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Aggregate Export and Food Crop Supply Response in Tanzania

by

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Abstract
Tanzania is among the many African countries that have engaged in agricultural liberalisation since the mid-1980s, in the hope that reforms which introduce price incentives and efficient marketing will encourage producers to respond. This paper assesses that claim by examining the supply response of agricultural output in Tanzania. Our estimates suggest that agricultural supply response is quite high so that the potential for agricultural sector response to liberalisation of agricultural prices and marketing may be quite significant. The long-run elasticity of food crop output to relative prices was almost unity; both food and aggregate short-run response was estimated at about 0.35. Liberalisation of agricultural markets, *where it increases the effective prices paid to farmers*, can be effective in promoting production, although complementary interventions, to improve infrastructure, marketing, access to inputs and credit, improved production technology etc, are probably necessary.

Outline

I INTRODUCTION

Like the majority of countries in sub-Saharan Africa (SSA), Tanzania has been following a structural adjustment programme since the mid-1980s. Significant progress was made during 1986-92: large and frequent devaluations; exchange rate liberalisation; reform and rationalisation of the tax system, especially tariffs; decontrol of agricultural prices and liberalisation of marketing (and the agricultural sector was the principal source of economic growth over 1986-92; see Morrissey, 1995). The rationale behind agricultural liberalisation is that the biases against agriculture inherent in protectionist policies, evident in Tanzania from the late 1960s, discourage production so that reforms which introduce price incentives and efficient marketing will encourage producers to respond (Bautista and Valdes, 1993). This paper assesses that claim by examining the supply response of agricultural output in Tanzania.

Section 2 reviews briefly the empirical evidence on aggregate agricultural supply response in a number of countries and presents the basic theoretical model, noting the inherent theoretical and empirical problems in traditional supply response models. Section 3 details the cointegration and error correction model (ECM) approaches, and argues that these offer a superior method for estimating aggregate supply response. Section 4 outlines the trends in performance and principal reforms in agriculture in Tanzania, and presents our data. Section 5 gives the econometric results from the aggregate supply response equations for all, food and export (cash) crops. Section 6 concludes by identifying the positive implications from the analysis but acknowledges the limitations of the data and approach.

II MODELLING AGGREGATE SUPPLY RESPONSE

In many developing countries, governments have been inclined, implicitly or explicitly, to tax the agricultural sector as part of a policy of industrialisation-led growth, justified by the belief that industry is the dynamic sector while the agricultural sector is static and unresponsive to incentives. If supply response is low, then taxing agriculture (ie. turning the internal terms of trade against agriculture) will generate resources for other sectors of the economy, without significantly affecting agricultural growth. But if, on the contrary, agricultural supply response is high, then taxing agriculture can retard agricultural growth, creating food and input supply bottlenecks which will eventually bring down the rate of growth of the entire economy (Chhibber, 1989), increase reliance on imports to meet food requirements and/or reduce agricultural exports (often the principal source of foreign exchange). In general, policies biased against agriculture have done more harm than good, reducing growth in the agricultural sector and consequently in the economy as a whole (Bautista and Valdes, 1993).

Table 1: Estimates of Aggregate Agricultural Supply Response

Studies	*Country*	*Supply Response*	
		Short-run	*Long-run*
Griliches (1959)[N]	US	0.28 - 0.30	1.20 - 1.32
Griliches (1960)[G]	US	0.10 - 0.20	0.15
Tweeten and Quance (1968)[G]	US	0.25	1.79
Rayner (1970)[G]	UK	0.34	0.42
Pandey *et al* (1982)[N,G]	Australia	0.30	0.6 - 1.00
Reca (1980)[N]	Argentina	0.21 - 0.35	0.42 - 0.78
Bapna (1980)[N]	India (Ajmer)	0.24	na
Krishna (1982)[N]	India	0.20 - 0.30	na
Chhibber (1989)[N]	India	0.20 - 0.30	0.40 - 0.50
Bond (1983)[N]:	Ghana	0.20	0.34
	Kenya	0.10	0.16
	Côte d’Ivoire	0.13	0.13
	Liberia	0.10	0.11
	Madagascar	0.10	0.14
	Senegal	0.54	0.54
	Tanzania	0.15	0.15
	Uganda	0.05	0.07
	Burkina Faso	0.22	0.24
	SSA (ave.)	0.18	0.21

Those studies indicated by N use the Nerlove model, those indicated by G use the Griliches approach.

