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African terms of trade and the commodity terms of trade: close cousins or distant relatives?

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This paper examines whether there is a relationship between the commodity terms of trade (the price of primary commodities relative to the price of manufactures) and the net barter terms of trade of 42 Sub-Saharan African countries. For most countries, there is little evidence of a stable, long-run relationship between the two terms of trade series. Accordingly, the practice in the literature of proxying for movements in any given country's terms of trade by using an aggregate index of relative commodity prices is inappropriate, and is likely to engender misleading policy conclusions.

I. Introduction

There are numerous reasons why it may be desirable for Sub-Saharan African (SSA) countries to encourage diversification of their exports (World Bank, 2000). However, this policy prescription is often taken to the extreme, resulting in pessimism regarding SSA's potential for exploiting its comparative advantage in primary products. Particularly in past decades, this led to policies that penalized the agricultural sector, and lowered economic growth. The pessimism regarding primary product exports for individual Sub-Saharan African countries largely rests on the Prebisch–Singer hypothesis (PSH), which postulates that there is a long-run negative trend in the prices of internationally-traded primary commodities relative to those of manufactured products.[1]

There are a two main reasons to remain sceptical about this line of argument. First, although few hypotheses in development economics have been studied more intensively than the PSH, findings have varied to such an extent that there is still no consensus in the literature. Second, due to the availability of long-run time-series data, the PSH has been investigated using the relative price commonly referred to as the commodity terms of trade (*CTT*), which is the price of primary commodities relative to manufactured goods.[2] However, a secular, negative trend in the *CTT* implies a secular negative trend in the terms of trade for *individual* Sub-Saharan African countries only if there is a stable long-run

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[1] See Prebisch (1950), Singer (1950) and Grilli and Yang (1988), among many others. In addition, the notion that commodity price booms have been generally harmful to SSA countries has supported scepticism regarding the growth potential of the production of primary commodities, although this claim has been challenged by Deaton and Miller (1996) and Collier and Gunning (1999).

[2] As noted by Grilli and Yang (1988, p. 2): '. . .a common problem of the analyses that have focused on the long-run trends in the terms of trade of developing countries . . .has been the inadequacy of the basic price data. Long-term movements in the [aggregate] terms of trade of developing countries were either inferred from those of certain industrial countries or from the movements in the prices of primary commodities relative to those of manufactured products (the so-called primary commodity terms of trade). . .'.

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relationship between these two series. The goal of this paper is to determine whether such a long-run relationship exists, using cointegration analysis. In doing so, it is examined whether there are stable cointegrating regressions for the terms of trade of individual African countries and the *CTT*, using stability tests based on the null hypothesis of no cointegration (Said and Dickey, 1984; Phillips and Ouliaris, 1990; Gregory and Hansen, 1996a) and those based on the null hypothesis of cointegration (Hansen, 1992).

Why is this an important question? Grilli and Yang (1988) asked: if a secular decline in the *CTT* has occurred, does this imply that the net barter terms of trade of developing countries has experienced a trend decline? They explore the question by estimating the relationship between the *CTT* and *aggregate* indices of the terms of trade for non-oil developing countries. Grilli and Yang, as well as Powell (1991), Bleaney and Greenaway (1993) and Lutz (1999), have found some evidence of a long run relationship between these two series. Grilli and Yang also estimate the relationship between the *CTT* and an aggregate index of the net barter terms of trade for different regions. Over the period 1953–1983, they find that while the *CTT* is becoming less important for Southeast Asian countries, as they have successfully diversified into exporting manufactures, the opposite is true for African countries, which they note have become increasingly dependent on primary commodity exports. Thus, for African countries, they conclude that this evidence supports claims of a secularly declining net barter terms of trade. The implication is that since there is a relationship between the *CTT* and the *aggregate* net barter terms of trade for Africa, the *CTT* is a useful proxy for the terms of trade of *individual* African countries.

In this paper it is argued that the correct way to analyse this issue is to examine the relationship between the *CTT* and the national terms of trade of individual African countries.[3] On the one hand, it may seem more likely that the relationship between the *CTT* and *national* terms of trade series would hold for Sub-Saharan Africa than for other regions, since most Sub-Saharan African countries remain reliant on the export of primary commodities, while many other developing countries now export larger shares of manufactures.[4] This is the assumption of Grilli and Yang (1988) and Lutz (1999). However, it is also true that Sub-Saharan African countries are relatively undiversified across the range of primary products, with only one or two commodities often accounting for the bulk of export earnings (Table 1). In addition, export diversification varies across Sub-Saharan African countries, but on average is much lower than for other regions (Table 2). In the construction of the *CTT*, the primary commodities included in the index reflect their share in world commodity exports, and the manufacturing goods used to deflate the series of primary commodity prices reflect the export unit values of a group of industrial countries. Thus, on the other hand it could also be argued that it is unlikely that the largely undiversified export baskets of many African countries would bear much relation to the aggregate price of primary commodities. Narrow production structures also lessen the likelihood that African countries' import prices closely follow the diversified manufacturing unit values series. Although both these arguments appear reasonable a priori, the question can only be answered empirically.

The present paper responds to the challenge in Lutz (1999) to investigate the 'middle hypothesis' in a chain of argument relating to the terms of trade, which goes as follows. (1) There is a long-run negative trend in the prices of primary commodities relative to those of manufactured products (the PSH). (2) This implies a negative trend in the net barter terms of trade for those developing countries that are largely primary commodity exporters. (3) Declines in the net barter terms of trade lead to declines in output in these countries. The middle link in this chain is examined, focusing on a comprehensive study of 42 Sub-Saharan African (SSA) countries. The focus is on SSA since although there is still no consensus in the literature on point (1), if point (2) is accurate, this would provide support for those who are pessimistic about Africa's potential for exploiting its comparative advantage in primary commodities.

[3] While there are numerous studies that use individual countries' terms of trade series to examine the impact of terms of trade on growth or another macroeconomic variable, there are few studies of the properties of the national terms of trade series. An important exception is that of Bidarkota and Crucini (2000), who find that the most important factor in determining variation in the national terms of trade is the price volatility of the particular commodities in which a country specializes. In addition, Cashin *et al.* (2004) examine the persistence of shocks to the terms of trade of Sub-Saharan African countries.
[4] The World Bank's *World Development Indicators* (1999) indicates that the ratio of primary commodity exports to total merchandise exports for low- and middle-income countries as a group was 38% in 1997, down from 75% in 1980 and 48% in 1991. However, for Sub-Saharan African countries this ratio fell less rapidly during the 1980s, moving from 64% in 1980 to 53% in 1991.

