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Price Incentives, Non-Price Factors, and Agricultural Production in Sub-Saharan Africa: A Cointegration Analysis

by
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Abstract:

This paper deals with the question of how responsive farmers in Sub-Saharan Africa (SSA) are to changes in incentives. Employing Johansen's multivariate cointegration approach, it investigates for ten selected SSA countries the long-run effect of pricing policies, macroeconomic distortions, and certain non-price factors on agricultural production. It turns out that – in those cases where cointegration relationships are found –estimated supply elasticities tend to lie between 0.20 and 0.50. Among the non-price factors, drought episodes have significantly impaired agricultural growth in six out of ten sample countries.

Keywords:	Agricultural Supply Response, Cointegration, Sub-Saharan Africa
JEL–classification:	C22, Q11

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CONTENTS

I. INTRODUCTION

Over the last decades, governments in Sub-Saharan Africa (SSA) have typically laid a substantial tax burden on agriculture, both directly via interventions in agricultural markets and indirectly via overvalued exchange rates and import substitution policies (Thiele 2002). Since the agricultural sector accounts for a large share of the region's employment and value added, its discrimination is likely to have entailed considerable welfare costs. The extent of these welfare costs can, however, only be guessed as there is no reliable evidence on the aggregate agricultural supply response (e.g. Schiff and Montenegro 1997) and hence no base for assessing how severely resources have been misallocated as a result of the distorted incentives.[1]

The objective of this paper is to provide a contribution towards filling the empirical gap. Employing the multivariate cointegration procedure developed by Johansen (1988), it investigates for ten selected SSA countries the long-run relationship between agricultural production and (direct and indirect) price incentives. Certain non-price factors such as the incidence of drought are included as control variables.

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[1] Only in the extreme case of perfectly inelastic supply the taxation of agriculture would have had no impact on resource allocation.

The remainder of the paper is structured as follows. Chapter II introduces the data and the methodology used in the econometric analysis, and explains the composition of the country sample. Chapter III discusses the empirical results, comprising tests for the time-series properties of each single variable and estimates of the equilibrium relationships between agricultural production and its potential determinants. The paper closes with some concluding remarks.

II. DATA AND METHODOLOGY

1. The Variables

Among the data required for the empirical analysis, those for aggregate agricultural supply are most readily available. The FAO yearly calculates a broad agricultural production index based on observations for more than 150 food crops and about 30 export crops. This index has now been compiled for the period 1961–2001 (FAOSTAT 2002), but the time series used here only comprise the years from 1965 to 1999 so as to be compatible with the data that are provided for other variables.

When it comes to estimating the price elasticity of supply, the real value added price is the most comprehensive price variable as it captures changes in producer prices, intermediate input costs, real exchange rates, and world market prices. It is defined as the ratio of the nominal value added price to the economy-wide price level. To arrive at an index of nominal value added prices, agricultural value added at current prices was divided by agricultural value added at constant prices. Both series are available from the World Bank Africa Database (World Bank 2001), as is the consumer price index (CPI) that is used to measure the general price level.

The real price can be decomposed into its component parts. Neglecting the costs of intermediate inputs, which in SSA only account for a small share of overall production costs and for which it is almost impossible to obtain reliable time series information (Thiele 2002), the decomposition of the real producer price is given as

$$(1) \qquad P^R = \frac{P^N}{CPI} = NPC \cdot RER \cdot p^*,$$

where P^R is the real producer price, P^N is the farmgate producer price, NPC is the nominal protection coefficient, RER is the real exchange rate, and p^* is the real world market price, i.e. the nominal world market price P^* deflated by an index of aggregate world prices.[2]

The decomposition implies that it is not only possible to analyze the net effect of real price movements, but that one can also trace the separate impact of direct incentives, macroeconomic policies, and variations in border prices, which is more revealing from an economic policy viewpoint. To do the latter, the following variables were constructed. First, in calculating NPCs, the nominal value added price index (see above) was taken as the numerator because farm gate price indices could not be derived in many instances. The denominator, i.e. the nominal border price, was approximated by the weighted average of aggregate import and export unit values, which are both given in FAOSTAT (2002). Second, to measure the indirect impact of macroeconomic policies on agricultural supply, three different options were considered. The most obvious choice is to rely on the RER itself. Changes in the RER may, however, reflect

