LONG-RUN PASS-THROUGH FROM THE EXCHANGE RATE TO IMPORT PRICES IN AFRICAN COUNTRIES

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This paper investigates the extent of pass-through from the nominal exchange rate to import prices for a sample of nineteen African countries. The methodology is based on panel data cointegration testing. Using annual data extending back to 1971, long-run pass-through can be best described as a fairly balanced combination of local-currency and producer-currency pricing. However, this paper offers additional insight from a moving window approach that indicates declining long-run pass-through, accompanied by decreasing inflation, occurring since the mid-1990s.

1. INTRODUCTION

An important policy dilemma facing less developed countries (LDCs) is the extent to which movements in the nominal exchange rate translate into movements in domestic inflation. While one might expect a nominal depreciation to be associated with a favourable movement in terms of price competitiveness, there is a potential danger that domestic inflation will increase thereby limiting, or indeed eliminating, any positive gains. Recent studies that have investigated the exchange rate pass through (ERPT) relationship between the effective nominal exchange rate and domestic inflation for various samples of LDCs have provided mixed conclusions on whether or not devaluations are inflationary (see, *inter alia*, Rana and Dowling (1985), Bahmani-Oskooee and Malixi (1992), Deme and Fayissa (1995), Barhoumi (2006)). With exception of Deme and Fayissa, these papers pay limited focused attention to the experience of African countries.

The purpose of this paper is to investigate the role of the nominal effective exchange

rate in the determination of imported goods prices in the case of nineteen African countries. Using effective exchange rate data recently made available by Bahmani-Oskooee and Gelan (2007), we measure the extent of long-run ERPT. For the purpose of investigating this relationship, this paper utilises recent econometric advances by Pedroni (1999, 2001, and 2004) that enable testing for long-run cointegrating relationships using panel data sets. A common problem with time series studies of LDCs is restricted data availability. Normally, a limited time series reduces the power of unit root and cointegration tests thereby making it harder to reject the null hypothesis of non-cointegration against cointegration alternatives. Creating a larger panel data set from the individual time series increases test power and lessens the likelihood that non-cointegration has been accepted because the sample size is small.

In recent years, a number of studies have observed that the response of domestic prices to changes in nominal exchange rates has been both incomplete and varied over time. Indeed, recent research has suggested that high (low) ERPT is associated with high (low) inflation (see, *inter alia*, Taylor (2000), Choudhri and Hakura (2001), Bailliu and Fujii (2004)) and this has been confirmed empirically by Barhoumi (2006) in a study of LDCs. The purpose of this paper is to analyse the changing nature of African ERPT over time. This paper provides a new insight into ERPT through the employment of a moving window approach to the Pedroni methodology in order to assess the extent to which ERPT has declined in recent decades.

The paper is organised as follows. The following section discusses the ERPT literature and considers studies relevant to LDCs. The third section discusses the methodological approach followed in this study. The fourth section describes the data and discusses the results. Drawing on the Bahmani-Oskooee and Gelan (2007) database, annual data for nineteen African countries are employed for the study period 1971 onwards. Evidence of complete ERPT is sparse and there is evidence of a significant decline in ERPT in Africa since the mid-1990s. The final section concludes.

2.   LITERATURE

The traditional ERPT literature is concerned with pass-through from the exchange rate to import prices stressing the role of market power and price discrimination in international markets (pricing to market). Goldberg and Knetter (1997) survey a literature where import price pass-through is essentially determined by microeconomic factors such as demand elasticities and market structure. Much of this work builds almost exclusively on the concept of market segmentation with discussion set in an oligopolistic framework with imperfect competition and third degree price discrimination.[1] Against this background, the debate over the nature of pass-through has

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[1] Given segmented markets, the seminal papers of Krugman (1987) and Dornbusch (1987) spawned

concerned the prevalence of producer-currency versus local-currency pricing of imports.