Table 1. African countries dependent on a single primary commodity for export earnings (annual average of export data, U.S. dollars, 1992–1997)

Commodity	For 50% or more of export earnings	For 20–49% of export earnings	For 10–19% of export earnings
Crude petroleum	Angola, Congo Gabon, Nigeria	Cameroon Equatorial Guinea	
Natural gas		Algeria	
Bauxite and alumina	Guinea		
Iron ore		Mauritania	
Rutile			Sierra Leone
Copper	Zambia		Congo, Dem. Rep.
Cobalt			Congo, Dem. Rep. Zambia
Gold		Ghana South Africa	Mali Zimbabwe
Diamonds	Botswana	Central African Rep. Sierra Leone	Congo, Dem. Rep.
Uranium	Niger		
Timber (African hardwood)		Equatorial Guinea	Central African Rep. Gabon, Ghana
Cotton		Benin, Chad Mali, Sudan	Burkina Faso
Tobacco	Malawi	Zimbabwe	
Arabica coffee	Burundi, Ethiopia	Rwanda	
Robusta coffee	Uganda		Cameroon
Cocoa	São Tomé and Principe	Côte d'Ivoire, Ghana	Cameroon
Tea			Kenya, Rwanda
Vanilla		Comoros	
Sugar		Mauritius	
Livestock		Mali	Niger, Sudan
Cashew nuts	Guinea Bissau		
Fish	Mauritania	Mozambique	Senegal
Oilseeds			Sudan

Sources: Cashin and Pattillo (2000) and International Monetary Fund.

The fact that a number of studies have found that declines in the net barter terms of trade lower output and growth, particularly for SSA, also underscores the importance of the 'middle hypothesis' (Ghura and Hadjimichael, 1996; Calamitsis *et al.*, 1999).

In summary, the contributions of this study are three-fold. First, it is examined whether there is a stable, long-run relationship between the commodity terms of trade (*CTT*) and the net barter terms of trade of individual SSA countries. Empirically investigating this issue has particular policy significance for SSA countries, as their reliance on primary commodities implies that a stable relationship between the two terms of trade series would be expected to be found. Second, multiple cointegration tests are used, in order to give the data the best chance of finding a long-run relationship between the two terms of trade series, if it exists. Third, it is examined whether any failure to find a long-run relationship between the *CTT* and the national net barter terms of trade is due to a structural break in the relationship between the two series.

The remainder of the paper is organized as follows. In Section II it is examined whether there is any empirical relationship between the commodity terms of trade and aggregate indices of the terms of trade (for both non-oil-exporting developing countries and African countries). The following section describes the national terms of trade data for the 42 Sub-Saharan African countries used in the study, and presents the main empirical findings regarding the presence of a long-run relationship between the commodity terms of trade and the national terms of trade series. Some concluding comments are contained in Section IV.

II. The Relationship Between the Commodity Terms of Trade and Aggregate Measures of the Terms of Trade

Beginning with Grilli and Yang (1988), the empirical relationship between the net barter terms of trade and the commodity terms of trade has largely been

Table 2. Sub-Saharan Africa: export diversification and concentration, 1962 and 1995

Country	1962 Diversification index[a]	1962 Concentration index[b]	1995 Diversification index[a]	1995 Concentration index[b]
Angola	0.840	0.442	0.914	0.913
Burkina Faso	0.875	0.615	0.929	0.555
Burundi	n.a.	n.a.	0.947	0.676
Cameroon	0.837	0.391	0.848	0.323
Central African Republic	0.890	0.476	0.859	0.425
Chad	0.907	0.766	n.a.	n.a.
Congo, Democratic Republic	0.824	0.478	n.a.	n.a.
Congo	0.860	0.539	0.882	0.557
Cote d'Ivoire	0.918	0.507	0.905	0.408
Ethiopia	0.912	0.543	0.912	0.601
Gabon	n.a.	n.a.	0.931	0.730
Gambia	0.981	0.945	n.a.	n.a.
Ghana	0.943	0.697	0.889	0.380
Kenya	0.867	0.356	0.735	0.231
Liberia	0.943	0.624	n.a.	n.a.
Madagascar	0.849	0.311	0.825	0.284
Malawi	0.918	0.460	0.899	0.680
Mali	0.945	0.502	n.a.	n.a.
Mauritania	0.914	0.819	0.966	0.610
Mauritius	0.977	0.982	0.858	0.313
Mozambique	0.840	0.339	0.825	0.443
Nigeria	0.837	0.368	0.906	0.897
Niger	0.938	0.635	n.a.	n.a.
Rwanda	0.949	0.603	0.946	0.664
Senegal	0.919	0.498	0.818	0.283
Seychelles	n.a.	n.a.	0.896	0.677
Sierra Leone	0.947	0.612	0.890	0.502
Somalia	0.956	0.582	n.a.	n.a.
South Africa	0.689	0.166	0.611	0.256
Sudan	0.936	0.621	0.921	0.375
Tanzania	0.898	0.363	0.844	0.257
Togo	0.931	0.514	0.871	0.393
Uganda	0.913	0.544	0.960	0.816
Zimbabwe	n.a.	n.a.	0.728	0.250
Average for all developing countries	0.826	0.478	0.731	0.352
Average for Sub-Saharan Africa	0.907	0.552	0.881	0.509

Notes: n.a. denotes not available.
[a] The export diversification index measures the absolute deviation of the country's share of commodity exports in its total exports from the share of commodity exports in total world exports, and ranges between 0 (most diversification) and 1 (least diversification).
[b] The Hirschmann concentration index measures the commodity concentration of a country's export structure, and is normalized to make values ranking from 0 (least concentration) to 1 (maximum concentration).
Source: UNCTAD (1996), *Handbook of International Trade and Development Statistics*.

examined by ordinary least squares estimation of the following regression:

$$TOTALL_t = \beta_0 + \beta_1 CTT_t + \beta_2 ROIL_t + \varepsilon_t \quad t = 1, \ldots, T \qquad (1)$$

where $TOTALL$ is the International Monetary Fund's aggregate index of the terms of trade for non-oil developing countries, CTT is the Grilli–Yang (1988) index of the price of non-fuel primary commodities divided by the price of manufactures, and $ROIL$ is the nominal price of oil deflated by the price of manufactures. All series have typically been expressed in natural logarithms. Grilli and Yang estimated the above equation over the period 1953–1983, and found:

$$TOTALL_t = 3.63 + 0.279\, CTT_t - 0.089\, ROIL_t + \varepsilon_t \qquad (2)$$

As pointed out by Powell (1991), Grilli and Yang (1988, p. 35) argue that their finding of a stable and declining commodity terms of trade, in tandem with the results of Equation 2, indicate that there is a stable and declining terms of trade for non-oil-exporting developing countries. In particular, their results imply that a 1% decline in the relative prices of non-fuel primary commodities results (in the long run) in a fall in the *aggregate* terms of trade of non-oil-exporting developing countries of 0.28%. These aggregate results were largely replicated in subsequent studies by Powell (1991), Bleaney and Greenaway (1993), and Lutz (1999).

Two questions arise from this work. First, are earlier findings of a stable, long-run relationship between the commodity terms of trade, and the aggregate terms of trade of non-oil-exporting developing countries, still applicable using more recent data? Second, does this relationship between the two measures of the terms of trade hold if the *aggregate* terms of trade series is replaced by the terms of trade of *individual* countries? In Section III this latter question will be analysed, using the national net barter terms of trade of the commodity-exporting countries of Sub-Saharan Africa.

