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Abstract
Plans are far advanced to form a second monetary union, the West African Monetary Zone (WAMZ), in Africa. While much attention is being placed on convergence criteria and preparedness of the five aspiring member states, less attention is being placed on how the dynamics of inflation in individual countries are (dis)similar. This paper aims to stimulate debate on the long term sustainability of the union by examining the dynamics of inflation within these countries. Using Fractional Integration (FI) methods, we establish that some significant differences exist among the countries. Shocks to inflation in Sierra Leone are non mean reverting; results for The Gambia, Ghana and Guinea-Bissau suggest some inflation persistence, despite being mean reverting. Some policy implications are discussed and possible outstanding policy questions are raised.

Keywords: Inflationary shocks, fractional integration, stationarity, West Africa, Monetary unions.
JEL Codes: C14, E31, E58

1. Introduction

In December 2000, heads of state and government of The Gambia, Ghana, Guinea, Nigeria and Sierra Leone signed a treaty to create a second monetary union, the West African Monetary Zone (WAMZ).[1] The ultimate objective is to establish a West African Central Bank, and a single currency, the *eco*. Entry into the union was made conditional on satisfying convergence criteria, among which is the attainment of single digit inflation.[2] However, macroeconomic indicators show that countries within the WAMZ are at different stages in their business cycle. At the last quarter of 2008 where data is available, the average annual inflation rate was 10%, but there is a great deal of variance. Ghana and Nigeria, at 17.6% and 14.8% respectively, exceeded the average by wide margins, while the lowest rates occurred in Sierra Leone (9%) and The Gambia (6.6%). At the same time GDP growth has averaged 3.9%. Heterogeneity in inflation rates across countries is useful in evaluating nominal rigidities relevant for the design of monetary policy. Most theoretical models, take as given that inflation is stationary. The seminal overlapping wage contract model of Phelps (1978) and Taylor (1980), posit that prices are sticky, and inflation rate is so flexible that monetary policy can drive a positive rate of inflation to zero with virtually no loss of output. However, forward-looking models are deemed inconsistent with empirical evidence of significant inflation persistence (e.g., see Ball, 1991; Fuhrer and Moore, 1995). Accordingly, a number of models have added backward-looking elements to enhance the degree of inflation persistence and to provide a better fit with aggregate data. Lags of inflation are typically introduced by postulating some form of price indexation (see Christiano *et al*. 2005) or rule-of-thumb behavior (e.g. Gali and Gertler 1999). Thus, the degree of persistence in inflation

[1] **Dateline:** West Africa already has a monetary union involving former French colonies, the *Communauté Financière Africaine* (CFA) zone (Benin, Burkina Faso, Côte d'Ivoire, Guinea-Bissau, Mali, Niger, Senegal, and Togo). The antecedents of a monetary union dates back to 1987 when heads of states of the Economic Community of West African States (ECOWAS) adopted the monetary cooperation programme to accelerate the process of integration. Following lack of progress in achieving, a harmonized monetary system the Lomé Summit of 1999 put forward a two-track approach to integration: a second monetary zone involving non-member of West African Economic and Monetary Union (WAEMU), which is part of the CFA zone. In the spirit of Lomé Summit WAMZ envisioned a full-blown monetary union by 2003. There have been several failed attempts since then following the failure of member states to achieve the various conversion criteria (see footnote 2).

[2] Other convergence criteria include fiscal deficit/GDP ratio of less than 4%; central bank financing of deficit to be less than 10% and gross external reserves of about 3 months of import cover.

