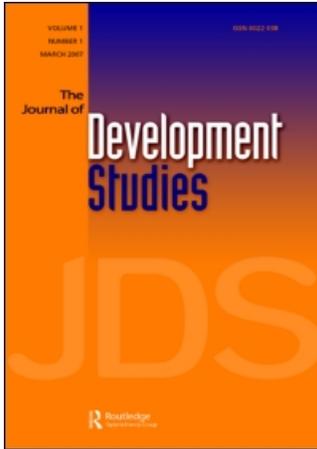


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Factor Proportions, Openness and Factor Prices in Kenya 1965-2000

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Factor Proportions, Openness and Factor Prices in Kenya 1965–2000

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ABSTRACT *This study analyses how changes in factor abundance and openness have affected relative factor prices in Kenya since 1965, using cointegration analysis and error correction models of relative factor prices. We find that factor proportions determined relative factor prices in the long run, while openness, measured by three different proxies, possibly had a short run effect on relative factor returns. The only deviation from this pattern occurred during the latter half of the 1990s when there was rapid wage growth, mainly due to labour market deregulation.*

I. Introduction

In recent years there has been a heated debate about the impact of globalisation or trade liberalisation on income distribution in low income countries. One crucial link between policy changes and inequality outcomes are changes in relative factor prices. A large amount of research has focused on the effect of openness on one particular factor price, namely the wage of skilled relative to unskilled labour, but we hardly know anything about how greater openness has affected other factor-price ratios, such as the returns to land and capital relative to labour. This is a potentially serious shortcoming, particularly in land-abundant African countries where the impact of liberalisation on overall inequality depends on what happens to these other relative factor prices. In this paper we provide what is probably the first analysis of an African economy of the long-term determination of the relative prices of the factors land, capital and labour.¹

The country chosen for this analysis is Kenya, which can be viewed as a representative sub-Saharan African country. Soon after independence in 1963 it started to pursue an import substitution policy approach to industrialisation. However, due to severe macroeconomic imbalances at the beginning of the 1980s, Kenya had to embark on a series of structural adjustment programmes that resulted in an uneven process of liberalisation of international trade and capital flows. In

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1993–94, full-scale trade liberalisation was implemented and Kenya became an open economy according to the criteria set up by Sachs and Warner (1995).

To evaluate the role of endowments and openness for the evolution of relative factor prices, we first compiled data on factor endowments and factor returns for land, capital and labour and indicators of openness for the period 1964–2000. We then used cointegration analysis and developed error-correction models for relative factor returns. In addition, the evolution of total factor productivity (TFP) was analysed, since it may influence factor returns.

Our analysis shows that factor proportions determine relative factor prices in the long run, while openness, measured by three different proxies, seems to play a small role in the determination of factor returns. The only deviation from this behaviour occurred during the latter half of the 1990s, when wages grew rapidly. This was primarily due to labour market deregulation and appears to be a temporary phenomenon: there is no evidence of productivity growth during this period and recent data show a slowdown in wage growth. Hence, the influence of the international economy on the three factor prices considered here has been weak, relative to the impact of changes in factor abundance, in spite of the steps taken to open up the Kenyan economy.²

The paper is structured as follows. In Section II, we present a simple theoretical framework, while Section III presents our data on changes in Kenya's factor endowments. Section IV discusses changes in Kenyan factor markets and describes our factor price estimates. Section V looks at changes in Kenya's trade policy, and introduces our measures of openness. Section VI provides estimates of productivity change, and section VII presents our econometric analysis of the links between changes in factor abundance and openness, on the one hand, and changes in factor prices on the other. Section VIII concludes and summarises the paper.

II. Theoretical Framework

The purpose of this paper is to analyse the evolution of factor prices in Kenya since independence in 1963. We have systematic information on the prices of labour, capital and land, and a three-factor model that can represent the Kenyan economy well is the specific factor model (Neary, 1978).

We assume that Kenya is a small open economy that takes world market prices as given. We further assume a competitive economy with two sectors, industry and agriculture, and that the goods provided are produced with constant returns to scale by the three factors: labour, capital and land. Industrial goods are produced with labour and capital, while agricultural goods are produced with labour and land. Capital and land are specific to their respective sectors, while labour is mobile between them. Unlike the simple two-by-two Heckscher-Ohlin model, there is no longer a one-to-one relationship between goods prices and factor prices. Factor prices now depend on more variables, including factor abundance.

Demand for labour in each sector is given by the value marginal product schedules, which depend on the amount of the fixed factor, technology and the goods price. In equilibrium, labour is allocated between the sectors so that wage rates are equalised and once we have determined the wage rate in this model we can easily determine the rental rates for the two specific factors.

Returns to factors are affected by changes in goods prices, which may be due to changes in international terms of trade or tariffs. Higher tariffs on industrial goods shift the industry value marginal product curve upwards and increases labour demand, and more labour is attracted into that sector. This means that wages in terms of industrial goods increase by less than the goods prices increase. Real wages thus fall in terms of industrial goods but, they rise in terms of agricultural goods. The change in the real wage is therefore ambiguous and depends on the structure of the labourer's demand for goods. As far as the specific factors are concerned the results are clear-cut. Capital rentals increase and land rentals decline.

With given goods prices (from the world market), the results of changes in factor endowments are also obvious. An increase in the endowment of a specific factor increases the returns to labour, while returns to the specific factors decline. An increase in the endowment of labour reduces returns to labour but increases returns to the specific factors, land and capital.

Finally, technical progress or increased productivity in one sector is beneficial for the factor specific to that sector as well as the mobile factor labour, while it is negative for the factor specific to the other sector.

The model as outlined here is a classical trade model that assumes full employment of factors of production and equilibrium, whilst in reality there are various distortions and imperfections in the economy. These may lead to deviations from the standard predictions of the model in the short to medium term but, for an analysis of long-run behaviour the model should be valid.

To conclude, within the framework of the model we distinguish four different types of influences on factor rewards. First, there are the effects that come from changes in goods prices that are either due to changes in trade policy, exogenous price changes in the world market for goods, or domestic shocks such as droughts. Secondly, there is the effect on factor returns due to changes in the factor endowments of the economy. Thirdly, we have effects from technical progress, and finally distortions.

III. Factor Endowments

Our model predicts that factor abundance is a crucial determinant of the pattern of specialisation and of factor prices, although market distortions may make factor prices diverge from their equilibrium values. We have factor estimates for the period from 1964 to 2000.³ Our capital stock estimates (K) are taken from the KIPPRA database (Ryan, 2002), while labour (L) and land (T) data are from World Development Indicators (2002).⁴ Using these data we have computed the endowment ratios, K/L , T/L and K/T , shown in Figure 1. The K/L ratio is very important for the pattern of specialisation and factor price outcomes. We see that the ratio increased until 1982 but that the trend then was reversed. Moreover, as expected, the capital-land ratio grows and the land-labour ratio declines continuously.

The pattern of factor growth up to 1982 should, according to our model, have led to increasing land rents and tended to increase wages and reduce capital rentals. The post-1982 pattern should have generated higher capital rentals and lower wages, while land rents should have continued to increase.