Empirical estimation

In this Section it is analysed whether there is a stable, long-run relationship between *aggregate* indices of the terms of trade, the commodity terms of trade (*CTT*) and real oil prices (*ROIL*). The approach to testing for the presence of such a relationship is to determine if the variables are cointegrated. In carrying out the cointegration analysis, it is first determined whether the three time series are integrated of order one (that is, possess a unit root), then the augmented Dickey–Fuller (ADF) test of Said–Dickey (1984) is used to determine if the residuals of the cointegrating regression are stationary. The annual data (for the period 1953–1998) are as described above; all series are expressed in logarithmic form, and their derivation is set out in Appendix B. Figure 1 plots the three series, and there is some evidence of a relationship between the *CTT*, the aggregate terms of trade series (especially *TOTALL*) and *ROIL* over the sample period. This hypothesis will be examined in greater detail below.

Using the Phillips–Perron (1988) unit root test, all series (*TOTALL*, the *CTT* and *ROIL*) were found to be nonstationary, and so the possibility of cointegration exists.[5] Updating the data used in Equation 2 to cover the period 1953–1998 yields the following least squares estimates of a potential cointegrating regression:

$$TOTALL_t = 3.66 + 0.183\,CTT_t - 0.081\,ROIL_t + \varepsilon_t \quad (3)$$

where an ADF test of the stationarity of the residuals from the regression fails to reject the null of no cointegration at the 5% level.[6] It is also examined whether cointegration existed between *TOTAF* (the International Monetary Fund's aggregate index of the terms of trade for African countries), the *CTT* and the *ROIL* over the period 1953–1998, which yielded:

$$TOTAF_t = 5.08 - 0.300\,CTT_t + 0.285\,ROIL_t + \varepsilon_t \quad (4)$$

where an ADF test of the non-stationarity of the residuals from the regression again fails to reject the null of no cointegration at the 5% level (the value of the ADF test is −3.28). In addition, the sign of the relationship between *TOTAF* and *CTT*, and between *TOTAF* and *ROIL*, is reversed from that found in Equation 3.[7,8]

[5] The critical value for the Phillips–Perron (1988) unit root test (including an intercept term) at the 5% level of significance is −2.93 (for $T = 50$). Results for the unit root test were: *TOTALL* (−1.88 (level); −6.50 (first difference)); *CTT* (−1.41; −6.14); and *ROIL* (−1.62; −7.59).

[6] The value of the ADF test is −3.41; from MacKinnon (1991), the 5 (10) per cent critical value (with a constant term) is (for $T = 46$) −3.93 (−3.59). In addition, using the same sample periods as Grilli and Yang (1988), Powell (1991), and Bleaney and Greenaway (1993), namely 1953–1983, 1953–1986 and 1955–1989, respectively, yields parameter estimates which are close to their studies, yet only for the latter period can the earlier finding of rejection of the null hypothesis of no cointegration at the 10% level of significance be confirmed.

[7] Whether cointegration existed between *TOTAFUN* (the World Bank's UNCTAD-based aggregate index of the terms of trade for African countries), the *CTT* and the *ROIL* over the period 1960–1996 was also examined, which yielded: $TOTAFUN_t = 5.03 + 0.030 CTT_t + 0.376\,ROIL_t$, where the ADF test of the stationarity of the residuals from the regression does reject the null of no cointegration at the 10% level (the value of the ADF test is −3.92). This result provides some evidence of a stable long-run relationship between the three series.

[8] The critical value for the Phillips–Perron (1988) unit root test (including an intercept term) at the 5% level of significance is −2.93 (for $T = 50$). Results for the unit root test were: *TOTAF* (−1.95 (level); −7.48 (first difference)); and *TOTAFUN* (−1.88; −7.12).

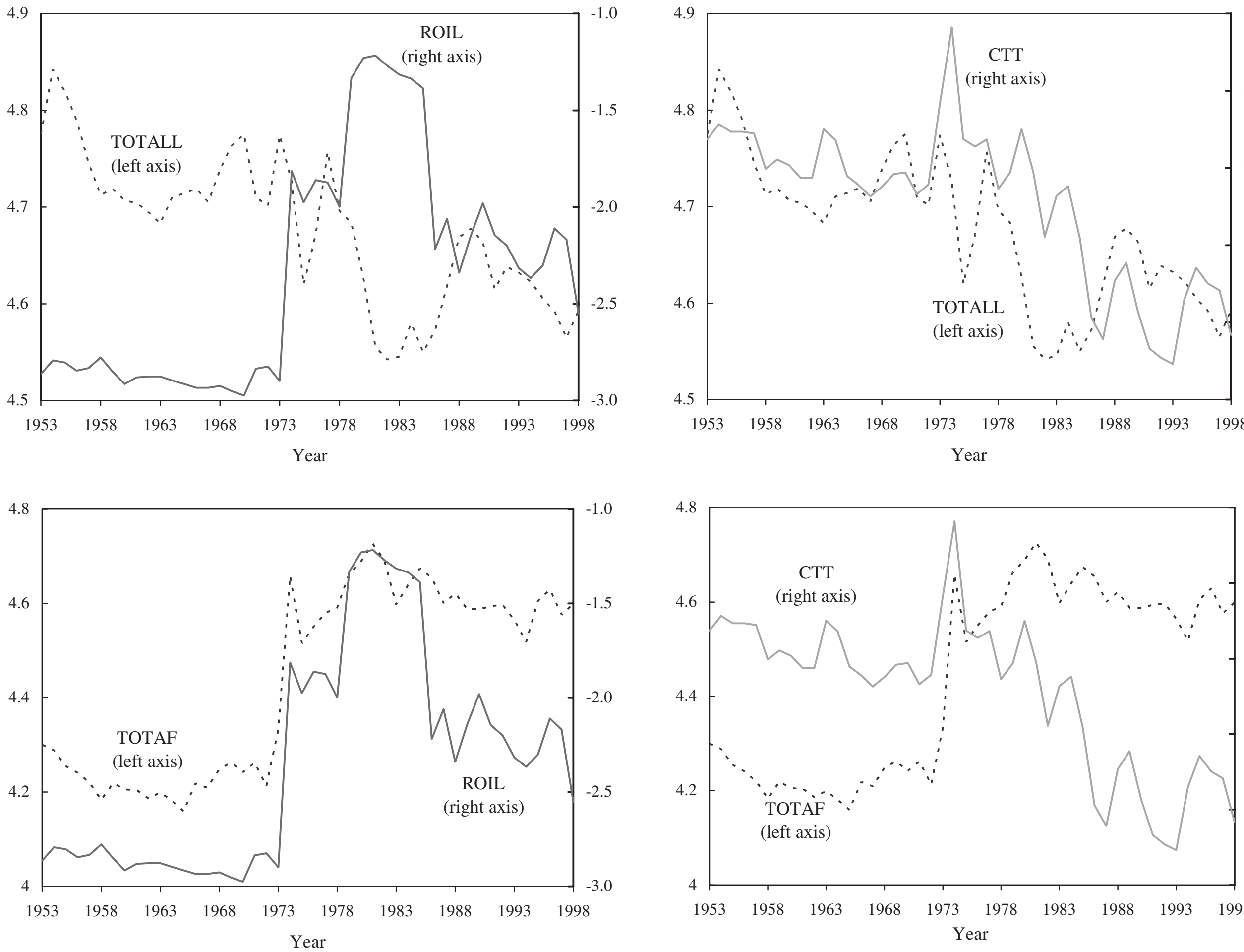


Fig. 1. Aggregate terms of trade, commodity terms of trade and real oil prices, 1953–1998 (logarithm)

Accordingly, the results from Equations 3 and 4 indicate that, extending the sample period into the 1990s, there is little evidence of a stable, long-run relationship between the commodity terms of trade and aggregate (or regional) terms of trade indices. However, even if such a relationship had been found, this would convey no information as to whether movements in the net barter terms of trade of *individual* African countries are associated with movements in the commodity terms of trade. This issue is examined in the following section.

