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Jemma Dridi and Anh D. M. Nguyen

International Monetary Fund, Bank of Lithuania

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# **Inflation Convergence In East African Countries**

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## **Abstract**

The paper investigates inflation convergence in five East African Countries: Burundi, Kenya, Rwanda, Tanzania, and Uganda, as they aspire to form a monetary union by 2024 under the umbrella of the East African Community. Based on various panel unit root tests, we find that inflation rates in these countries have been converging. An explanation for the convergence is also provided from the perspective of a Global Vector Autoregressive (GVAR) model, which attributes this convergence to a similarity in terms of the nature of shocks affecting EAC countries as well as the role of foreign factors as drivers of inflation given that inflation has been low and less volatile in industrial and emerging countries since the early 1990s.

JEL Classification: C32, C33, E31, F40

Keywords: Inflation, Global VAR (GVAR), Panel Unit Root Tests, Spillovers, East African Community.

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<sup>1</sup> Dridi: International Monetary Fund, 700 19th Street, N.W., Washington, D.C., Email: [jdridi@imf.org](mailto:jdridi@imf.org); Nguyen: Bank of Lithuania, 4 Totorių Street, Vilnius, Lithuania, Email: [anguyen@lb.lt](mailto:anguyen@lb.lt). The views expressed are those of the authors and do not necessarily represent those of the International Monetary Fund and the Bank of Lithuania. We thank Oral Williams for useful comments and suggestions.

## 1. Introduction

The treaty to revive the East African Community (EAC) came into force on July 2000 with the objective of fostering a closer cooperation in political, economic, social, and cultural fields. In November 2013, the five EAC countries, including Burundi, Kenya, Rwanda, Tanzania, and Uganda, signed a protocol outlining their plans for launching a monetary union in 2024. To reap the maximum benefits and minimize costs of a monetary union, member countries need to achieve a sufficient degree of macro-economic convergence, and financial integration among them ahead of the union. Like other regional economic communities elsewhere, EAC countries have put in place macro convergence criteria to be met by each country prior to entry into the monetary union. These convergence criteria were formulated to accommodate the developmental desires of EAC while at the same time continuing to safeguard macroeconomic stability. The focus is on price stability, sustainable fiscal deficit, and maintaining desirable levels of foreign exchange reserves.<sup>2</sup> Given that inflation convergence is one of the key requirements for the succession of a currency union, it is important to understand the dynamics of inflation across the EAC members.

This paper aims to contribute to the discussion on inflation convergence across EAC and its implications for the establishment of a monetary union in the region. First, we test for the existence of inflation rates' mean-reverting behavior, thus allowing us to address whether existing differentials in inflation rates should be a major concern for policymakers. To do so, we use various panel-based unit root tests, because of the known low power of univariate unit root tests, including two generations of the tests with respect to the feature of cross-sectional dependences. Taking into account the cross-sectional dependences is important as ties between EAC economies have been increasing, especially after the Treaty came into force in 2000. Second, we investigate the causes of convergence (or divergence) in inflation, using a novel, recently developed method called Global VAR (see Chudik and Pesaran, 2014). Based on such an approach, we can explicitly account for linkages among economies such that impacts of regional and global shocks on domestic economies (Dees et al., 2007).

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<sup>2</sup> The performance convergence criteria which each of the EAC countries must achieve are: headline inflation of no more than 8 percent; fiscal deficit, including grants of no more than 3 percent of GDP; gross public debt of no more than 50 percent of GDP in Net Present Value terms; and maintenance of official foreign reserves equivalent to no less than 4.5 months of imports (EAC, 2012).

Our results find broad support for inflation convergence in the EAC countries in the post-treaty period. Panel unit root tests suggest that inflation differentials in the five EAC countries are not persistent, implying that inflation rates in these countries have been converging. Such a convergence in inflation rates can be explained, based on the results of the GVAR model, by a similarity in terms of the economic nature of shocks and by a larger role of foreign factors compared with domestic factors in the variations of inflation. Supplemented with the larger role of foreign factors is that inflation has been low and less volatile in industrial and emerging countries since the early 1990s as documented by Helbling, Jaumotte, and Sommer (2006).

The remainder of the paper is organized as follows. Section 2 reviews the literature. Section 3 discusses the development history of the East African Community, the importance of inflation convergence for the region and then describes key features of the EAC countries' inflation. Section 4 presents the panel unit root tests. Section 5 presents the GVAR model and its results. Finally, Section 6 concludes the paper.

## **2. Literature Review**

The participation in a monetary union can help to eliminate currency conversion costs and exchange rate uncertainties between member countries, thereby spurring intra-regional trade, a hypothesis being empirically supported by Rose (2000) and Rose and Stanley (2005). Moreover, by delegating the monetary policy tool to a supranational authority, it helps enhance the credibility of monetary policy by restricting domestic political interference. According to Guillaume and Stasavage (2000), this benefit is potentially important for African countries given the role of common fiscal pressure/dominance in the region.<sup>3</sup> On the other hand, the economic cost for a country that joins a monetary union is the abnegation of using exchange rate and monetary policies to stabilize shocks-induced output and employment fluctuations. According to Mundell (1961), the magnitude of the cost depends on the degree of asymmetry of the shocks to the member countries' economies. Attempts to measure whether the business cycles of the EAC countries have

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<sup>3</sup> Grilli, Masciandaro, and Tabellini (1991) and Alesina and Summers (1993) show that central bank independence likely promotes price stability. Keefer and Stasavage (2000), however, argue that this only occurs under specific institutional and political conditions, particularly the existence of checks and balances within political institutions, which prevent the reversal of legal central bank independence. Guillaume and Stasavage (2000) argue that only few African countries satisfy those conditions; therefore, joining a regional agreement might act as a substitute mechanism to establish credibility.

synchronized, following the main approach being to use the Blanchard and Quah (1989)'s method to identify supply and demand shocks in a VAR framework. Examples include Drummond et al. (2015), Mafusire and Brixiova (2013), Kishor and Ssozi (2011), and Buigut and Valev (2005).

Another important issue, which has received considerable attention within the countries of the European Monetary Union (EMU), but is still scarcely explored in the EAC context, is inflation persistence (see, e.g., Estrada, Galí, López-Salido, 2013; Buseti et al., 2007; Weber and Beck, 2005; Kočenda and Papell, 1997 among others for the EMU).<sup>4</sup> Given the one-size-fits-all monetary policy, persistent differences in inflation pattern among the member countries may cause disparities in real interest rates, leading to unfavorable impacts for some countries in the union. For instance, a country whose economic activity is relatively subdued will probably have lower inflation pressures in comparison to other members. Hence, it will face a relatively high real interest rate, causing more difficulties for economic activities and making inflation become more divergent in the union (Busetti et al., 2007).