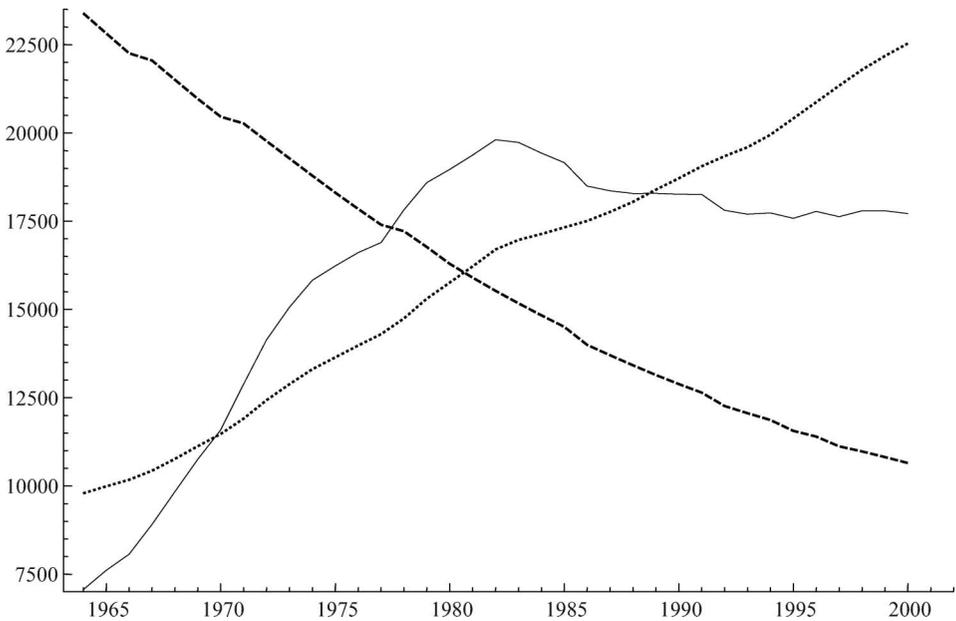


Figure 1. Relative factor endowments in Kenya 1964–2000. *Note:* $K/L = \text{—}$, $T/L = \text{---}$, $K/T = \text{.....}$. The variables T/L and K/T have been mean and variance adjusted to increase the readability of the graph.

IV. Factor Markets

Before independence in 1963, there was a system of minimum wages in place that generated substantial real-wage increases (Bigsten, 1984, 1986). This process continued for a period after 1963 but at the same time there was a very high rate of population growth, resulting in a rapid increase in the labour force. Initially, most of this increase was absorbed by the smallholder sector. However, as labour supply growth continued to outstrip formal sector job creation and the use of minimum wages to push up real formal sector wages became increasingly ineffective, private sector real wages fell. As a result, the period from 1968 onwards can be characterised as one of competition as far as the labour market is concerned (Collier and Lal, 1986). Still, some labour market controls remained until the 1990s when several major labour market reforms took place. Trade unions were allowed to seek full compensation for price increases from 1994 and the relaxation of wage guidelines made it possible for firms and employees to negotiate wages on the basis of productivity considerations rather than the cost of living indices that had earlier been the case (Ikiara and Ndung'u, 1997). In addition, the redundancy laws were amended in 1994, making it easier for firms to shed excess labour.

The evolution of real wages is depicted in Figure 2. It is measured as earnings, including allowances, of employees in the private sector divided by the gross domestic product (GDP) deflator. Figure 2 shows that real return to labour increased by about 25 per cent from the mid-1960s until the beginning of the 1970s.

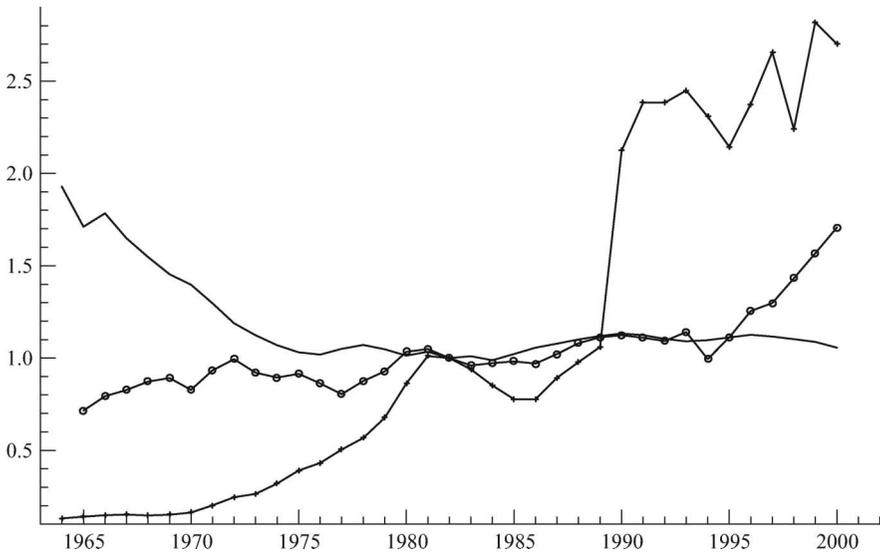


Figure 2. Indexes of real returns to factors in Kenya 1964–2000. *Note:* Real return to capital = —, real return to labour = ○—○—○, real return to land = +—+—+. The GDP deflator was used to calculate the real values of earnings and land prices. The base year is 1982 = 1. The series for land prices is the moving average of the actual series.

This was followed by stagnation until 1995 when real wages started to increase rapidly, a process that continued for the rest of the decade; the growth in private sector real earnings between 1994 and 2000 was 65 per cent.

One might suspect that wages in agriculture, which can be taken as an imperfect proxy for unskilled wages, would have increased at a slower rate or even declined with deregulation but, real wages in agriculture also grew at approximately the same rate as private sector wages outside agriculture. Hence, the choice of wage series does not affect the conclusions drawn from our analysis.

The rapid wage growth during 1994–2000 was probably due to the labour market reforms, as argued by IMF (2003), but it was amplified by wage-price dynamics. Prior to the reform, consumer-price inflation increased sharply, peaking at close to 40 per cent in 1993. As a result, wages, measured in terms of consumer goods, declined every year between 1989 and 1994. The real-wage decline and labour-market liberalisation generated collective agreements that led to substantial nominal wage increases, at the same time as inflation dropped abruptly, that is, to 1.5 per cent in 1995, and then stayed at a single-digit level.⁵ In Figure 2, the wage series is fairly stable until 1996 because the GDP deflator lagged behind consumer prices, growing more or less at the same speed as nominal wages.⁶

When looking at the evolution of wages post-2000, we see that there was another two years of rapid increases but then growth in real earnings came down to more modest levels (3–4 per cent per year in 2003 and 2004). It thus seems plausible that the wage increases from the mid-1990s were part of a temporary adjustment process.