has potentially important implications, not least in theoretical modeling in monetary economics and in applied macroeconomics, but also more generally, the real cost inflation imposes on the economy. For countries in a monetary union, this is even more relevant as it has welfare implications in the sense of informing the adjustments process within the union. Some authors have provided empirical evidence establishing the relationship between monetary policy and well-being of the poor (e.g. Romer and Romer 1998), Easterly and Fisher 2000), Fielding 2004). West African countries, typically dominant raw material exporters, are susceptible to terms of trade shocks. This in turn raises a number of interesting policy conundrums. For instance, how should a potential West African Central Bank conduct monetary policy in a variegated environment of different inflation rates and output gaps? Moreover, how should policymakers stimulate growth in different countries without jeopardizing the goal of price stability? Should monetary policy rule be fully optimal or geared towards output gap stabilization? Answers to these questions depend largely on the speed with which inflation returns to baseline after a shock. Interestingly, the issue of inflation rate dispersion and convergence has played only a minor role in the literature in WAMZ. To the best of our knowledge, the existing literature either concentrates on the economics of monetary unions in general such as the work of Mason and Debrun (2005), and/or address inflation persistence in disaggregated prices, such as Coleman (2010). The former, for instance, argues that pressure exerted by the ruling elite on central banks to extract seigniorage, and the inability of authorities to precommit to price stability affect the incentives of fiscally heterogeneous countries to form a currency union. For West African countries, Mason and Debrun (2005) found fiscal heterogeneity appears critical in shaping regional currency blocs that would be mutually beneficial for all their members. Coleman (2010) examines inflation persistence in food and non-food prices for 12 *Communauté Financière Africaine* (CFA) member states. He argues that both inflation series are characterized by mean-reversion and finite variance, with significant asymmetries in inflation persistence across member states in both sectors and highlights some possible difficulties this is likely to pose for policymaking.

In this paper, we contribute to this literature by focusing exclusively on WAMZ by examining the dynamics of annual changes to inflation using standard unit root tests, and then exploring the mean-reverting behavior using Fractional Integration methods before discussing policy implications of the results.

The rest of the paper is constructed as follows. Section 2 presents the data and reviews the methodologies. Section 3 presents the empirical estimates. Section 4 concludes with a summary of the evidence and some policy implications of our results.

2. Data and econometric methodology

2.1 Data

We analyze monthly data on percentage change in the annual consumer price index (CPI) for each of the five candidates of membership to the (WAMZ) monetary union. The annual percentage changes in CPI have been obtained from the International Financial statistics (IFS) database of the IMF. The summary statistics are shown in Table 1, where we observe great variability in inflation rates across countries. Inflation rates for Ghana, Guinea-Bissau and Sierra Leone appear to have been more volatile over the sample period and the mean inflation ranged from a low of 8.5% to 30.9%

First, we apply a set of panel unit root tests. We note that although these tests take into account cross-sectional information, it is not possible to distinguish which series are $I(0)$

when the null is rejected. Thus, we apply tests proposed by Levin, Lin and Chu (2002) (LLC), Im, Pesaran and Shin (2003) (IPS), Maddala and Wu (1999) and Choi (2001) (MWC).

Table 1: Summary Statistics

Country	Sample	No. of observations	Mean	Standard deviation	Minimum	Maximum
Gambia	1962:1 – 2010:1	577	8.506	10.352	-10.910	75.642
Ghana	1964:3 – 2009:6	544	30.933	31.727	-12.085	174.144
Guinea-Bissau	1987:2 – 2010:2	277	28.165	32.309	-18.231	136.188
Nigeria	1961:1 – 2010:2	590	16.921	17.331	-7.812	89.566
Sierra Leone	1987:10 – 2009:9	264	30.255	34.538	-21.756	145.707

LLC impose a common unit root under the null hypothesis, against the alternative of individual stationarity, whereas the latter allow for individual stationarity under the alternative hypothesis. This implies a less restrictive framework since in the former case the assumption of a common unit root under the null, or general stationarity under the alternative, may be too strong.

LLC consider the following Augmented Dickey Fuller (ADF) regression for panel data:-

$$\Delta y_{it} = \rho y_{i,t-1} + \sum_{L=1}^{p_i} \theta_{iL} \Delta y_{it-L} + \alpha_{mi} d_{mt} + \varepsilon_{it} \qquad (1)$$

where d_{mt} is a vector of deterministic components, so that d_{it} implies no deterministic component, d_{1t} incorporates a drift and d_{3t} includes a constant and a time trend. The null hypothesis is formulated as $\rho = 0$, against the alternative $\rho < 0$.

IPS base their test on the assumption of different autoregressive parameters (ρ) for every individual in Equation (1) i.e. ρ_i . Hence, the null hypothesis is formulated as $\rho_i = 0$ for all *i*, against the alternative that $\rho_i < 0$ for $i = 1,2,\ldots,N_1$ and $\rho_i = 0$ for $i = N_{1+1},\ldots,N$.