We contribute to the literature by investigating inflation convergence across EAC and its implications for the establishment of a monetary union in the region. First, we test for the existence of inflation rates' mean-reverting behavior, thus allowing us to address whether existing differentials in inflation rates should be a major concern for policy-makers, which should be the case if we find no or only very weak indications of mean-reverting behavior. A popular approach to test the mean-reverting behavior is to use standard univariate unit-root tests, i.e. Dickey-Fuller based tests (e.g., Nelson and Plosser, 1982; Charemza et al., 2005). Nonetheless, these tests are known to have low power, i.e. it is difficult to reject the null hypothesis of a unit root when it is in fact false. To overcome such a problem, several methods have been proposed. Among those, using panel-based unit root tests, such as those developed in Levin, Liu and Chu (2002) and Im, Pesaran and Shin (2003), is one of the most popular approach.<sup>5</sup> In this aspect, Kočenda and Papell (1997) and Weber and Beck

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<sup>4</sup> With respect to inflation, the convergence condition for the EMU requires that a country can only join the Union if its inflation rate is no more than 1.5 percentage points higher than the rate of the three best performing member states. Meanwhile, the convergence criterion for the EAC is below or equal 5 percent in core inflation and 8 percent in headline inflation.

<sup>5</sup> Other approaches to investigate inflation- or economic growth- convergence include the fractional integration (e.g. Robinson, 1995, Carcel et al., 2015), the unobserved component model (e.g. Hall and Lagoa, 2014), or distribution dynamics (e.g. Quah, 1996 and Weber and Beck, 2005).

(2005) test inflation convergence within European Union, Cecchetti et al. (2002) investigate price index convergence among US cities.

Several panel unit root tests have been proposed in the literature. The main differences between those tests lie in the homogeneity assumption under the alternative hypothesis, the existence of cross-sectional dependencies and the specification of the cross-sectional dependencies. In general, the literature distinguishes two generations of panel unit root tests based on the feature of cross-sectional dependences (see Breitung and Pesaran, 2005 and Hurlin and Mignon, 2007 for surveys). The first-generation tests assume that all cross-sections are independent. In this generation, there are two different groups: one assumes homogeneity under the alternative, the other allows heterogeneity. However, the assumption of independent cross-sections appears to be too restrictive given the increasing ties between EAC economies, especially after the Treaty came into force in 2000. Therefore, we also use the second-generation tests that take into account cross-unit dependencies by different approaches. In summary, we consider a battery of panel unit root tests in both generations to ensure that our results are not driven by the choice of a certain type of test.

Although conducting unit root tests enables us to assess inflation convergence, it does not reveal what causes convergence (or divergence) in inflation. To address this question, we use a novel, recently developed method called Global VAR (see Chudik and Pesaran, 2014). Based on such an approach, we can explicitly account for linkages among economies such that impacts of regional and global shocks on domestic economies (Dees et al., 2007). The objective is, therefore, to clearly identify both what factors drive inflation and the extent to which inflationary pressures are caused by foreign versus domestic sources.<sup>6</sup> In addition, this approach supplements the convergence test in the sense that if inflation rates among member countries appear to converge, their drivers of inflation should not be too different and vice versa.

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<sup>6</sup> This approach is similar to Nguyen et al. (forthcoming) which uses a GVAR model to examine inflation dynamics in Sub-Saharan Africa, although the EAC countries' inflation is not the focus of the paper. In addition, unlike the Nguyen et al. (forthcoming) paper, we model global variables by the dominant unit as in Chubik and Pesaran (2013) and expand the sample size to 2013Q4. Our paper also contributes to the recent literature which applies the GVAR model to African countries, including Gurara and Ncube (2013), which analyzes the global growth spillover effects on Africa, and Canales-Kriljenko et al. (2014), which discusses the spillovers from global financial variables to economic activity.

### **3. The East African Community**

#### **3.1 History of the EAC**

The East African Community aims at deepening cooperation among its member states in the political, economic, and social domains for their mutual benefits. The EAC came into force in 2000 following its ratification by the original founding three partner states – Kenya, Tanzania and Uganda. Rwanda and Burundi joined the EAC seven years later in 2007, and in March 2016 South Sudan was admitted as the sixth member of the regional bloc.<sup>7</sup> The Customs Union (CU), which is the first protocol underpinning the integration process, was signed in 2005 initially by Kenya, Tanzania and Uganda with Rwanda and Burundi joining in 2009. The CU aims to liberalize intra-EAC trade and promote efficiency in production through facilitating the free movement of goods within the community. The second is the common market (CM) protocol, which was signed in 2010 and aims to form a single area in which there is free movement of goods, people, capital, labor, services, and right of establishment and residence amongst the partner states. The East African Community Monetary Union (EAMU) represents the third stage of integration to maximize the benefits of the single market. The EAMU Protocol was signed in 2013 and ratified by all five partner states in early 2015. It sets out the process, including macroeconomic convergence criteria, and legal and institutional framework for the establishment by 2024 of a single currency.

The EAC has made progress in implementing the CU and CM together with improved macroeconomic management as part of the integration process in recent years has helped EAC partner states macroeconomic performance. The EAC has also made progress in establishing an EAC Monetary Union. The critical areas of harmonization include: monetary and exchange rate policy harmonization, statistic harmonization, fiscal policy coordination and harmonization, financial market coordination, banking supervision and financial stability, harmonization of payments and settlement systems, and cohesive accounting and financial standards. EAC also made the decision to establish the East African Monetary Institute and the East African Central Bank to fulfill these goals. Successful implementation of the proposed monetary union would help promote trade through the enhancement of the payment system for goods and services between the states,

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<sup>7</sup> South Sudan is not covered in this paper due to lack of data.

create a larger regional market and broaden business and trade-related income earning opportunities for the sub-region, support labor mobility, strengthen cooperation, and promote competitiveness and efficiency in production.

### **3.2 The Importance of Inflation Convergence**

As noted in earlier sections, inflation convergence is one of the critical requirements for the suitability of currency unions among different countries. If the member countries experience asymmetric inflation rates, a EAC regional central bank that primarily aims to stabilize inflation across the region will find it challenging to apply a single nominal interest rate. The EAC regional bank's monetary policy will be too tight for a member country with inflation below the regional inflation average; while the regional bank's monetary policy will be too loose for a country with inflation above the EAC average. In other words, countries with below average inflation rates will face above average real interest rates, while those with above average inflation rates will face below average real interest rates.

Inflation convergence is a key indicator of the structural synchronization between countries. Differences in inflation could be due to regional heterogeneities in the relative productivity growth of the tradeable versus then non-tradeable sectors (Balassa-Samuelson effect). Exchange rate movements create different pass-through effect in importing countries. Honohan and Lane (2003, 2004) and Buseti et al. (2007) found that exchange rate fluctuations can have strong effects on inflation.

