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# The Output-Inflation Trade-off in African Less Developed Countries

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This paper investigates the nature of the short-run output-inflation trade-off and policy effectiveness in African less developed countries. Using a sample of thirteen countries over the period 1960-98, cointegration and error correction modelling suggest that the impacts on inflation and real output growth from a shock to nominal aggregate demand will be of the ratio one-sixth to five-sixths. Furthermore, this study finds that the short-run potency of demand-side policy on inflation (real output growth) is positively (inversely) related to the variability of nominal income shocks rather than the underlying rate of inflation. While the speed of adjustment towards long-run equilibrium between price and nominal output is fairly sluggish, it is concluded that the New Classical perspective on the trade-off is applicable in the case of African economies. The New Keynesian perspective, which emphasises wage and price rigidities and policy effectiveness, is probably of lesser relevance.

## I. Introduction

One of the most contentious areas of macroeconomic debate is the role of demand management in influencing real output. Whereas New Classical macroeconomic theory argues that rational expectations combined with market clearing eliminates the scope for short-run policy effectiveness (Lucas (1973), Sargent and Wallace (1976)), the introduction of wage and price rigidities provides some scope for short-run policy effectiveness (Fischer (1977)). Furthermore, New Keynesian macroeconomics offers a range of optimising models that formally justify assumptions of wage and price rigidities (Mankiw and Romer (1991)) An area of empirical investigation that draws on this debate is concerned with measuring the short-run trade-off between inflation and real output and therefore assessing the real impact of nominal demand shocks. Lucas (1973) argues that the trade-off becomes steeper with increases in the volatility of nominal aggregate demand shocks. However, the New Keynesian viewpoint, represented by Ball *et al*. (1988) and Ball and Mankiw (1994), employs menu cost models to argue that it is higher rates of inflation, rather than the volatility of nominal demand shocks, that reduces the short-run effectiveness of demand management because agents have greater incentives to change prices rather than quantities. While numerous studies have found general empirical support for the Lucas model, the samples of countries used have tended to lump developed- and less developed countries (LDCs) together (see, *inter alia*, Alberro (1981), Attfield and Duck (1983), Fernandez (1977),

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Jung (1985), Parkin *et al.* (1981), Ram (1984), Taylor (1980)). Among the few notable exceptions is Odedokun (1991) who explicitly distinguishes between developed and LDCs to reject the Lucas proposition in the latter case. This paper contributes to this literature by investigating the slope of the trade-off and policy effectiveness in the case of African LDCs. Using annual data for a sample of thirteen African economies over periods of study between 1960 and 1998, the focus is on the short-run relationship between inflation and nominal demand shocks to provide measures of policy effectiveness. The methodology employed is cointegration analysis where the short-run impact of nominal demand shocks is modelled within an error correction framework.

Such a study is of interest for a number of key reasons. First, we offer a hitherto unexplored area of macroeconomic research that focuses explicitly on African LDCs. If policy effectiveness is dependant upon price rigidities, this study offers the opportunity to resolve a debate over the extent of price rigidities in LDCs. On the one hand, Collier and Gunning (1999) highlight the types of price rigidities that are prevalent in African LDCs which, they argue, serve to hinder the achievement of market clearing pricing- restrictions on traders, high taxation and poor infrastructure. However, Hossain and Chowdhury (1998) argue that price rigidities in LDCs are limited because agricultural prices are flexible showing considerable variation and fluctuations between peak and slack times. Since the agricultural sector constitutes a significant proportion of African national output, particularly in the case of low-income LDCs, this contributes towards a significant degree of wage and price variability.[1] Second, this study contributes to the New Classical versus New Keynesian debate concerning policy effectiveness. Policy effectiveness is measured and there is an assessment on whether the experience of African LDCs complies with the Lucas or Ball *et al.* view on what determines the slope of the trade-off. Third, the methodology employed is based on cointegration and error correction modelling. This overcomes a shortcoming of the methodologies employed in the above-mentioned studies that do not pay close attention to the statistical properties of the data they employ. For completeness and comparison, estimates based on the 'original' Lucas and Ball *et al.* methodologies are also reported.

The paper is organised as follows. The following section discusses the relevant literature and derives the model used for investigation. The third section discusses the data and analyses the results obtained. It is found that in the case of African LDCs typically five-sixths of the short-run effect of a positive nominal demand shock will be felt in terms of increased real output growth. Furthermore, there is evidence in favour of the New Classical perspective on policy effectiveness because the slope of the short-run trade-off between inflation and real output growth is positively related to the volatility of nominal demand shocks. The final section concludes.

## II. Relevant Literature and the Model

In his original work, Lucas (1973) argues that the responsiveness of real output to aggregate demand movements is negatively related to the size of aggregate demand shocks.

1. Hossain and Chowdhury (1998, pp. 185-7).

Using a two step procedure, the real effects of nominal demand shocks are captured by the slope of the short-run Phillips curve ($\tau$) which is derived from a regression where the natural logarithm of real output ($y$) is regressed on the change in the natural logarithm of nominal output ($\Delta x$). The use of $\Delta x$ as a proxy for a nominal demand shock is based on the assumptions that the price elasticity of aggregate demand is unity where nominal output is determined by the demand-side of the economy while real output is determined by the supply-side. Using this framework, $\tau = 1$ ($\tau = 0$) means that all (none) of the change in nominal demand is reflected in $y$. In the second step, regression analysis finds that $\tau_i$ is inversely related to the standard deviation of $\Delta x$, denoted as $\sigma_{\Delta x,i}$, which is used as a measure of the variability of nominal demand shocks. Using this framework, general support for the Lucas model is offered by a range of subsequent studies that include Fernandez (1977), Jung (1985) and others.