An alternative approach is followed by MWC, who combine the different *p*-values of the individual auxiliary regressions, either for the ADF and Phillips-Perron tests, to obtain the following Fisher (1932)-type test,

$$-2\sum_{i=1}^{N} \ln p_i \rightarrow \chi^2_{2N} \qquad (2)$$

where p_i is the asymptotic *p*-value of a unit root test for each individual *i*. Additionally, Choi (2001) proposes the following test, based on the combination of individual *p*-values:-

$$Z = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1}(p_i) \rightarrow N(0,1) \qquad (3)$$

where Φ is the standard normal cumulative distribution function.

In order to analyze the order of integration of the inflation rates in our target countries we consider two groups of unit root tests: Ng and Perron (2001), which are based on linear models, and Kapetanios, Shin and Snell (2003) (KSS).

On the one hand, Ng and Perron (2001) combine a Modified Information Criterion for the lag length and a Generalized Least Squares method for detrending the data. In particular, they propose the following tests: MZ_a and MZ_t that are the modified versions of Phillips' (1987) and Phillips and Perron's (1988) Z_a and Z_t tests; the *MSB*, that is related to Bhargava's (1986) R_1 test; and, finally, the *MPT* test that is a modified version of Elliot, Rothenberg and Stock's (1996) Point Optimal Test. On the other hand, and within the nonlinear framework, KSS develop a unit root test that takes into account the possibility of a globally stationary exponential smooth transition autoregressive (ESTAR) process under the alternative hypothesis. This makes it possible to characterize the target variable as a two regime process for which the change in regimes is smooth rather than sudden. Therefore, the variable may behave as a stationary process in the outer regime, but a unit root in the inner regime. This implies that the autoregressive parameter gets smaller and the variable tends to revert faster to its fundamental equilibrium the further it deviates from the equilibrium. The unit root hypothesis can be tested against the alternative of a globally stationary ESTAR process using the following regression:-

$$y_t \;=\; \varphi_{t-1} \;+\; \phi y_{t-1}\, F(\theta; y_{t-1}) \;+\; \varepsilon_t\,, \qquad (4)$$

where ε_t is $iid(0,\sigma^2)$ and $F(\theta; y_{t-1})$ is the transition function, which is assumed to be exponential (ESTAR),

$$F(\theta; y_{t-1}) \;=\; 1 \;-\; exp\{-\theta\, y_{t-1}^2\}, \qquad (5)$$

with $\theta > 0$. In practice, it is common to rewrite equation (7) as

$$\Delta y_t \;=\; \alpha\, y_{t-1} \;+\; \gamma\, y_{t-1}(1 - exp\{-\theta\, y_{t-1}^2\}) \;+\; \varepsilon_t, \qquad (6)$$

KSS impose that the variable is a unit root process in the central regime so that $\alpha = 0$, although the process is globally stationary. The null hypothesis $H_0 : \theta = 0$ that the process is a unit root in the outer regime is then tested against the alternative $H_1 : \theta > 0$ of stationarity. However, this test cannot be performed directly over θ, since in practice the parameter ϕ cannot be identified under the null. KSS propose the use of a Taylor approximation for equation (6) of the form:-

$$\Delta y_t \;=\; \beta y_{t-1}^3 + error \qquad (7)$$

Testing $H_0 : \beta = 0$ against $H_1 : \beta < 0$ is equivalent to testing for unit roots in the outer regime in equation (4). Equation (7) may incorporate lags in order to control for autocorrelation in the residuals, which lag length may be chosen by means of information criteria. KSS consider three possibilities regarding the deterministic components in their test: applying the test to the raw data, to the demeaned data and to the demeaned and detrended data. Since we are analyzing current account ratio to GDP against convergence to an equilibrium value, it is applied the KSS test to the demeaned data.

As pointed out by KSS amongst many others, traditional (linear) unit root tests may suffer from important power distortions in the presence of nonlinearities in the data generating process (DGP). If the DGP is nonlinear, traditional unit root tests may point to the non-rejection of the null of unit root, when in fact the series are nonlinear and globally stationary. In our case, let us suppose an inner regime and an outer regime, where the inflation rate may behave in a different manner. That is, for small deviations (inner regime), the monetary authorities may not be interested in applying any kind of policy in order to correct these deviations, given that the costs of those policies may overwhelm the benefits, and the variable may behave as a unit root process. However, when deviations from the fundamental equilibrium are significant, authorities may apply policies to reduce those deviations, and therefore the variable may behave as a mean reverting and stationary process in the outer regime. In this situation, we may observe that the further the variable deviates from the equilibrium value, the faster will be the reversion towards it. This implies an autoregressive parameter, which is dependent on the values of the variable.