During the colonial period there were land market interventions in favour of the settler community but these were eliminated at independence in 1963. Still, the land

market in Kenya has remained fairly inflexible and it has continued to be difficult for farmers to buy and sell land in some areas. Nevertheless, the land that is actually traded has been sold at freely negotiated prices. Consequently, even if those are not full equilibrium prices, development over time should give a reasonable reflection of how the scarcity value of land has changed. We use real land prices as a proxy for land rentals following Williamson (2002) and O'Rourke and Williamson (2002). Land prices are not an ideal proxy for land rent but they would reflect the return on land in a perfect market. We collected prices from agricultural land sales in two districts (Kakamega and Kiambu) and constructed indexes that are used for land rentals.⁷ Although the districts and their land price levels are very different, both series have similar growth rates. This indicates that our measurement of the evolution of land prices, which is used in the empirical analysis, is fairly accurate. Figure 2 shows a smoothed version of real land prices in Kiambu, which are available for a longer period than the Kakamega prices. Real land prices and return to land have risen sharply during the last 35 years; over the whole period they increased fifteen fold. Hence, in our case they behave as predicted by theory, at least in the sense that they have grown faster than returns to labour and capital.

The third factor return depicted in Figure 2 is the real return to capital. During most of the period covered in our subsequent analysis, there were extensive capital market interventions with controlled interest rates and administrative interventions in the allocation of credit (Bevan et al., 1990; Isaksson and Wihlborg, 2002). Financial liberalisation was part of the structural adjustment efforts that started in the early 1980s but the process was gradual and it is only recently that interest rates have approached equilibrium rates. Therefore, we do not use interest rates to measure capital rentals. Instead, our measure of return to capital is derived with the approach developed by Sarel (1997) and described in Appendix 2. It uses the national accounts and a production function to compute the marginal return to capital. The novelty of the approach is that the calculations of the technical capital shares for the individual sectors (at the one-digit GDP level) are averages obtained from a number of other countries. In this way we can correct for the classification of incomes of self-employed as capital income, something that cannot be done with Kenyan national accounts data alone. We obtain capital shares that vary between 0.245 and 0.280 over the period 1964–2000. Our values are in line with international averages and, for instance, close to those obtained for the Ivory Coast where data on compensation for self-employment are available (see Gollin, 2002).

As Figure 2 shows, return to capital declined continuously from 1964 to the mid-1970s, in total a reduction of 50 per cent. After 1975 it fluctuated slightly, with an increase from 1985 to 1990, a period of stability, and a small decline after 1995.

V. Openness

Soon after independence in 1963 Kenya started to pursue an import substitution policy, gradually raising the level of protection of the industrial sector during the 1960s. This policy continued until about 1980, when macroeconomic imbalances had become so severe that the country had to embark on a structural adjustment programme, initiating a gradual liberalisation of international trade. This process

lasted until 1993–94 when the liberalisation of the current and capital accounts was more or less completed.

Sachs and Warner (1995) constructed an openness indicator based on five variables, namely tariff level, extent of non-tariff barriers, black market premium, presence of state monopolies in exports of major crops and whether countries had a socialist economic system. By this approach they concluded that Kenya was open by 1993. A recent paper by Wacziarg and Welch (2003) updated Sachs and Warner's data set and redid the exercise, and they also conclude that Kenya opened up in 1993 and remained open during the rest of the 1990s (see also Ryan, 2002). It is of course not straightforward to judge when a gradual liberalisation process has reached the point where the country changes from closed to open but it is clearly fair to say that trade interventions remained extensive into the 1990s.

The policy of import substitution should have pushed up the relative price of industrial goods, and it should have had an impact on factor rewards separate from the factor endowment effect. Our theory predicts that Kenyan trade policy interventions in the 1960s tended to increase capital rentals and reduce land rentals, while we would expect the reverse effects during the 1980s and the first half of the 1990s, given relative resource endowments. Changes in world market prices would also have an effect. One such major external event during the period was the coffee boom in the 1970s, which temporarily pushed up the returns to land.

To give a good description of changes in the level of protection in Kenya we would need consistent time series on tariffs and detailed information about quantitative restrictions. What is available for part of the period are estimates of tariffs collected as shares of imports.⁸ These estimates were 13–14 per cent in the mid-1970s, peaking at close to 23 per cent in 1982, before declining to the 13–14 per cent range again in the late 1990s (World Development Indicators, 2003). However, the actual tariff rates on the books were considerably higher than what this measure shows, but some importers were able to get their import taxes waived. There were also extensive quantitative restrictions on imports, which had a large effect on domestic prices. We have no means to quantify these effects.

Given the above complications we use three different variables as proxies for trade policies in the econometric analysis of the determinants of factor prices in the next section. Figure 3 depicts a dummy that is based on the assumption that Kenya was a relatively closed economy from independence until the beginning of the 1980s, that there was a process of opening up between 1982 until 1992 and, that Kenya was open after that. The dummy, denoted, *OPEN*, is a stylised description of the opening-up process but it should capture major changes in protection levels.

The second proxy measures the impact of trade-policy reform by relating the development of domestic market prices of manufactured goods with the development of world market prices of industrial goods (converted to Kenya Shillings using the official exchange rate).⁹ More specifically, we use the GDP deflator for manufacturing in Kenya and the price index for industrial production in the UK, the latter being a major trading partner.¹⁰ This indicator should capture changes in the level of protection of the manufacturing sector. The development of the log of the indicator, denoted *RPM*, is shown in Figure 4. It provides more or less the same information as *OPEN*, but the trade liberalisation has already started in the mid-1970s and there is a small reversal after 1993.



Figure 3. Openness dummy. *Note:* The dummy is based on the detailed information about trade liberalisation provided in the KIPPRA database (Sachs and Warner, 1995; Ryan, 2002; Wacziarg and Welch, 2003).

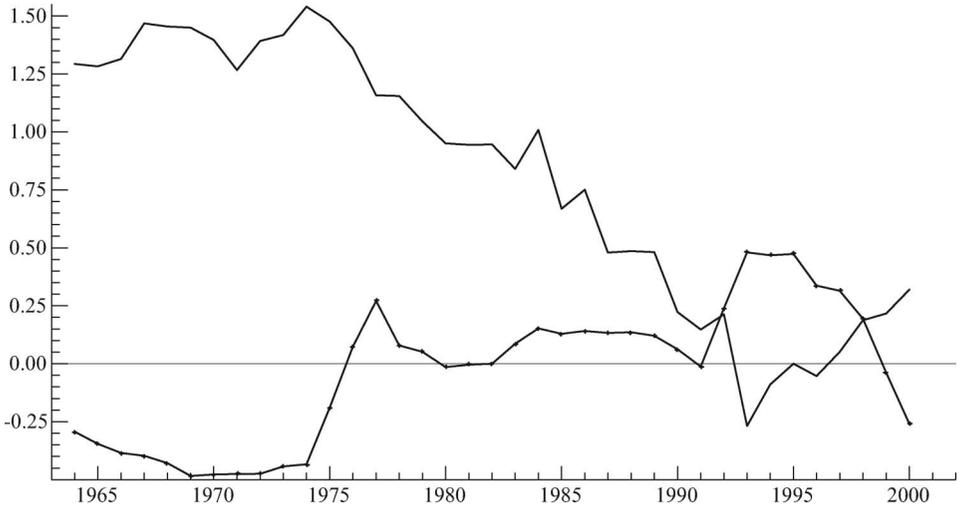


Figure 4. Log of relative goods prices (*PAPM*) and the ratio of Kenyan to UK manufacturing prices (*RPM*), 1964–2000. *Notes:* *RPM* = ———, *PAPM* = + + +. *RPM* is the log of the ratio between the deflator for the manufacturing sector in Kenya and the price index of industrial production in UK, converted to Kenya Shillings using the official exchange rate. *PAPM* is the log of the ratio of the deflators for agriculture and manufacturing.

