weak exogeneity*

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

** Significance at the 1% level.

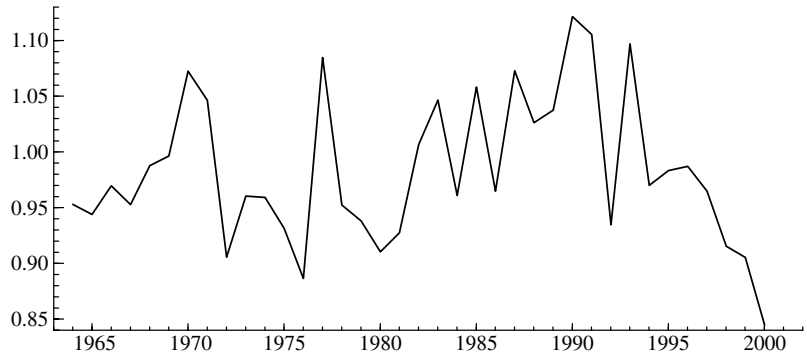

Figure 6. *The cointegrating vector, 1964–2000.*

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 & Doornik, 2001, for details).

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 toward 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. [15] 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 5 shows results from diagnostic tests on omitted variables. It reports *F*-statistics from Lagrange multiplier tests for adding a variable to Eqn. (1), showing that none of the omitted variables were significant. The variables are the log differences of the relative price of Kenyan agricultural and manufacturing goods, Δ*pmpa*, the exchange rate, Δ*e*, terms of trade, Δ*tot*, coffee prices in US$, Δ*pcoffUS*, a measure of educational attainment, Δ*educ*, relative labor productivity in agriculture and manufacturing, Δ*relprod*, and the capital–labor ratio in first and second difference, Δ*caplab*

Table 5. *Diagnostic tests for omitted variables, 1966–2000*

Variables	*F*-tests
$\Delta pmpa_t\ \Delta pmpa_{t-1}$	$F(2,29) = 0.24$ [0.79]
$\Delta e_t\ \Delta e_{t-1}$	$F(2,29) = 0.17$ [0.84]
$\Delta tot_t\ \Delta tot_{t-1}$	$F(2,29) = 1.23$ [0.31]
$\Delta pcoffUS_t\ \Delta pcoffpUS_{t-1}$	$F(2,29) = 0.45$ [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]

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; *educ* is a educational attainment; *relprod* is the relative labor productivity between agriculture and manufacturing; *caplab* is the capital–labor ratio; and *lanlab* is the land–labor ratio.

[a] The sample is 1967–2000.