In the case of LDCs, the econometric evidence on the role played by the effective exchange rate in determining domestic inflation has been mixed. Bautista, using cross-section data, finds that exchange rate depreciations were inflationary in the case of twenty two LDCs over the period 1972-79. Rana and Dowling examine nine Asian LDCs using quarterly data for the period 1973Q4-79Q4 and conclude that depreciation of the nominal effective exchange rate has not been the most inflationary influence on the economy. Boadu and Gyawu (1992) employ single equation techniques and find that devaluation with respect to the US dollar exchange rate was inflationary for a small number of Sub-Saharan African countries over the period 1970-86. Bahmani-Oskooee and Malixi (1992) examine a sample of thirteen LDCs (which includes Egypt as the only African country) using quarterly data for the sample period 1973Q1-85Q4 and find that depreciations are inflationary in nine countries. Deme and Fayissa (1995) examine Granger-causality between domestic inflation and nominal exchange rate depreciation in the cases of Egypt, Morocco, and Tunisia using annual data for the period 1964-93 and find that devaluations are inflationary in the latter two cases. More recent evidence is provided by Alba and Papell (1998) who, using quarterly data for 1979Q1-95Q2, find that effective exchange rate variation affects inflation in their Southeast Asian sample of the Philippines, Malaysia and Singapore though the effects are small in the latter two cases. Finally, Barhoumi (2006) investigates the exchange rate pass-through into import prices in a sample of 24 developing countries (comprising 10 African countries) over the period from 1980 to 2003. Using tests for cointegration and non-stationary panel estimation techniques, long run homogeneity of pass-through rates across countries is rejected.

An interesting direction in the literature has recently argued that the extent of ERPT has declined with the transition towards a credible low inflationary environment. Against a background of staggered price setting and monopolistic competition, Taylor (2000) presents a dynamic general equilibrium open economy model where firms set prices for several periods in advance, but their prices respond more to price increases (due to exchange rate depreciation or other sources) if cost changes are perceived to be more persistent. Regimes with higher inflation appear to have more persistent costs and will therefore increase the extent of ERPT with firms experiencing greater pricing power. Choudhri and Hakura (2001) emphasise a similar channel to Taylor (2000) where ERPT is reflected in the expected effects of monetary shocks on current and future costs, while

oligopolistic models based on the variations in mark-ups in response to exchange rate changes. Another line taken in the literature has its roots in Baldwin (1988), Dixit (1989) and Baldwin and Krugman (1989). This line of literature emphasises hysteretic effects arising from the sunk costs of entering a market that firms cannot recoup when they leave the market. A third direction in the literature concentrates on institutional settings such as the effects of non-tariff barriers or the role of multinational corporations and intra-firm trade (see Menon (1995, 1996)).

Devereux and Yetman (2003) argue that the frequency of price changes of importing firms declines with the degree of credibility attached to monetary policy. Finally, Devereux and Yetman (2005) develop a theoretical framework to show how sticky prices represent a key determinant of exchange rate pass-through to consumer prices.

A number of empirical studies indicate a reduction in ERPT. For example, event studies by Cunningham and Haldane (1999) show a remarkably small pass-through of exchange rate changes to retail prices in the cases the UK, Sweden and Brazil. Using a cross-sectional approach applied to various samples of countries, Choudhri and Hakura (2001) and Devereux and Yetman (2003) conduct a two-stage methodological approach. In the first stage, the ERPT coefficient is estimated for each country using time-series data. The second stage entails regressing the ERPT coefficients against explanatory variables that include inflation. Using data for the Bretton Woods period, these studies find that inflation significantly explains the differences in the ERPT coefficients. Gagnon and Ihrig (2002) use a similar two-stage approach for industrialised countries over the period 1971-2000. However, they subdivide their study period on the basis of inflationary experience and find that ERPT declined in the regime of low inflation. In most candidate countries the high inflation environment of the early 1990s gradually changed into a single-digit inflation rate episode. Consequently, we might expect this development to influence the pass-through relationship.