Achieving inflation convergence across EAC countries is important given weaknesses of traditional adjustment channels to macroeconomic shocks. The main cost of currency unions is the loss of monetary policy independence and the possibility of macroeconomic adjustments through exchange rate movements. This usually raises economic and political tensions, which can be eased if economies can adjust quickly to their long-run equilibrium after a macroeconomic shock. In the short-term, however, there is a tradeoff between inflation and unemployment. The faster economies adjust and return to their long-run equilibrium the better. The speed of adjustment to the long-run equilibrium is higher if there is a higher degree of wage flexibility and/or mobility labor mobility in the region. Notwithstanding recent reforms, rigidities persist in EAC markets and these two conditions are far from being satisfied. Hence, it is very important that EAC exhibits convergent rates of inflation prior to the establishment of an EAC currency union. On the other hand, sharing a common currency and a regional exchange rate, inflation differentials may work

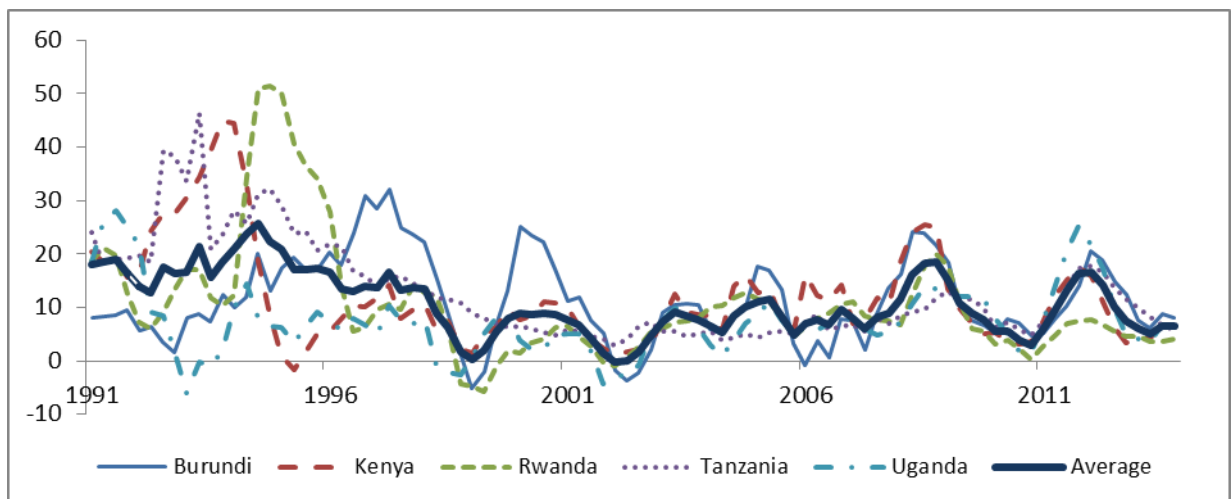


as an adjustment mechanism: countries with higher productivity or lower wage growth than others would experience a depreciation of the real exchange rate (i.e. a fall in relative prices) and thus a gain in trade competitiveness (Yilmazkuday, 2009).

### 3.3 Descriptive Analysis

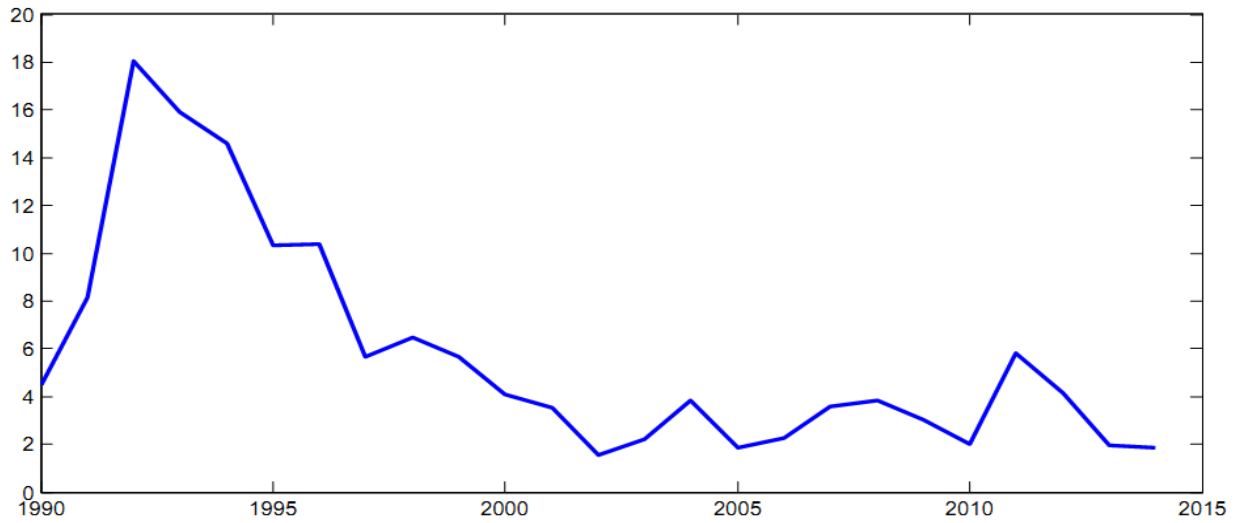
Figure 1 presents the inflation rates over 1990–2014 for five countries in EAC, as measured by the percentage change in consumer price index. The figure shows that the differentials of inflation during 1990s were substantial, with large spikes in Kenya, Rwanda and Tanzania. However, since the late 1990s inflation rates between these countries appear less volatile and move closer to the EAC average, suggesting a possibility of nominal convergence. This is also supported by Figure 2 which shows the evolution of the cross-section standard deviation. The statistics in Table 1 confirms that inflation was high and more dispersed in the pre-2000 era than the post-2000 era. However, this phenomenon is not restricted to EAC countries; as shown in Figure 3, it has occurred in other developing countries as well.

**Figure 1: Inflation in EAC Countries: 1990–2014 (percent)**



Sources: IMF and authors' calculations.

**Figure 2: Inflation Dispersion between EAC Countries**



*Notes:* Inflation dispersion is calculated as the standard deviations of inflation between five EAC countries.

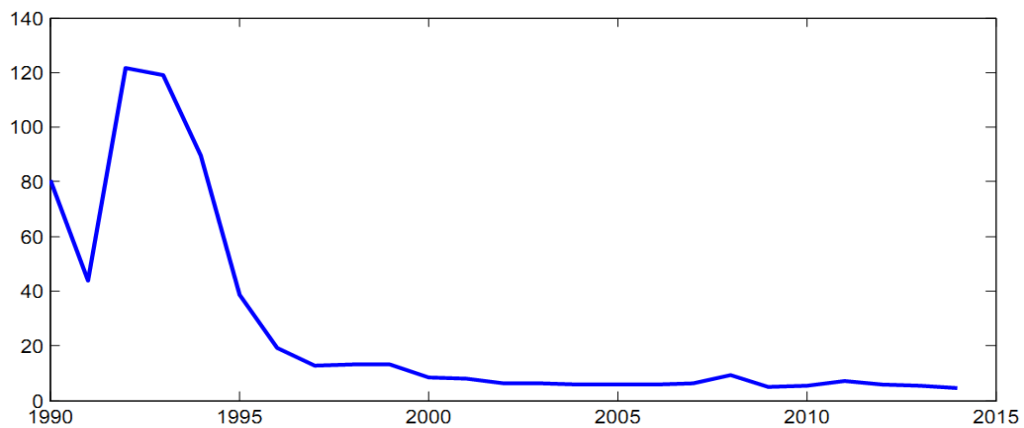
*Sources:* IMF and authors' calculations.