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[2] Under the standard assumption (see, for example, Edwards 1989) that the real exchange rate, i.e. the relative price of tradables to non-tradables, is approximated by $RER = \frac{e \cdot WPI^{US}}{CPI}$, where e is the official nominal exchange rate measured in domestic currency units per US$, and WPI^{US} is the US wholesale price index, equation (1) can be derived via the following simple manipulations:

$$P^R = \frac{P^N}{CPI} = \frac{P^N}{CPI} \cdot \frac{P^* \cdot e}{P^* \cdot e} \cdot \frac{WPI^{US}}{WPI^{US}} = \frac{P^N}{P^* \cdot e} \cdot \frac{WPI^{US} \cdot e}{CPI} \cdot \frac{P^*}{WPI^{US}}$$
$$= NPC \cdot RER \cdot p^*$$

either policy changes or shifts in exchange rate equilibria which, for example, can come about through technical progress (Edwards 1989). To isolate the effect of macroeconomic policies, two measures of exchange rate misalignment were chosen as additional indicators: the black market premium (BMP) and a model-based indicator of misalignment. The BMP is simply defined as the ratio of the parallel market to the official exchange rate. It is usually high in the presence of an overvalued exchange rate. The model-based indicator follows Edwards' (1989) distinction between the equilibrium exchange rate that is only affected by real variables such as the terms of trade, and exchange rate disequilibria caused by inappropriate macro policies. The calculation of this indicator for a number of SSA countries is described in Thiele (2002). Finally, the nominal border price was deflated by the US wholesale price index (IMF 2001) to obtain the real border price.

Apart from the direct and indirect price incentives, there are various non-price factors that affect the agricultural supply response. These factors can roughly be subsumed under four different categories (Mamingi 1997): physical capital, human capital, technology, and agroclimatic conditions. The first three of them are main target areas of public investment. Government expenditures on physical infrastructure (roads, irrigation facilities), education (schooling, agricultural extension), and agricultural research are supposed to be associated with higher agricultural production. In addition, access to credit and secure land tenure should have a positive impact on output by providing essential preconditions for capital formation. Among the agroclimatic factors, soil quality and the intensity and regularity of rainfall are likely to be most decisive for the supply response.

In deriving appropriate indicators for the empirical analysis from this set of factors, two criteria have to be taken into account: data must exhibit some degree of variance, and they must be available over time. In the area of public investment, only irrigation meets these requirements, whereas indicators of road infrastructure are either highly persistent (e.g. the share of roads that are paved), or follow a negative trend when measured relative to population (e.g. road density).[3] As a consequence, the share of irrigated land in the total area devoted to annual and permanent crops was calculated as a proxy for investment in infrastructure, using data from FAOSTAT (2002). Time series for land tenure and access to credit, the two key determinants of private agricultural investment, could not be constructed because information turned out to be too sketchy. Among the different education indicators, adult literacy and the provision of extension services arguably are most closely associated with agricultural

[3] Indicators of road infrastructure are much better suited for cross-country studies of agricultural supply as they vary substantially between countries.

production as both help farmers to adopt technical advances such as new seed varieties. Neither of them can be used here. While adult literacy tends to rise monotonously over time and thus lacks variance, only very few countries in SSA report separate extension statistics on a yearly basis. The latter is also true for public expenditures on agricultural research. Instead of this variable, a simple time trend will serve as a crude proxy for technical progress in the empirical analysis.

Finally, given that soil quality appears to be rather stable over time, despite the degradation that can be observed in various places, rainfall was chosen to represent agroclimatic conditions.[4] It was approximated by a dummy variable to which a value of 1 was assigned if a significant shortage of rain unfavorably affected agricultural production in a particular year, and a 0 otherwise. Such a dummy variable, which mainly captures short-run fluctuations of output, is to be prefered over data showing annual rainfall because the latter fail to consider seasonal and regional characteristics. Delayed rain in a particular year, for example, might lead to a bad harvest even if the overall quantity of rainfall is normal. The information required to create the dummy variable was collected

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[4] Like road infrastructure, soil quality differs markedly between countries and would thus constitute a suitable variable for cross-country regressions.

from various national sources for the period 1965–79, and taken from the World Bank (2002) for the period 1980–99.