III. The Relationship Between the Commodity Terms of Trade and National Measures of the Terms of Trade

Previous work on the terms of trade has largely concentrated on examining the properties of the commodity terms of trade (*CTT*), given that there was perceived to be a long-run relationship between aggregate measures of the terms of trade (such as the *CTT*) and the net barter terms of trade of individual countries. Unlike that work, in this study data on each country's actual terms of trade index are used, which has been constructed using information on import and export unit values. Accordingly, and unlike the *CTT* index, the country-specific data on the net barter terms of trade do not assume that all developing countries export primary commodities and import manufactures in the same proportion as underpins the construction of the aggregate *CTT* index.

Unit root and cointegration results

In examining whether there is a long-run relationship between the net barter terms of trade of individual countries and the commodity terms of trade, annual terms of trade data are used for the period 1960–1996 for 42 Sub-Saharan African countries, taken from the indices provided by the World Bank in its World

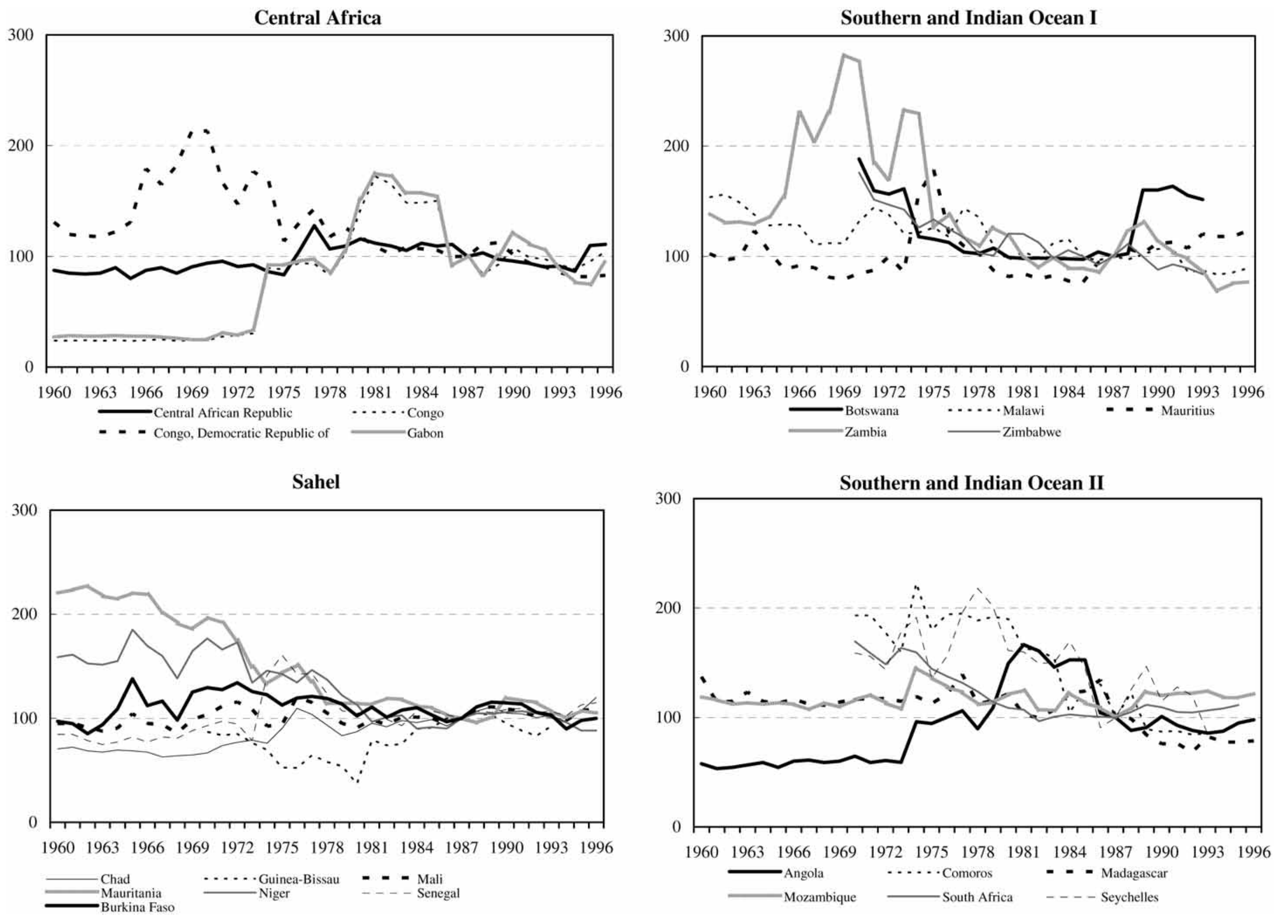


Fig. 2. Net barter terms of trade, sub-Saharan Africa, 1960–1996 (1987 = 100)