A New Keynesian perspective on the trade-off debate is offered by Ball *et al.* (1988) who develop a menu cost model of price adjustment. Firms operate in an imperfectly competitive market aiming to maximise profits, however the adjustment of prices is subject to a menu cost. The size of $\tau$ depends on how often firms adjust their prices. The greater is the speed of adjustment, the smaller is $\tau$, i.e., the smaller are the real effects of nominal demand shocks.[2] However, the higher is inflation, the greater is the frequency of price adjustment and therefore the smaller is the value of $\tau$.[3] The following equation is used to estimate $\tau$.

$$y_t = \alpha + \tau\Delta x_t + \eta y_{t-1} + \gamma time + u_t, \tag{1}$$

where *time* is a time trend and $u$ is a residual. Estimating (1) across 43 countries for the period 1948-86 yields $\bar{\tau} = 0.242$ which is indicative of some degree of nominal price rigidity. In the second step, Ball *et al.* confirm a negative (non-linear) relationship between $\tau_i$ and average inflation ($\bar{\pi}_i$). It is possible that a negative relationship has been identified because $\bar{\pi}_i$ and $\sigma_{\Delta x,i}$ are positively correlated, i.e., countries with high inflation generally have more variable aggregate demand. Ball *et al.* estimate an equation for $\tau_i$ that nests both $\bar{\pi}_i$ and $\sigma_{\Delta x,i}$ finding that the latter is insignificant at any reasonable degree of confidence.

Numerous studies have empirically tested the income variability hypothesis (see, *inter alia*, Lucas (1973), Alberro (1981), Parkin *et al.* (1981), Ram (1984)) and are generally supportive. However, these studies generally lump developed countries and LDCs together thereby removing the opportunity to examine differences between these countries that might be due to their respective economic structures. This point is acknowledged by Odedokun (1991) whose study incorporates ninety LDCs using a variety of study periods within the

2. This model can be used to show that $\tau$ and $\sigma_{\Delta x,i}$ are negatively related insofar as the higher is the variance of aggregate demand then less certain is the firm about what its optimal price should be and therefore the shorter is time interval between price changes.

3. The argument that an inflationary regime increases the penalty of not adjusting prices in response to a demand shock is further analysed in theoretical contributions that include Tssidon (1993).

boundaries 1958 and 1985. Using OLS, Odedokun finds little evidence of support for the natural rate theory. Moreover, it is only in the case of developed countries where the variability of aggregate demand (proxied by nominal income growth) significantly affects real cyclical output. In explaining the result for LDCs, Odedokun appeals to mass illiteracy, small scale peasant primary production and gross market imperfections. Addison *et al.* (1986) also finds that the potential role of demand management is larger in LDCs than in developed countries whereas Jung (1990) and Katsimbris (1990a, 1990b) reject this claim.[4]

The empirical approach employed in this paper is based on two methodologies. First, the 'original' approach used by Lucas (1973), Ball *et al.* (1988) and others is followed where Equation (1) is estimated by OLS to gauge the slopes of the trade-off curves ( $\boldsymbol{t}_i$ 's).[5] Across countries, $\boldsymbol{t}_i$ is then regressed on $\bar{\boldsymbol{p}}_i$ and $\boldsymbol{s}_{\Delta x,i}$

$$\boldsymbol{t}_i = \boldsymbol{k}_0 + \boldsymbol{k}_1 \bar{\boldsymbol{p}}_i + u_{1t}\,, \tag{2a}$$

$$\boldsymbol{t}_i = \boldsymbol{y}_0 + \boldsymbol{y}_1 \boldsymbol{s}_{\Delta x,i} + u_{2t}\,, \tag{2b}$$

to confirm or reject the New Classical or New Keynesian view of the determinants of $\boldsymbol{t}_i$ according to whether $\boldsymbol{k}_1 > 0$ or $\boldsymbol{y}_1 > 0$.

The second methodological approach addresses a major shortcoming of the original Lucas and Ball *et al.* methodologies. If (1) is to constitute a valid estimating relationship, then $y$ and $\Delta x$ should be integrated to the same order, i.e., both should be first difference stationary (I(1)) if a cointegrating relationship is present, or both should be stationary (I(0)) if OLS can be applied and a valid interpretation made of the $F$ and $t$ statistics. In the earlier literature, little has been said about statistical properties of $y$ and $\Delta x$. Indeed, Equation (1) could well contain a combination of stationary and non-stationary variables particularly if one needs to assume that $\Delta x$ is stationary if it is to be interpreted as a demand shock. This analysis considers the short-run impact of changes in nominal income through a cointegration framework that sees the short-run inflation-real output trade-off in the context of an error correction model that comprises stationary variables only.

Let us consider this alternative two-step methodology in more detail. The long-run relationship from which the error correction model is derived is between the natural logarithm of the price level $p$ and $x$. Assume that in the long-run, movements in $x$ are associated with movements in $p$ and movements in the level of real output, $y$. Further assume

4. An alternative methodology for investigating policy effectiveness is provided by Barro (1977, 1978). This two-step procedure first estimates monetary policy shocks using forecast errors derived from a hypothesised money supply equation. The second step looks at how these shocks along with systematic policy affect real output. Odedokun (1993) investigates whether unanticipated and systematic monetary policy influences real output in the case of sixteen LDCs over the period 1957-89 and finds little support for the rational expectations hypothesis in (particularly low income) LDCs.