**Table 1: Average Inflation (%) in East African Countries**

	Burundi	Kenya	Rwanda	Tanzania	Uganda	Average	Dispersion
Pre-2000	13.44	15.23	13.49	19.96	8.38	14.10	3.72
Post-2000	9.00	7.85	6.45	7.43	6.87	7.52	0.88

*Notes:* The table shows the average of inflation of the East African countries in two sub-samples. Columns 7 and 8 calculate the mean and standard deviation of inflation rates in EA countries shown in Columns 2-5 for each sub-sample.

**Figure 3: Average Inflation of Emerging and Developing Countries (percent)**



*Source:* WEO

Although inflation dispersion between the five EAC members has reduced substantially on average, there have been still nontrivial differentials happening sporadically. Notably, in 2011 the inflation gap between the largest and the smallest rate - Uganda against Rwanda - is about 15 percent. In addition, inflation rates of the member countries often go beyond 10 percent, which is higher than the convergence criterion of inflation, as shown in Figure 1. For these reasons, it is important to investigate inflation convergence and identify the driving factors of inflation in these countries from a statistic perspective.

#### **4. Inflation convergence: Panel Unit Root Tests**

The panel unit root tests are categorized in two different groups: First and second generation tests. In the former, all cross-sections are assumed to be independent, while the second generation relaxes this assumption to allow for the cross-sectional dependences. A detailed description of the tests in both generations is presented in Appendix. Specifically, we consider a total of nine panel unit root tests, with four tests in the first generation and five tests in the second generation. Regarding the former, the first test is proposed by Levin, Li and Chu (2002) (LLC thereafter) which is based on a homogeneous alternative assumption, while the other three tests allow for heterogeneity, including Im, Pesaran and Shin (2003) (IPS thereafter) and two Fisher type tests of Maddala and Wu (1999) and Choi (2001). However, the assumption of independent cross-sections appears to be restrictive in many empirical applications, particularly in our study because the integration of EAC economies has significantly increased as discussed above. Therefore, in the second-generation tests, this assumption is relaxed to allow for dependent cross-sections. In this context, it is necessary to specify the cross-sectional correlations. Several approaches have been proposed, for instance, using a factor structure model as in line with Pesaran (2007), Bai and Ng (2004) and Moon and Perron (2004); an error-component model following Choi (2006); and a nonlinear instrumental variable approach as in Chang (2002).

Let  $\pi_{j,t}$  denote the series of inflation rate in country  $j$ ,  $j=1, \dots, 5$ , defined as the monthly percentage change in headline consumer price index. According to Buseti et al., (2007), the convergence properties between countries  $j$  and  $k$  can be studied from the time-series properties of inflation differential between them defined by:

$$y_{jk,t} = \pi_{j,t} - \pi_{k,t}, \quad j, k = 1, \dots, 5, \text{ and } t = 1, \dots, T.$$

With five countries, we construct ten series of  $y_{jk,t}$  of inflation differentials and then test if these differentials converge by using nine panel unit-root tests described above.<sup>8</sup> We use the sample 2000M1–2015M2 for our panel unit-root tests.

The first-generation panel unit root tests are presented in Table 2 including the Levin, Liu and Chu (2002) test, the Im, Pesaran and Shin (2003) test, the Maddala and Wu (1999) test and the Choi (2001) test. The LLC test rejects the null hypothesis of a unit root, therefore suggesting an evidence of inflation convergence in EAC. However, the LLC test assumes the homogeneity in the alternative, which implies that all panel members are forced to be stationary under the alternative hypothesis. Then there may be the case that with as few as  $I(0)$  series, the rejection rate rises above the normal size of the test, and continues to increase with the number of stationary series in the panel (Hurlin, 2010).

Relaxing the assumption of homogeneity in LLC, Im, Pesaran and Shin (2003) considers heterogeneous panel unit root tests. Both the statistics  $W_{tbar}$  and  $Z_{tbar}$  find that the null hypothesis of a unit root is rejected. This result is also confirmed by the two Fisher type tests: Maddala and Wu (1999) test and the Choi (2001). Therefore, the first generation unit root tests do not suggest that there are persistent inflation differentials between EAC countries.<sup>9</sup>

It should be noted that the first-generation tests are based on the assumption of independence across units. However, this assumption is restrictive, and if violated, can cause over-rejections of the null hypothesis (Bai and Ng, 2004). Banerjee, Marcellino, and Osbat (2001) argue against the use of first-generation panel unit root tests because of this potential problem. Hence, we consider the second generation panel unit root tests which take the cross-sample dependence into account. Five tests in this category are considered, including Bai and Ng (2004), Moon and Perron (2004), Pesaran (2007), Choi (2006), and Chang (2002), whose results are presented in Tables 3 and 4.

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<sup>8</sup> We obtain similar results when testing the convergence using the difference between inflation rates and cross-sectional mean as in Kočenda and Papell (1997).

<sup>9</sup> The results are robust if we include the deterministic trend in the unit root tests.

**Table 2: First generation panel unit root tests**

	<b>Levin, Liu and Chu (2002)</b>	<b>Im, Pesaran and Shin (2003)</b>		<b>Maddala and Wu (1999)</b>	<b>Choi (2001)</b>
	$t_{\rho}^*$	$W_{tbar}$	$Z_{tbar}$	$P_{MW}$	$Z_C$
$y_{ij,t}$	-38.19*	-36.88*	-36.91*	92.10*	11.40*

Notes:  $t_{\rho}^*$  denotes the adjusted  $t$ -statistic calculated with a Bartlett kernel function and a common lag truncation parameter  $\bar{K} = 3.21T^{1/3}$  (Levin et al., 2002);  $W_{tbar}$  and  $Z_{tbar}$  are the standardized  $t$ -barNT statistics based on the moments of the Dickey Fuller distribution and the simulated approximated moments, respectively (Im et al., 2003);  $P_{MW}$  and  $Z_C$  are the Fisher's test statistics suggested by Maddala and Wu (1999) and Choi (2001), respectively, which are based on a combination of the different  $p$ -values of the individual auxiliary regression from ADF tests. \* indicates significant at 5% level.

**Table 3: Bai and Ng (2004) unit root tests**

	<b>Number of common factors</b>	<b>Idiosyncratic shocks</b>		<b>Common Trends</b>	
	$\hat{r}$	$Z_{BN}$	$P_{BN}$	$MQ_c$	$MQ_f$
$y_{ij,t}$	4	5.25*	53.20*	0	0

Notes: For the idiosyncratic component, two Fisher-type statistics are reported:  $Z_{BN}$  and  $P_{BN}$ . \* indicates significant at 5% level. The number of stochastic trends in common trends is presented in the last two columns based on two statistics:  $MQ_c$  and  $MQ_f$  using 5% as the level of these tests.