Estimates of aggregate agricultural supply response for a number of countries are presented in Table 1. Time-series studies produce rather low estimates for the price elasticity of aggregate supply. Short-run estimates range between 0.1 and 0.3, consistent with the general belief that the aggregate response to price changes is small in the short-run, as aggregate production can increase only if more resources are devoted to agriculture or if technical change is adopted. However, long-run elasticities are only slightly higher, with most ranging between 0.1 and 0.5. Thus Tanzania, like many SSA countries, has very low estimated supply response in both the short- and long-run. It may be the case that response is indeed very low; even if this is so it may reflect other factors such as a lack of public investment in agriculture and the disincentives associated with anti-agricultural bias (in which case aggregate price data do not accurately represent the incentives facing farmers). Thus, structural adjustment programmes generally include agricultural liberalisation, and complementary sectoral reforms are necessary if agriculture is to reap any benefits from trade liberalisation (McKay *et al*, 1997). However, the lack of evidence for supply response may also be a feature of the model and estimation techniques used; we argue that this is the case, hence the potential for agricultural growth in response to liberalisation (of agricultural prices and marketing) may be greater than suggested by this evidence.

Modelling Supply Response

Most studies of supply response, aggregate or individual crop, are based on time-series data, and either use the Nerlove (1958) model devised for single commodities or the method developed by Griliches (1960) for aggregate supply response. Seven studies in Table 1 are based on the Nerlove model and four on the Griliches model. Both models are usually applied by estimating a single equation independently for each commodity, or group of commodities (if aggregate), without characterising linkages between them via a matrix of cross-price elasticities. Both are also partial equilibrium as they do not model the non-agricultural sector and thus implicitly assume that the interactions between the two sectors are insignificant (although sometimes a non-agricultural relative price may be included). In other respects however, the models are very different. The Nerlove model involves a one-stage procedure and directly regresses production on prices and other relevant variables. The Griliches model involves a two-step procedure, with the supply function being derived from the profit maximising marginal conditions of a Cobb-Douglas production function (Colman, 1983). In this section, we first analyse the theoretical framework of both models and then present the main statistical issues encountered when using them.

Nerlove's model describes the dynamics of agricultural supply by incorporating price expectations and/or adjustment costs. The general form of this supply function is:

$$X_t^* = a + bP_{xt}^e \tag{1}$$

where X_t^* is 'desired' or equilibrium output X at time t and P_{xt}^e is the expectation of price P_X in time t, formed at time t-1. We first assume that the dynamics of supply is driven by price expectations only so that $X_t^* = X_t$. In Nerlove's model price expectations are generally assumed to be adaptive:

$$P_{xt}^e - P_{x,t-1}^e = \delta(P_{x,t-1} - P_{x,t-1}^e)$$

or

$$P_{xt}^e = \delta P_{x,t-1} + (1-\delta) P_{x,t-1}^e \tag{2}$$

hence

$$P_{xt}^e = \delta \sum_{i=1}^{T} (1-\delta)^{i-1} P_{x,t-i}$$

Substituting (2) into (1) and rewriting gives:

$$X_t = a\delta + b\delta P_{x,t-1} + (1-\delta) X_{t-1} \tag{3}$$

where δ $(0 < \delta < 1)$ is the price expectation coefficient, b is the long-term elasticity of X with respect to P_x (long-run supply response), and $b\delta$ is the short-term elasticity (immediate response).

Adjustment costs can also cause lags in the response of output to price changes. This is especially relevant for supply response at an aggregate level, as moving factors across sectors is likely to be a long and costly process. Indeed, most of the cited studies on aggregate supply response ignore farmer's price expectations and concentrate on the partial adjustment hypothesis, whereby the actual change in output X_t from X_{t-1} is only some fraction of the change required to achieve the optimal level X_t^*. Following a similar logic to that for price expectations:

$$X_t = \phi X_t^* + (1-\phi) X_{t-1} \tag{4}$$

where ϕ $(0 < \phi < 1)$ is the partial adjustment coefficient. Assuming now that the expected price is the lagged price, $P_{x,t-1}$, substituting (4) into (1) and rearranging:

$$X_t = \phi a + \phi b P_{x,t-1} + (1-\phi) X_{t-1} \tag{5}$$

Comparing (3) and (5), one can see that adaptive price expectations and the partial adjustment hypothesis result in the same dynamic specification. As Nerlove acknowledged, this is one difficulty of the model: when both partial adjustment and adaptive expectations are present, it becomes impossible to distinguish between their respective coefficients, ϕ and δ, unless certain (arbitrary) restrictions on one or other are imposed. Moreover, the theoretical assumptions used in partial adjustment models are often considered inadequate (Nerlove, 1979), as modelling the dynamics of supply comes down to an *ad hoc* assumption that each period a fraction of the difference between the current position, X_t, and the long-run position, X_t^*, is eliminated.