5. Some of these studies, for example Odedokun (1991), employ a measure of cyclical real output as the dependent variable. As an alternative to this, Equation (1) employs a time trend as an explanatory variable.

that $p$ and $x$ are first difference stationary. The long-run relationship between these two variables may be written as

$$p_t = g_0 + g_1 x_t + u_t \,, \tag{3}$$

where $g_1 \geq 0$ and $u_t$ is the deviation from long-run equilibrium. The extent of price rigidities will be instrumental in governing the long-run relationship between $x$ and $p$. If there is zero price flexibility then movements in $x$ are not associated with movements in $p$ so there are corresponding movements in $y$. If $g_1 = 1$ then movements in $x$ are matched by movements in $p$ therefore $y$ is unchanged. Finally, it might be the case that $g_1 > 1$ which means that the long-run relationship between $p$ and $x$ is such that a given movement $x$ gives rise to an even larger response of $p$ such that $y$ is reduced. This last scenario may reflect a situation where, say, an increase in nominal income ultimately creates a wage-price spiral.

The Johansen (1988) multivariate approach is particularly suited to the analysis of the relationship expressed in Equation (3) because it is a multivariate approach that allows for potential endogeneity between the variables concerned. As well as estimating the long-run equilibrium relationship between $p$ and $x$, it also allows the simultaneous estimation of the speed at which equilibrium is re-established following some disturbance (Banerjee *et al*. (1993), Maddala and Kim (1998)). Moreover, the Granger representation theorem demonstrates that cointegrated variables have a valid error correction model representation. It is the short-run dynamics behind the long-run relationship between $p$ and $x$ provides the perspective on the short-run output-inflation trade-off. Define $\mathbf{z}$ as an $(n\text{x}1)$ vector of stochastic variables ($p$ and $x$). The unrestricted vector autoregression (VAR) can be written as $\mathbf{z}_t = \sum_{i=1}^{k} \mathbf{A}_i \mathbf{z}_{t-i} + \mathbf{u}_t$ where $\mathbf{A}$ is an $(n\text{x}n)$ matrix of parameters and $\mathbf{u}_t \sim IN(0, \Sigma)$. The Johansen procedure entails the maximum likelihood estimation of the following vector error correction model (VECM) which is derived from the VAR.

$$\Delta \mathbf{z}_t = \sum_{i=1}^{k-1} \Gamma_i \Delta \mathbf{z}_{t-i} + \Pi \mathbf{z}_{t-k} + \mathbf{u}_t \,. \tag{4}$$

The hypothesis of cointegration between $p$ and $x$ can be formulated as a hypothesis about the rank $(r \leq n-1)$ of matrix $\Pi$ which can be decomposed into a matrix of long-run coefficients ($\boldsymbol{b}$) and the matrix of speed of adjustment coefficients ($\boldsymbol{a}$), i.e., $\Pi = \boldsymbol{a}\boldsymbol{b}'$. Thus $\boldsymbol{b}'\mathbf{z}_{t-k}$ represents the cointegration relationship between $p$ and $x$, namely $p_t - g_0 - g_1 x_t$, while the elements of $\boldsymbol{a}$ provide information on the short-run responsiveness of inflation to a nominal income shock that has disturbed the long-run equilibrium. The closer are the elements of $\boldsymbol{a}$ to zero then the less flexible are prices in the short-run. If we denote the relevant speed of adjustment as $\boldsymbol{l}$, then it can be said that the absolute value of the speed of adjustment $|l|$ is positively related to the slope of the trade-off. If $k > 1$ then the short-run trade-off will also be influenced by any lagged $\mathbf{z}$ that appears in the VECM. We can therefore define the short-run impact of a nominal demand shock ( $\boldsymbol{t}_0$ ) as the sum of $|l|$ and any significant coefficient on $\Delta x_{t-i}$ .

Using the information obtained from the estimation of (4), a key proposition of the New Classical versus New Keynesian debate can be formally tested. New Classical economists would argue that the sensitivity of real output to a nominal demand shock is negatively related to the volatility of demand shocks. Using the estimates for $\boldsymbol{t}_0$ , we can test whether the sensitivity of inflation (and therefore real output growth) to a nominal demand shock is positively (negatively) influenced by average inflation $\bar{\boldsymbol{p}}$ (the standard deviation of demand shocks $\boldsymbol{s}_{\Delta x,i}$ ). In other words,

$$\boldsymbol{t}_{0,i} = \boldsymbol{k}_0 + \boldsymbol{k}_1 \bar{\boldsymbol{p}}_i + u_{3t} , \quad (5a)$$

$$\boldsymbol{t}_{0,i} = \boldsymbol{y}_0 + \boldsymbol{y}_1 \boldsymbol{s}_{\Delta x,i} + u_{4t} , \quad (5b)$$

where $\boldsymbol{k}_1 > 0$ ( $\boldsymbol{y}_1 > 0$ ) may lend support the New Keynesian (New Classical) view of the inflation-output trade-off in the policy ineffectiveness proposition.

### III. Data and Results

Annual data for *p*, *x* and *y* are obtained from *International Financial Statistics* for Congo (1964-96), Ethiopia (1966-97), Gabon (1963-96), Ghana (1963-97), Ivory Coast (1960-97), Kenya (1968-97), Mauritius (1962-98), Niger (1968-97), Nigeria (1960-98), Seychelles (1971-97), South Africa (1960-98), Swaziland (1967-97) and Togo (1969-97). Generally the total period of study covers 1960-98 where the number of observation ranges from 27 observations in the case of the Seychelles to 39 observations in the cases of Nigeria and South Africa. The price level *p* is based on consumer price data (line 64) and GDP data are taken from line 99.