Finally, Campa and Goldberg (2005) provide cross-country and time series evidence on the extent of exchange rate pass-through into the import prices of 23 OECD countries. They find compelling evidence of partial pass-through in the short run, especially within manufacturing industries. Over the long run, producer currency pricing is more prevalent for many types of imported goods. Countries with higher rates of exchange rate volatility have higher pass-through elasticities, although macroeconomic variables have played a minor role in the evolution of pass-through elasticities over time. Indeed, Campa and Goldberg find that dramatic shifts in the composition of country import bundles have been an important influence on pass-through changes.

Other approaches to the analysis of OECD data include McCarthy (2000) who uses a vector autoregression model to show a decline in exchange rate pass-through to consumer prices for all nine of the OECD countries examined in the period 1983-98 compared with the period 1976-82. According to these estimates, pass-through declined by 50% or more in the US, UK, France, and Japan and by a smaller amount in Germany, Belgium, Netherlands, Sweden, and Switzerland. In a further study, Bailliu and Fujii (2004) employ a generalised method of moments panel data approach rather than cross sectional approach. Using data for eleven industrialised countries over the study period 1977-2001, they confirm the positive relationship between ERPT and inflation.

In the case of LDCs, Barhoumi (2006) finds that countries with a fixed exchange rate and relatively low tariff barriers exhibit a higher long-run exchange rate pass-through into import prices than countries with higher tariff barriers and a floating exchange rate. In addition to this, Barhoumi finds a small difference in long-run ERPT across countries characterized according to varying inflationary regimes. This paper develops and

extends the study of Barhoumi (2006) in a number of important ways. First, there is a focus on African countries rather than on LDCs in general where the sample of African countries is almost doubled. Second, the study period is longer where the 1970s is incorporated into the analysis. Third, this study develops an innovative application of Pedroni panel cointegration techniques through a moving window approach. This approach avoids the imposition of an arbitrary division of the sample into sub-periods and enables us to consider how ERPT has changed over the past few decades.

3.   METHODOLOGY

The micro-foundations of pricing behavior by exporters are a useful starting point for understanding the dynamics of exchange rate pass through into import prices. The import prices for any country i (P_{it}^m) are a transformation of the export prices of that country's trading partners (P_{it}^x) using the exchange rate (E_{it}) measured as the domestic currency per unit foreign currency:

$$P_{it}^m = E_{it} P_{it}^x . \tag{1}$$

Export prices, in turn, are a markup ($mkup_{it}^x$) over exporter marginal costs (mc_{it}^x). Using lowercase letters to reflect logarithms, we may therefore rewrite Equation (1) as

$$p_{it}^m = e_{it} + mkup_{it}^x + mc_{it}^x . \tag{2}$$

We may further allow markups to have both a country-specific fixed effect and a component that is sensitive to macroeconomic conditions, expressed for simplicity as a function only of the exchange rate,

$$mkup_{it}^x = \phi_i + \Phi_i e_{it} . \tag{3}$$

Exporter marginal costs can be specified as rising with export market wages (w_{it}^x) and destination market demand conditions (y_{it}) as

$$mc_{it}^x = c_{0i} y_{it} + c_{1i} w_{it}^x . \tag{4}$$

Substituting Equations (3) and (4) into (2) means that import prices can be written in general form as

$$p_{it}^m = \phi_i + \left(1 + \Phi_i\right)e_{it} + c_{0i}y_{it} + c_{1i}w_{it}^x. \tag{5}$$

This structure permits exchange rate pass-through, represented by $\beta_i = \left(1 + \Phi_i\right)$, to depend on the structure of competition in the economy. This is consistent with the large literature on explaining cross-sectional differences on exchange rate pass through, which has been summarized by Dornbusch (1987) and Marston (1990), among others, and supported empirically by Knetter (1993) and Yang (1997). This structure also has a direct analogy in the discussion of producer- versus local-currency pricing. If $\Phi_i = 0$, producer-currency pricing takes place; if $\Phi_i = -1$, local-currency pricing does, and exporters fully absorb the fluctuations in exchange rates in their own markups.