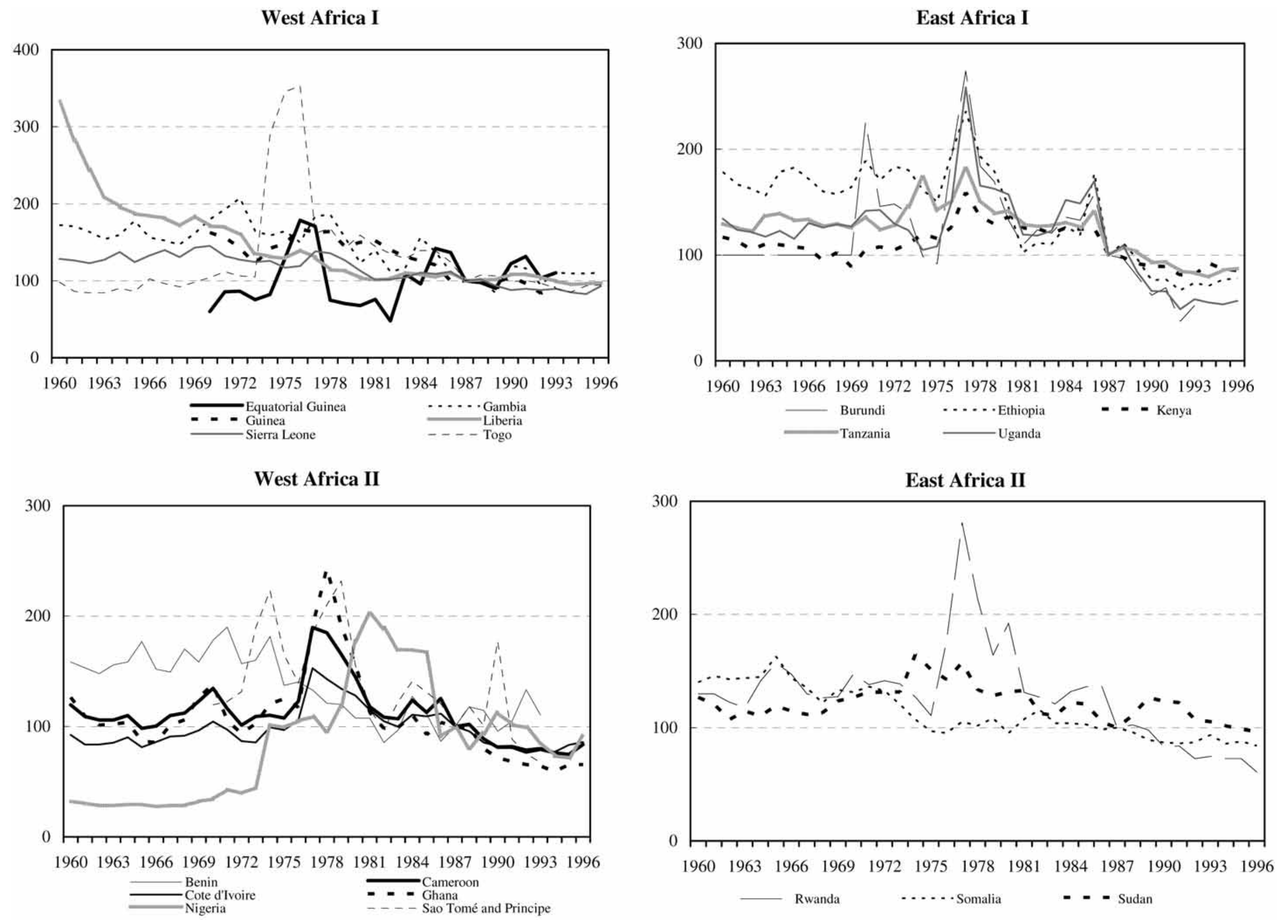
Development Indicators database (see Appendix B for details). The data are for each country's net barter terms of trade (*TOT*, an index number formed as the price index of exports of goods divided by the price index of imports of goods), with base year 1987. Data on *CTT* and *ROIL* are as described previously. All data are expressed in logarithmic form.[9]

Figure 2 displays the *TOT* indices (base 1987 = 100) for the countries in the sample, grouped by region. For many of the countries the figures show a gradual decline in their *TOT* since the mid-1970s to early 1980s. For a number of Southern and Indian Ocean countries as well as Somalia and Sudan, however, *TOT* have fluctuated around a flat trend for most of the period. Focusing on shocks to a number of commodities that are important for many Sub-Saharan African countries, the figure illustrates the beverage boom of the late 1970s–coffee for many East African countries and cocoa for a number West African and Sahelian countries. The *TOT* for oil-exporters reflects the 1974 and 1979 oil price increases, as well as the sharp fall in oil prices in the mid-1980s. The cotton price increases of the mid-1970s and mid-1990s register for only some of the cotton-exporting countries. The copper price boom of 1964–1974 is evident from Zambia's *TOT*, followed by a series of shocks to a low level thereafter. Similarly, the pattern for Togo is dominated by the huge phosphate boom of 1973–1978. The increase in iron ore prices from 1973–1981 is not reflected in Mauritania's *TOT*, although some of the smaller shocks to iron ore prices later in the period are evident.

Using the Phillips–Perron (1988) unit root test, each country's *TOT*, the *CTT* and *ROIL* were found to be nonstationary, and so the possibility of cointegration exists.[10] Accordingly, it is examined

[9] The countries in the sample are listed in Appendix A. The data for all countries run from 1960–1996, excepting: Botswana, Burundi, Comoros, Equatorial Guinea, Guinea, Guinea–Bissau, São Tomé and Principe, Seychelles, South Africa, and Zimbabwe (all 1970–1993).

[10] The results of the unit root tests are available from the authors.



Source: World Bank, World Development Indicators

Fig. 2. Continued

whether there is cointegration between these series for the period 1960–1996, using empirical techniques which allow for structural shifts in the relationship between time series.

Gregory and Hansen (1996a) demonstrate that the power of standard tests for cointegration falls when no allowance is made for structural shifts in the relationship between nonstationary series. Accordingly, the first step in the estimation procedure is to allow for the possibility that the cointegrated (long-run) relationship between the national net barter terms of trade (TOT), the commodity terms of trade (CTT) and the real price of oil ($ROIL$) has shifted at an unknown point in the sample. The possibility of a structural shift is allowed for because the period 1960–1996 has been marked by some significant policy framework changes in many countries, which lead to shifts in the composition of exports and imports, and by rapid fluctuations in the world prices of many primary commodities. Moreover, the timing of any such shift is likely to be unknown, because there is not necessarily a one-to-one correspondence between potential causes of a regime shift and its occurrence in the data. Use of the Gregory–Hansen (1996a) test for cointegration is therefore helpful in this instance, since it allows for the timing of any regime shift to be unknown *a priori*.

Gregory and Hansen (1996a) commence with the standard model for cointegration in the presence of no structural change, namely:

$$TOT_t = \beta_0 + \beta_1\, CTT_t + \beta_2\, ROIL + \varepsilon_t \quad t = 1, \ldots, T \tag{5}$$

where TOT, CTT and $ROIL$ are I(1) variables, and the residual ε_t is I(0). As an alternative to Equation 5, Gregory and Hansen propose a model where structural change occurs with a shift in the intercept term:

$$TOT_t = \beta_0 + \beta_1\, CTT_t + \beta_2\, ROIL + \beta_3\, \varphi_{t\pi} + \varepsilon_t \quad t = 1, \ldots, T \tag{6}$$

where β_0 denotes the cointegrating intercept coefficients before the shift, β_3 denotes the change in the intercept coefficients, and CTT, $ROIL$ and ε_t are as described above. Importantly, structural change is modelled using the following dummy variable:

$$\varphi_{t\pi} = \begin{cases} 0 & \text{if } t \leq [T\pi] \\ 1 & \text{if } t > [T\pi] \end{cases} \tag{7}$$

where the unknown parameter $\pi \in (0,1)$ denotes the timing of the change point in terms of a fraction of the sample and [] denotes integer part. Given that the timing of shifts ($T\pi$) in the relationship between macroeconomic series is unlikely to be known *a priori*, the Gregory–Hansen test for shifts in cointegrated models is useful as it does not require information on the timing of the such events.