The methodology developed by Griliches (1959, 1960) is specifically for estimating aggregate supply response and based on the aggregation of input demand elasticities. A constant returns Cobb-Douglas production function with a vector of n inputs can be differentiated with respect to producer price, and reformulated in terms of elasticities. Assuming profit maximisation, the elasticity of output with respect to input i can be estimated by the value share of input i in total revenues (or total costs assuming that zero economic profits prevail at the equilibrium). The aggregate supply elasticity is then obtained by aggregating the input demand price elasticities in concordance with their factor shares in total costs or revenues. Lags in the input demand functions are introduced to estimate short-run and long-run elasticities. If we assume that the output reaches equilibrium only if all inputs are in equilibrium, then the short-run aggregate supply response to prices is obtained from the weighted aggregation of the short-run input demand elasticities, and the long-run supply response from the weighted aggregation of the long-run input demand elasticities.

To date, the Griliches method has been used for developed countries only as it demands an extensive dataset on input and output prices and quantities, not usually available for developing countries. Furthermore, it assumes that the increased use of purchased inputs would increase aggregate supply, but does not account for essential inputs which are not purchased by farmers, such as family labour. The model also presupposes that resources supplied to the agricultural sector are perfectly elastic and always meet the farmer's demand at going prices, which is certainly not true for most developing countries. For example, when fertilisers and pesticides are imported, their supply is often constrained by shortages in foreign exchange. Finally, the method rests on the assumption that the underlying production function is of the Cobb-Douglas form, implying a unit elasticity of substitution between factors of production (one could argue that a Leontief specification, having a zero elasticity of substitution, is more appropriate for peasant farmers, at least in the short run).

Issues in Application

Peterson (1979) argues against using time-series data in estimating long-run elasticities, because only short-run year-to-year fluctuations are observed. The output response to annual fluctuations is likely to be small (even after full adjustment) because farmers will respond strongly to price changes only if they are perceived to be permanent. This implies that long-run elasticity estimates from time-series data are biased downward. Similarly, Schiff and Montenegro (1995) argue that time-series estimates are subject to the 'Lucas critique'. The long-run supply response to prices is properly evaluated when the transition to a new price regime is detected. Likewise, any permanent change in price policy affects the decision rules and thus any results drawn from past observations become obsolete (unless the estimation is based on a full behavioural model, which is unlikely to be practical in most cases). In other words, the values of the parameters estimated from time-series data are specific only to a given policy regime, so one cannot forecast the impact of a policy reform. While this is strictly true, we take the view that long-run elasticities derived from time series analysis are at least indicative and, while we do not forecast, we infer that if there is evidence of a long-run response, reforms that enhance the ability of (and increase the positive incentives for) farmers to respond may increase the long-run price elasticity further, so promoting increased growth in agricultural output.

In time series analysis of a single commodity, the main difficulty lies in selecting the correct data. Apart from identifying the correct output measure (planted area, marketed production, crop yields, etc.), researchers must determine which (relative) price variable should be used: the choice of deflator (consumer price index, input prices, other crop prices, etc.) is essential in formulating price responses (Askari and Cummings, 1977). For example, if farmers formulate their price expectations using relative prices and yet an absolute price is used in estimation, results may incorrectly present farmers as being not responsive to prices. Some of these problems are alleviated by addressing aggregate supply response.

At an aggregate level, a major problem relates to the method of indexation (Colman and Ozanne, 1988). Typically, studies of aggregate supply response use the Laspeyres index. However, there are serious reservations about this index and the Divisia index has often been proposed as a better alternative. The Laspeyres index holds all weights fixed at their base period levels, whereas the Divisia index uses weights from both the base and comparison periods. Thus, the comparative advantage of the Divisia index is that it reflects changes in the composition of agricultural output. Of course, in situations where the weights (output shares of the different commodities) have changed little over the period considered, there should not be any significant divergence in the value of the two indices and then the choice of the index is not so critical.

The root problem with the Laspeyres index is that it is equivalent to using a linear production function which restricts all factors of production to be perfect substitutes. On the other hand, the Divisia index is equivalent to using an homogeneous translog production function which does not imply that inputs are perfect substitutes. Both indexes are subject to the ‘output mix’ critique (Schiff and Montenegro, 1995): whichever econometric method used, one cannot forecast the response to a price reform on products which were not produced before the reform took place, such as when farmers in developing countries have been given incentives to diversify production towards higher value added export products, such as fruits and flowers. This however is an intrinsic, hence unavoidable, index number problem.

Last but not least, both the Nerlovian and Grilichian approaches use OLS to estimate the dynamic specification of their supply response. This means that the estimates of aggregate agricultural supply response are based on the assumption that the underlying data processes are stationary. However, most economic variables, including agricultural time-series, tend to be non-stationary, ie. their first two moments, mean and variance, are not constant. Using OLS with non-stationary variables may result in spurious regressions (Granger and Newbold, 1974). To ensure stationary variables, equations could be reformulated in terms of differences, but this loses important information conveyed by the levels, such as information on long-run elasticities. These issues are addressed in the next section.