2.3 Fractional integration tests

There is ample empirical evidence suggesting that inflation are better characterized by fractional integration (FI) rather than unit root behavior (see among others Baum *et al.* 1999b, Baillie *et al.* 2002, Gadea and Mayoral 2006 and Zagaglia 2009). In the light of such findings, and the policy implications for these potential member states, it is instructive that we also explore the possibility of FI in the inflation data for each country. In order to verify the robustness, we apply a battery of tests aimed at investigating the phenomenon for each country. The non-uniform results obtained from the preliminary unit root tests is further hint, if one needs any, that we investigate the potential for FI behavior in each individual series.

In brief, a stationary stochastic process, say y_t, is said to be a long-memory process or fractionally integrated (FI) if there exist a real number H and a finite constant C such that the autocorrelation function $\rho(\tau)$ decays at the rate $C\tau^{2(H-1)}$ as $\tau \rightarrow \infty$ The fractional degree of integration d is related to the parameter H by the equality $d = H - 0.5$. Table 2 provides a summary of the implications of the parameter values for Fractional Integration.

Table 2: Parameter values and implications for Fractional Integration

d	**Variance**	**Shock duration**	**Stationarity**
$d=0$	Finite	Short-lived	Stationary
$0<d<0.5$	Finite	Long-lived	Stationary
$0.5\leq d<1$	Infinite	Long-lived	Nonstationary
$d=1$	Infinite	Infinite	Nonstationary
$d>1$	Infinite	Infinite	Nonstationary

Source: Tkacz (2001)

For $-0.5<d<0$, the series is called *antipersistent.* The long-memory process can be characterized by the behavior of its spectrum $f(\lambda_j)$ estimated at the harmonic frequencies $\lambda_j=(2\pi j/T)$ where $j=1,2,\ldots,[T/2]$ near the zero frequency:

$$\lim_{\lambda_j \rightarrow 0^+} f(\lambda_j) = C\lambda_j^{-2d}$$

where C is a strictly positive constant.[3]

Summary of specific FI tests

First, we implement four separate tests to investigate long run dependence in each series – the Robust Rescaled Range Statistic (Lo, 1991), the KPSS Statistics (Kwiatkowski *et al.* 1992), the Rescaled Variance Test of Long-Memory (Kokoszka and Leipus, 1998) and a Semiparametric test for $I(0)$ of a time series against fractional alternatives (Lobato and Robinson, 1997). We provide a brief description of the FI statistics we use. Lo proposed a statistic, $T^{-0.5}Q_n$, with

$$Q_T = \frac{1}{\hat{\sigma}_T(q)}\left[\max_{1\le k<T}\sum_{j=1}^{k}\left(X_j - \bar{X}_T\right) - \min_{1\le k<T}\sum_{j=1}^{k}\left(X_j - \bar{X}_T\right)\right] \qquad (8)$$

which incorporates the HAC variance estimator in the denominator of the statistic and is able to detect nonperiodic cycles. Under the null hypothesis of no long-memory, Lo's statistic converges to a Brownian motion based distribution, which is tabulated in his paper. It is, however, extremely sensitive to the order of truncation q and following suggestions by Taqqu *et al.* (1999), we use this statistic with other tests.

Next, we compute the two KPSS statistics, denoted by $KPSS_\mu$ and $KPSS_t$, which are respectively based on the residuals of two regression models: on a constant μ, and on an intercept and a trend t. By denoting the partial sums by S_t i.e. $S_t = \sum_{i=1}^{t}\hat{e}_i$, where $\hat{e}_i$ are the residuals of the regressions, the KPSS statistic is defined as:

$$T^{-2}\sum_{t=1}^{T}\left[{S_t^2}\big/{\hat{\sigma}_T^2(q)}\right] \qquad (9)$$

where $\hat{\sigma}_T^2(q)$ is the HAC estimator of the variance of the residuals and defined as $\hat{\sigma}_T^2(q) = \hat{\gamma}_0 + 2\sum_{j=1}^{q}\left[1 - \frac{j}{1+q}\right]\hat{\gamma}_j, q < T.$ where $\hat{\gamma}_0$ is the variance of the process, and the sequence

[3] For an excellent and comprehensive survey on long-memory, see Beran (1994), Robinson (1994a), Baillie (1996), Baum *et al.* (1999a) and Arize *et al.* (2005).