Table 1 reports the estimates for Equation (1) which is based on the methodologies employed by Lucas and Ball *et al*. In all cases we find that the slope of the inflation-output trade-off, $\boldsymbol{t}$ , is non-negative with $\bar{\boldsymbol{t}} = 0.584$ . According to the Ball *et al*. interpretation of $\boldsymbol{t}$ , this means that the effects of a given nominal demand shock will be distributed as 58.4% towards increased real output and 41.6% towards higher inflation. This suggests that there is a substantial degree of policy effectiveness in African LDCs. The higher is the value of $\boldsymbol{t}$ then the flatter is the trade-off. In their original study which embodied developed countries and LDCs, Ball *et al*. calculated $\bar{\boldsymbol{t}} = 0.242$ thus, using Equation (1) at least, it would appear that policy effectiveness is greater in African LDCs than in developed countries. There is considerable variation across the estimates for $\boldsymbol{t}$ . In particular, the response of real output to nominal demand shocks is greatest in the cases of Ethiopia, Swaziland and Congo whereas it is insignificantly different from zero in the cases of Ghana and Kenya and rather low in the case of Niger. The next step in this methodology is to estimate Equations (2a) and (2b) to see if the slope of the trade-off is driven by average inflation or the variability of demand shocks. The results reported in Table 2 (Part A) indicate that neither explanatory variable plays a significant role in influencing $\boldsymbol{t}$ .

**Table 1 Initial Estimation of the African Output-Inflation Trade-Off**

| | $t$ | $h$ | $g$ | $\bar{R}^2$ | $SE$ | $Q$ | $s^2_{\Delta x,i}$ | $\bar{p}$ |
|---|---|---|---|---|---|---|---|---|
| Congo | 0.803*** (0.102) | 0.965*** (0.052) | −0.000 (0.002) | 0.969 | 0.079 | 2.690 | 0.018 | 0.072 |
| Ethiopia | 1.076*** (0.333) | 1.019*** (0.134) | −0.000 (0.002) | 0.660 | 0.094 | 1.593 | 0.003 | 0.066 |
| Gabon | 0.775*** (0.067) | 0.848*** (0.042) | 0.009*** (0.003) | 0.986 | 0.078 | 6.049 | 0.040 | 0.063 |
| Ghana | 0.195 (0.195) | 0.892*** (0.123) | −0.001 (0.003) | 0.660 | 0.185 | 4.710 | 0.045 | 0.281 |
| Ivory Coast | 0.519*** (0.078) | 0.915*** (0.032) | 0.001 (0.001) | 0.983 | 0.046 | 9.031 | 0.010 | 0.068 |
| Kenya | −0.308 (0.301) | 0.809*** (0.090) | 0.001 (0.002) | 0.864 | 0.061 | 7.757 | 0.002 | 0.126 |
| Mauritius | 0.656*** (0.096) | 0.822*** (0.079) | 0.010*** (0.005) | 0.991 | 0.060 | 5.692 | 0.011 | 0.082 |
| Niger | 0.291** (0.125) | 0.837*** (0.130) | 0.004 (0.004) | 0.919 | 0.066 | 6.513 | 0.010 | 0.059 |
| Nigeria | 0.636*** (0.127) | 0.909*** (0.060) | −0.004 (0.002) | 0.910 | 0.116 | 11.139 | 0.026 | 0.165 |
| Seychelles | 0.905*** (0.135) | 0.974*** (0.096) | 0.008 (0.006) | 0.987 | 0.051 | 0.761 | 0.012 | 0.061 |
| South Africa | 0.461*** (0.099) | 0.834*** (0.038) | 0.003** (0.001) | 0.997 | 0.018 | 10.694 | 0.001 | 0.092 |
| Swaziland | 0.891*** (0.114) | 0.823*** (0.060) | 0.006* (0.003) | 0.986 | 0.047 | 3.533 | 0.006 | 0.112 |
| Togo | 0.389*** (0.125) | 0.832*** (0.184) | 0.003 (0.003) | 0.839 | 0.063 | 4.285 | 0.012 | 0.073 |

Estimation of the equation $y_t = a + t\Delta x_t + h y_{t-1} + g time + u_t$ where $\bar{R}^2$ is the adjusted goodness of fit, SE is the standard error of the regression, $Q$ is the Box-Pierce test for zero coefficients on five lagged autocorrelation coefficients, standard errors are reported in parentheses. ***, ** and * denotes rejection of the null hypotheses at the 1, 5 and 10% significance levels respectively.

**Table 2 Sensitivity of the Output-Inflation Trade-Off**

**Part A Estimates based on 'Original' Approach**

| | $k_0$ | $k_1$ | $y_0$ | $y_1$ | $\bar{R}^2$ | SE |
|---|---|---|---|---|---|---|
| $t_i$ | 0.807** (0.160) | −2.196 (1.359) | | | 0.118 | 0.292 |
| $t_i$ | | | 0.594** (0.213) | −0.090 (1.734) | 0.000 | 0.326 |

**Part B Estimates based on Error Correction Modelling Approach**

| | $\kappa_0$ | $\kappa_1$ | $\psi_0$ | $\psi_1$ | $\bar{R}^2$ | SE |
|---|---|---|---|---|---|---|
| $\tau_{0,i}$ | 0.136** (0.057) | 0.234 (0.485) | | | 0.000 | 0.104 |
| $\tau_{0,i}$ | | | 0.033 (0.054) | 1.145** (0.443) | 0.321 | 0.083 |
| $\tau_{0,i}$ | 0.045 (0.059) | −0.034 (0.450) | | 1.309** (0.516) | 0.285 | 0.085 |