III ERROR CORRECTION AND COINTEGRATION MODELS

The traditional approach used for estimating aggregate supply response has been criticised on both empirical and theoretical grounds. The Nerlove and Griliches techniques seem unable to give an adequate clear-cut distinction between short-run and long-run elasticities, while the use of OLS may produce spurious results. The *ad hoc* behavioural assumptions of the Nerlove empirical approach are by no means satisfactory and estimating supply response from the Griliches model is often not feasible given the data requirements. Using cross-country data might seem to offer an alternative, but the restrictions of assuming that coefficients and specifications are the same for each country renders cross-country analysis unreliable.

Time-series analysis remains the most widely used approach for estimating supply response. Modern time series techniques offer new promise. Cointegration analysis can be used with non-stationary data to avoid spurious regressions (Banerjee *et al*, 1993). When combined with error correction models, it offers a means of obtaining consistent yet distinct estimates of both long-run and short-run elasticities. Hallam and Zanouli

(1992), Townsend and Thirtle (1994), Abdulai and Rieder (1995), and Townsend (1996), have used cointegration analysis and ECMs to estimate supply response at a commodity level, on the basis that they are preferable to the traditional partial adjustment model.

The first step in cointegration analysis is to test the order of integration of the variables. A series is said to be integrated if it accumulates some past effects, so that following any perturbance the series will rarely return to any particular 'mean' value, hence is non-stationary. The order of integration is given by the number of times a series needs to be differenced so as to make it stationary. If series are integrated of the same order, a linear relationship between these variables can be estimated, and co-integration can be tested by examining the order of integration of this linear relationship. Formally, variables are said to be co-integrated (m,n) if they are integrated of the same order, n, and if a linear combination exists between them with an order of integration, m-n, which is strictly lower than that of either of the variables.

In practice, economists look for the existence of *stationary* cointegrated relationships, since only these can be used to describe long-run stable equilibrium states. Indeed, if there is a linear combination between the variables which is stationary, I(O), then any deviation from the regressed relationship is temporary. Although the variables may drift apart in the short-run, an equilibrium or stationary relationship is guaranteed to hold between them in the long-run. Typically, then, economists look for variables which are co-integrated (1,1). The concept of equilibrium in cointegration analysis is quite different from that in economics; no assumptions are made on market-clearing conditions nor of optimisation behaviour as, for example, in the Nerlovian model. In the theory of cointegration, an equilibrium is simply an observed *relationship* between variables which has, on average, been maintained over time. In fact, the definition of equilibrium in the cointegration literature is intended to be general and therefore incorporates any economic equilibrium which may be described using cointegration.

There is a strong case for believing that cointegration analysis could describe the dynamics of supply better than the Nerlovian methodology; indeed, the dynamics of supply is directly observed with cointegration, whereas in the Nerlove model it can only be asserted by recourse to theoretical assumptions which are not explicitly tested. Furthermore, the optimal output is not observable and only the reduced form of the Nerlovian model can be estimated (as equations (3) and (5) are identical). On the other hand, the theory of cointegration allows for testing the existence of equilibrium relationships between the output and other relevant variables such as prices. Hence, the optimal or 'normal' output position can be estimated from historical time-series data.

The dynamics of supply in the Nerlove model is driven by the partial adjustment hypothesis that farmers move closer to their equilibrium position by some fraction each year. When variables are cointegrated (1,1), there is a general and systematic tendency for the series to return to their equilibrium value; short-run discrepancies may be constantly occurring but they can not grow indefinitely. This means that the dynamics of adjustment is intrinsically embodied in the theory of cointegration, and in a more general way than encapsulated in the partial adjustment hypothesis. The Granger representation theorem states that if a set of variables are co-integrated (1,1), implying that the residual of the cointegrating regression is of order I(0), then there exists an error correction mechanism (ECM) describing that relationship. This theorem is a vital result as it implies that cointegration and ECMs can be used as a unified empirical and theoretical framework for the analysis of both short- and long-run behaviour. The ECM specification is based on the idea that adjustments are made so as to get closer to the long-run equilibrium relationship. Hence, the link between cointegrated series and ECMs is intuitive: error-correction behaviour induces cointegrated stationary relationships and *vice-versa*. The Granger theorem can be represented formally. If two variables x and y are I(1) and if there is a linear combination,

$$z_t = y_t - \beta x_t \tag{6}$$

which is I(0), then x and y are said to be cointegrated (1,1) and there exists an ECM describing the relationship. Assuming that x 'causes' y, then the ECM can be written:

$$\Delta y_t = \alpha \Delta x_t - \lambda \left(y_{t-1} - \beta x_{t-1} \right) + u_t \tag{7}$$

The estimated residuals from the cointegrating regression, z_t, represent the divergence from equilibrium or the 'equilibrium errors' that are going to influence changes in y in the following period. The coefficient β measures the long-run elasticity of x with respect to y and is estimated from (6); α measures the short-run effect on y of changes in x; λ measures the extent to which changes in y can be attributed to the ECM. If $z_t > 0$, that is, if y_t is above its equilibrium value, then y decreases in the following period ($\Delta y_{t+1} < 0$) and errors at time t are corrected by the proportion λ.