$\{\hat{\gamma}_0\}_{j=1}^{q}$ denotes the autocovariances of the process up to the order q. The statistic $KPSS_\mu$ tests for stationarity against a long-memory alternative, while the statistic $KPSS_t$ tests for trend-stationarity against a long-memory alternative.

Following this, we make use of the Rescaled Variance Test of Long-Memory Statistic (RVLM), which seeks to improve on the KPSS statistics and is based on the partial sum of the deviations from the mean. The statistic is computed as:

$$RVLM = T^{-1}\left[\frac{\hat{V}(S_1,\dots,S_T)}{\hat{\sigma}_T^2(q)}\right] \qquad (10)$$

where $S_k = \sum_{j=1}^{k}(Y_j - \bar{Y}_n)$ are the partial sums of the observations. Simply, the statistic is, therefore, the sample variance of the series of partial sums $\{S_t\}_{t=1}^{T}$. Importantly though, the RVLM statistic has uniformly higher power than the KPSS, and is less sensitive than the Lo statistic to the choice of q.

The next alternative statistic (*Lobrob*) we compute, proposed by Lobato and Robinson (1997), is also a nonparametric test for $I(0)$ against $I(d)$ and has a t-statistic

$$t = m^{0.5}\left({}^{\hat{C}_1}\!/\!_{\hat{C}_0}\right) \qquad (11)$$

where $\hat{C}_k = m^{-1}\sum_{j=1}^{m}\left(\ln(j) - \frac{1}{m}\sum_{i=1}^{m}\ln(i)\right)^k I(\lambda_j)$

with $I(\lambda) = (2\pi T)^{-1}|\sum_{t=1}^{T} y_t e^{it\lambda}|^2$ being the periodogram estimated for Fourier frequencies $\lambda_j = 2\pi j T^{-1}, j = 1,2,\dots,m \ll [T/2]$ and m, the bandwidth parameter. Under the null hypothesis of an $I(0)$ or *short-memory* time series, the t-statistic is asymptotically normally distributed. Hence, the *Lobrob* statistic has the advantage of being a two-sided test allowing discrimination between $d > 0$ and $d < 0$.[4]

Following these four preliminary tests of $I(0)$ versus $I(\mathrm{d})$ alternatives, we then employ the spectral regression (Geweke and Porter-Hudak, 1983) method to determine the d parameter, which estimates the long-memory parameter d with the following spectral regression:

$$\log\left[I(\lambda_j)\right] = c - d\log\{4\sec^2(\lambda_j/2)\} + \varepsilon_j \qquad (12)$$

where λ_j represents the harmonic frequencies.

3. Results

The results of the tests for the order of integration of the variables are reported in Tables 3 and 4. With the LLC test, we cannot reject the null hypothesis of a common unit root. However, with the rest of the tests we are able to reject the null hypothesis that all the countries' inflation contain a unit root. This implies that for some of them the unit root hypothesis is rejected.

Table 3: Panel unit root tests

Test	***p-value***
LLC	1.0000
IPS	0.0217
ADF-Fisher	0.0480
ADF-Choi	0.0218
PP-Fisher	0.0000
PP-Choi	0.0000

In order to distinguish which countries' inflation rates are $I(1)$ or $I(0)$ we report, in Table 4, the unit root test estimates for each individual country. It appears that the unit root hypothesis

[4] If the *t*-statistic is in the lower fractile of the standardized normal distribution, the series exhibits *long-memory*, whilst if the series is in the upper fractile of that distribution, the series is *antipersistent*.

can be rejected for all the countries in our sample, with the exception of Sierra Leone. For the latter, shocks to inflation will have permanent effects, whereas for the rest, shocks will only have transitory effects and the variable will revert to its long run equilibrium eventually.

In order to take into account the possibility of structural changes in the drift, in Table 5 we display the results of the LS unit root test with two structural changes. The results point to the rejection of the unit root hypothesis in all cases. In the case of Sierra Leone, the existence of structural changes may have prevented the Ng and Perron (2001) and the KSS test from rejecting the null hypothesis.