In theory, the correct measure of trade policy is the ratio between the domestic prices of exportables, that is agricultural goods, and importables, which are manufacturing goods (O'Rourke and Williamson, 2002). We denote the log of this indicator *PAPM*; it is measured as the ratio between the sectoral deflators. In the

Kenyan context, an increase in *PAPM* may, *ceteris paribus*, be interpreted as an indication of increased integration in the world economy and vice versa. This is based on the fact that import substitution policies, which have been a dominant obstacle to international integration, aimed at increasing prices of manufactured goods relative to agricultural goods.

Figure 4 shows the evolution of *PAPM*. It remains low until 1975–77 when the coffee boom pushed up agricultural prices dramatically. Then there is small downward correction. During the 1980s, *PAPM* changes little, in spite of a strong decline in terms of trade (not reported). This could indicate that the negative effect on *PAPM* was counteracted by liberalisation measures in a series of structural adjustment programmes pushing up relative agricultural prices. The pattern of the 1980s may also have been affected by the introduction of purchasing power parity-based pricing of agricultural goods in the 1980s (Mwega and Ndung'u, 2002). In the early 1990s there was again a drought increasing *PAPM*, while the later part of the 1990s is a period of falling *PAPM*. During this period there were exogenous shocks such as falling coffee prices and, a decline in terms of trade by 16 per cent between 1998 and 2000 (Economic Survey, 2003).

The recent behaviour of *PAPM* seems to contradict the claim that Kenya was an open economy from about 1993; it declines during the latter half of the 1990s and in 2000, *PAPM* is at the same level as in the mid-1970s. However, the Kenyan relative goods' prices have been strongly affected by non-policy factors. Changes in goods' prices in the domestic market may either be due to changes in the international goods market or to changes in domestic policies or to exogenous shocks. What we observe is the aggregate outcome of all the influences and it is hard to separate out the effects of the various influences. Still, changes to relative goods' prices should in any case have an effect on factor prices in an open economy.

VI. Measuring Total Factor Productivity Growth

Before estimating models of relative factor prices, we analysed the evolution of total factor productivity (TFP) since changes in productivity are likely to affect factor returns. Three different approaches were used to estimate TFP growth. They all show that there was hardly any TFP growth during the period 1964–2000, which is in line with the results of the detailed study by Gerdin (1997) for the period 1964–94. To save space, we provide only a brief summary of the results.

To measure TFP growth, we first used growth accounting with the same income shares for capital and labour as in the derivation of the return to capital (see Appendix 2). The analysis shows that TFP declined by 2 per cent during the period 1964–75, then grew by 5 per cent over 1975–85 and 1 per cent over 1985–95, and then dropped by 2 per cent between 1995 and 2000. All in all, TFP rose by 1.8 per cent between 1964 and 2000. Next, we estimated a log-linear production function with capital and labour. Tests for cointegration based on both the Engle-Granger Two-Step approach and Fully Modified OLS also indicate that there was hardly any growth in TFP during the period 1964–2000, that is the level of TFP is a stationary process. Moreover, the cointegration analyses support the assumption of constant returns to scale; the estimates of the elasticity for capital ranged from 0.23 to 0.37, depending on the specification. Finally, based on the assumption of a

constant-returns-to-scale Cobb-Douglas production technology, we calculated TFP using the elasticities (0.5, 0.5), (0.3, 0.6) and (0.2, 0.8) for capital and labour respectively. In no case is the TFP growth rate more than 1 per cent during the period 1964–2000. Hence, although we have not analysed sectoral technical progress, our findings indicate that productivity growth was of second order importance in Kenya during the period of our study.

VII. Empirical Analysis

In this section we look at the determinants of relative factor returns. Since we are interested in their long-run determinants and several of the variables are non-stationary, we first test for cointegration. We then estimate error correction models (ECMs) to evaluate the adjustment process towards the long-run equilibrium. By developing empirically stable ECMs we are able to substantiate the results from the cointegration analyses and show that Granger causality runs in the expected direction, which is from endowments to factor returns. Moreover, it allows us to evaluate the stability of each parameter and gain insight into the role of individual variables.

The specification of the long-run model is,

$$f_j = c_j + \beta_{1j}LIB + \beta_{2j}KL + \beta_{3j}KT \quad (1)$$

where f_j is the log of the wage-capital rental ratio (WR) or land-capital rental ratio (LRR); LIB is either $OPEN$, the openness dummy, RPM , the log of the ratio of Kenyan to UK manufacturing prices, or $PAPM$, the log of the price ratio of agricultural goods and manufactured goods. KL and KT are the log of capital-labour and capital-land ratios. We do not analyse the wage-land rental ratio, since it is a linear combination of the other two ratios. In principle, the tests for long-run relations should be done in a system but all variables do not matter for each relative factor price, and we have few observations. Hence, we estimate single-equation models.

The ECM is specified as,

$$\Delta f_{j,t} = c_j + \Phi_j D_{j,t} + \sum_{i=1}^n \pi_{j1,i} \Delta f_{t-i} + \sum_{i=0}^n \pi_{j2,i} \Delta x_{t-i} + \alpha_j [f_j + \beta_j x]_{t-1} \quad (2)$$

where Δ is the first difference operator, c_j is the constant, D_j is a vector of deterministic (dummy) variables, and x is a vector of explanatory variables.

Data Description

In this subsection, we plot some of the data series to show potential long-run relations and the stochastic properties of the variables. Figure 5 depicts the log of the capital-labour ratio (KL) and the mean and variance adjusted log of the wage-capital rental ratio (WR). The upper panel of Figure 5 shows that WR and KL evolve in a similar way until the beginning of 1990s and, that WR increases while KL stays flat during the rest of the period. The lower panel plots the two series for the period 1965–92. It highlights the close relationship between the two variables during the

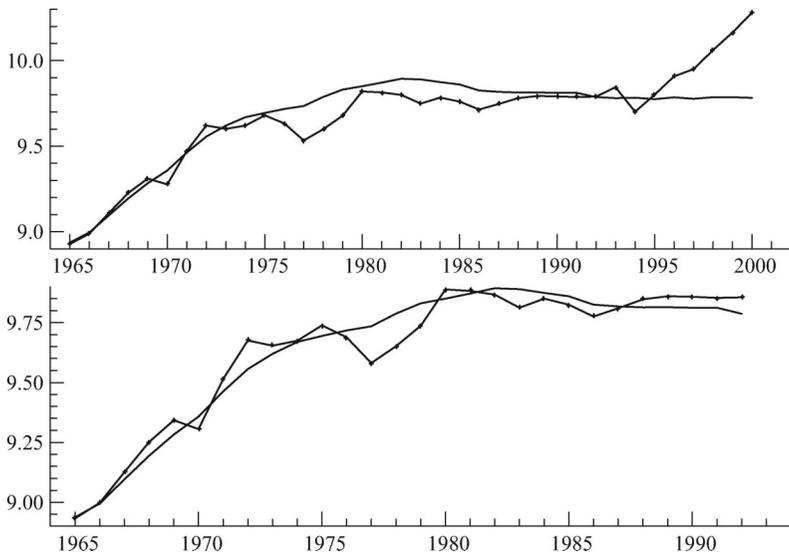


Figure 5. Log of capital-labour and wage-capital rental ratios, 1965–2000, 1965–92. *Note:* $KL = \text{—}$, $WR = + + +$. WR has been mean and variance adjusted to increase the readability of the graphs.