**Table 4: Other second generation panel unit root tests**

	<b>Moon and Perron (2004)</b>		<b>Choi (2006)</b>			<b>Pesaran (2007)</b>		<b>Chang (2002)</b>
	$t_a^*$	$t_b^*$	$P_m$	$Z$	$L^*$	CIPS	CIPS*	$t_{IV}^*$
$y_{ij,t}$	-343.05*	-39.97*	25.96*	-11.76*	-16.06*	-8.60*	-6.21*	-32.32*

Notes:  $t_a^*$  and  $t_b^*$  in the Moon and Perron (2004) test are calculated from de-factored panel data using a Quadratic Spectral kernel function.  $P_m$ ,  $Z$  and  $L^*$  are the three different Fisher type statistics suggested in Choi (2006). For the Pesaran (2007) test, CIPS denotes the mean of individual cross-sectionally augmented ADF statistics for the optimal lag length which is one, and CIPS\* is the mean of truncated individual CADF statistics.  $t_{IV}^*$  is the average IV  $t$ -ratio statistic in the Chang (2002) test. \* indicates significant at 5% level.

Table 3 shows the results of PANIC approach proposed by Bai and Ng (2004). This approach decomposes the data into idiosyncratic and common components and then conducts unit root tests for each component. As shown in Table 3, we identify four common components, suggesting the importance of dealing with the dependencies between units in the panel. Regarding the idiosyncratic components, both  $Z_{BN}$  and  $P_{BN}$  statistics show that the null hypothesis of a unit root in inflation differentials in country-specific factors can be rejected. Although the rejection of nonstationarity of idiosyncratic components does not guarantee that the series are stationary because nonstationarity can arise if one or more of common factors are nonstationary. We therefore test the number of independent common stochastic trends among common factors based on two statistics  $MQ_c$  and  $MQ_f$ . The results show that both statistics reject the null hypothesis of a unit root at 5 percent significant level, suggesting that there is no pervasive divergence among inflation differentials between EAC countries.

Moon and Perron (2004) also rely on a factor model to tackle cross-section dependence, but use a slightly different testing strategy from that of Bai and Ng (2004) as documented in Appendix. Moon and Perron (2004) propose two test statistics  $t_a^*$  and  $t_b^*$  whose values are presented in Table 4. The results suggest that the null hypothesis of a unit root can be rejected at 5 percent significant level, which is also confirmed by the Pesaran (2007) test with a one-factor model.

Instead of using factor models, Choi (2006) and Chang (2002) propose alternative approaches to model the dependences in cross sections. The Choi (2006)'s test considers error-component models and suggests three statistics:  $P_m$ ,  $Z$  and  $L^*$ . We find that the nonstationarity is rejected no matter what the choice of the statistics is. Chang (2002) introduces the average IV  $t$ -ratio statistic which is based on a nonlinear IV estimation of the augmented Dickey-Fuller type regression. The result in Table 4 also indicates that the null hypothesis of a unit root can be rejected.

In summary, all the unit root tests in both generations suggest that inflation differentials in the five EAC countries are not persistent. In other words, inflation rates in EAC appear to converge. The similarity among a battery of tests therefore confirms the robustness of the finding.

## 5. Global VAR

This section aims to shed light on the factors that can help explain the convergence of inflation between EAC countries as found in the panel unit root tests. A natural approach to this issue is to identify what shocks underline inflation dynamics in each country and then make comparisons between them. In addition, given the fact of increasing economic integration of these countries as discussed in Section 2, it is important to consider spillover effects as well as the origins of shocks, domestic vs. foreign, to each country. To do so, we use a Global Vector Auto-regression (GVAR) model which has proven to be a useful tool in exploring the various channels and interlinkages through which shocks are transmitted and how countries are interconnected through spillovers.<sup>10</sup> Specifically, we consider a GVAR model covering 65 countries which account for more than 90 percent of world output. We expand the core set of 33 countries often considered in the GVAR literature, such as Dees et al. (2007) and Galesi and Lombardi (2013), with the inclusion of 32 additional SSA countries, including 5 EAC countries. The list of countries is reported in Table 5.<sup>11</sup>

The GVAR approach can be regarded as a two-step approach. In the first step, small scale country-specific models are estimated conditional on the rest of the world. These models feature domestic variables and (weighted) cross section averages of foreign variables, which are treated as weakly exogenous (or long-run forcing). In the second step, individual country models are stacked and solved simultaneously as one large global VAR model. Dees et al. (2007) provide a theoretical framework where the GVAR is derived as an approximation to a global unobserved common factor model. In a nutshell, when N is relatively large, unobserved factors can be proxied by the cross-sectional averages of

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<sup>10</sup> The framework allows for the construction and use of weakly exogenous country-specific foreign variables and global variables in the estimation of individual country models. In other words, trade (and/or financial) linkages are exploited to allow for a coherent inclusion of national models into a global model that deals with the “curse of dimensionality problem” associated with large models.

<sup>11</sup> To deal with the modelling issues arising from the creation of the euro area in the post 1999, 8 Euro area countries are grouped together as a single economy based on their PPP-GDP weights (see Dees et al., 2007). In addition, 14 countries in the African Financial Community franc zone, which have a fixed exchange rate to the euro, are also grouped together. The other 43 countries are modelled separately. To group the countries, we use GDP in Purchasing Power Parity terms in current international dollars from the World Bank’s World Development Indicators database.

country-specific variables and the observed common effects. Thus, the individual country VARX\*( $p_i, q_i$ ) model can be written as follows:

$$\Phi_i(L, p_i)\mathbf{x}_{it} = \mathbf{a}_{i0} + \mathbf{a}_{i1}t + \mathbf{Y}_i(L, q_i)\mathbf{d}_t + \mathbf{\Lambda}_i(L, q_i)\mathbf{x}_{it}^* + \mathbf{u}_{it}, \quad (2)$$

for  $i = 0, 1, 2, \dots, N$ ;  $t = 1, 2, \dots, T$ , where  $\mathbf{a}_{i0}$  and  $\mathbf{a}_{i1}$  are the coefficients of the deterministic trend time trend.  $\Phi_i(L, p_i)$ ,  $\mathbf{\Lambda}_i(L, q_i)$ , and  $\mathbf{Y}_i(L, q_i)$  are the matrix lag polynomial of the associated coefficients;<sup>12</sup>  $\mathbf{x}_{it}^*$  a set of country-specific foreign variables, and  $\mathbf{d}_t$  denotes global variables such as oil and food prices;  $\mathbf{u}_{it}$  is a  $k_i \times 1$  vector of idiosyncratic, serially uncorrelated, country-specific shocks with  $\mathbf{u}_{it} \sim \text{iid}(0, \mathbf{\Sigma}_{ii})$ , for  $i = 0, 1, 2, \dots, N$  and  $t = 1, 2, \dots, T$ , where  $\mathbf{\Sigma}_{ii}$  is nonsingular. The idiosyncratic shocks  $\mathbf{u}_{it}$  are correlated across countries/regions.