2. The Country Sample

The choice of countries for the empirical analysis was guided by three basic principles. First, and foremost, it was attempted to cover a wide range of geographical zones with varying natural conditions for growing crops. Second, countries that experienced regime shifts during the period under consideration were preferably chosen as only sustained policy changes are likely to bring about a significant long-run response of agricultural supply. Finally, some countries, notably those that went through longer phases of civil war, were excluded from the analysis because of severe data limitations. The resulting sample comprises ten countries from four different regions:

(i) Burkina Faso and Niger represent the Sahel Zone, a drought-prone region, where millet and sorghum are the main food crops, and cotton is the main cash crop. In both countries, the taxation of cotton and the degree of macroeconomic distortions has been reduced in the 1980s and 1990s (Thiele 2002).

(ii) Cameroon, Ghana and Nigeria represent West and Central Africa, where rice, maize, millet and sorghum are the main food crops, and cocoa and coffee are the main cash crops. Cameroon is an interesting case as the government has managed to improve direct incentives for farmers growing cash crops, but at the same time the macroeconomic situation has worsened continuously, even after the 50 percent devaluation of the CFA franc in 1994. Ghana is the SSA country that, beside Uganda, has been most successful in restoring macroeconomic equilibrium after a dramatic crisis. By contrast, in Nigeria macroeconomic disequilibria have proved persistent, with a slight change for the worse over the last two decades (ibid.).

(iii) Kenya and Tanzania represent East Africa, where maize is the main food crop, and coffee and tea are the main cash crops. While Kenya has over the whole sample period more or less held on to a strategy of taxing agriculture only moderately, Tanzania has traditionally laid a high tax burden on farmers, both directly and indirectly, but has recently turned from heavy to moderate indirect discrimination of agriculture (ibid.). Due to data shortages, Uganda, Rwanda and Burundi were not considered, although Uganda is one of SSA's rare examples of a genuine regime shift.

(iv) Malawi, Zambia and Zimbabwe represent Southern Africa, another drought-prone region, where maize is the main food crop, and cotton and tobacco are the main cash crops. All three countries have experienced marked changes in macroeconomic policies. While Malawi and Zimbabwe have achieved steady improvements since the early 1970s, except for the last few years in Zimbabwe, the pattern that emerges for Zambia is more volatile, with a high level of distortions over the whole time period that has peaked in the late 1980s. As for the direct price incentives, only Zambia exhibits a trend towards stronger discrimination (ibid.).

With such a diverse country sample, one objective of this paper is to reveal whether there are common patterns of supply response across SSA, or whether country or region specific features prevail.

3. The Estimation Method

Most of the few existing empirical studies on the aggregate agricultural supply response have applied the so-called Nerlove-method (Nerlove 1979). This method involves the estimation of a partial adjustment model of agricultural production for a particular country. The overwhelming majority of the regression analyses along these lines obtains low, or even zero, long-run price elasticities of agricultural supply. The widely-cited paper by Bond (1983) constitutes an illustrative example. She estimates a significantly positive long-term supply elasticity for only two out of nine SSA countries examined, and even for those two countries – Ghana and Kenya – the elasticity values are as low as 0.23 and 0.16, respectively. It has to be taken into account, however, that these estimates are likely to be downward-biased as the Nerlove-method specifies the dynamics of supply in a very restrictive way.

As a response to the limitations of the partial adjustment framework, two different directions can be taken.[5] One option is to estimate the supply elasticity for a cross-section of countries rather than for single countries over time. The other option, which will be pursued here, is to use more sophisticated time-series techniques. Basically, two such techniques are available: dynamic general equilibrium models and cointegration analyses. A distinctive advantage of dynamic general equilibrium models is that they explicitly consider factor movements and technology adoption, two important channels through which agricultural expansion may occur (e.g. Coeymans, Mundlak 1993). This advantage has to be weighed against one serious disadvantage, namely the high requirements in terms of data and modeling effort. Given the notorious data problems in much of SSA, the cointegration approach appears to be more promising than dynamic general equilibrium models.