period when Kenya was closed according to Sachs and Warner (1995) and shows that they are likely to be cointegrated. Figure 5 also shows that it is not straightforward to describe WR and KL as first order difference stationary series, that is integrated of order one (denoted $I(1)$). The stochastic properties seem to change over the sample, maybe being $I(1)$ around a deterministic trend during the first 10 to 15 years, and then $I(1)$ or $I(0)$ (stationary) during the rest of the period. Nevertheless, both series appear to have the same stochastic properties and form a stationary relation during 1965–92.

The land-capital rental (LRR) and capital-land (KT) ratios are plotted in Figure 6. They follow each other closely over the whole sample, although LRR is much more volatile than KT . It seems obvious that the two series are cointegrated and follow a common stochastic trend. The two extreme observations for land prices in 1990 and 1994 are probably due to measurement errors, or special circumstances in the Kiambu sample, since they are not present in the series obtained in Kakamega. We treat them as outliers when estimating the ECM.

Data Analysis

Before testing for cointegration we carried out a series of unit root tests on the individual series. Since the finding of unit roots can be due to structural breaks, we used the tests that allow the breakpoint to be unknown.¹¹ However, no breaks were detected, possibly because of our small sample (results not reported).¹² Nevertheless, one finding of the unit root analysis was that no variable appears to have two unit roots when allowing for trends. Hence, we proceed under the assumption that there is at most one unit root in each series.



Figure 6. Log of capital-land (KT) and land-capital rental ratios (LRR), 1964–2000. *Note:* $KT = \text{—}$, $LRR = + + + +$. LRR has been mean and variance adjusted to increase the readability of the graph.

To test for cointegration, we first applied the two-step Engle-Granger procedure. However, since trade liberalisation should lead to changes in the long-run relations, we also tested for structural breaks. Then we used the Fully Modified OLS (FMOLS) estimator developed by Phillips (1988) and Phillips and Hansen (1990) which allows for cointegration tests and estimates of coefficients without requiring regressors that are weakly exogenous. Although the ideal is to test for cointegration within a well-specified full system using the Johansen (1988) approach, that would be demanding too much from our limited data set.

We start by regressing WR on KL and a constant for the period 1965–2000, and then add the indicators of trade liberalisation one at the time. Columns 1 to 4 in Table 1 show that the Engle-Granger test fails to reject the null of no cointegration in all cases. Since this result could be due to structural breaks we apply the Gregory and Hansen (1996) test, which is designed to detect cointegration when there is a shift in the long-run relation. It has a null hypothesis of no-cointegration and an alternative of cointegration with or without a break, where C allows for a level shift and C/S allows for a shift in both the level and the slope. All test statistics are insignificant except for the model with $OPEN$ where the null is rejected at the 10 per cent level for both the C and C/S models. According to the test, the break is in 1995. In spite of this result, we are reluctant to believe we have found a long-run relation for the wage-rental ratio. The period after 1995 consists of only six observations and they can easily be modelled by shifts in the constant and the coefficients of KL and $OPEN$. Moreover, the rapid growth in real wages that took place during 1990s was probably caused by labour market liberalisation implemented in 1994, as noted earlier, and not by trade liberalisation (IMF, 2003).

Since Kenya implemented several reforms in the mid-1990s, we also test for cointegration between WR and KL for the period 1965–94 (see column 5). The ADF

Table 1. Cointegration tests for wage-capital rental ratio: two-step Engle-Granger test and Gregory-Hansen cointegration test

	1	2	3	4	5
Time period	1965–2000	1965–2000	1965–2000	1965–2000	1965–94
Dependent var.	<i>WR</i>	<i>WR</i>	<i>WR</i>	<i>WR</i>	<i>WR</i>
Other variables	<i>KL</i> ,	<i>KL, OPEN</i>	<i>KL, RPM</i>	<i>KL, PAPM</i>	<i>KL</i> ,
ADF	0.588	-1.104	-0.607	0.723	-3.420*
ADF ^r , C	-3.862	-3.862	-3.862	-3.994	-4.069
ADF ^r , C/S	-4.062	-5.287*	-4.856	-4.693	-4.039

Notes: All ratios are in logs. ADF is the augmented Dickey Fuller test with a null of no cointegration; ADF^r is the Gregory and Hansen (1996) test where the null is no cointegration and the alternative hypothesis is cointegration with or without a break, where C allows for a level shift and C/S allows for a shift in the level and the slope. One lag was used in all unit root tests. * indicates significance at the 10 per cent level.

test is significant at the 10 per cent level, and although the Gregory-Hansen test fails to reject the null, the test statistics are probably close to being significant at the 10 per cent level.¹³

Since trade reform should generate structural change, we investigate the stability of the *WR-KL* relation further by applying the Andrews-Quandt and Andrews-Ploberger tests for breaks in linear regressions (Andrews and Ploberger, 1994; Hansen, 1997). For the period 1965–2000 they clearly reject stability (not reported); for instance the *p*-value for the Andrews-Ploberger test statistic for a break in all coefficients in 1995 is 0.02. However, there does not appear to be a break during the period 1965–92; the *p*-values for the test of all coefficients are 0.39 for the Andrews-Quandt test and 0.14 for the Andrews-Ploberger test. We thus conclude that *WR* and *KL* cointegrate over the period 1965–94 and that there is no structural break during the 1980s. A corollary of this is that we fail to find evidence of trade liberalisation affecting the wage-rental ratio before 1994.

Table 2 reports the results from the analysis of the land-capital rental ratio. Both the Engle-Granger and Gregory-Hansen tests reject the null of no cointegration at the 1 per cent level between *LRR* and *KT*. Tests for the presences of structural breaks using the Andrews-Quandt and Andrews-Ploberger tests under the assumption of cointegration do not reject the null of stability: the *p*-values for the joint test of all coefficients are 0.74, for the Andrews-Quandt test and 0.45, for the Andrews-Ploberger test. Hence, as Figure 6 illustrates, the land-capital rental ratio and the capital-land ratio are cointegrated over the period 1965–2000.

It is possible that the land-capital rental ratio together with a trade liberalisation indicator form a cointegrating vector without the capital-land ratio. We test for this but all the standard ADF-test statistics are insignificant (see Table 2). However, according to the Gregory-Hansen test, both *OPEN* and *RPM* cointegrate with *LRR* when we allow for structural breaks. According to the tests, there is a break in the intercept in 1973 and 1972, respectively. Although it is not obvious why openness should influence one factor-return ratio and not the other, we proceed under the assumption that there are three cointegrating vectors including *LRR* and investigate their explanatory power in ECM models.