Table 4: Individual unit root tests

Country	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MP_{T}^{GLS}	$\hat{t}_{NL}$	$\hat{t}_{NLD}$
Gambia	-7.1855*	-1.8901	0.2630*	3.4296*	-1.7379	-2.0569
Ghana	-9.541**	-2.183**	0.2288**	2.5697	-4.5362**	-5.477**
Guinea-Bissau	-0.3092	-0.1947	0.6296	24.6927	-3.6260**	-5.498**
Nigeria	-9.594**	-2.183**	0.2275**	-2.2512**	-2.2349**	-2.5195
Sierra Leone	0.2147	0.3265	1.0373	64.4222	-1.4281	-3.979**

Notes: The order of lag to compute the tests has been chosen using the modified AIC (MAIC) suggested by Ng and Perron (2001). The Ng-Perron tests include an intercept, whereas the KSS test has been applied to the raw data, $\hat{t}_{NL}$ say, and to the demeaned data, $\hat{t}_{NLD}$ say. The symbols * and ** mean rejection of the null hypothesis of unit root at the 10% and 5% respectively. The critical values for the Ng-Perron tests have been taken from Ng and Perron (2001), whereas those for the KSS have been obtained by Monte Carlo simulations with 50,000 replications:

Fractile	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MP_{T}^{GLS}	$\hat{t}_{NL}$	$\hat{t}_{NLD}$
5%	8.100	1.980	0.233	3.170	2.196	2.906
10%	5.700	1.620	0.275	4.450	1.908	2.636

Table 5: LS unit root tests

Country	**TB1**	**TB2**	**Statistic**
Gambia	1985:6	1987:12	-5.97844**
Ghana	1976:4	1984:5	-6.11532**

Guinea-Bissau	1996:2	1998:6	-4.40329**
Nigeria	1992:1	1996:6	-5.92056**
Sierra Leone	1992:11	1999:12	-3.86491**

Notes: The critical values are -3.842 and -3.504 at the 5% and 10% significance levels, respectively, and have been obtained from Lee and Strazicich (2003, Table 2). The symbol ** means rejection of the null hypothesis of unit root at the 5%. The lag length has been obtained by following a general-to-specific approach (10% significance level) from a maximum of 12 lags.

Results from our FI analyses, summarized in Tables 6, support the proposition that standard unit root tests alone that produce $I(0)/I(1)$ outcomes may, indeed, be overly restrictive.

Table 6: Individual test statistics for Fractional Integration

Country	**Fractional Integration test statistic**					
	Lo	$KPSS_\mu$	$KPSS_t$	***RVLM***	***Lobrob***	***GPH***
Gambia	2.6098**	3.9052*	3.8958*	3.7800**	-6.1972*	0.5179
Ghana	2.7753**	4.2558*	3.8672*	4.0643**	-5.7576*	0.6698
Guinea-Bissau	2.8889**	18.534*	1.3890*	3.7632**	-13.313*	0.8337
Nigeria	2.7160**	7.7516*	3.5706*	4.5761**	-6.9639*	0.4575
Sierra Leone	2.3811**	13.308*	1.4641*	3.1268**	-12.143*	1.0289

Notes: *Lo* statistic has been computed using an expansion order of 5, with (0.809, 1.862) being the 95% confidence interval for no long-memory. $KPSS_\mu$ and $KPSS_t$ based on a regression on a constant mu, and on a constant and a time trend t, respectively. Critical values for $KPSS_\mu$ and $KPSS_t$ at 10%, 5% and 1% are 0.347, 0.463, 0.739 and 0.119, 0.146, 0.216 respectively. The 95% critical value for *RVLM* is 0.1869, and entries have been computed for zero truncation lag. *Lobrob* entries have been computed for bandwidth of 50, and critical values follow the standardized normal distribution. *GPH* entry indicates scalar which is Geweke, Porter-Hudak (1983) degree of long memory.

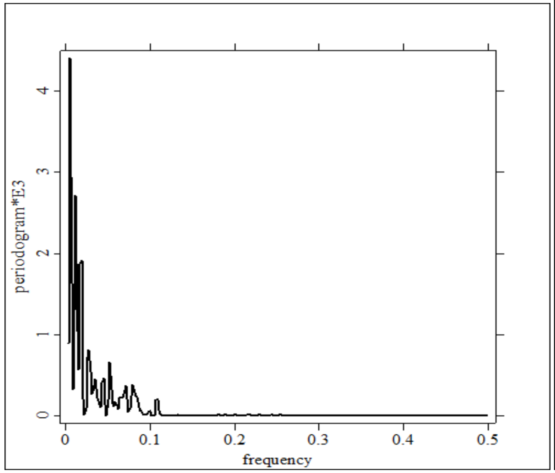
The four initial tests each reject the null hypothesis of no long memory, confirming the importance of FI analyses.[5] Following this, the estimated d parameter becomes crucial and two points are worthy of note here. On the one hand, the estimated d confirms the mean-reverting nature of changes in inflation for four of the countries, except for Sierra Leone which appears to have infinite variance and is not mean reverting. In fact, a casual look at the (raw) periodogram, presented in Figures 2-6, reinforce the point that Sierra Leone's indicates