Part A: OLS estimation of $\tau_i = \kappa_0 + \kappa_1 \bar{\pi}_i + u_{1t}$ and $\tau_i = \psi_0 + \psi_1 \sigma_{\Delta x,i} + u_{2t}$ where $\tau_i$ 's are taken form the estimates for Equation (1). Part B: OLS estimation of $\tau_{0,i} = \kappa_0 + \kappa_1 \bar{\pi}_i + u_{3t}$ and $\tau_{0,i} = \psi_0 + \psi_1 \sigma_{\Delta x,i} + u_{4t}$ where $\tau_{0,i}$ 's are taken from the estimates of Equation (4). $\bar{R}^2$ is the adjusted goodness of fit, SE is the standard error of the regression and standard errors are given parentheses and ** denotes rejection of the null at the 5% significance level.

We can now consider the second methodological approach which models the trade-off parameter through cointegration analysis and an error correction framework. For Equation (3) to constitute a long-run cointegrating relationship, we must first satisfy ourselves that $p$ and $x$ are I(1) variables. Augmented Dickey-Fuller (ADF) tests are used to investigate the time-series properties of $p$ and $x$. Where appropriate, the relationship between $p$ and $x$ defined in Equation (4) is estimated using the Johansen (1988) maximum likelihood cointegrating procedure.

Table 3 reports the unit root tests for $p$ and $x$. For all series except Ethiopia, Seychelles and South Africa we find that $p$ and $x$ ~ I(1) while $\pi$ and $\Delta x$ ~ I(0) at the 5% significance level. Thus in the majority of cases Equation (4) can be estimated using the Johansen maximum likelihood procedure. In the cases of Ethiopia and South Africa, there is evidence that $p$ and $x$ ~ I(2) while $\pi$ and $\Delta x$ ~ I(1) and for the Seychelles $p$ and $x$ ~ I(0). For these three countries, such statistical properties of the data mean that finding a cointegrating relationship between $p$ and $x$ using the Johansen procedure is problematic. Instead, a cointegrating relationship between $\pi$ and $\Delta x$ is sought in the cases of Ethiopia and South Africa while we may concentrate on estimating the VAR in levels in the case of the Seychelles.

**Table 3 ADF Unit Root Tests**

| | Period | $p$ | $\Delta p$ | $x$ | $\Delta x$ |
|---|---|---|---|---|---|
| Congo | 1965-96 | −2.006 | −4.251*** | −1.691 | −3.661** |
| Ethiopia | 1967-97 | −0.895 | −2.911* | −0.145 | −2.878* |
| Gabon | 1964-96 | −1.010 | −4.042*** | −1.214 | −4.140*** |
| Ghana | 1964-97 | −3.204 | −4.224*** | −2.114 | −5.519*** |
| Ivory Coast | 1961-97 | −1.720 | −4.039*** | −1.361 | −3.491** |
| Kenya | 1969-97 | −2.161 | −3.353** | −3.057 | −3.503** |
| Mauritius | 1963-98 | −2.163 | −3.251** | −2.635 | −4.145*** |
| Niger | 1969-97 | −3.310 | −3.045** | −0.755 | −3.891*** |

**Table 3 (Continued)**

| | Period | $p$ | $\Delta p$ | $x$ | $\Delta x$ |
|---|---|---|---|---|---|
| Nigeria | 1961-98 | −1.744 | −3.346** | −2.172 | −3.941*** |
| Seychelles | 1972-97 | −6.603*** | −3.779*** | −3.913** | −3.137** |
| South Africa | 1961-98 | −2.965 | −1.811 | −2.264 | −2.861* |
| Swaziland | 1968-97 | −2.326 | −6.344*** | −2.191 | −4.707*** |
| Togo | 1970-97 | −2.285 | −4.170*** | −0.745 | −6.494*** |

Lag lengths in the ADF unit root tests are determined by the Schwarz Information Criteria. Tests on levels, i.e., *p* and *x*, include a time-trend with relevant critical values −4.38, −3.60 and −3.24 for significance at the 1, 5 and 10% levels respectively. The remaining tests, which are on first differences, exclude a time trend and have critical values of −3.75, −3.00 and −2.63 for significance at the 1, 5 and 10% levels respectively. Critical values are taken from Fuller (1976). ***, ** and * denotes rejection of the null hypotheses at the 1, 5 and 10% significance levels respectively.

Table 4 reports the cointegration results. In all cases, there is a single cointegrating vector between *p* and *x*. Following Pesaran and Shin (1999), a single normalising restriction is sufficient to exactly identify the cointegrating relationship. The positive estimates for $g_1$ confirm the priors with values ranging from 0.528 to 1.507 in the respective cases of Gabon and South Africa. The majority of values for $g_1$ are less than unity which initially suggests that there exist long-run price rigidities to the extent that nominal income movements have real effects. However, at the 5% significance level, likelihood ratio tests indicate that the null $g_1 = 1$, which constitutes an over-identifying restriction, is accepted in all cases except Congo, Gabon, Ivory Coast and South Africa. Thus, in the long-run, the majority of African LDCs experience nominal income movements that are accompanied by equivalent price level movements. There is strong evidence that Congo, Gabon and the Ivory Coast are characterised by $g_1 < 1$ which is indicative of long-run price rigidities, while the likelihood ratio tests of restrictions confirms the South African result that $g_1 > 1$ which suggests that ultimately, given nominal income movements lead to even greater price increases (say, through a wage-price spiral) which puts downward pressure on real output.[6]