A test of the null hypothesis of no cointegration is run, against the alternative hypothesis given by Equation 6. In doing so, the usual Phillips–Perron (1988) $Z(t)$ cointegration test statistic is computed for each possible shift $\pi \in \Pi$, using the residuals from the cointegrating regression of Equation 6. The π is chosen so that $Z(t)$ takes the smallest value (largest negative value) across all possible break points, where Π is any compact subset of $(0, 1)$ since the smallest $Z(t)$ gives the least favourable result for the null hypothesis (that is, the greater chance of rejecting the null of no cointegration). The smallest of these $Z(t)$ statistics will be denoted as $Z(t)^*$.[11]

While the Gregory–Hansen (1996a) test was designed to investigate if there is a cointegrating relation after allowing for a structural shift, the test also has power to detect cointegration when there is no structural shift. Consequently, a rejection of the null hypothesis of no cointegration may not be indicative of changes in the cointegrating vector, as the existence of a stable cointegrating relationship could also induce such a rejection. Accordingly, Gregory and Hansen (1996b) recommend that it is also necessary to test for cointegration using standard statistics that assume a stable cointegrating relation.

Haug (1996) finds that the ADF test has the least size distortions among a range of cointegration tests, and that the residual-based $Z(\alpha)$ cointegration test of Phillips and Ouliaris (1990) is one of the more powerful tests. Accordingly, in examining cointegration using standard statistics that assume a stable cointegrating relation, the data is given the best chance of finding a long-run relationship between the three variables, by following the rule that there is evidence of cointegration if either the ADF test or the $Z(\alpha)$ test reject the null of no cointegration at the 5% level of significance.

The Phillips–Ouliaris (1990) and ADF cointegration statistics test the null hypothesis of no cointegration against the alternative hypothesis of a stable cointegration relationship. The null of the Gregory–Hansen (1996a) model is also no cointegration, while the alternative hypothesis is cointegration with a one-time structural shift of unknown timing in the cointegrating relationship (change in cointegrating intercept coefficients). Note that if the conventional cointegration test (such as the Phillips–Ouliaris $Z(\alpha)$ and ADF tests) does not reject the null of no cointegration but the Gregory–Hansen $Z(t)^*$ test does, then there is evidence of a structural shift in the cointegrating relationship (Gregory and Hansen, 1996a).

The results of the Gregory–Hansen (1996a) cointegration test are set out in Appendix A. For four countries (Burkina Faso, Côte d'Ivoire, Uganda and Zimbabwe), the Gregory–Hansen statistics are consistent with a long–run cointegrating relationship between *TOT*, *CTT* and *ROIL* (allowing for a structural shift), as conventional cointegration tests cannot reject the null of no cointegration but the Gregory–Hansen test does. It is also found that for all but five of the 42 countries, the Phillips–Ouliaris $Z(\alpha)$ statistic is too small to reject the null of no cointegration. Similarly, the ADF test rejects the null of no cointegration for only eight of the 42 countries, and so provides little evidence of cointegration between each country's *TOT*, the *CTT* and the *ROIL* (see Appendix A).[12]

Importantly, if *both* conventional cointegration tests and the Gregory–Hansen test reject the null hypothesis of no cointegration (as occurs for Mali), then while it is clear that there is strong evidence in favour of a long-run relationship, it is unclear whether a structural shift has occurred because (as noted above) the Gregory–Hansen test is powerful against conventional cointegration. In this case, further investigation is necessary to enable a distinction to be drawn between cointegration with stable parameters and cointegration with a structural shift, as the null hypothesis of no cointegration is rejected in comparison with either alternative hypothesis. Gregory and Hansen (1996b) suggest using

[11] Following Gregory and Hansen (1996a), Π is here taken to be the compact subset $\Pi = [0.15T, 0.85T]$.
[12] In an extension of the work of Section II, it was also re-examined whether there was a long-run relationship between the aggregate terms of trade indices (*TOTALL*, *TOTAF* and *TOTAFUN*), the *CTT* and *ROIL* over the 1960–1996 sample period. As reported at the bottom of Appendix A, only for *TOTAFUN* was the null hypothesis of no cointegration able to be rejected in favour of cointegration with stable parameters.

Hansen's (1992) parameter instability tests (which are based on the residuals of a FM least squares regression), where the null hypothesis is cointegration with stable parameters, to determine whether there has been a shift in the cointegration relationship. For all three Hansen (1992) tests, the null hypothesis is that the cointegrating parameters are constant, while the alternative hypothesis is no cointegration due to a change in the parameters at some unknown point in the sample.[13] In particular, under the alternative hypothesis of parameter instability, the SupF test is focused on any abrupt shift in the cointegrating vector; the MeanF and Lc tests detect any gradual changes in the regression coefficients.[14] Using the Hansen (1992) tests evidence is found of unstable parameters for the case of Mali.[15] Accordingly, for the five countries displaying evidence of a structural shift (Burkina Faso, Côte d'Ivoire, Mali, Uganda and Zimbabwe), the cointegrating regression was augmented with an appropriate dummy variable to take account of the shift in the intercept.

Cointegration and structural change

For those five countries where the null hypothesis of no cointegration could be rejected using the $Z(t)^*$ test, the cointegrating relationship between each country's *TOT*, the *CTT* and *ROIL* (as set out in equation (6) above) was estimated using Phillips and Hansen's (1990) Fully Modified (FM) method. FM estimation is a semiparametric procedure that modifies least squares regression to account for potential endogeneity of the regressors and serial correlation caused by cointegrating relationships.[16] The FM method yields an asymptotically correct variance–covariance estimator when estimating cointegrating vectors in the presence of serial correlation and endogeneity–the results are set out in the lower panel of Table 3.[17,18]

For those eight countries where the null hypothesis of no cointegration could be rejected using the Phillips–Ouliaris $Z(\alpha)$ or ADF tests, the cointegrating relationship between each country's *TOT*, the *CTT* and *ROIL* (as set out in Equation 5 above) was again estimated using Phillips and Hansen's (1990) FM method. The results are set out in the upper panel of Table 3.

One potential problem with time series regression models is that the estimated parameters may be unstable. In particular, the many exogenous shocks and policy changes that significantly affect small economies may cause the parameter estimates in the cointegrating relationship between each country's *TOT*, the *CTT* and *ROIL* to change over time. Accordingly, in interpreting the relationship between these variables it is important that the long-run parameter estimates be structurally stable. To examine the hypothesis of parameter instability in the context of FM estimation of a cointegrated regression model, the tests suggested by Hansen (1992) are again used. The results indicate that there is instability in the relationship between each country's *TOT*, the *CTT* and *ROIL* (at the 5% level of significance) for two of the eight countries found to have a cointegrating relationship, as for these countries (Benin and Guinea) the null of parameter stability is rejected by at least one of the tests (see columns (3)–(5) of the upper panel of Table 3). Accordingly, evidence of a stable cointegrating relation between the three series is found for six countries: Angola, Central African Republic, Gambia, Malawi, Niger, and Seychelles.[19]

[13] The asymptotic critical values (with a constant term) at the 5% level of significance for the Lc, MeanF and SupF tests are 0.78, 7.69 and 17.3, respectively – computed values for the stability tests greater than these critical values indicate acceptance of the alternative hypothesis (see Hansen, 1992).