We can capture the arguments of Equation (5) through a log linear regression specification similar to that tested throughout the exchange rate pass-through literature:

$$p_{it} = \alpha_i + \delta_i w_{it} + \beta_i e_{it} + \varphi_i y_{it} + \varepsilon_{it}, \tag{6}$$

where p_{it} denotes local currency import prices, e_{it} is the exchange rate, w_{it} is a primary "control" variable representing exporter costs relevant to country i, and y_{it} is a vector of other controls, including the real GDP of the destination market.

A necessary but not sufficient condition for long-run ERPT is that p_{it}, w_{it}, e_{it} and y_{it} are cointegrated. The procedure for computing the test statistics for panel data cointegration involves estimating the hypothesized cointegration regression described in (6) and using the residuals to estimate the appropriate autoregression. Pedroni advocates two statistics both based on a group-mean approach. *Group PP* is non-parametric and analogous to the Phillips-Perron t statistic and *Group ADF* is a parametric statistic and analogous to the ADF t statistic.[2] These two statistics are referred to as *between-dimension* statistics that average the estimated autoregressive coefficients for each country. Under the alternative hypothesis of cointegration, the autoregressive coefficient is allowed to vary across countries. This allows one to model an additional source of potential heterogeneity across countries.[3] Following an appropriate standardization, both of these statistics tend to a standard normal distribution as $N, T \to \infty$ diverging to negative infinity under the alternative hypothesis and consequently, the left tail of the normal distribution is used to reject the null hypothesis of non-cointegration.

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[2] This latter statistic is analogous to the Im *et al*. (2003) test for a panel unit root applied to the estimated residuals of a cointegrating regression.

[3] Pedroni also proposes four *within-dimension* statistics (*panel v, panel* ρ, *panel t* and *panel ADF*) that effectively pool the autoregressive coefficients across different countries during the unit root tests. In these tests, a common value for the autoregressive coefficient is specified under the alternative hypothesis of cointegration.

Following Pedroni (2001), an FMOLS procedure can be employed to obtain the panel data estimates for β_i. Using a dynamic modelling procedure results in a more powerful test for cointegration as well as giving generally unbiased estimates of the long-run relationship and standard t-statistics.[4] FMOLS amounts to the application of non-parametric adjustment to the OLS estimates of both the long-run parameter β_i and associated t-statistic, on account of any bias due to autocorrelation or endogeneity bias that shows up in the OLS residuals (Phillips and Hansen (1990)).

Following on from (6), let $\xi_{it} = \left(\hat{\mu}_{it}, \Delta e_{it}\right)'$ be a stationary vector comprising the estimated residuals and the differences in the nominal effective exchange rate. Also, let $\Omega_i = \lim_{T \to \infty} E\left[T^{-1}\left(\sum_{t=1}^{T} \xi_{iT}\right)\left(\sum_{t=1}^{T} \xi_{iT}\right)'\right]$ be the long-run covariance for this vector process which can be decomposed into $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i'$ where Ω_i^0 is the contemporaneous covariance and Γ_i is a weighted sum of autocovariances. Pedroni shows that the group mean panel FMOLS estimator is given as

$$\hat{\beta}_{GFM}^* = N^{-1} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \left(e_{it} - \bar{e}_i\right)^2\right)^{-1} \left(\sum_{t=1}^{T} \left(e_{it} - \bar{e}_i\right) p_{it} - T\hat{\gamma}_i\right), \tag{8}$$

where $p_{it} = \left(p_{it} - \bar{p}_i\right) - \dfrac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}} \Delta e_{it}$ and $\hat{\gamma}_i \equiv \hat{\Gamma}_{21i} + \hat{\Omega}_{21i}^0 - \dfrac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}}\left(\hat{\Gamma}_{22i} + \hat{\Omega}_{22i}^0\right)$. The between-dimension estimator is calculated as $\hat{\beta}_{GFM}^* = N^{-1} \sum_{i=1}^{N} \hat{\beta}_{FM,i}^*$ where $\hat{\beta}_{FM,i}^*$ is the conventional FMOLS estimator applied to the ith member of the panel. The associated t-statistics are calculated as

$$t_{\hat{\beta}_{GFM}^*} = N^{-0.5} \sum_{i=1}^{N} t_{\hat{\beta}_{FM,i}^*} , \tag{9}$$