Conducting a cointegration analysis is a straightforward way to overcome the restrictive dynamic specification of the Nerlove model. This method does not

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[5] For a critical discussion of the approaches available for the estimation of the long-run agricultural supply elasticity, see Thiele (2000).

impose any restrictions on the short- run behavior of variables. It only requires a stable long-run equilibrium relationship, which formally implies that there exists a linear combination of variables that is stationary even though each single variable may be non-stationary. To test for cointegration, two main approaches have been developed: one involves the estimation of a static model where all variables enter in levels (Engle and Granger 1987), the other the estimation of an error correction model (Johansen 1988). While the Johansen procedure is somewhat more difficult to apply, it is to be prefered over the Engle-Granger approach for two major reasons. First, in the multivariate case considered here, it avoids the identification problems one may encounter with the Engle-Granger approach if there is more than one co-integrating vector. Second, the Dickey-Fuller test employed to test for cointegration in Engle-Granger regressions too often rejects the existence of equilibrium relationships (Kremers et al. 1992).

The Johansen procedure is based on the maximum likelihood estimation of the error correction model

(2) $$\Delta X_t = \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-1} + \Pi X_{t-1} + \delta t + \theta D_t + u_t ,$$

where X denotes the vector of endogenous variables,[6] Γ_i the matrix of short-run coefficients, Π the matrix of long-run coefficients, δ the vector of coefficients of the linear deterministic time trend, θ the vector of coefficients of the drought dummy, and u_t the vector of independently normally distributed errors.

The matrix Π contains the co-integrating vectors and a set of so-called loading vectors which determine the weight of the co-integrating vectors in each single equation. By means of a normalization, the co-integrating vectors can be identified from the estimated Π matrix. To determine the number of co-integrating relationships r, the Johansen procedure provides two likelihood ratio tests: the trace and the maximum eigenvalue test. The trace statistic tests the null hypothesis of r co-integrating relations against the alternative of k co-integrating relations, where k is the number of endogenous variables, for $r = 0, 1, ..., k - 1$. It is computed as

$$(3) \qquad TR\,(r/k) = T\sum_{i=r+1}^{k} \ln\left(1-\lambda_i\right),$$

where λ_i is the i-th largest eigenvalue of the Π matrix.

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6 In the cointegration framework, all variables except deterministic trends and dummies are assumed to be endogenous.

The maximum eigenvalue statistic tests the null hypothesis of r co-integrating vectors against the alternative of $r+1$ co-integrating vectors. It is computed as

$$(4) \qquad MAX(r/r+1) = -T\ln\left(1-\lambda_{r+1}\right).$$

III. EMPIRICAL RESULTS

The first step in the empirical analysis is to identify the order of integration of each single time series. Only if the variables are integrated of the same order (usually of the order 1, denoted as I (1)), a linear combination of them may be stationary.

1. Tests for the Order of Integration

The order of integration equals the number of unit roots a series contains, or the number of differencing operations it takes to make the series stationary. To test for it, two widely used unit root tests were performed: the Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests, and the Phillips-Perron (PP) test. The DF and PP tests are based on estimating the univariate expression

$$(5) \qquad \Delta x_t = \gamma\, x_{t-1} + \varepsilon_t,$$

where ε_1 is a stationary random disturbance term. If indicated by the data, a constant and a deterministic trend may be included in equation (5). Both tests take the unit root as the null hypothesis, i.e. $H_0 : \lambda = 0,$ which is tested against the alternative $H_1 : \lambda < 0.$ This simple unit root test is valid only if the series follows a first-order autoregressive process. To control for higher-order serial correlation, the ADF and the PP tests use different methods. While the ADF test adds lagged differences on the right-hand side of equation (5), the PP test makes a correction to the t-statistic of the γ coefficient to account for the serial correlation in ε (Phillips and Perron 1988).

Unit root test results are given in Table 1. The reported statistics are confined to the DF (ADF) approach in levels, because in most instances the DF (ADF) and the PP test led to the same conclusions, and because integration of higher order, which requires a test in differences, was only detected in one single case (the real exchange rate in Zimbabwe is I(2)). For most series, the null hypothesis of a unit root could not be rejeced at the 5 percent level. Being I(1), these variables can be fed into a cointegration regression. In Niger, Zambia and Zimbabwe, however, agricultural production and some of its potential determinants seem to follow a stationary process. Hence, for these countries one has to perform a

Table 1 — Unit Root Tests[a]