**Table 5: Countries in the GVAR model**

<b><u>NCFA-SSA</u></b>	<b><u>CFA-SSA</u></b>	<b><u>Rest of the World</u></b>	
Botswana	Benin	USA	<b>Asia</b>
Burundi	Burkina Faso	UK	Australia
Cape Verde	Cameroon	Sweden	China
Ethiopia	Central AFR Rep	Switzerland	India
Gambia	Chad	Norway	Indonesia
Ghana	Congo Rep	Canada	Japan
Kenya	Cote d'Ivoire		Korea
Madagascar	Equatorial Guinea	<b>Euro Area</b>	Malaysia
Malawi	Gabon	Austria	New Zealand
Mauritius	Guinea-Bissau	Belgium	Philippines
Nigeria	Mali	Finland	Singapore
Rwanda	Niger	France	Thailand
Seychelles	Senegal	Germany	
Sierra Leone	Togo	Italy	<b>Others</b>
South Africa		Netherlands	Turkey
Swaziland		Spain	Saudi Arabia
Tanzania			
Uganda		<b>Latin America</b>	
Zambia		Argentina	
		Brazil	
		Chile	
		Mexico	
		Peru	

<sup>12</sup> The lag orders,  $p_i$  and  $q_i$ , are respectively related to the domestic variables and to both the foreign-variables and the global variables. Following Dees et al. (2007), for each country  $i$ , they are selected by the AIC, where the maximum lag order is set equal to 2 due to data limitations.



As it can be seen, the model (2) includes  $\mathbf{x}_{it}^*$  a set of country-specific foreign variables as regressors, therefore capturing the contemporaneous interrelation of domestic variables  $\mathbf{x}_{it}$  with country-specific foreign variables  $\mathbf{x}_{it}^*$  and with their lagged values. This set of country-specific foreign variables is constructed by  $\mathbf{x}_{it}^* = \sum_{j=0}^N w_{ij} \mathbf{x}_{jt}$  with  $w_{ii} = 0$  and  $\sum_{j=0}^N w_{ij} = 1$ , where  $w_{ij}$  is the trade share of country  $j$  in total trade of country  $i$ .<sup>13</sup> This implies that country with higher trade share with country  $i$  will have more influences on macroeconomic fluctuations in country  $i$  than the one with lower trade share. In addition, the GVAR model allows for interdependence through (i) the dependence of domestic variables  $\mathbf{x}_{it}$  on global variables  $\mathbf{d}_t$  and their associated lagged values and (ii) the contemporaneous dependence of shocks in country  $i$  on the shocks in country  $j$  because of the cross-country covariances captured by correlated across countries/regions in  $\mathbf{u}_{it}$ .

Our dataset include consumer prices index (CPI), real GDP (RGDP), nominal effective exchange rate (NEER), broad money (M), nominal interest rates (either deposit or discount rates) (NIR) and global oil and food prices.<sup>14</sup> The choice of these variables is based on the literature on inflation dynamics in African economies, e.g. Loungani and Swagel (2001), Barnichon and Peiris (2008), Thornton (2008), Baldini and Poplawski-Ribeiro (2011), and Durevall and Sjö (2012) among others. For each country, the model includes five country-specific variables for each country-VARX\* model  $\mathbf{x}_{it} = (dCPI_{it}, dRGDP_{it}, dNEER_{it}, dM_{it}, NIR_{it})$ .<sup>15</sup> Note that the model allows for the case that some country-VARX\* models do not include the whole set of country-specific variables due to limited data availability. The set of country-specific foreign variables is given by  $\mathbf{x}_{it}^* = (dCPI_{it}^*, dGDP_{it}^*, dM_{it}^*, NIR_{it}^*)$ . However, in the case of U.S., the foreign variables are  $\mathbf{x}_{it}^* = (dCPI_{it}^*, dGDP_{it}^*)$  implying that monetary variables of other countries do

<sup>13</sup> Trade weights are calculated using the data from the IMF's Direction of Trade statistics. Trade shares were used as weights to construct country-specific foreign variables which sum up to one for a given country.

<sup>14</sup> The main data sources are the IMF's International Financial Statistics (IFS), World Economic Outlook (WEO), and Smith and Galesi (2014)'s dataset. The chosen sample is 1995Q1-20013Q4, slightly different from the one in panel unit root tests. This is because the data on real GDP of many countries are only available till 2013Q4. Meanwhile, the starting point at 1995Q1 is to guarantee an appropriate sample size to obtain reasonable results in the GVAR model (starting with 2000Q1, some of the eigenvalues of the GVAR model are greater than one, therefore causing the model not stable) while still controlling for possible structural breaks in inflation dynamics as shown in Section 3. However, it is worth noting that the GVAR is more robust to the possibility of structural breaks as compared to standard VAR models or reduced-form single equation models (Dees et al., 2007)

<sup>15</sup> Lower case letter 'd' denotes the first difference.

not influence on the U.S.<sup>16</sup> Moreover, the oil and food prices are exogenous to all countries and modeled by the dominant unit as in Chudik and Pesaran (2013).<sup>17</sup> In addition, model specification is selected to satisfy the stability condition in which all the eigenvalues of the GVAR model are not greater than one.

**Table 6: Drivers of Inflation: Geographic Origin ( percent)**

	<b>Burundi</b>	<b>Kenya</b>	<b>Rwanda</b>	<b>Tanzania</b>	<b>Uganda</b>	<b>EAC-Mean</b>
1. Domestic factors	37.29	34.73	36.15	43.83	28.33	36.06
2. Foreign factors	62.71	65.27	63.85	56.17	71.67	63.94

*Note:* Generalized forecast error variance decomposition (GFEVD) over 10 quarters for inflation of each country. Last column shows the average across EAC countries. Domestic factors refer to the impact on domestic inflation of domestic shocks. Foreign factors refer to the spillover effects of shocks to other Sub-Saharan African (SSA) economies, shocks to the non-SSA economies of the model and the global oil and food shocks.

Table 6 shows that both domestic and foreign factors have been important drivers of inflation in the five East African countries. Domestic factors refer to the impact on domestic inflation of domestic shocks. Foreign factors refer to the spillover effects of shocks to other Sub-Saharan African (SSA) economies, shocks to the non-SSA economies of the model and the global oil and food shocks. It happens in all EAC countries that foreign factors appear to contribute more to inflation fluctuations than domestic ones, as in line with Nguyen et al. (2017). This can be explained by increases in trade and financial openness in the area, making the economy more exposed to foreign factors as discussed in Section 2. Also, linking this results with the fact that inflation has been low and less volatile in industrial and emerging countries since the early 1990s (Helbling, Jaumotte, and Sommer, 2006) helps explain the convergence of inflation in the five EAC countries as suggested by the panel unit root tests.