Since we cannot rule out that the endowment ratios and trade liberalisation indicators are endogenous, we re-estimate the cointegrating relations using the Fully Modified OLS (FMOLS) estimator developed by Phillips (1988) and Phillips and Hansen (1990). In contrast to the two-step Engle-Granger cointegration test, which uses OLS, the FMOLS estimator is unbiased even when the regressors are endogenous and the t-statistic has an asymptotically normal distribution. The estimated coefficients and their t-values are reported in Table 3, where deterministic

Table 2. Cointegration tests for land-capital rental ratio: two-step Engle and Granger test and Gregory-Hansen cointegration test

	1	2	3	4
Time period	1965–2000	1965–2000	1965–2000	1965–2000
Dependent var.	<i>LRR</i>	<i>LRR</i>	<i>LRR</i>	<i>LRR</i>
Other variables	<i>KT</i> ,	<i>OPEN</i>	<i>RPM</i>	<i>PAPM</i>
ADF	−4.927***	−2.067	−2.442	−2.799
ADF ^r , C	−5.374***	−4.695**	−4.886**	−3.345
ADF ^r , C/S	−5.462***	−4.695	−4.903*	−4.125

Note: All ratios are in logs. ADF is the augmented Dickey Fuller test with a null of no cointegration; ADF^r is the Gregory and Hansen (1996) test where the null is no cointegration and the alternative hypothesis is cointegration with or without a break. C allows for a level shift and C/S allows for a shift in the level and the slope. One lag was used in all unit root tests. * indicates significance at the 10 per cent level, ** at the 5 per cent level, *** at the 1 per cent level.

Table 3. Fully modified OLS estimation of cointegrating vectors

	1	2	3	4	5
Time period	1965–2000	1965–94	1965–2000	1965–2000	1965–2000
Dependent var.	<i>WR</i>	<i>WR</i>	<i>LRR</i>	<i>LRR</i>	<i>LRR</i>
<i>KL</i>	0.852 (14.15)	0.825 (9.56)			
<i>DUMT95</i>	0.081 (6.10)				
<i>KT</i>			1.826 (14.705)		
<i>OPEN</i>					0.106 (5.56)
<i>RMP</i>				−1.034 (7.05)	
<i>DUMS72</i>				1.630 (7.93)	
<i>DUMS73</i>					1.811 (8.47)
<i>Constant</i>	−8.413 (14.48)	−8.02 (9.81)	−15.591 (14.59)	−0.404 (1.57)	−1.849 (10.93)

Notes: All ratios are in logs. DUMT95 is a trend dummy for the period 1995–2000 and DUMS72 and DUMS73 are step dummies that are zero until 1972 and 1973, respectively, and then unity. Absolute values of t-statistics are in parenthesis.

dummies capture the structural breaks revealed by the previous analysis. All the coefficients are statistically significant at the 1 per cent level, supporting our previous findings and they have the expected signs. To obtain a cointegrating relation for WR - KL for the whole sample, a time trend starting in 1995 was included to capture the change in the evolution of WR that cannot be explained by the capital-labour ratio or trade liberalisation (see column 1). It should not be viewed as a permanent part of the long-run relation but as a dummy that picks up a temporary deviation at the end of the sample period. We also estimated the WR - KL relation for the period before the break 1965–94, to check whether the inclusion of a trend dummy affects the coefficient of KL . As shown in column 2, the coefficients are very similar. The cointegrating vectors reported in Table 3 are used when estimating the ECMs in the following section.

Modelling the Dynamics of Relative Factor Prices

In this subsection we estimate ECMs for WR and LRR to test to what extent relative factor returns are driven by resource endowments and our measures of openness. By estimating ECMs we are also able to test for Granger causality: for our analysis to make sense the endowment ratios and openness indicators should Granger-cause relative factor returns. The time period for the analysis is 1967–2000, two observations are lost due to differencing and the use of lags. Since the analysis produced a lot of output we only give the main results in this and the following subsections (an Appendix with details is available on request).

First, we estimate a general unrestricted ECM for WR . The general ECM has one lag of ΔWR , contemporaneous and lagged values of ΔKL , the first differences of the trend dummy, and WR and KL and the dummy lagged one period. In addition, first differences of the liberalisation indicators also enter the general model. One lag of the differenced variables is sufficient to capture the dynamics and to pass all the misspecification tests (an Appendix with details is available on request). After estimating the general ECM, insignificant variables are excluded to obtain an unrestricted parsimonious ECM, and then the long-run restrictions are imposed based on the estimated parameters of the variables in levels. The preferred restricted ECM is,

$$\begin{aligned} \Delta WR_t = & -4.7 + 1.97\Delta KL_t + 0.08\Delta OPEN + 0.13\Delta DUMT95 \\ & (-3.70) \quad (6.20) \quad (3.01) \quad (5.08) \\ & -0.49[WR - 0.96KL - 0.08DUMT95]_{t-1} \\ & (-3.66) \end{aligned}$$

$$R^2 = 0.673 \quad \hat{\sigma} = 0.045 \quad T = 1966 - 2000 \quad DW = 1.84 \quad F_{ar}(2, 28) = 2.10 [0.14]$$

$$F_{arch}(1, 28) = 0.62 [0.44] \quad F_{het}(6, 23) = 0.71 [0.65] \quad \chi^2_{norm}(2) = 4.96 [0.83]$$

$$F_{RESET}(1, 29) = 0.00 [0.98] \quad \text{Tests of model reduction: } 1967 - 2000 : F(4, 25) = 0.28 [0.89]$$

where t -values are shown in parentheses and p -values in brackets, $\hat{\sigma}$ is the residual standard deviation, T is the sample period, and DW is the Durbin Watson test statistic. The diagnostic tests are approximate F -tests against, serial correlation of order 2, F_{ar} ; autoregressive conditional heteroscedasticity of order 1, F_{arch} ;

heteroscedasticity, F_{het} ; non-linearity, the RESET test, F_{RESET} ; and a chi-square test for normality, $\chi^2(2)$ (see Hendry and Doornik, 2001).

The preferred ECM is well specified since no diagnostic test is significant and all the t -values are high. Moreover, the test for reducing the general model to the parsimonious model is insignificant. Since the long-run coefficients are close to those reported in Table 3, the use of a single-equation ECM seems to be valid.

According to our model, WR is determined by KL in the long run but there is a structural break in 1995. The adjustment process is fairly rapid, 50 per cent of the disequilibrium is corrected within a year. Furthermore, growth in the capital-labour ratio increases growth in the wage-capital rental ratio, as expected. And there seems to be a short-run effect from trade liberalisation, since the change in $OPEN$ increased the growth rate of WR over the period 1982–92. Note that $\Delta DUM95$ enters significantly because $DUM95$ is restricted to be in the error correction term.