Variable / Country	Production (Q)	Real price (P^R)	Nominal protection coefficient (NPC)	Border price (p*)	Real exchange rate (RER)	Black market premium (BMP)	Exchange rate misalignment (EMIS)	Irrigation (IR)
Burkina Faso	-1.70 (c,t)	-2.13 (c)	-3.22 (c)*[b]	-1.66	-2.06 (c,t)	/[c]	-1,10 (c,t)	-5.94 (c,t)*
Cameroon	-1.85 (c,t)	-1.72 (c)	-0.92	-1.00	-1.79 (c)	/[c]	1.58	-2.60 (c,t)
Ghana	-1.72	-2.38 (c)	-3.40*[b]	-1.78	-0.49	-1.29	-0.57	-4.40 (c)*
Kenya	-3.05 (c,t)	-2.65 (c)	-2.30 (c)	-1.52 (c,t)	-2.37 (c,t)	-2.96 (c)*[b]	-0.64	-2.26 (c,t)
Malawi	-1.56 (c,t)	-2.30 (c)	-2.02	-2.04	-2.99 (c,t)	-1.08	-2.69 (c,t)	-1.93 (c)
Niger	-3.02*	-2.31 (c,t)	-3.04 (c)*	-3.94 (c)*	-1.82 (c,t)	/[c]	-0.47	-1.23 (c,t)
Nigeria	-1.17 (c,t)	-2.74 (c)	-2.42 (c)	-1.48 (c,t)	0.20	-0.20	0.55 (c)	-1.38 (c,t)
Tanzania	-3.49 (c,t)	-3.41 (c,t)	-2.74 (c,t)	-1.60	-0.28	-0.65	-0.92 (c,t)	-1.74 (c)
Zambia	-3.55 (c,t)*	-2.86 (c)	-3.85 (c,t)*	-2.85*	-3.49 (c,t)	-3.47 (c)*[b]	-1.79 (c,t)	-2.57 (c,t)
Zimbabwe	-4.75 (c,t)*	-3.97 (c)*	-5.30 (c,t)*	-4.63 (c)*	2.62	-1.21 (c)	0.85	-2.06 (c)

[a] DF (ADF) tests in levels. Significant γ coefficients are indicated by a '*' and the inclusion of a constant or a trend is shown in brackets. Critical values at the 5 percent level are –1.95, -2.95 (with a constant), and –3.55 (with trend and constant). The lag length of the DF (ADF) test equation was chosen so as to minimize the Akaike and Schwarz information criteria. In the few cases where the two criteria suggested different lag structures and where this mattered for the order of integration, the decision was based on the result of the PP test. – [b] The PP test does not reject the null hypothesis of a unit root. – [c] Country belongs to the CFA zone, where the free convertibility of domestic currencies vis-à-vis the French franc implies that no parallel market for foreign exchange has developed.

Source: Own estimations based on the data described in Chapter II.

regression in levels instead of a cointegration analysis. Given the tendency of the unit root tests to be somewhat biased against rejecting the null hypothesis, the same might also be true for Tansania where test statistics for production and the real price are close to the critical value.

2. Tests for Cointegration

The cointegration analysis discussed in this section focusses on the long-run relationship between variables and leaves out the short-run dynamics of supply which are jointly estimated in the error-correction framework (see equation (3)). Results of the analysis are reported in Tables 2 and 3. Table 2 displays the trace and maximum eigenvalue tests. Niger, Zambia, and Zimbabwe are not included in this table because in these countries agricultural production does not contain a unit root. Table 3 lists the cointegrating vectors, covering all cases where at least one of the test statistics turned out to be significant at the 5 percent level. In addition, regressions in levels are presented for Niger, Zambia and Zimbabwe.

Table 2 — Cointegration Tests[a]