<sup>16</sup> Following Dees et al. (2007), given the importance of the U.S. financial variables in driving the global financial variables, U.S. specific foreign financial variables would be unlikely to be weakly exogenous with respect to the U.S. domestic financial variables. The U.S. specific foreign output and inflation variables,  $dGDP_{it}^*$  and  $dCPI_{it}^*$ , are however included in the U.S. model in order to capture possible spillover of external shocks to the U.S. economy.

<sup>17</sup> We test the weak exogeneity assumption for the country-specific foreign variables  $x_{it}^*$  and global variables (oil and food prices) based on the methodology outlined in Johansen (1992) and Harbo et al. (1998) and find that exogeneity assumption cannot be rejected in most cases.

**Table 7: Drivers of Inflation: Types of shocks (percent)**

	<b>Burundi</b>	<b>Kenya</b>	<b>Rwanda</b>	<b>Tanzania</b>	<b>Uganda</b>	<b>EAC-Mean</b>
<b>Oil price, Food price, CPI</b>	47.31	47.05	52.09	46.30	40.97	46.74
<b>GDP, M, NIR, NEER</b>	52.69	52.95	47.91	53.70	59.03	53.26

*Note:* Generalized forecast error variance decomposition over 10 quarters for inflation of each country. Last column shows the average across EAC countries.

Table 7 provides another perspective on the drivers of inflation dynamics based on the nature of shocks. The use of term "shock" might not be entirely appropriate in this framework, as the structural shocks in the system are not identified. However, in the rest of the paper we refer to one standard error shifts to the observable variables as shocks for the ease of interpretation. We follow the classification of shocks considered in Osorio and Unsal (2013). Specifically, we group oil and food price shocks as well as idiosyncratic inflation shocks (i.e. weather-related/political shocks) into one group (relating to shocks to the supply side), whereas, shocks to real activities, nominal effective exchange rate, and monetary variables including money supply and nominal interest rates belong to the other groups (relating to shocks to the demand side).

As shown in the table, both types of shocks are important in determining inflation. First, oil and food price shocks as well as idiosyncratic inflation shocks account for about 45 percent of inflation fluctuations. This result is in line with Barnichon and Peiris (2008) who argue that rainfall has a significant negative impact on inflation, indicating that a drought would lead to an increase in prices. Moreover, Aisen and Veiga (2006) document that a higher degree of political instability is associated with a higher level of inflation. Our finding also supports Alper et al. (2016), Walsh, (2011) and Caceres (2011) that emphasizes the influences of food inflation. Second, the other group, including shocks to real activities, nominal effective exchange rate, and monetary variables, explains about 55 percent of inflation variations. Portillo et al. (2017) also document the contribution of output shocks to inflation variation has increased over the last 15 years in Sub-Saharan African countries. Meanwhile, Berg et al (2013) argue for the importance of monetary policy in stabilizing the

economy. Interestingly, there is a noticeable similarity in terms of the contributions of shocks to inflation fluctuations between these five economies, therefore helps to explain for the observed convergence of inflation rates. This result therefore supports for Kishor and Ssozi (2010) who point out that the speed of inflation convergence has increased significantly in the post-Treaty period.

## **6. Conclusions**

This paper investigated the issue of inflation convergence in five East African Countries, Burundi, Kenya, Rwanda, Tanzania, and Uganda. Inflation convergence is a key prior to the viability of a single currency under the proposed monetary union by 2024. If significance divergence exists, it will be problematic for the EAC central bank to apply a single monetary policy. It can lead to too loose monetary policy for high inflation countries and too tight monetary policy for low inflation countries. Inflation convergence is also a key indicator of structural synchronization. Furthermore, since the traditional channels of adjustment, namely, wage flexibility and labor mobility, are weak in the EAC, it is crucial for member states to exhibit inflation convergence prior to the establishment of the monetary union.

In order to investigate inflation convergence in the EAC, we used two generations of panel unit root tests and supplemented the above tests with a Global VAR, which has proven to be a useful tool in exploring the various channels and interlinkages through which shocks are transmitted and how countries are interconnected through spillovers. Our findings from applying panel unit root tests suggest inflation differentials between the EAC countries are not persistent, therefore implying an inflation convergence. An explanation for this convergence is also provided from the perspective of a global VAR model, which attributes this convergence to a similarity in terms of the nature of shocks affecting EAC countries as well as the role of foreign factors as drivers of inflation given that inflation has been low and less volatile in industrial and emerging countries since the early 1990s.

Given the importance of regional and global shocks, policymakers in the region should be more cautious to the regional and global inflation and growth developments, hence supporting for a cooperative approach in managing inflation between these countries. Interpreted in the context of historical developments in monetary policy, the evidence

suggests that rather than just being a mechanical occurrence, the convergence in national inflation rates experienced after the EAC treaty was brought about by monetary policy becoming more similar across countries, with authorities becoming more focused on achieving low inflation. This improvement has included greater use of market-based instruments, along with more clarity and transparency with respect to monetary objectives and instruments as well as exchange rate flexibility as documented in Berg et al. (2013). Further progress in this direction helps to further stabilize inflation in the area, thus facilitating the establishment of the EAC monetary union.

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## Online Appendix: Panel Unit Root Tests

### A.1 First generation tests

This category includes four different tests. The first is proposed by Levin, Li and Chu (2002) (LLC thereafter) which is based on a homogeneous alternative assumption. Meanwhile, three other tests allow for heterogeneity: Im, Pesaran and Shin (2003) (IPS thereafter) and two Fisher type tests of Maddala and Wu (1999) and Choi (2001).

*Levin, Li and Chu (2002) unit root test*

The LLC test is one of the most popular first generation unit root tests. To examine the mean-reverting behavior of inflation, the LLC test is based on the following equation:

$$\Delta y_{i,t} = c_i + \rho y_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t},$$

where  $k_i$  is the number of lags, for  $i=1, \dots, N$ , and  $t=1, \dots, T$ . We then test the null hypothesis that all  $\rho_i$  are equal to zero against the alternative hypothesis that they are all smaller than zero, i.e., we test the null hypothesis -  $H_0: \rho = 0$ , against  $H_1: \rho = \rho_i < 0$  for all  $i=1, \dots, N$ . The restriction in  $H_1$  implies that the autoregressive parameters are identical across the panel.

LLC show that the inclusion of individual effect causes a downward bias, making the standard  $t$ -ratio statistic  $t_\rho$  of the pool estimator  $\hat{\rho}$  diverge to negative infinity. For this reason, LLC propose an adjusted  $t$ -statistic, denoted  $t_\rho^*$ , that converges to a standard normal distribution. Specifically, the adjusted  $t$ -statistic  $t_\rho^*$  is calculated by

$$t_\rho^* = \frac{t_\rho - N\tilde{T}\hat{S}_N\hat{\sigma}_{\tilde{\varepsilon}}^{-2}STD(\hat{\rho})\mu_{m\tilde{T}}^*}{\sigma_{m\tilde{T}}^*},$$

where  $\mu_{m\tilde{T}}^*$  and  $\sigma_{m\tilde{T}}^*$  are the mean and standard deviation adjustments which can be found in Levin, Li and Chu (2002) and  $\tilde{T} = T - \frac{1}{N}\sum_{i=1}^N k_i - 1$ . Also,  $\hat{S}_N = \frac{\sum_{i=1}^N \frac{\hat{\sigma}_{y_i}}{\hat{\sigma}_{\varepsilon_i}}}{N}$  is the average standard deviation ratio where  $\hat{\sigma}_{\varepsilon_i}$  is the regression standard deviation and  $\hat{\sigma}_{\pi_i}$  is the long-

run variance which is estimated based on the choice of kernel and the selection of the bandwidth parameter.