The presence of $DUM95$ could be due to opening-up, since it shows that labour earnings grow quicker than returns to capital while the capital-labour ratio was constant. However, there were relative-price changes that indicated reduced openness during the latter half of the decade; RPM increased, $PAPM$ decreased and, the real effective exchange rate appreciated by about 40 per cent from 1993–94 to the late 1990s (IMF, 2004). The most likely explanation is, thus, that labour market liberalisation and the large rise and fall in inflation during the 1990s have created a medium-term deviation from a long-run equilibrium: several years of stagnant labour productivity and recent slow wage growth are signs of a reversal back to long-run equilibrium (Economic Survey, 2005). However, this is occurring outside our study period and can only be substantiated by future research.

Next, we estimate a general ECM for the land-capital rental ratio. It contains ΔKT_t , ΔKT_{t-1} and ΔLRR_{t-1} , and impulse dummies for the outliers in 1990 and 1994 (see Figure 6). Since the cointegration tests indicated that there are three different cointegrating vectors, as reported in columns 3–5 in Table 3, we estimate the ECM with three different error correction terms. However, only the one derived from column 3, $ECMLRR = LRR - 1.83KT$, is significant (an Appendix with details is available on request). After removing insignificant variables the following parsimonious ECM is obtained,

$$\begin{aligned} \Delta LLR_t = & -12.2 + 2.98\Delta KT_t + 1.28DUMI90 + 0.97DUMI94 - 0.78[LLR - 1.83KT]_{t-1} \\ & (-6.34) \quad (2.40) \quad (4.73) \quad (3.64) \quad (-6.25) \\ R^2 = & 0.742 \quad \hat{\sigma} = 0.260 \quad T = 1965 - 2000 \quad DW = 1.81 \quad F_{ar}(2, 29) = 0.47[0.63] \\ F_{arch}(1, 29) = & 0.21[0.65] \quad F_{het}(6, 24) = 0.81[0.58] \quad \chi^2_{norm}(2) = 0.07[0.96] \\ F_{RESET}(1, 30) = & 0.05[0.83] \quad \text{Tests of model reduction: } 1966 - 2000, F(2, 28) = 0.41[0.66] \end{aligned}$$

The reduction from the general model is accepted and none of the diagnostic tests are significant. Again, resource endowments – here the capital-land ratio – drive relative returns in the long run, as shown by the significance of the error correction term. The adjustment is 78 per cent per year, which is faster than for WR . Moreover,

the growth rate in *KT* has a positive impact on the growth rate of *LRR* in the short run. None of the openness indicators enters significantly.

The findings that *KL* drives *WR* and *KT* drives *LRR* are supported by Granger causality tests: *WR* and *LRR* adjust to deviations from the long-run equilibrium relations; while *KL* and *KT* do not adjust, that is, they Granger-cause *WR* and *LRR*. Furthermore, recursive estimation and omitted variable tests show that the final models are empirically well specified and that both ECMs are reasonably stable over the period 1978–2000 (an Appendix with details is available on request). The LM-tests for omitted variables, which summarise the results from general-to-specific analyses, show that none of the excluded variables enter the models significantly.

VIII. Conclusion

Kenya started to dismantle restrictions on international trade in the early 1980s and, according to the definition of openness from Sachs and Warner (1995), it can be characterised as an open economy from 1993. Moreover, international trade has increased significantly, which may indicate that the Kenyan economy has taken a large step towards becoming integrated with the world economy. At the same time, Kenya has been in a process of structural transformation, characterised by rapid labour force growth and increasing land scarcity. Both these processes can be expected to affect factor prices.

Our analysis shows that the evolution of relative factor prices was essentially determined by changes in relative resource endowments during the period 1965–2000; the wage-capital rental ratio was driven by the capital-labour ratio while the land-capital rentals ratio was driven by the capital-land ratio. We fail to find convincing evidence that that increased openness has significantly affected relative factor prices.

There was rapid growth in wages during the latter part of the 1990s that cannot be explained by resource endowments but we do not interpret this as an effect of trade liberalisation. It is not picked up by any of our openness indicators and two of them even indicate reduced openness. The most likely explanation is that wages rose quickly because of labour market deregulation, which had been preceded by a period of unusually high inflation and when slow wage adjustments had eroded real wages. Consequently, nominal wages continued to grow rapidly during the rest of the decade while inflation dropped abruptly in 1995 and then stayed at a low level. Wage growth declined to modest levels in 2003 and 2004 and there have been no improvements in labour productivity, so it seems likely that the wage increases constitute a medium-term deviation from long-run equilibrium; this is an issue for future research.

The most dramatic factor price change during the period analysed was the increase in land prices, reflecting the increased land scarcity. Since much of the land in Kenya is owned by relatively poor smallholders, this may have had beneficial effects on household income distribution.

Our conclusion is that factor price changes were driven by changes in factor endowments during the period 1965–2000, in spite of the removal of many trade restrictions in the later part of the period. Consequently, the changes in income

inequality, that have taken place in Kenya since independence, have been influenced much more by the long-term process of structural change than by changes in international economic policy. This suggests that African governments need not be so concerned about the distributional implications of their trade policy but rather let it be decided on the basis of growth and efficiency considerations. Still, the effect of openness on factor prices may have been dampened by a lack of full integration between Kenyan and international markets. Much of the economy is not trading internationally and even sectors that are trading may be partially isolated from external competition. This could mean that the impact of international economic integration could increase in the future if Kenya becomes more fully integrated with the world economy.

In our analysis we have treated factor accumulation as exogenous but, at least as far as capital is concerned, this can be questioned. With trade liberalisation, the price of (largely imported) capital goods went up and the profitability of investments in the manufacturing sector fell as international competition increased. It would be of interest to extend the analysis by making capital accumulation endogenous and, by doing so, one might possibly find that the liberalisation policies had a larger indirect impact on factor prices than that which our analysis finds.

Another possible explanation to the surprisingly low impact of trade liberalisation is that there remain restrictions to trade that inhibit the impact of international prices. These can be lack of competition, poor infrastructure and high transport costs, both domestically and internationally, political uncertainty and so forth. Finally, by doing the analysis on a highly aggregate level we may have concealed some effects, but to go beyond this we need industry-level data and to use more elaborate measures of openness. Further studies are needed to analyse these issues.

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Notes

1. The recent review by Anderson (2005), notes that 'there has been a large amount of research into the effect of openness on one particular factor price, the wage of skilled relative to unskilled labor' (Anderson, 2005: 1051) and, 'We still know little about whether, and if so how, greater openness has affected other factor-price ratios, such as the return to land (and other natural resources) relative to labor, skilled or unskilled, in low- and middle income countries. Movements in this latter ratio may have a much larger impact on overall inequality in Africa and Latin America, with their relatively abundant supplies of land and other natural resources' (Anderson, 2005: 1057).
2. Manda and Sen (2004) found that trade liberalisation had a negligible impact on the structure of the manufacturing sector, which points in the same direction as our results.
3. See Appendix 1 for details about the data used in this section.
4. The variable used for land is arable land (in hectares), which includes land defined by the FAO as land under temporary crops, temporary meadows for mowing or for pasture, land under market or kitchen gardens, and land temporarily fallow.