Country/relationship \ Test statistic	Trace test	Maximum eigenvalue test
Burkina Faso		
lnQ, $\ln P^R$, T, D	17.92 (25.32)	12.06 (18.96)
lnQ, lnp*, lnRER, T, D	41.25 (42.44)	24.80 (25.54)
Cameroon		
lnQ, $\ln P^R$, T	15.48 (25.32)	14.28 (18.96)
lnQ, lnNPC, lnp*, lnEMIS, T	64.39 (62.99)*	24.04 (31.46)
Ghana		
lnQ, $\ln P^R$, T	15.17 (25.32)	12.13 (18.96)
lnQ, lnNPC, lnBMP	31.47 (29.68)*	22.88 (20.97)*
lnQ, lnEMIS, T	26.56 (25.32)*	22.55 (18.96)*
Kenya		
lnQ, $\ln P^R$, T, D	24.67 (25.32)	18.61 (18.96)
lnQ, lnNPC, lnEMIS, T, D	38.68 (42.44)	23.31 (25.54)
Malawi		
lnQ, $\ln P^R$, T, D	26.90 (25.32)*	23.31 (18.96)*
lnQ, lnNPC, lnBMP, T, D	42.41 (42.44)	32.09 (25.54)*
lnQ, lnNPC, lnEMIS, T, D	46.21 (42.44)*	32.81 (25.54)*
Nigeria		
lnQ, $\ln P^R$, T	16.35 (25.32)	9.36 (18.96)
lnQ, lnNPC, lnp*, lnRER, T	58.57 (62.99)	28.61 (31.46)
Tanzania		
lnQ, $\ln P^R$, T	27.59 (25.32)*	16.92 (18.96)
lnQ, lnNPC, lnp*, lnRER, T	59.98 (62.99)	31.50 (31.46)*
lnQ, lnp*, lnBMP, T	40.51 (42.44)	28.40 (25.54)*
lnQ, lnNPC, lnp*, lnEMIS, T	73.59 (62.99)*	35.07 (25.54)*

[a]Critical values at the 5 percent level are given in brackets. A '*' indicates support for cointegration. Since in no case two or more co-integrating vectors were detected, only the statistics for r = 1 are reported.

Source: Own estimates based on the data described in Chapter II.

Table 3 — Cointegrating Vectors[a]

Country	Cointegrating vector
Cameroon	$\ln Q = 0.63 \ln NPC + 0.85 \ln p^* - 1.93 \ln EMIS + 0.03\ T$
Ghana	$\ln Q = 0.38 \ln NPC - 0.45 \ln BMP$
	$\ln Q = -0.25 \ln EMIS + 0.005\ T$
Malawi	$\ln Q = 0.55 \ln P^R + 0.01\ T - 0.09\ D$
	$\ln Q = 0.30 \ln NPC - 0.21 \ln BMP + 0.01\ T - 0.09\ D$
	$\ln Q = 0.31 \ln NPC + 0.14 \ln EMIS + 0.01\ T - 0.08\ D$
Niger[b]	$\ln Q = 0.23 \ln NPC + 0.33 \ln p^* - 0.36 \ln EMIS - 0.11\ D$
Tanzania	$\ln Q = 0.72 \ln P^R + 0.02\ T$
	$\ln Q = 0.23 \ln NPC + 0.48 \ln p^* + 0.38 \ln RER + 0.01\ T$
	$\ln Q = 0.38 \ln p^* - 0.13 \ln BMP + 0.01\ T$
	$\ln Q = -0.16 \ln NPC + 0.27 \ln p^* - 0.05 \ln EMIS + 0.01\ T$
Zambia[b]	$\ln Q = 0.19 \ln p^* + 0.02\ T - 0.09\ D$
Zimbabwe[b]	$\ln Q = 0.30 \ln P^R + 0.004\ T - 0.15\ D$
	$\ln Q = 0.20 \ln NPC + 0.18 \ln p^* + 0.01\ T - 0.16\ D$

[a]All co-integrating vectors were normalized by setting the coefficient of Q equal to –1. – [b]Since agricultural production and some of its potential determinants turned out to follow a stationary process, a model in levels was estimated.

Source: Own calculations based on the data described in Chapter II.

Since production, prices, and macroeconomic variables are given in logs, the respective parameters can be interpreted as elasticities. With two exceptions (lnEMIS in Malawi and lnNPC in Tanzania), the estimated coefficients have the expected signs. As regards the net impact of real prices on output, supply elasticities for Malawi (0.55) and Tanzania (0.72) clearly exceed those commonly obtained by means of the Nerlove method, while for Zimbabwe (0.30) this is not the case.[7] The decomposition of the real price reveals that all components parts – the nominal protection coefficient, border prices, and macroeconomic indicators – affect agricultural output in various countries. In Ghana, for example, the continuous removal of macroeconomic distortions has contributed to higher agricultural growth. The supply elasticities with respect to the three component parts tend to lie between 0.20 and 0.50, with Cameroon being a positive outlier. It has to be noted, however, that all elasticity estimates reported in Table 3 may be downward-biased because of the inherent problem of time-series analyses that in case of regime shifts parameters are likely to be affected (Lucas-critique). Errors in variables, which cannot be excluded in SSA, would compound the downward bias.