The advantage of using ECMs is twofold. First, spurious regression problems are by-passed, as Δx, Δy and z are all I(0). Second, ECMs offer a means to incorporate the levels of the variables x and y alongside their differences. This means that ECMs convey information on both short-run and long-run dynamics. Nickell (1985) demonstrates that the ECM specification represents forward-looking behaviour, such that the solution of a dynamic optimisation problem can be represented by an ECM. The ECM can thus be interpreted as describing farmers reacting to 'moving' targets and optimising their objective function under dynamic conditions.

IV AGRICULTURAL OUTPUT IN TANZANIA

Agriculture performed relatively well in Tanzania in the 1980s. In real terms the sector has grown at rates above those of the non-agricultural sectors, about five per cent per annum in the latter half of the 1980s, and has been the major source of GDP growth: in 1981-83, when real GDP fell by one per cent agricultural GDP rose by two per cent; in 1984-85 real GDP rose by 2.6 per cent but agricultural GDP increased six per cent; even over 1986-92 agricultural GDP increased by 4.7 per cent compared to 4.2 for GDP. Consequently, agriculture rose from some 45 per cent of GDP in 1980 to about 60 per cent by 1990 (World Bank, 1994: 3-4). Within agriculture the best performance was in foods (notably pulses and starches, but with good growth in cereals and sugar products) and non-traditional exports (which expanded rapidly to be worth $40m a year in the early 1990s, equivalent to the traditional exports, cotton or coffee). Traditional export crops performed poorly, reflecting the effect of unfavourable terms of trade on Tanzania: real export prices for coffee, cotton and tea in 1990 were less than half their value in 1984. Real agricultural exports in 1989-91 were worth 88 per cent of their 1984-86 value but only half the 1979-81 value, and traditional crops had fallen to less than half of the total (World Bank, 1994: 9).

Data on prices and quantities of crop production for Tanzania were obtained from World Bank (1992). Price and quantity indices were constructed using the Tornqvist formula, which is a discrete approximation of the Divisia index. A Tornqvist quantity index may be expressed in logarithms as:

$$\log Q_t - \log Q_{t-1} = \Sigma_{\mathrm{i}} \left[(s_{it} + s_{i,t-1}) (\log x_{it} - \log x_{i,t-1}) \right] / 2 \qquad (8)$$

$$\text{where } s_{it} = (p_{it} x_{it}) / (\Sigma_{\mathrm{i}}\, p_{it} x_{it})$$

where, in (8): Q_t is the quantity index at time t, x_{it} is the quantity of the ith crop at time t, p_{it} is the price of the ith crop at time t, and s_{it} is the value share of the ith crop at time t. An analogous equation is used for aggregating prices.

The period covered is 1964 - 1990, with indices benchmarked to 1964 ($Q_{1964} = P_{1964} = 1$). The quantity indices are expressed per capita (dividing by an index of total population) and the price indices are in real terms, using the Agricultural GDP deflator (data are reported in the Appendix). The commodities in the indices for food crops, denoted Q^f and P^f, are: maize, paddy, sorghum, millet, cassava and beans. Those in the export crop indices, denoted Q^e and P^e, are cashew, coffee, tea, pyrethrum, tobacco and cotton. The aggregate indices, Q^a and P^a, include all these food and export crops. The econometric analysis of the next section is based on these data, but general trends are outlined here.

Per capita production of food, export and all crops are shown in Figure 1, covering 1964-90. Since the early 1970s, per capita production of export crops has declined steadily but this has been more than compensated by the steady rise in food production, so that aggregate agricultural production per capita has risen slightly over the period. This switching supports a food versus cash crops argument, that as relative prices/profitability alter farmers substitute one crop for the other (World Bank, 1992). This is also supported by the trend in relative producer prices of export and food crops (Figure 2). As mentioned above, unfavourable terms of trade movements led to a gradual but almost continuous decline in export producer prices, whereas the *relative* official producer price of food crops rose fairly steadily in the 1970s and maintained its level for most of the 1980s, suggesting that farmers respond to relative prices.