5. Minimum wages are still set by the Ministry of Labour but since they are increasingly falling behind average wages their contribution to the general wage increase seems to be rather modest.
6. The deviation between the price indexes lasted until 2003, when the GDP deflator caught up with the consumer price index. One reason for the deviation was a series of devaluations during the period 1989–93 that accumulated to 180 per cent for the Shilling-US dollar exchange rate. From 1993 to 2004 the Shilling was only devalued by 36 per cent (World Development Indicators, 2005).
7. See Appendix 1 for a description of the data on land prices.
8. Export taxes have not been important in Kenya.
9. Athukorala and Rajapatirana (2000) use the same measure for an analysis of the Sri Lankan experience.
10. We also calculated the index with US prices but the overall pattern was very similar. Since imports from the UK are much larger than those from the US, we prefer to use UK prices. In principle, a trade-weighted index would be the best measure but annual revisions of the trade weights generate abrupt changes in the index due to the high level of time aggregation.
11. We use the tests of Zivot and Andrews (1992) and Perron (1997) that allow for unknown breakpoint. See Patterson (2000: 277–283) for details.
12. All the unit root and cointegration tests were implemented with RATS procedures. The results from the unit roots tests can be obtained from the authors upon request.
13. Unfortunately, we do not have p -values for the Gregory and Hansen (1996) test.

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Appendix 1

Data Sources

Land Prices: we collected data for land prices from all registered sales of land in two provinces, Kiambu and Kakamega, and calculated the average price per hectare for each year. The data runs from 1960–2000 for Kiambu and 1968–2000 for Kakamega. Kiambu is close to Nairobi so prices are much higher than in Kakamega. However, when turned into indexes of real prices, the evolution of the two prices series is very similar apart from two outliers in each series. We use the Kiambu series in the empirical analysis since it spans a longer time period.

Labour Earnings: the sources are various issues of Statistical Abstract of Kenya and the KIPPRA database (Ryan, 2002). Although different series of labour earnings in the private sector are available, they all follow a similar pattern. However, there is a difference between labour earnings in commercial agriculture and earnings in the rest of the private sector. Earnings in commercial agriculture, when measured in real terms using the GDP-deflator, grew faster than earnings in the rest of the economy from the end of the 1980s until the mid-1990s. In our regressions, returns to labour are measured by earnings, including allowances, of all employees in the private sector, including agriculture. However, we also carried out the analysis with earnings in commercial farming, since farm workers' wages are likely to be close to wages in the informal sector. Moreover, they should have been less affected by changes in the

skill composition of the labour force over the last decades and, thus, constitute a better measure for unskilled wages. Empirically, the only difference between using wages in agriculture and private sector wages is that the break in the relation between the wage-capital rental ratio and the capital-labour ratio is in 1991–92, for the former series, and in 1995 for the latter series.

Commodity Prices: we used the deflators for agricultural production and manufacturing production. They are from the KIPPRA database (Ryan, 2002). The price indexes for manufactured goods in the UK and the US were obtained from World Development Indicators 2003.

Land and Labour: both are from World Development Indicators 2003. Land is defined as arable land (in hectares), which includes land classified by the FAO as land under temporary crops, temporary meadows for mowing or for pasture, land under market or kitchen gardens, and land temporarily fallow. Since land is only measured irregularly, it was smoothed using nine-year moving averages. When measured in percentages, the yearly changes in the land series are very small.

Capital Stock: the total capital stock series is from the KIPPRA database and covers the period 1964–2000. It was originally constructed by the Long Range Planning Unit at the Government of Kenya for the period 1972–94 and based on seven different kinds of capital and 18 sectors, with depreciation rates varying according to the kind of capital. This series was revised, updated and extended backwards by KIPPRA. Another estimate of the capital stock is available. It was originally compiled by Nehru and Dhareshwar (1993) and updated by Bosworth and Collins (2003). This series is based on aggregate fixed capital formation and one depreciation rate for the whole capital stock. Overall, both capital stocks follow a similar pattern; for instance, when measured as capital per worker, the capital-labour ratios peak in the beginning of the 1980s. However, the Bosworth and Collins estimates of the capital stock do not seem to grow fast enough. For instance, the capital-labour ratio is the same in the latter half of the 1990s as it is during the 1960s and the capital-output ratio declines almost continuously between 1960 and 1990, and then settles at a level 40 per cent below the one in the 1960s. Moreover, the capital-output ratio declines from 2 to 1.5 between 1968 and 1986, while gross fixed capital formation averages 22 per cent of GDP. To find out why the Bosworth and Collins capital stock series behaves in such a way, we used the value of the capital stock in 1960 as a starting value and altered the depreciation rate. To obtain an almost stable capital-output ratio, the depreciation had to be reduced to 0.02, which is low. Hence, it seems as if the capital stock for 1960 is over-estimated: too large an investment share of GDP is needed to maintain a reasonable growth rate of the capital stock.

Appendix 2

Calculation of Return to Capital

This appendix explains how capital rentals were computed. The problem of calculating the capital rentals is to find values for R in $Y = RK + WL$, where Y is real

GDP, R the real return to capital, K the capital stock, W the wage rate and L the labour supply. Sarel (1997) and Gollin (2002) review the difficulties of using national account data on factor income to obtain R . Here we use the fact that with a Cobb-Douglas production function and perfect competition, R is also the marginal product of capital.

In the first step the capital shares, $\alpha_t = R_t K_t / Y_t$, for Kenya were calculated with the help of information about technical capital shares provided by Sarel (1997) and used by, among others, Fernald and Neiman (2003). Sarel estimated sectoral capital shares for a number of countries at the one-digit level of GDP and calculated their averages. Although these calculations do not give the true capital shares, they are likely to be closer to them than what can be obtained for Kenya using factor incomes from national accounts data. The most serious problem with national accounts data is that income from self-employment is included in capital income, which we cannot do anything about with the Kenyan data (see Gollin, 2002). In Sarel's estimates, employment compensation is corrected for self-employment. The capital shares of Sarel were applied to each one-digit sector of GDP in Kenya and the weighted averages were calculated for each year over the period 1964–2000. The capital shares were slightly above 0.24 in the 1960s, and then increased slowly to close at 0.27 at the end of the 1990s.

In the second step R was calculated as $\alpha_t(Y_t/K_t)$, that is the marginal product of capital. R starts at 0.18 and declines to 0.1 in the middle of the 1970s and hovers around this value until 2000. Since the choice of capital stock influences R , we investigated the sensitivity of the wage-rental ratio to the evolution of the capital stock. Our measure of capital stock comes from the KIPPRA database (see Appendix 1). The alternative capital stock is from Nehru and Dhareshwar (1993) and Bosworth and Collins (2003). To compare the behaviour of R we redid the calculations using data from Bosworth and Collins. Starting with a depreciation rate of 0.05, which Bosworth and Collins use, R grows until the mid-1970s and then remains stable. However, as noted in Appendix 1, such a high depreciation rate leads to an almost continuous decline in the capita-output ratio, which seems unrealistic. When the depreciation rate is reduced to 0.02, a level where there is an almost stable capital-output ratio, we obtain an R that behaves as our measure of R . Hence, our measure of R is robust to changes of capital stock as long as the depreciation rate is not so high that the capital-output ratio declines continuously.