[5] Here, we are counting the two *KPSS* tests as one.

longer shock duration than the other countries.[6] On the other hand, it also points to some differences in the shock duration or persistence even among the four countries with mean reverting changes in inflation. Such differences are often muted by the standard unit root tests. Only Nigeria appears to have a finite variance; The Gambia, Ghana and Guinea-Bissau although mean reverting, each appear to have infinite variance and are non-stationary. Shocks to the series are more persistent in Guinea-Bissau, then Ghana and then The Gambia. This finding further highlights the importance of investigating FI when testing for unit roots, especially when dealing with countries that envisage implementing a one-size-fits-all monetary policy (*via* a common central bank) in a monetary union.

On the economic front the policy importance of this result cannot be overemphasized, as there may be net gainers and net losers in such a union, increasing the possibility of reneging. Furthermore, the largest economy, Nigeria, which may arguably have more clout in such an economic union, appears to have the least persistence. How sensitive it will be to policies that benefit the majority, but that are less favorable to smaller economies e.g. Sierra Leone, with the proclivity for more persistence, will be an important factor for optimality.

[6] The raw- (as opposed to the log-) periodogram values can be interpreted in terms of variance (sums of squares) of the data at the respective frequency or period.



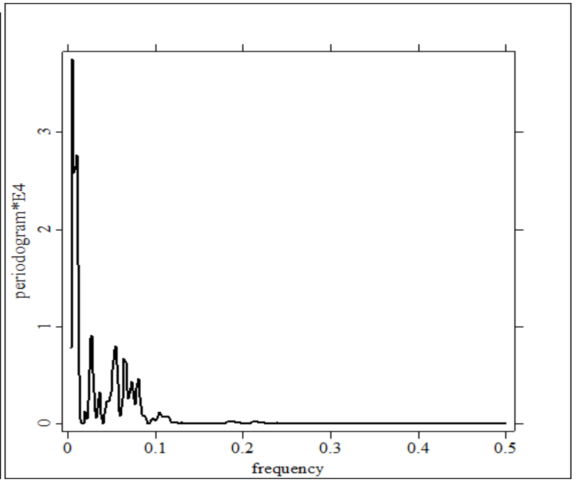


Figure 2: Periodogram for the Gambia

Figure 3: Periodogram for Ghana

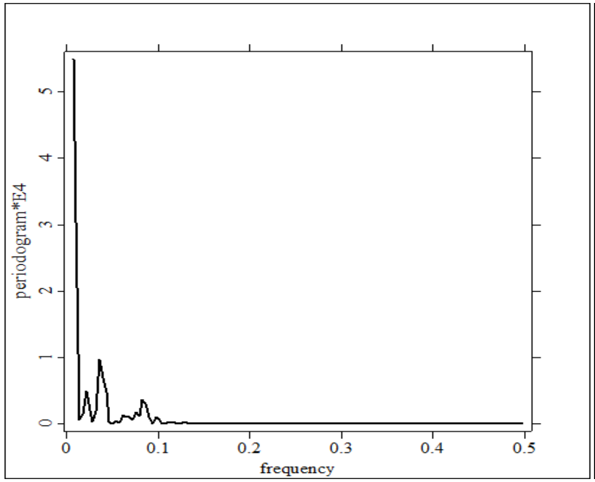


Figure 4: Periodogram for Guinea-Bissau

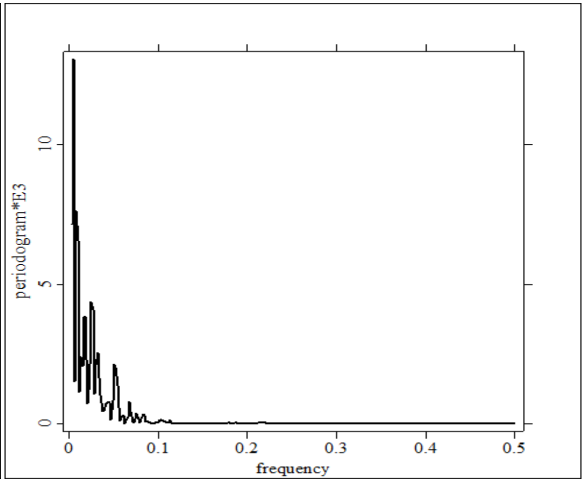


Figure 5: Periodogram for Nigeria

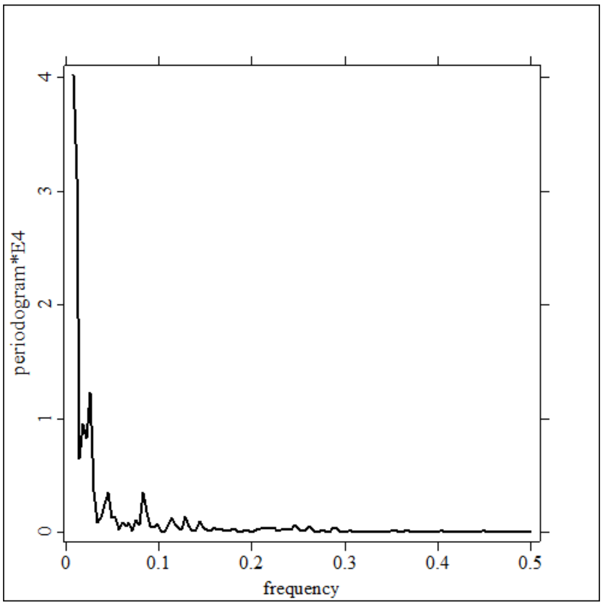


Figure 6: Periodogram for Sierra Leone

4. Policy Implications and Concluding Remarks

In this paper, we empirically investigate the persistence properties of annual change in the CPI of five countries that are working towards forming a second monetary union, the West African Monetary Zone (WAMZ) in Africa. Previous literature have largely focused on the more established CFA Franc zone, and most of those that focus on the WAMZ have focused on the convergence criteria and the state of preparedness of the members – in short, the focus has been on how the proposed union can actually take off. This paper aims to highlight potential policy formulation issues that may seriously hamper the long-term survival of such a union, and thereby stimulate debate on the constitution, the criteria and possibly the buffering measures that should be put in place to support potential net losers when the one-size-fits-all policy measures start to take effect.