**Table 4 Cointegration Analysis**

| | $g_0$ | $g_1$ | $H_0 : r = 0$ | $H_0 : r \leq 1$ | $k$ | $c^2(1)$ |
|---|---|---|---|---|---|---|
| Congo | 1.061 | 0.565 | 43.991** | 7.141 | 1 | 34.063*** |
| Ethiopia | 0.009 | 0.715 | 25.384** | 6.914 | 2 | 0.152 |
| Gabon | | 0.528 | 30.097** | 2.752 | 1 | 22.774*** |
| Ghana | −3.604 | 0.968 | 52.529** | 2.269 | 1 | 1.240 |
| Ivory Coast | −0.772 | 0.670 | 56.190** | 7.738 | 1 | 37.317** |

6. For completeness, the results for Ethiopia and South Africa are included but it should be remembered that these are based on long-run relationships between $p$ and $\Delta x$.

**Table 4 (Continued)**

| | $g_0$ | $g_1$ | $H_0 : r = 0$ | $H_0 : r \leq 1$ | $k$ | $c^2(1)$ |
|---|---|---|---|---|---|---|
| Kenya | −3.174 | 0.928 | 101.577** | 7.407 | 1 | 0.114 |
| Mauritius | −2.273 | 0.803 | 63.261** | 5.186 | 1 | 2.158 |
| Niger | −1.052 | 0.870 | 27.750** | 2.465 | 1 | 0.728 |
| Nigeria | −8.327 | 1.242 | 51.124** | 3.824 | 1 | 3.939 |
| South Africa | −0.089 | 1.507 | 26.254** | 3.700 | 1 | 6.918*** |
| Swaziland | −0.529 | 0.800 | 76.690** | 5.945 | 1 | 2.715* |
| Togo | −1.672 | 1.018 | 32.799** | 5.779 | 1 | 0.009 |

Estimation of $p_t = g_0 + g_1 x_t + u_t$ is by the Johansen (1988) maximum likelihood cointegration approach where $r$ refers to the number of cointegrating vectors according to the Trace test, $H_0$ is the null hypothesis concerning *r*, *k* is the lag length of the VAR determined by the Schwarz Information Criteria, $c^2(1)$ is the test statistic for the over-identifying restriction $H_0 : g_1 = 1$ where ***, ** and * denotes rej ection of the null hypotheses at the 1, 5 and 10% significance levels respectively. Critical values for the Trace test are taken from Osterwald-Lenum (1992). Following the application of the Pantula Principal, all vectors include restricted intercepts and no trend except Gabon which features no intercept and no trend.

Using the error correction terms associated with the estimated cointegrating vectors, Table 5 reports estimates of each country' s vector error correction model as defined in Equation (4). These estimates permit us to examine the short-run effect of nominal demand shocks on inflation. The estimates of $t_0$ define the short-run effect of $\Delta x$ on inflation and we require that $l < 0$ and significantly different from zero if the estimates in Table 4 are to constitute valid cointegrating relationships. Generally, there is a reasonable goodness of fit where the residuals satisfy autocorrelation and normality tests. The estimates for $t_0$ and $l$ conform to the priors. Across the sample we find that $\bar{t}_0 = 0.160$ which means that on average, the effects of a given positive nominal demand shock will be distributed approximately one-sixth towards higher inflation and five-sixths towards higher real output growth. This would suggest that central authorities have considerable scope to engage in short-run demand management. This is result may be compared to the earlier calculation of $\bar{t} = 0.584$ (and therefore $(1-\bar{t}) = 0.416$ ) which suggests that the effects of a nominal demand shock are less strongly distributed towards real output. The results reported in Table 5 can also be compared with Odedokun (1991) who employs the Lucas and Ball *et al.* methodologies on ten of the African LDCs used here to find that the output-inflation trade-off is steeper (implied degree of price flexibility is greater).[7] Indeed, this same conclusion applies when a comparison is made between the Table 5 results and the results that Odedokun obtains for twenty-one developed countries, i.e., real output is more responsive to nominal demand shocks in African LDCs than in developed economies. It should be noted that $\bar{t}_0 = 0.160$ masks a considerable degree of variation across African LDCs. At the extremes, the Ethiopia, Gabon and Ghana results indicates considerable

7. The countries excluded are Gabon, Seychelles and Swaziland.

short-run price flexibility with high speed of adjustment coefficients while the opposite applies in the cases of Kenya, Mauritius and Swaziland where price rigidities are of greater prominence suggesting that the almost the full impact of a nominal demand shock will be reflected in real output growth.