Based on other evidence as well, the auguries for supply response in Tanzania are favourable. A recent study of supply response in food and traditional export crops found: the own price supply elasticity for foods on aggregate was 0.34; the cross-price supply elasticity of foods with respect to export crops (at official producer prices) was minus unity, confirming that crop substitution responses are significant; the own price supply elasticity for annual export crops on aggregate was 1.24; and the own price supply elasticity for perennial export crops on aggregate was in the range 0.55-0.83, consistent with the long time lags required to substitute into such crops so that annual production response will reflect harvesting changes rather than planting changes (World Bank, 1994: 117). The evidence for Tanzania is that producers will respond to prices (obviously subject to constraints, notably access to technology and credit as long-run determinants of output growth); of course this is not necessarily beneficial: if real official prices are falling output will decline. The trade liberalisation undertaken since 1986 did little to encourage agricultural exports: the principal reforms only made it easier to import and the benefits of devaluation were not transmitted to producers (McKay *et al*, 1997); real producer prices for export crops did not rise.

V ECONOMETRIC ESTIMATES FOR TANZANIA

The first step in the analysis is to identify the order of integration of the variables. The Dickey-Fuller approach can be applied to test the null hypothesis that a series contains a unit root (is non-stationary). This involves estimating (9) for each variable y_t and testing the null hypothesis H_o: $\rho = 1$ against the alternative H_1: $\rho < 1$.

$$\Delta y_t = \gamma + \tau t + (\rho - 1)y_{t-1} + \nu_t \tag{9}$$

If the variable does not follow an AR(1) process but is AR(n), then the Augmented Dickey Fuller (ADF) test should be used and in place of (9) one estimates:

$$\Delta y_t = \gamma + \tau t + (\rho - 1)y_{t-1} + \Sigma_i \lambda_i \Delta y_{t-i} + \nu_t \tag{10}$$

If H_o cannot be rejected, then y_t contains a unit root and hence is not stationary. If its first difference is then tested and found stationary, y_t is I(1). If not, y_t needs to be differenced further. Unit root test results are given in Table 2. These show that, for all variable except Q^e (export quantities), the hypothesis of a unit root cannot be rejected at the 5 per cent level. On the other hand, when the first differences are tested, the null hypothesis can be rejected. This suggests that all price and quantity indices except Q^e are I(1) and that we can test for cointegration between corresponding price and quantity variables for food and all crops; in the case of aggregate output of export crops, it appears that the decline in quantities is indistinguishable from a secular trend decline (see Table 3).

As producer prices are generally fixed and announced by the Tanzanian government before the harvesting season, aggregate producer price is likely to cause agricultural production and not *vice-versa*. Without running tests on causality, we can reasonably assume that P is the explanatory variable and Q the dependent variable. The supply function of the agricultural sector can be written as:

$$Q^* = \alpha_0 + \alpha_1 P^* + \varepsilon \tag{11}$$

In (11), Q^* is expected aggregate agricultural production and P^* is expected aggregate real producer price. Following Hallam and Zanouli (1992), we assume price expectations to be rational. This means that the expected future values of P and Q are reflected in their generation process. The laws of motion of P^* and Q^* are characterised by:

$$Q^* = A(L)\, Q_t + u_t$$

$$P^* = B(L)\, P_t + v_t \qquad (12)$$

Table 2: Unit-Root Tests

Variables	*ADF test in levels*	*differences*
Q^f	-1.18 (-3.59)	-5.60 (-3.01)
P^f	-3.24 (-3.59)	-3.95 (-3.01)
Q^e	-4.24 (-3.59)	-6.48 (-3.01)
P^e	-2.42 (-3.59)	-4.13 (-3.01)
P^f/P^e	-2.03 (-3.59)	-4.78 (-3.01)
Q^a	-2.58 (-3.59)	-7.41 (-3.01)
P^a	-3.09 (-3.59)	-4.12 (-3.01)

Critical values at the 5 per cent level are in brackets. The DF tests reported here are based on estimation of an equation including a trend, corresponding to equation (10); the results are supported by F-tests on the nulls of parameters jointly being zero. The tests on differences are based on second order integration tests and in most cases support a zero lag.

With A(L) and B(L) being the polynomial lag operators of order 1 for Q and P. If naive expectations were used, as is common in supply response models, one has $Q_t^* = Q_{t-1}$ and $P_t^* = P_{t-1}$, and one would estimate: $Q_t = bP_t + e_t$. However, if Q_t^* and P_t^* follow AR(1) processes, a static model would suffer from serial correlation as the omitted short-run dynamics are captured in the residual. Thus, to test cointegration we first estimate a dynamic specification of the model derived from (11) and (12), using OLS :

$$A(L)\, Q_t = B(L)\, P_t + e_t \qquad (13)$$

If significant, a trend is included in equation (13). We then test for cointegration through testing the stationarity of e_t using the Augmented Dickey-Fuller Test. A more robust test is the maximum likelihood approach proposed by Johansen (see Banerjee *et al*, 1993), which specifically provides a means to analyse the number of co-integrating vectors in a multivariate case (when there are more than two variables involved, it is possible to have more than one cointegrating vector). In these circumstances, the cointegrating vector estimated from OLS may not be unique and may be incorrectly interpreted as an estimate of the long-run relationship between the variables. The Johansen approach provides two test statistics for the number of cointegrating vectors: the trace and maximum eigenvalue tests. The trace statistic tests whether the number of cointegrating vectors is less than k (k = 0,1,2,...) whereas the maximum eigenvalue statistic tests whether the number of cointegrating vectors is k=0 (or k=1, k=2, ...) against the alternative k=1 (or k=2, k=3, ...). Results for the various tests are in Table 3.