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[4] Pedroni (2000) discusses the asymptotic properties of the group mean dynamic OLS (DOLS) estimator and shows how asymptotic normality is applicable. It should be noted that DOLS estimators always have the same asymptotic properties as fully modified OLS estimators. This is because the only difference between them is in terms of how the nuisance parameters are treated (parametrically versus non-parametrically). Therefore, the asymptotic properties for the between-dimension group mean DOLS estimator are identical to those of the between-dimension group mean FMOLS estimator.

where $t_{\hat{\beta}^*_{FM,i}} = \left(\hat{\beta}^*_{FM,i} - \beta_0\right)\left(\hat{\Omega}^{-1}_{11i}\sum_{t=1}^{T}\left(e_{it} - \bar{e}_i\right)^2\right)^{0.5}$.

In the empirical analysis, we focus on the between-dimension panel FMOLS tests. Pedroni (2001) highlights several advantages over the within-dimension approach. In particular, it is argued that the between-dimension approach allows for greater flexibility in the presence of heterogeneity across the cointegrating vectors where β_i is allowed to vary. Under the within-dimension approach, β_i would be constrained to be the same value for each country under the alternative hypothesis. Also, under the within-dimension approach the point estimates of the between dimension estimator can be interpreted as the mean value of the cointegrating vectors which is helpful in interpreting the results. Finally, it can be shown that the between-dimension estimator suffers from lower small-sample size distortions than is the case with the within-dimension estimator.

4. DATA AND RESULTS

This study incorporates a sample of nineteen African countries. These are Burkina Faso, Burundi, Cameroon, Cote d'lvoire, Egypt, Ethiopia, Gabon, Ghana, Kenya, Madagascar, Mauritius, Morocco, Nigeria, Rwanda, Senegal, Seychelles, South Africa, Tanzania and Togo. Most African countries peg their currencies to a major currency or a basket of currencies.[5] Despite the reluctance of central banks to facilitate depreciation of the nominal exchange rate in a regime of managed floating (Hossain and Chowdhury (1998)), nominal effective exchange rates will still vary so long as major currencies fluctuate against each other. In this respect, effective exchange rates may exhibit the same characteristics as the major currencies (Warner and Kreinin (1983)). While the International Monetary Fund constructs and publishes the index for all industrial and some newly industrialized countries, the African nations receive very little attention. Bahmani-Oskooee and Gelan (2007) have addressed this concern by providing quarterly indices of real and nominal effective exchange rates over the 1971Q1-2004Q3 period for twenty one African countries. Using this data set, the nominal effective exchange rate data are expressed as the domestic price of foreign currency and are in natural log form (e_{it}). With regard to the other variables in Equation (6), data for log import prices (p_{it}) and log real GDP (y_{it}) are only available annually and are taken from the World Bank data base. Data limitations regarding these series mean that Niger and Sierra Leone are excluded from the original list of countries in the Bahmani-Oskooee and Gelan sample. Finally, we follow Campa and Goldberg (2005) and proxy exporter costs relevant to

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[5] See, for example, various issues of *International Finance Statistics*.

each African country (w_{it}) by taking the log real effective exchange rate and subtracting both the log nominal effective exchange rate and the log domestic consumer price index. This provides a measure of trading-partner costs (over all partners of the importing African country), with each partner weighted by its importance in the importing country's trade.

Table 1. Panel Data Unit Root Tests

Variable	1971-2003
p_{it}	-0.882
y_{it}	-1.244
e_{it}	-1.141
w_{it}	-1.479

Notes: These are panel data unit root tests advocated by Im *et al*. (2003). These statistics tend to a standard normal distribution as $N,T \to \infty$. Common time dummies are included and individual lag lengths are based on the Akaike information criterion. In all cases, the null of joint non-stationarity is accepted at the 5% significance level with a critical value of -1.64.