[14] The MeanF and SupF tests require truncation of the sample of size T – Hansen (1992) is followed and the subset [$0.15T$, $0.85T$] used.

[15] For Mali, the computed values of the Lc, MeanF and SupF parameter stability tests (0.37, 7.06 and 25.76) reveal that at least one of them exceeds their 5% critical values, indicating rejection of the null hypothesis of cointegration with stable parameters.

[16] For the Phillips–Hanson FM estimation the Bartlett kernel, Andrews' (1991) automatic bandwidth selector and the pre-whitened kernel estimator of Andrews and Monahan (1992) were employed.

[17] Ordinary least squares estimation could be used to yield consistent estimates of the cointegrating parameters. However, least squares estimation is inefficient and yields non-standard distributions of the estimators, making standard inference tests problematic in the least squares framework, while these difficulties are overcome in the FM method (Phillips and Hansen, 1990). Importantly, FM-based estimates are robust to any potential endogeneity of the commodity terms of trade.

[18] The cointegrating regressions include a level shift dummy variable (φ_t, parameter value not reported) where the Gregory–Hansen test indicated Equation 6 was the appropriate cointegrating regression. See Appendix A for the dates of the structural shift.

[19] According to Hansen (1992), the stability tests can also be used as tests of the null of cointegration against the alternative of no cointegration, and so the results for these six countries corroborate earlier finding of cointegration among each country's *TOT*, the *CTT* and *ROIL*.

Table 3. Cointegration and stability tests, commodity and national terms of trade, Sub-Saharan Africa, 1960–1996

Country	Cointegrating parameters		Hansen tests		
Country (1)	*CTT* (2a)	*ROIL* (2b)	Lc (3)	MeanF (4)	SupF (5)
I. Countries rejecting the null hypothesis of no cointegration in favor of cointegration					
Angola	−0.144 (0.055)	0.579 (0.019)	0.32	4.64	10.90
Benin	0.710 (0.114)	−0.254 (0.039)	0.18	8.05*	22.37*
Central African Republic	0.001 (0.062)	0.151 (0.021)	0.34	4.74	13.52
Gambia	0.905 (0.118)	−0.129 (0.040)	0.16	2.21	4.46
Guinea	0.835 (0.093)	0.028 (0.042)	0.74	8.81*	13.62
Malawi	0.631 (0.105)	−0.114 (0.036)	0.17	2.33	4.69
Niger	0.740 (0.095)	−0.278 (0.032)	0.16	3.25	6.21
Seychelles	0.763 (0.129)	0.060 (0.058)	0.29	3.87	7.40
II. Countries rejecting the null hypothesis of no cointegration in favor of cointegration with a structural shift					
Burkina Faso	0.261 (0.095)	−0.080 (0.035)			
Côte d'Ivoire	0.491 (0.116)	0.143 (0.043)			
Mali	−0.007 (0.061)	0.084 (0.027)			
Uganda	0.404 (0.232)	0.075 (0.057)			
Zimbabwe	0.509 (0.069)	−0.027 (0.035)			

Notes: The data (described in Appendix B) for all countries are annual and are expressed in logarithmic form. The estimated cointegrating parameters are from the Fully Modified (FM) cointegrating regression (Phillips and Hansen, 1990): $TOT = \beta_0 + \beta_1\ CTT + \beta_2\ ROIL + \varepsilon$, where *TOT* is the country's terms of trade; *CTT* the commodity terms of trade; and *ROIL* the real price of oil, are reported in columns 2a (for *CTT*) and 2b (for *ROIL*); the asymptotically-correct standard error of this estimate is in parentheses. All cointegrating regressions have been run using the Bartlett kernel, Andrews (1991) automatic bandwidth selector and the pre-whitened kernel estimator of Andrews and Monahan (1992). Columns (3–5): the 5 (10) per cent critical values for the Hansen (1992) tests of parameter stability (Lc, MeanF and SupF) are 0.78, 7.69 and 17.3 (0.63, 6.58 and 15.3), respectively. For columns 3–5, an asterisk (*) denotes statistical significance at the 5 per cent level. Gregory–Hansen (1996a) tests for the presence of a regime shift in the cointegrating vector (reported in Appendix A) reveal that the null hypothesis of no cointegration can be rejected, indicating a significant level shift in the cointegrating relation, for: Burkina Faso and Côte d'Ivoire (both in 1967), Mali (in 1972), Uganda (in 1990) and Zimbabwe (in 1974). Accordingly, level shift dummy variables have been included in the estimation of the cointegrating regressions for these countries.

For the 11 countries that indicate that there is a long-run relationship between the *CTT*, each country's *TOT* and *ROIL* (the six countries exhibiting cointegration with stable parameters and the five countries exhibiting cointegration with a one-time structural shift), the value of the elasticity of each country's net barter terms of trade with respect to the commodity terms of trade is of particular interest. Estimates of this elasticity range from −0.144 (for Angola) to 0.905 (for Gambia), with most (the exceptions being Central African Republic and Uganda) being significantly different from zero. These elasticities are also considerably higher than previous point estimates obtained by Grilli and Yang (1988), Powell (1991), and Bleaney and Greenaway (1993) for aggregate non-oil developing countries (Table 3). For example, for Malawi the estimated elasticity implies that a 1% increase in the commodity terms of trade is associated with a 0.63 percentage point rise in its net barter terms of trade. As expected, the elasticity of each country's *TOT* with respect to the *ROIL* is positive (and significant) for several net-oil-exporting countries (particularly for Angola).

However, it is important to note that even for the minority of countries (11 of 42) where there is evidence of a stable relationship between the two terms of trade series, on average only a fraction of any given movement in the commodity terms of trade is associated with a corresponding movement in national terms of trade. Importantly, for almost three-quarters of the countries in our sample (31 of 42), it is found that there is no evidence of a stable, long-run relationship between each country's terms of trade and the commodity terms of trade, even after allowing for a structural shift of unknown timing in the cointegrating relationship.[20]

[20] This finding of a lack of a long-run relationship between movements in the commodity terms of trade and movements in the net barter terms of trade is consistent with the results of Lutz (1999), who found that in comparison with Asian and Latin American countries, African countries exhibited the weakest association between the commodity terms of trade and each country's net barter terms of trade.

As argued by Cashin and Pattillo (2000) and Lutz (1999), it is perhaps unsurprising that movement in the *CTT* reveals little about movements in the *TOT* of individual countries, given that the former is by construction an aggregated index. In particular, the primary commodities included in the index reflect their share in world commodity exports, and the manufacturing goods used to deflate the series of primary commodity prices reflects the export unit values of a group of industrial countries. For industrial countries, which import a wide range of primary commodities, aggregate commodity price indices such as the *CTT* may be useful an indicator of general movements in commodity prices. However, many developing countries (and especially those in Sub-Saharan Africa) often export only a limited range of commodities, and the results indicate it is unlikely that the *CTT*'s aggregate series of exports and imports bear much relation to the export and import baskets of individual commodity-exporting countries.