Table 3: Cointegration Tests

Relationship	*ADF /EG*	*ADF/DR*	*Eigen value*	*Trace test*
Q^f, P^f, P^e, T	-4.93*	-5.44*	21.3 (25.5)	32.9 (42.4)
Q^f, P^f, T	-3.28	-3.01		
$Q^f, P^f/P^e, T$	-5.09*	-5.62*	19.6 (19.0)*	23.5 (25.3)
Q^e, P^e, P^f, T	-4.09*	n/a	35.8 (25.5)*	57.6 (42.4)*
Q^e, P^e, T	-4.31*	n/a		
$Q^e, P^f/P^e, T$	-4.08*	n/a		
$Q^e, P^f/P^e$	-1.47			
Q^a, P^a, T	-3.54	-1.73	12.9 (19.0)	15.3 (25.3)

Critical values for small samples: ADF/EG test, based on Engle-Granger two step method, for one price variable with trend, 3.58, and for two price variables with trend, 4.08 (* indicates support for cointegration); ADF/DR test, based on residuals from dynamic model, 4.17 and 4.60 respectively (* indicates support for cointegration). Critical values for Eigenvalue and Trace tests in parentheses. The ADF tests were applied using PC-GIVE and the trace and eigenvalue tests using PC-FIML. The constant was set *unrestricted* (i.e. not forced to lie in the cointegration space only), whereas the trend was *restricted* to the cointegrating relationship.

The results of the cointegration tests are decidedly mixed. There is fairly consistent evidence that the cointegrating relationship for Q^f is with respect to relative prices and trend. While there is some support for a cointegrating relationship for Q^e with P^e, P^f and *Trend*, it transpires that this appearance is due solely to the trend. There is no cointegrating relationship for Q^e excluding trend; there is no solved long-run equation and, as discussed below, the long-run determinant of Q^e appears to be a trend, although prices have a short-run influence. Perhaps for this reason, there is no support for a cointegrating relationship for total quantity and price indices. The balance of the various tests suggest that food producers respond to relative rather than own prices in the long-run. The solved static long-run equation is:

$$lnQ^f = -0.31 + 0.78 lnP^f - 0.93 lnP^e + 0.02 Trend$$

The two elasticities tend to offset each other, in that a reduction in export price has a similar impact to a similar proportional increase in food prices. We can test this by imposing the restriction that the sum of the elasticities is zero; a χ^2 test confirms that this restriction cannot be rejected. Inferring that food crop production responds to relative prices yields:

$$lnQ^f = -0.36 + 0.92\, ln(P^f/P^e) + 0.02 Trend$$

The evidence suggests that production of food crops has increased almost in proportion to the increase in their price relative to export crops, although there is a residual slight trend increase; that the trend decline in export crop production over this period is consistent with, but not apparently statistically explained by, the long run decline in producer prices for these crops. Over the long run then, farmers have been responsive to prices in switching between these two categories of crops. But what of the response in aggregate? There was no evidence for a long-run cointegrating relation for all crops, which may be due to the trend features of export crops, a major component of the total. For this reason we cannot report a long-run relationship, although we can consider the short-run dynamics.

Table 4 reports estimates of the short-run relationships: for food crops this is an ECM, but for exports and all crops the tests do not support estimating an ECM. Our most encouraging results (in a statistical sense) are for food crops: the short-run elasticity with respect to the price ratio P^f/P^e is significant and about 0.4 (roughly equivalent to own-price elasticity which is offset by cross-price elasticity); the significant coefficient on the ECM suggests that about 70% of deviation from long-run equilibrium is made up within one time period; and the positive coefficient on lagged output is consistent with the positive long-run trend. As exports appeared to be trend stationary, we estimated a short-run model in levels: the results confirm that the trend dominates; the positive sign on P^f/P^e (which is lagged three periods) is perhaps indicating some inter-cropping or complementarity in production in the short-run.