**Table 5 Error Correction Analysis of Short-Run African Trade-Off**

| | $l$ | $\Delta x_{t-1}$ | $t_0$ | $\overline{R}^2$ | *Se* | *B-J* | *LM* |
|---|---|---|---|---|---|---|---|
| Congo | $-0.226^{***}$ (0.032) | | 0.226 | 0.670 | 0.046 | 2.643 | 0.198 |
| Ethiopia | $-0.653^{**}$ (0.245) | $-0.438^{*}$ (0.231) | 0.215 | 0.252 | 0.100 | 0.971 | 0.031 |
| Gabon | $-0.327^{***}$ (0.072) | | 0.327 | 0.743 | 0.051 | 0.235 | 1.867 |
| Ghana | $-0.322^{***}$ (0.056) | | 0.322 | 0.035 | 0.231 | 0.924 | 0.588 |
| Ivory Coast | $-0.188^{***}$ (0.023) | | 0.188 | 0.511 | 0.048 | 1.324 | 0.739 |
| Kenya | $-0.017^{***}$ (0.002) | | 0.017 | 0.686 | 0.047 | 0.430 | $3.431^{*}$ |
| Mauritius | $-0.024^{***}$ (0.003) | | 0.024 | 0.672 | 0.040 | 1.678 | 0.444 |
| Niger | $-0.140^{***}$ (0.031) | | 0.140 | 0.686 | 0.054 | 1.743 | 1.602 |
| Nigeria | $-0.084^{***}$ (0.012) | | 0.084 | 0.498 | 0.109 | 2.473 | 0.269 |
| Seychelles | | $0.171^{*}$ (0.088) | 0.171 | 0.992 | 0.024 | 0.034 | 0.126 |
| South Africa | $-0.151^{*}$ (0.081) | | 0.151 | 0.058 | 0.020 | 1.571 | 0.206 |
| Swaziland | $-0.039^{***}$ (0.003) | | 0.039 | 0.188 | 0.049 | 1.846 | 0.950 |
| Togo | $-0.177^{***}$ (0.035) | | 0.177 | 0.600 | 0.057 | 0.676 | 2.389 |

See notes for Table 1. Estimation of $\Delta\mathbf{z}_t = \Gamma_1\Delta\mathbf{z}_{t-1} + \ldots + \Gamma_{k-1}\Delta\mathbf{z}_{t-k+1} + \Pi\mathbf{z}_{t-k} + \mathbf{u}_t$. Dependant variable is $\Delta p$ while $l$ is the speed of adjustment derived from $\Pi = ab'$. The short-run impact of a nominal demand shock ($t_0$) is calculated as the sum of $|l|$ and any significant coefficient on $\Delta x_{t-1}$. Figures in parentheses are White heteroscedasticity-adjusted standard errors, *B-J* is the Jarque-Bera test for normal residuals distributed as $c^2(2)$ on the null, *LM* is the LM test for first order serial correlation of the residuals distributed as $c^2(1)$ on the null. Estimates for Seychelles employ data on *p* and *x* while Ethiopia and South Africa employ data on $\Delta^2 p$ and $\Delta^2 x$.

Unlike the Lucas and Ball *et al*. methodologies, using the error correction modelling framework allows for the impact of the previous period's disparity from long-run equilibrium between the levels of $p$ and $x$ to impinge on inflation. Further analysis of the error correction results indicate that $\bar{\lambda} = -0.196$, i.e., following a shock to long-run equilibrium between $p$ and $x$, 19.6% of the required adjustment back to equilibrium is completed each year. The mean half-life of a shock to long-run equilibrium is of the order $\ln(0.5)/\ln(1+\bar{\lambda}) = 3.177$ years. Considerable variability is masked by the mean result. Following the above discussion, we would expect the half-life to be shorter in Ethiopia, Gabon and Ghana than in the remaining countries.

Following Lucas and Ball *et al*., we can now consider whether the value of $\tau_0$ is affected by the average inflation. As prescribed in Equations (5a) and (5b), the values of $\tau_0$ taken from Table 4 are regressed on the average values for $\Delta p$ and $\sigma_{\Delta x,i}$ reported in Table 1.[8] The results are reported in Table 2 (Part B). The positive and significant values for $\psi_1$ lend some support to the New Classical perspective that proposes a negative (positive) relationship between the variability of nominal demand shocks and the sensitivity of real output (inflation) to demand shocks. This finding can be contrasted with the New Keynesian view that it is the average rate of inflation that is crucial in determining the nature of the short-run trade-off by making inflation (real output growth) more (less) sensitive to a nominal demand shock. The estimation of Equation (5a) reveals that the coefficient on $\pi_i$ is correctly signed but insignificantly different from zero. Indeed, even at very generous confidence levels, we are unable to reject the null hypothesis that inflation does not influence the slope of the trade-off since the p-value associated with $\kappa_1$ is equal to 0.639. These findings can be seen in the context of studies which examine long-run purchasing power parity (PPP) in LDCs and find that relative PPP through price adjustment is more likely to hold in the case of high inflation countries (Holmes (2000), Liu (1992), Mahdavi and Zhou (1994), McNown and Wallace (1989)). Clearly, it takes time for price adjustment in African LDCs to respond to the underlying rate of inflation.

## IV. Summary and Conclusion

This study has estimated the inflation-output trade-off for African LDCs through a cointegration and error correction modelling framework that assesses the impact effect of nominal demand shocks on inflation. It has been argued that this methodology has a distinct advantage over other studies of the trade-off in developed and LDCs which have paid less attention to the time-series properties of the data employed. Furthermore, several of these studies have tended lump developed and less developed countries together thereby side-stepping important structural differences that might affect the nature of the trade-off. Using data on inflation and nominal output for a sample of thirteen African LDCs over the period 1960-98, it is estimated that the short-run effect of a given nominal aggregate demand shock will be distributed one-sixth towards inflation and five-sixths towards real output

8. Each $\tau_{0,i}$ is assigned a value of zero where significance at the 10% level or better has not been achieved.

growth. This result suggests that central authorities have considerable scope for effective short-run demand management. When compared to earlier studies of the trade-off, this result indicates that the slope of the output-inflation trade-off is flatter for African LDCs than for developed countries. The high sensitivity of real output growth to nominal demand shocks, and therefore policy effectiveness, may reflect considerable price inflexibility in African LDCs. However, there is support for New Classical macroeconomics insofar as the variability of aggregate demand and the sensitivity of inflation to nominal demand shocks are positively related. A useful avenue for continued research might be to examine the effectiveness of demand-side policy, and monetary policy in particular, paying attention to the possibility of asymmetries in any short- and long-run real effects.