IV. Conclusion

This paper has examined the validity of the purported link between the commodity terms of trade (the price of primary commodities relative to manufactured goods) and the net barter terms of trade (the price of exports relative to imports) of individual developing countries, using data on 42 commodity-dependent Sub-Saharan African countries. Previous empirical work on the terms of trade has largely examined trends and volatility in the commodity terms of trade, while analysis of the properties of countries' actual net barter terms of trade has been neglected. The main reason for this neglect is most likely the lack of long time series data on national terms of trade. In addition, following the Prebisch–Singer hypothesis, it has been assumed that for commodity-exporting developing countries, the trend in each country's net barter terms of trade generally reflects the trend in the commodity terms of trade.

It is found that the above assumption does not hold –there is little evidence of a stable, long-run relationship between the commodity terms of trade and national net barter terms of trade for the majority of the countries of Sub-Saharan Africa. Accordingly, proxying for movements in any given country's net barter terms of trade by using the commodity terms of trade (an aggregate index of relative commodity prices) is inappropriate, and is likely to engender misleading policy conclusions. In particular, the results indicate that any finding of a secularly declining commodity terms of trade does not typically imply that there has been a corresponding trend decline in the net barter terms of trade of individual African countries.

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Appendix A

Table A1. Cointegration tests: commodity and national terms of trade, Sub-Saharan Africa, 1960–1996

Country (1)	*ADF* (2)	$Z(\alpha)$ (3)	$Z(t)$* (4)
Angola	−4.25*	−25.37*	−3.53
Benin	−4.85*	−26.18*	−3.37
Botswana	−2.75	−10.93	−3.16
Burkina Faso	−3.01	−15.64	−6.34* [1967]
Burundi	−3.07	−10.22	−3.65
Cameroon	−3.62	−15.72	−5.43
Central African Republic	−4.40*	−24.95*	−4.07
Chad	−2.82	−11.83	−4.67
Comoros	−4.08	−15.54	−2.98
Congo, Democratic Republic	−2.30	−7.92	−2.72
Congo	−2.76	−14.73	−3.72
Côte d'Ivoire	−3.76	−16.50	−4.95* [1967]
Equatorial Guinea	−2.77	−13.55	−4.01
Ethiopia	−3.31	−15.74	−4.54
Gabon	−2.89	−18.90	−3.60
Gambia	−4.89*	−26.80*	−4.28
Ghana	−3.61	−15.98	−4.63
Guinea	−4.57*	−12.55	−4.37
Guinea Bissau	−2.25	−15.51	−2.79
Kenya	−3.53	−19.21	−4.38
Liberia	−3.93	−13.95	−4.82
Madagascar	−2.70	−13.30	−4.56
Malawi	−4.72*	−17.63	−4.66
Mali	−3.99*	−17.99	−6.51* [1972]
Mauritania	−3.79	−16.03	−4.85
Mauritius	−2.58	−11.31	−2.91
Mozambique	−3.34	−19.74	−3.79
Niger	−3.77	−22.91*	−4.68
Nigeria	−2.50	−17.61	−4.27
Rwanda	−3.78	−17.01	−4.26
São Tomé and Principe	−3.20	−16.45	−4.32
Senegal	−2.38	−9.95	−3.52
Seychelles	−4.16*	−15.34	−3.40
Sierra Leone	−3.87	−16.09	−4.29
Somalia	−2.19	−13.02	−3.61
South Africa	−3.57	−18.84	−4.04
Sudan	−2.33	−9.81	−3.15
Tanzania	−2.98	−16.34	−4.78
Togo	−3.01	−14.07	−4.22
Uganda	−2.74	−11.31	−4.93* [1990]
Zambia	−2.46	−10.20	−3.02
Zimbabwe	−3.89	−14.69	−5.12* [1974]
African Countries (*TOTAF*)	−3.09	−20.11	−4.08
African Countries (*TOTAFUN*)	−3.92	−22.97*	−4.07
Non-Oil Developing Countries (*TOTALL*)	−3.08	−16.29	−2.80

Notes: The data (described in Appendix B) for all countries is annual, and are expressed in logarithmic form. The estimated regression from which the residuals are derived is: $TOT = \beta_0 + \beta_1 CTT + \beta_2 ROIL + \varepsilon$, where TOT is the country's terms of trade; CTT the commodity terms of trade; and $ROIL$ the real price of oil. Column (2): the 5 (10) per cent critical values for Said and Dickey's (1984) ADF test (with a constant) of stationarity of ε (for $T=37$) are −3.98 (−3.62) and (for $T=24$) are −4.11 (−3.72), based on MacKinnon (1991). Column (3): the 5 (10) per cent critical value (for $T=50$) for the Phillips–Ouliaris (1990) residual-based $Z(\alpha)$ test (with a constant) is −22.68 (−19.60), taken from Haug (1992). Column (4): the 5 (10) per cent critical value for the Gregory–Hansen (1996) $Z(t)$* test for the presence of a level shift in the cointegrating vector is −4.92 (−4.69); the date in which the regime change is estimated to occur is given in square brackets. For columns 2–4, an asterisk (*) denotes statistical significance at the 5% level, indicating that the null hypothesis of no cointegration can be rejected.

Appendix B: Data Description and Derivation

The definitions of all variables used in the paper are given below. All variables are expressed in natural logarithms, and all data are annual in frequency over the relevant sample period.

TOTALL: Index of the terms of trade of non-oil developing countries, base 1995 = 100, taken from the International Monetary Fund's (IMF), *International Financial Statistics* (*IFS*) (*Supplement on Price Statistics No.2,* 1981 *and No.12,* 1986) for 1953–1985; extended using the IMF's *IFS* for 1986–1998.

TOTAF: Index of the terms of trade of Africa, base 1995 = 100, taken from the same sources as *TOTALL* for 1953–1998.

TOTAFUN: Index of the net barter terms of trade of Africa, base 1987 = 100, taken from the World Bank's *World Development Indicators* database and United Nations (1999, and earlier issues) for 1960–1996. The World Bank obtains the terms of trade data from the United Nations Conference on Trade and Development's (UNCTAD) publication *Handbook of International Trade and Development Statistics,* and publishes it in the *World Development Indicators* volume (previously *World Tables*). Thus, we obtained the data from a file sent by the World Bank Data group and from various issues of these published sources.

TOT: Index of the net barter terms of trade for each of the 42 countries listed in Appendix A, base 1987 = 100, taken from the same sources as *TOTAFUN* for 1960–1996.

CTT: Index of the commodity terms of trade: calculated as *GYCPI* divided by *MUV*, for 1953–1998.

GYCPI: Price index of non-fuel commodities, base 1977–1979 = 100, taken from Grilli–Yang (1988) for 1953–1986; extended using Grilli–Yang weights for individual commodities and the IMF's *IFS* commodities data for 1987–1998.

ROIL: The real price of oil: calculated as the index of the nominal price of oil (US$ per barrel, taken from IMF's *IFS*), deflated by MUV; for 1953–1998.

MUV: The United Nation's index of unit export values of manufactures of industrial countries, base 1977–1979 = 100, taken from Grilli–Yang (1988) for 1953–1986; extended using the IMF's *IFS* series on manufactured unit export values for 1987–1998.