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[7] For Tanzania, a somewhat different result was obtained by McKay et al. (1999). They detect a cointegration relationship with a relatively high supply elasticity for food crops, but not for aggregate agriculture. This difference may partly be due to the fact that their sample period does not cover the 1990s. Danielson (2002), by contrast, argues that structural features of Tanzania's agricultural sector prevented the improved price incentives over the past 15 years from being translated into high agricultural growth.

Among the three non-price factors considered, irrigation proved insignificant across the board, probably due to two reasons: first, the share of irrigated land in the total area cultivated is very small (partly below 1 percent) in all sample countries so that moderate changes in that ratio are not likely to affect *aggregate* agricultural supply. Second, the irrigation variable is highly correlated with the time trend and therefore any existing separate effect on output may not show up statistically. The deterministic trend (T) was found to be significant for eight countries, and in some specifications also for Ghana. The only exception is Niger, where the agricultural production index was even slightly lower in 1999 than in 1965. Assuming that the trend can be regarded as a crude proxy for technical progress, its coefficients reveal that – apart from Cameroon – agricultural productivity growth has been moderate at best, and sometimes even below 1 percent per year. The dummy variable (D) that captures the incidence of drought has a negative impact on agricultural production in 6 out of 10 sample countries, including those from the Sahel Zone and Southern Africa, plus Kenya.

In a number of cases, no cointegrating vector was found. The most obvious explanation for this outcome is a lack of agricultural supply response. But there are also indications that the existence of cointegrating vectors depends on how strongly governments have reformed their policy regime over the period under consideration. Ghana, Malawi, Tanzania, and Zimbabwe achieved marked shifts in macroeconomic policies (Thiele 2002) and at the same time performed well in the regressions. In Kenya, by contrast, a remarkably stable policy framework corresponds with the non-existence of any long-run relationship. With moderate policy changes, the three CFA countries cover the middle ground, while in Zambia the lack of a sustained output response to policies may be explained by the volatility of the macroeconomic framework. Only Nigeria does not fit into this pattern as its deteriorating macroeconomic situation has not had a measurable effect on agricultural supply. Beside the absence of regime shifts, a further reason for not detecting cointegration may lie in the relatively low power of the cointegration tests (e.g. Rapach 2002).

IV. CONCLUDING REMARKS

Using Johansen's cointegration procedure, this paper has estimated the long-run relationship between agricultural production, direct and indirect price incentives, and non-price factors, for ten selected SSA countries over the period 1965–99. Two basic results of the analysis stand out. First, estimated supply elasticities tend to be well below unity, but they appear to be high enough to imply that the remaining discrimination against agriculture in SSA entails substantial welfare costs, thus indicating the need of further agricultural and macroeconomic reforms. Second, among the non-price factors, the coefficient of the time trend points towards low productivity growth in all but one sample countries, while a dummy variable shows that the agricultural growth of six countries has been impaired significantly by drought episodes. If one combines these results with the recent finding by Voortman et al. (2000) that SSA might experience some sort of Green Revolution if agricultural research was carefully tailored to local conditions, there seems to be a rationale for intensified international cooperation that aims, for example, at the development of more drought-resistant seed varieties.

Another important feature of the empirical analysis is that in a number of cases no long-run relationship could be detected. This result may, of course, reflect the absence of a supply response, but it may as well be due to two methodological problems encountered here, namely the relatively low power of the cointegration tests in small samples, and the rarity of marked regime shifts throughout SSA. In future research, the first problem may be resolved at least partly by applying a cointegration analysis with panel data as it has been demonstrated recently that the use of panel data increases the power of cointegration tests compared to pure time series (e.g. Rapach 2002). As for the persistence of variables over time, the only promising alternative is to carry out cross-country studies. Since not only price incentives but also institutional factors such as road infrastructure and educational attainment vary substantially between countries, both are well-suited to be integrated into cross-country regressions. It has to be kept in mind, however, that cross-country regressions face their own problems, most notably the biased estimators resulting from unobserved country characteristics.

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