## References

Addison, J., H. Chappell, and A. Castro (1986), "Output-Inflation Trade-Offs in 34 Countries," *Journal of Economics and Business*, 38, 353-60.

Alberro, J. (1981), "The Lucas Hypothesis on the Phillips Curve: Further International Evidence," *Journal of Monetary Economics*, 7, 239-50.

Attfield, C., and N. Duck (1983), "The Influence of Unanticipated Money Growth on Real Output: Some Cross Country Estimates," *Journal of Money, Credit and Banking*, 15, 442-54.

Ball, L., and N. Mankiw (1994), "Asymmetric Price Adjustment and Economic Fluctuations," *Economic Journal*, 104, 423, 247-261.

Ball, L., N. Mankiw, and D. Romer (1988), "The New Keynesian Economics and the Output-Inflation Trade-off," *Brookings Papers on Economic Activity*, 1, 1-65.

Banerjee, A., J. Dolado, J. Galbraith, and D. Hendry (1993), *Cointegration, Error Correction, and the Econometric Analysis of Non-stationary Data*, Oxford: Oxford University Press.

Barro R. (1977), "Unanticipated Money Growth and Unemployment in the United States," *American Economic Review*, 67, 101-15.

_____ (1978), "Unanticipated Money, Output and the Price Level in the United States," *Journal of Political Economy*, 86, 549-80.

Collier, P., and J. Gunning (1999), "Explaining African Performance," *Journal of Economic Literature*, XXXVII, 64-111.

Fernandez, R. (1977), "An Empirical Enquiry on the Short Run Dynamics of Output and Prices," *American Economic Review*, 67, 595-609.

Fischer, S. (1977), "Long-Term Contracts, Rational Expectations, and the Optimal Money Supply Rule," *Journal of Political Economy*, 85, 191-205.

Fuller, W. (1976), *Introduction to Statistical Time Series*, London: John Wiley.

Holmes, M.J. (2000), "Does Purchasing Power Parity hold in African Less Developed Countries? Evidence from a Panel Data Unit Root Test," *Journal of African Economies*, 9, 63-78.

Hossain, A., and A. Chowdhury (1998), *Open-Economy Macroeconomics for Developing Countries*, Cheltenham: Edward Elgar.

Johansen S. (1988), "Statistical Analysis of Cointegrating Vectors," *Journal of Economic Dynamics and Control*, 12, 2/3, 231-254.

Jung, W. (1985), "Output-Inflation Trade-Offs in Industrial and Developing Countries," *Journal of Macroeconomics*, 7, 101-13.

Katsimbris, G. (1990a), "Intracountry Evidence on the Lucas Variance Hypothesis," *Economic Inquiry*, 28, 756-73.

_____ (1990b), "Output-Inflation Trade-Offs in Industrial and Developing Countries: A Comment and Additional Evidence," *Journal of Macroeconomics*, 12, 483-99.

Liu, P. (1992), "Purchasing Power Parity in Latin America: A Cointegration Analysis," *Weltwirtschaftliches Archiv*, 128, 662-80.

Lucas, R.E. (1973), "Some International Evidence on Output-Inflation Trade-Offs," *American Economic Review*, 63, 326-34.

Maddala, G., and I. Kim (1998), *Unit Roots, Cointegration and Structural Change*, Cambridge: Cambridge University Press.

Mahdavi, S., and S. Zhou (1994), "Purchasing Power Parity in High Inflation Countries: Further Evidence," *Journal of Macroeconomics*, 16, 403-22.

Mankiw, N.G., and D. Romer (eds.) (1991), *New Keynesian Economics, Volume 1: Imperfect Competition and Sticky Prices*, Cambridge: MIT Press.

McNown, R., and M. Wallace (1989), "National Price Levels, Purchasing Power Parity, and Cointegration: A Test of Four High Inflation Economies," *Journal of International Money and Finance*, 8, 533-45.

Odedokun, M. (1991), "Evidence on Inflation-Output Trade-Off in Developing and Industrial Countries," *Applied Economics*, 23, 731-42.

_____ (1993), "The Impact of Monetary Surprise and Systematic Monetary Policy in Developing Countries: Empirical Results from an International Perspective," *Social and Economic Studies*, 42, 1, 47-74.

Osterwald-Lenum, M. (1992), "A Note with Quantiles of the Asymptotic Distribution of the ML Cointegration Rank Test Statistics," *Oxford Bulletin of Economics and Statistics*, 54, 461-72.

Parkin, M., B. Bentley, and C. Fader (1981), "Some International Evidence on Output-Inflation Trade-Offs: A Reappraisal," in *Development in an Inflationary World*, eds. by Flanders, M. and A. Razin, New York: Academic Press.

Pesaran, H., and Y. Shin (1999), "Long-Run Structural Modelling," Department of Applied Economics Working Paper, University of Cambridge.

Ram, R. (1984), "Further International Evidence on Inflation-Output Trade-Off," *Canadian Journal of Economics*, 17, 523-40.

Sargent, T.J., and N. Wallace (1976), "Rational Expectations and the Theory of Economic Policy," Studies in Monetary Economics, Federal Reserve Bank of Minneapolis, Parts 1 and 2.

Taylor, J. (1980), "Output and Price Stability: An International Comparison," *Journal of Economic Dynamics and Control*, 2, 109-32.

Tsiddon, D. (1993), "The (Mis)Behaviour of the Aggregate Price Level," *Review of Economic Studies*, 60, 889-902.