Before reporting tests for a cointegrating relationship between the variables included in Equation (6), Table 1 reports Im *et al*. panel data unit root tests for each of the variables. In each case, we are unable to reject the null hypothesis that all series are non-stationary in favour of the alternative hypothesis that at least one series from the panel is stationary. Table 2 reports the Pedroni cointegration tests. These tests include time-specific dummies to allow for the possibility that the residuals are correlated across countries. The null of non-cointegration for the panel is comfortably rejected at the 5% significance level or better by both the *Group PP* and *Group ADF* tests in favour of the alternative of cointegration for all countries. Table 3, reports the FMOLS estimates of Equation (6). An examination of the estimated long-run pass through coefficient β^*_{GFM} indicates a long-run elasticity of 0.61 that is both significantly different from zero and unity with *t*-statistics of 28.02 and -15.53 respectively. This study confirms $0 < \hat{\beta}^*_{GFM} = (1 + \Phi_i) < 1$ thereby rejecting the pure local-currency pricing (where $\hat{\beta}^*_{GFM} = 0 \Rightarrow \Phi_i = -1$) and producer-currency pricing (where $\hat{\beta}^*_{GFM} = 1 \Rightarrow \Phi_i = 0$) scenarios. In the case of African LDCs, pricing would appear to be a hybrid of these cases which is slightly more disposed towards producer currency pricing.[6] This finding may be contrasted with Barhoumi (2006) whose implied long-run pass through elasticity of 0.7 for the panel comprising 24 African and non-African LDCs suggests a greater

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[6] Further tests led to the rejection of the null hypothesis $\beta^*_{GFM} = 0.5$ with an associated t statistic of 6.25.

leaning towards producer currency pricing. A further feature to note about this result is the evidence that $\hat{\delta}^*_{GFM} = 1$ which suggests that changes in exporter costs relevant to African countries are fully passed through to import prices in the long-run.

Table 2. Panel Data Cointegration Tests

Test	1971-2003
Group PP	-3.592***
Group ADF	-2.224**

Notes: These are the Pedroni tests for panel cointegration (discussed in Pedroni (1999, 2004)) between $p_{it}, w_{it}, e_{it}, y_{it}$. These estimates include common time dummies. Individual lag lengths are based on the Akaike information criterion. These statistics tend to a standard normal distribution as $N,T \to \infty$. *** and ** denote rejection of the null of non-cointegration at the 1 and 5% significance levels with critical values of -2.33 and -1.64 respectively.

Table 3. FMOLS Group Mean Estimation

	1971-2003
$\hat{\delta}^*_{GFM}$	1.21 (23.71) [-0.67]
$\hat{\beta}^*_{GFM}$	0.61 (28.02) [-15.53]
$\hat{\varphi}^*_{GFM}$	0.24 (1.59) [-12.91]

Notes: This table reports fully modified OLS (FMOLS) panel data group mean estimates using the Pedroni panel data cointegration methodology. Each slope estimate is accompanied by two *t*-statistics. The *t*-statistics in parentheses are based on the null of a zero slope. The *t*-statistics in square brackets are based on the null of a minus unity slope. All *t*-statistics tend to a standard normal distribution as $N,T \to \infty$.

In a new line of investigation, we move on to consider how ERPT has changed over time. As highlighted earlier, the existing literature has pointed towards a significant decline in ERPT for developed countries, but little has been said about LDCs in general or African countries in particular. To address this issue, we adapt the Pedroni methodology towards testing for cointegration within a twenty-year moving window panel. This provides panel estimates of ERPT starting with the window 1971-90 and finishing with 1984-2003. Figure 1 reports the moving window *Group PP* and *Group ADF* cointegration tests and indicates that import prices, the nominal effective exchange rate, GDP and foreign costs have remained cointegrated over virtually the entire study

period. Figure 2 reports the moving window estimates for β^*_{GFM}. Here we see a decline in value with African import prices becoming less sensitive to movements in the nominal effective exchange rate over time. Despite the fall in β^*_{GFM}, the moving window estimates of the *t* statistics reported in Figure 3 still reject the null of zero pass through and local currency pricing throughout.