Initial results from standard unit root tests, although less informative, suggest that some asymmetries in dynamic behavior of inflation is likely. This view is confirmed by the less restrictive Fractional Integration tests, where the *d* parameter is not restricted to 0 or 1. We find evidence that the dynamics of the annual change in inflation in these countries is best characterized by fractional integration, and we find that d takes values other than 0 or 1. A shock to inflation, possibly due to natural disasters or even central bank policy, is likely to have different implications for the countries due to the differences in the time the shock will take to decay to the baseline. Policy implications of these results are important, keeping in mind that common monetary policy will apply for these countries. First, the assumption that meeting the convergence criteria is in itself the making of the union may be misplaced. Second, with the link between monetary policy and welfare empirically established in previous literature, increased knowledge about the duration of a shock should better inform policymakers' projections about macroeconomic aggregates, and welfare of the population.

Specifically, of the five countries in our sample, the results for Sierra Leone suggest that shocks to inflation will persist longest in that country and is *non*-mean-reverting, and if policymakers are concerned about the detrimental effects of inflation on welfare, as suggested in previous literature, this should be a policy concern. Persistence estimates for Gambia, Ghana and Guinea-Bissau also suggest some inflation persistence, though less than in Sierra Leone, which should also be considered by policymakers. Finally, on the assumption that the size and wealth of Nigeria will imply a central role within the union, our finding that shocks to inflation are least persistent in that country raises the question whether self interest may be problematic. Some outstanding policy questions arise: Will there be a continuous need for monetary support for the countries with a higher proclivity for higher inflation persistence? If there are domestic policy responses aimed at smoothing the path of national consumption in response to inflationary shocks, the implications for both household and national savings should be an important consideration. In effect, how committed will member states be to upholding the merits of the union if they continue to be net losers? Serious *a priori* consideration of these questions should assist in shaping a more formidable, beneficial and sustainable monetary union.

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