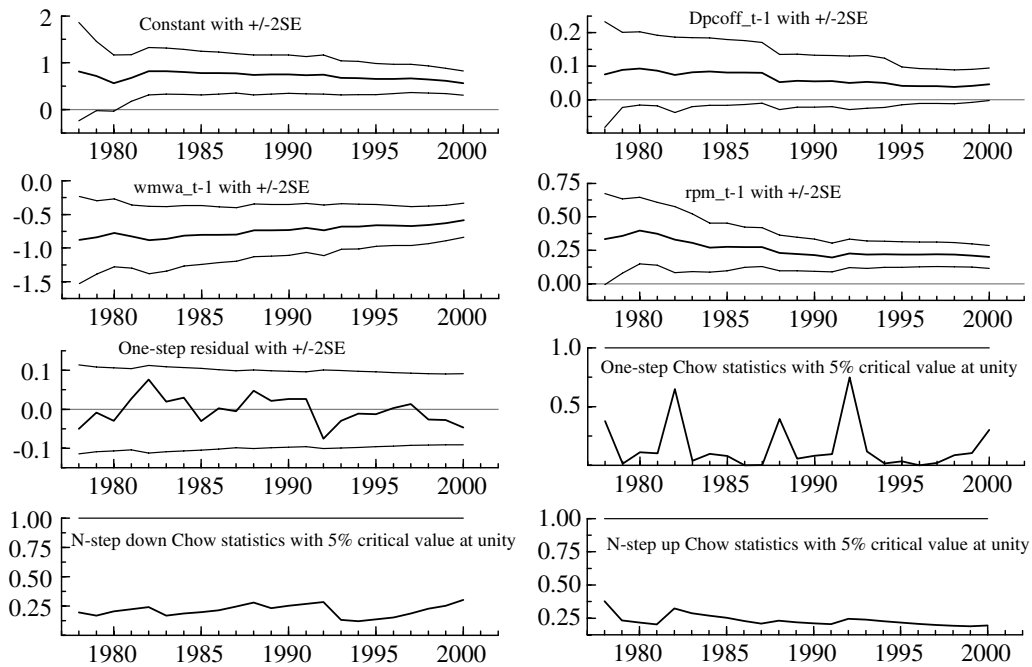

Figure 7. *Recursive estimates of the coefficients with ±2 standard error (top four graphs), one-step residual with ±2 estimated standard errors, One-step, N-step down (break-point) and N-step up (forecast) Chow statistics scaled with their 5% critical values. The straight line at unity shows the 5% critical level. See Hendry and Doornik (2001) for detailed description of the tests.*

and $\Delta\Delta caplab$, and the land–labor ratio, $\Delta lanlab$.

The model's empirical constancy was assessed for the period 1978–2000 using recursive estimation. The output from this exercise is summarized in the graphs in Figure 7, showing recursive estimates of the four coefficients with ±2 standard errors; One-step residuals with ±2 estimated standard errors; and sequences of One-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.

7. DISCUSSION

The only other study dealing with changes in wage-income inequality in Kenya is, to our knowledge, Manda and Sen (2004). They look at wage differences between skilled and unskilled workers in manufacturing using firm level data for 1995 and 2000 and estimates based labor force survey data for 1978 and 1986 (from an earlier study by Appleton, Bigsten, & Manda, 1999). They conclude (p. 42) that "less skilled workers experienced losses in earnings during the period 1986–2000, along with an increase in inequality between highly skilled and less skilled workers." This conclusion is based on Table 10 in their article, which we have recast in index-form to see better what has happened over time (see Table 6).

In the samples, reported in Table 6 in Manda and Sen (2004), the share of those with no education falls from 15% to 2% during the period of analysis, which means that comparisons using estimates for this category are really not meaningful. Moreover, the share of those with university education is very small; it goes from 3% to 4%. It is thus mainly the groups with primary and secondary education that can be reliably compared; the share of those with primary school falls from 46% to 33%, while the share of those with secondary school increases from 36% to 61%.

Table 6. *Indexes of real earnings in manufacturing by education level, 1978–2000*

Education	1978	1986	1995	2000
Uneducated	100	124.0	88.5	34.9
Primary 1–4	100	112.7	77.2	79.3
Primary 5–8	100	90.0	64.0	72.0
Secondary 1–2	100	83.2	63.8	61.8
Secondary 3–4	100	59.6	38.7	48.2
Secondary 5–6	100	91.1	26.8	61.1
University	100	95.2	52.7	88.3

Source: Computed from Table 10 in Manda and Sen (2004).

If we look at the period from 1978 to 2000, we see that those with primary education did better than those with secondary education, and that during the period 1978–86 those with primary education did clearly better. Furthermore, in the period 1986–95, only those with 1–2 years of secondary education did better than those with primary education, while in the final period, 1995–2000, two of the secondary education categories did better than the primary education ones. Since our measure of wage inequality declined until 1994, when there was a reversal, this seems to be consistent with the data presented by Manda and Sen. Their conclusion that globalization in Kenya has been associated with increasing wage inequality is therefore debatable. We would argue that the surge in skilled wages in the late 1990s was not due to trade liberalization; our indicator shows a reduction in openness during this period, so the increase in relative skilled wages is what the theory predicts. Yet this effect alone is unlikely to be strong enough to explain the rapid increase in skilled wages. The deregulation of the labor market in 1994 probably played a considerable role, making it possible for the more skilled workers to bounce back and reclaim some of the losses in relative income that they had experienced since the 1970s (see IMF, 2003).

8. SUMMARY AND CONCLUSIONS

We have analyzed how relative wages in Kenya have been affected by changes in 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 international industrial prices proxied by those in the United Kingdom, 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. Our results thus indicate that liberalization has reduced wage inequality in Kenya.