Pedroni Group Statistic
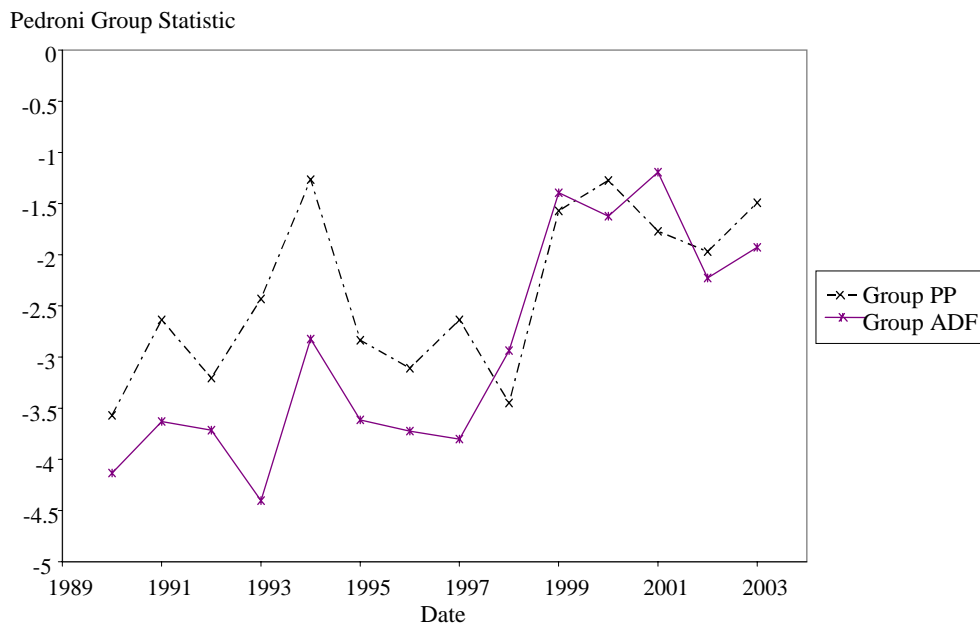
Notes: These are the Pedroni tests for panel cointegration conducted for a twenty-year moving window. Each test is distributed as asymptotically normal as $N,T\rightarrow\infty$ with a 5% critical value of -1.64 against the null hypothesis of non-cointegration for all panel members.

**Figure 1.**   Moving Window Panel Cointegration Tests

An examination of Figure 2 suggests that a positive association between twenty-year average inflation for the panel and ERPT is generally observed in the case of African countries. In contrast to Barhoumi (2006), who finds little difference in long-run exchange rate pass-through across inflation regimes, the results presented here apply specifically to African countries and allow for changes in mean inflation rate over the study period. While there are episodes during where inflation and ERPT have moved in opposite directions, most notably 1990-93, 1996 and 2000, the generally positive association can still be observed particularly from the 1997 onwards. Nonetheless, the interaction between inflation and ERPT is probably more complex than hitherto thought. For example, when considering the long-run adjustment of prices, one can point to the