Table 4: ECM and Short-run Dynamic Models

Quantity	D$lnQlag$	DlnP^f	DlnP^e	D$ln(P^f/P^e)$	*ECM*	R^2	F
DlnQ^f	0.439	0.372	-0.415		-0.723		
	(2.18)	(1.85)	(-2.5)		(-3.13)	0.49	4.16
DlnQ^f	0.394			0.391	-0.715		
	(2.08)			(2.62)	(-3.44)	0.50	5.99

	Constant	D$ln(P^f/P^e)_{t-3}$	*Trend*	R^2	F
lnQ^e	0.496	0.272	-0.043		
	(9.70)	(1.89)	(-14.96)	0.92	114.18

	D$lnQlag$	D$lnP^a{}_t$	D$lnP^a{}_{t-1}$	*ECM*	R^2	F
DlnQ^a	-0.140	-0.177	0.268	-0.301		
	(-0.55)	(-1.12)	(1.67)	(-1.02)	0.43	3.20
DlnQ^a	-0.425		0.354			
	(-2.63)		(2.63)		0.47	9.58

Notes: ECMs based on residuals from a dynamic model; t-statistics in parentheses.

Turning to aggregate production, we first included an ECM as an indirect test of whether the failure to find cointegration reflected the low power of tests given the problematic statistical features of the export indices; none of the estimated parameters were significant (Table 4). A more simple specification produced more significant results: the short-run price elasticity of aggregate production is 0.35, higher than previous estimates, and the negative coefficient on lagged quantity is consistent with the downward trend in export crop production.

VI CONCLUSIONS

Many previous studies found evidence of very low long-run and short-run aggregate supply response for Tanzania and other SSA countries. Such evidence has been used to argue that farmers are not responsive, and to justify government policies biased against agriculture. We have argued that one possible reason for the low estimates is that the model and econometric techniques used were inappropriate. Based on what we consider more appropriate econometric techniques, our estimates suggest that agricultural supply response is quite high (and support World Bank, 1994). Moreover, our results are consistent with the conclusions of a recent qualitative study of the impact of economic reform based on village level fieldwork (Booth *et al*, 1993), which suggested that liberalisation had a beneficial impact on village level production (though not on the production of traditional export crops). Our results suggest that the potential for agricultural sector response to liberalisation of agricultural prices and marketing may be quite significant:

- The performance of export crop production (measured as official purchases), in both the long-run and short-run, can be fully explained by a secular downward trend. The dominance of the trend prevented estimation of price elasticities, although the trend in production is in line with that in prices. The failure to find a short-run response is consistent with the time lags inherent in export crop response, as many of the major crops are perennials.
- A 10 per cent increase in the relative price of food will lead to a 3.9 per cent increase in per capita food production in the short-run (the short-run own-price elasticity is similar) and a 9.2 per cent increase in the long-run.
- The error correction coefficient indicates that more than 70 per cent of the adjustment towards long-run equilibrium for food crops is completed in one period.
- Unlike earlier estimates (Table 1), although consistent with some more recent estimates, our results suggest that Tanzanian farmers are quite responsive to prices: the short-run elasticity of aggregate output is 0.35.

There are a number of limitations which must be stressed. Data availability restricted our analysis to data on official prices and purchases only. One obvious omission is that we cannot account for subsistence consumption of food, nor for diversion of output to informal markets. In the case of food, this implies that a potentially large proportion of production is excluded; almost our entire period of analysis is pre-liberalisation, when official prices were often held below the market level. Similarly, with respect to export crops, the Tanzanian Shilling was overvalued, often by large amounts, for most of the period up to the mid-1980s. Consequently, incentives for smuggling were often great. Our results should be interpreted as measuring the responsiveness of officially marketed surplus, and as such may over-estimate aggregate output response. Another major omission is that we are unable to incorporate non-price factors which influence supply response; weather conditions will affect short-run response but we did not have the time series data required to account for this (we did not incorporate the possibility of structural breaks, although examining the trends does not suggest there are any). The principal non-price influences, however, are likely to be constraints: limited access to credit and a steady supply of quality inputs will reduce responsiveness. In this sense our long-run elasticities may be under-estimates. Furthermore. liberalisation was too late in the period of analysis for any possible effects to be evaluated. These caveats notwithstanding, our results indicate higher price elasticities of supply than suggested by previous studies.

Our analysis therefore suggests that farmers are indeed responsive, which is consistent with the evidence of agricultural sector growth following adjustment policies in Tanzania in the mid-1980s. Liberalisation of agricultural markets, *where it increases the effective prices paid to farmers*, can be effective in promoting production, and is consistent with the observed improved performance of the sector following liberalisation in the 1980s. Complementary interventions, to improve infrastructure, marketing, access to inputs and credit, improved production technology etc, can be expected to make producers even more responsive. This latter point is especially important if the objective is to expand total agricultural output; our evidence is consistent with the view that much of the response is substitution between (export and food) crops, although there is a strong suggestion that total production will respond if constraints are relaxed and incentives improved.

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