There was a rapid increase in skilled wages in the second half of the 1990s, which appears to contradict this conclusion. However, the degree of openness peaked in 1993/94 according to our measure, and during the latter half of the decade several factors contributed to relative price changes that indicated reduced openness. For example, at the end of the 1990s, the real effective exchange rate was about 40% higher than in the beginning of the decade. Moreover, extensive liberalization of the labor market in 1994 appears to have generated rapid growth in real manufacturing wages, probably due to increased bargaining power of the labor unions. Yet, the mechanisms behind the development of real and relative wages since the mid-1990s are not well understood and merit further research.

In a recent study, Manda and Sen (2004) argue that wage inequality in Kenya increased as a result of trade liberalization. However, a close scrutiny of their data reveals that this conclusion is questionable. The increase in wage inequality occurred mainly during the period 1995–2000, while there was strong decrease in wage equality during 1978–2000.

Our results thus suggest that the reduction of protection of the manufacturing sector led to a fall in the ratio of wages in manufacturing to wages in agriculture, which can be seen as an indicator of sectoral wage inequality or as proxy for skilled to unskilled wages. Globalization or opening up to trade seems to have reduced wage inequality in Kenya.

NOTES

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

2. During 1963–65 public sector real wages increased by 48%, while real wages increased by merely 6% in the private sector (Collier & Lal, 1986, p. 62).

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. For example, inward FDI during the period 1981–1989 was in the range of 1–3% of total investment (Manda & Sen, 2004, p. 32).

5. Glenday and Ryan (2003) discuss the stages economic liberalization in Kenya and the links between trade liberalization and economic growth.

6. Domestic producers of specific products could object to import of competing goods.

7. Export taxes have not been very important in the case of Kenya.

8. A similar measure was used by Athukorala and Rajupatirana (2000) for an analysis of the Sri Lankan experience.

9. In principle, a trade-weighted index would be the best measure. However, annual revisions of the trade weights would create abrupt changes in the index due to the high level of time aggregation. The major alternative to the UK price index is the one for the United States, but we preferred to use the UK price index since the United Kingdom is Kenya's major trading partner; imports from the United Kingdom in 2000 were 2.5 times those from the United States, and exports to the United Kingdom were 6.6 times those to the United States (Kenya, 2005). We also calculated the index with prices from some other countries but the overall pattern was the same.

10. The data on exports of manufactured goods are from World Bank (2004). The US price index on producer goods was used to calculate the real values. The series was obtained from IMF (2004).

11. The real effective exchange rate was taken from IMF (2004).

12. It is very hard to find data on inequality in Kenya that is comparable over time. The WIDER (2005) data on household inequality give Gini coefficients for incomes in 1977 at 57.0, and for consumption in 1992 at 56.9, in 1994 at 44.3, and in 1997 at 44.5. These estimates do not contradict our results.

13. The land–labor ratio comes out significant in some specifications. It has a downward trend coinciding with that of *wmwa* over the period 1977–93 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 Eqn. (1), does not alter the interpretation our results, though the absolute value of some coefficients becomes lower.

14. We report the attempts to add other variables to the error correction model in the form omitted-variables tests, see Table 5.

15. The study of the coffee-boom of 1975–78 by Bevan, Collier, Gunning, Bigsten, and Horsnell (1990) showed that over time much of the effect of the boom spread to the urban areas. It is noted (p. 205) that "in spite of the decision of the Kenyan government to pass the coffee prices intact on to farmers the bulk of the income gain does not accrue to them. More than 75% of the boom income eventually ends up in the hands of urban households, a group that forms a small proportion of the population." This clearly supports our argument for the sign of the coefficient.

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(*See Overleaf*)

APPENDIX A. GENERAL ERROR CORRECTION MODEL

Table 7. *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
$[wmwa - 0.35rpm]_{t-1}$	−0.549	0.185	−2.970

$R^2 = 0.47$,

$\hat{\sigma} = 0.047$,

$T = 1966\text{–}2000$,

$F_{\mathrm{AR}}(2, 26) = 0.704[0.50]$,

$F_{\mathrm{ARCH}}(1, 26) = 0.752[0.39]$,

$F_{\mathrm{Het}}(12, 15) = 1.29[0.32]$,

$\chi^2_{\mathrm{Norm}}(2) = 2.47[0.29]$,

$F_{\mathrm{RESET}}(1, 27) = 1.32[0.26]$.

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