presence of transactions costs that inhibit international goods arbitrage, product markets that are characterised by extensive government intervention through taxation, price setting and public trading monopolies (Collier and Gunning (1999)).
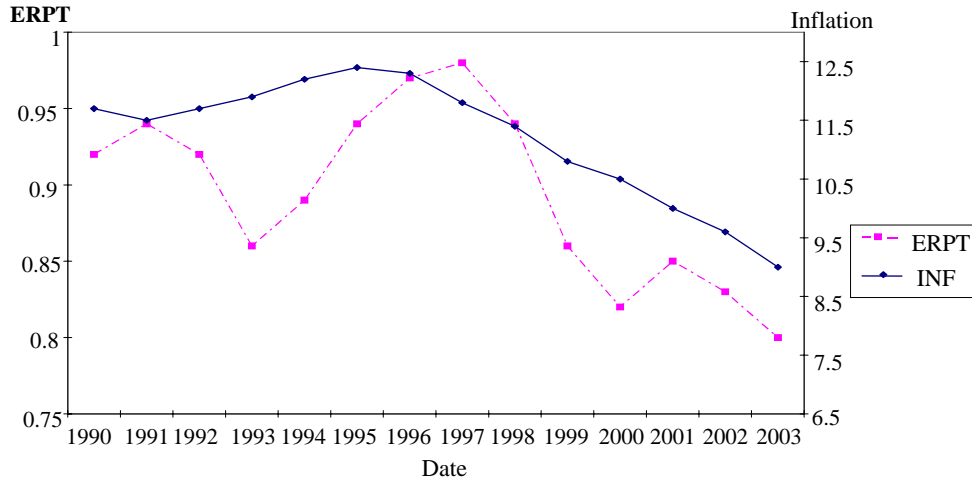
Notes: ERPT refers to the panel FMOLS estimate $\hat{\beta}^*_{GFM}$ for Equation (6) obtained from a twenty-year moving window. INF is average inflation for the panel over the previous twenty years.

Figure 2. Moving Window ERPT and Inflation
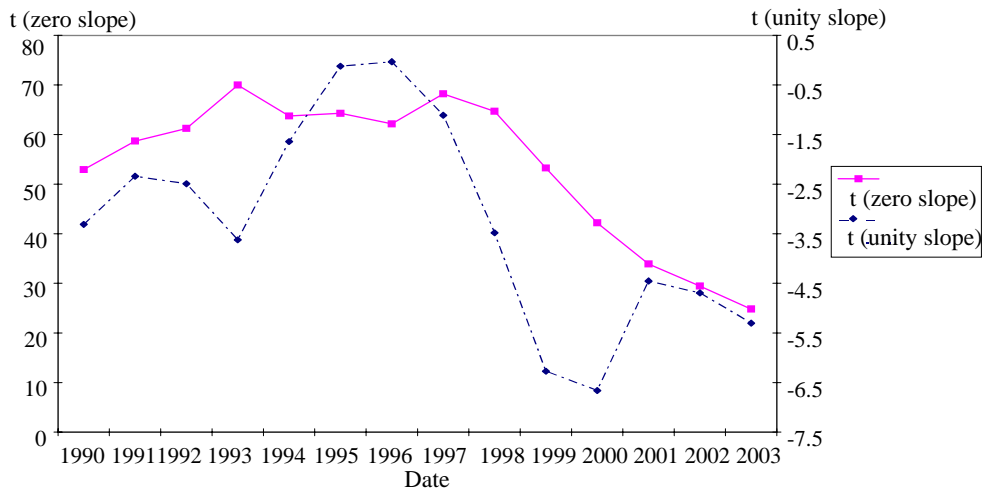
Notes: The calculations of the panel *t* statistics for $\hat{\beta}^*_{GFM}$ are based on a twenty-year moving window. All *t*-statistics tend to a standard normal distribution as $N,T \rightarrow \infty$.

Figure 3. Moving Window Tests of Significance

5.   SUMMARY AND CONCLUSION

Using annual data for a panel comprising nineteen African countries over a study period spanning four decades, we find that the extent of long-run exchange rate pass through lies between the scenarios of local-pricing versus producer-pricing behaviour. However, the new evidence presented here suggests that the degree of pass through emanating from a change in the nominal exchange rate has declined over time and is probably about one-quarter less than what is was in the mid-1990s. This finding has important implications for the inflationary effects of nominal devaluations and macroeconomic independence. For African countries, a declining ERPT coefficient points towards increasing emphasis on local currency pricing behaviour. It also implies a greater degree of macroeconomic independence where central authorities may worry less than was previously the case about the long-run inflationary impact of devaluations in the nominal exchange rate. A number of avenues for future research lead on from this particularly with regard to modeling and quantifying the potential explanations that are chiefly responsible for the decline in African pass through.

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