*Im, Pesaran and Shin (2003) unit root test*

Im, Pesaran and Shin (2003) relax the homogeneous assumption in LLC to allow for heterogeneity in the value of  $\rho_i$  under the alternative hypothesis. The corresponding model specification for the IPS test is described as:

$$\Delta y_{i,t} = c_i + \rho_i y_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t}.$$

We test the null hypothesis-  $H_0: \rho_i = 0$ , for all  $i=1, \dots, N$ , against  $H_1: \rho_i < 0$  for  $i=1, \dots, N_1$ , and  $\rho_i = 0$  for  $i= N_1 + 1, \dots, N$ , with  $0 < N_1 \leq N$ . The IPS statistic is constructed as

$$t_{barNT} = \frac{1}{N} \sum_{i=1}^N t_{iT}(\rho_i, \phi_i),$$

with  $\phi_i = (\phi_{i,1}, \dots, \phi_{i,k_i})$  and  $t_{iT}(\rho_i, \phi_i)$  is the t-statistic for testing unit root in the  $i^{\text{th}}$  country. IPS show that  $t_{barNT}$  sequentially converges to a normal distribution and then propose two corresponding standardized  $t$ -bar statistics,  $Z_{tbar}$  and  $W_{tbar}$ . These standardized tests are based on the asymptotic moments of the Dickey-Fuller distribution and the means and variances of  $t_{iT}(\rho_i, 0)$  evaluated by simulations under the null  $\rho_i = 0$ , respectively.

*Fisher type tests: Maddala and Wu (1999) and Choi (2001)*

The null and alternative assumptions in Maddala and Wu (1999) and Choi (2001) are the same as in IPS. However, their tests are based on a combination of the different p-values of the individual auxiliary regression based on ADF tests. For Maddala and Wu (1999), the proposed statistic is defined as:

$$P_{MW} = -2 \sum_{i=1}^N \log(p_i),$$

which has chi-square distribution with  $2N$  degrees of freedom, when  $T$  tends to infinity and  $N$  is fixed. Meanwhile, Choi (2001) proposes a similar statistic  $Z_{MW}$  as follows:

$$Z_{MW} = \frac{-\sum_{i=1}^N \log(p_i) + N}{\sqrt{N}}.$$

which converges to a standard normal distribution under the unit root hypothesis.

## A.2 Second generation tests

In the second generation tests, the cross-sections are allowed to be dependent. In this context, it is necessary to specify the cross-section correlations. Several approaches have been proposed, for instance, using a factor structure model as in line with Pesaran (2007), Bai and Ng (2004) and Moon and Perron (2004); an error-component model following Choi (2006); or a nonlinear instrumental variable approach as in Chang (2002). In what follows, we describe the properties of these tests.

### *Bai and Ng (2004) unit root test*

Bai and Ng (2004) propose the PANIC – Panel analysis of non-stationarity in idiosyncratic and common components- approach that use a factor structure to investigate the nature of non-stationarity in a series. This is particularly useful in our context because we can identify whether the non-stationarity of inflation among EAC countries (if exist) is pervasive or specific or both. To conduct this test, we consider the following factor model:

$$y_{i,t} = c_i + \lambda_i' F_t + e_{i,t}$$

where  $F_t$  is an  $r \times 1$  vector of common factors,  $\lambda_i$  a vector of factor loadings, and  $e_{i,t}$  an idiosyncratic error. In  $F_t$ , we allow  $r_0$  stationary factors and  $r_1$  stochastic common trends ( $r = r_0 + r_1$ ). The first difference of the model is:

$$\Delta y_{i,t} = \lambda_i' f_t + z_{i,t},$$

where  $f_t = \Delta F_t$  and  $z_{i,t} = \Delta e_{i,t}$ . Applying the method of principal components to  $\Delta \tilde{\pi}_{i,t}$  yields  $r$  estimated factors  $\hat{f}_t$ , the associated loadings  $\hat{\lambda}_t$ , and the estimated residuals  $\hat{z}_{i,t}$ . Based on these results, the estimates of  $F_t$  and  $e_{i,t}$  are obtained by accumulation:  $\hat{F}_t = \sum_{s=2}^t \hat{f}_s$  (an  $r \times 1$  vector) and  $e_{i,t} = \sum_{s=2}^t \hat{z}_{i,s}$ .

The series  $\tilde{\pi}_{i,t}$  is nonstationary if one or more of common factors are nonstationary, or the idiosyncratic error is nonstationary, or both.<sup>18</sup> Thus, the approach in the Bai and Ng test is to test the idiosyncratic components and the common factors separately. For the former, Bai and Ng propose to use Fisher-type pooled ADF tests (see above) in a model with no deterministic term. For the later, the number of common factors is estimated based on  $IC_2$  or  $BIC_3$  (see Bai and Ng, 2002). When there is only one common factor ( $r = 1$ ), they use a standard ADF test in a model with an intercept. If  $r > 1$ , Bai and Ng test the number of stochastic common trends  $r_1$  based on two statistics designed to test if the real part of the smallest eigenvalue of an autoregressive coefficient matrix is unity. The first statistic,  $MC_f$ , assumes the nonstationary components of  $F_t$  to be finite order vector-autoregressive processes. Meanwhile, the second statistic,  $MC_c$ , allows the unit root process to have more general dynamics. If the number of stochastic common trends is zero ( $r_1 = 0$ ), it implies that all factors are stationary.

*Moon and Perron (2004) unit root test*

Similar to Bai and Ng (2004), Moon and Perron (2004) use a factor model to capture the cross-section dependence. Nonetheless, they assume that the error terms follow an approximate factor model instead of the original series as in Bai and Ng (2004). In our context, the dynamic panel model of inflation for the Moon and Perron test is given as follows:

$$y_{i,t} = c_i + x_{i,t}$$

$$x_{i,t} = \rho_i x_{i,t-1} + \mu_{i,t}$$

$$\mu_{i,t} = \lambda_i' F_t + e_{i,t}$$

where  $F_t$  is an  $r \times 1$  vector of common factors,  $\lambda_i$  a vector of factor loadings, and  $e_{i,t}$  are idiosyncratic shocks. The null hypothesis is defined as:  $H_0: \rho_i = 1, \forall i = 1, \dots, N$ , against the stationary alternative hypothesis:  $H_1: \rho_i < 1$  for some  $i$ .

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<sup>18</sup> Except by assumption, there is nothing that restricts  $F_t$  to be all I(1) or all I(0). In addition, it is possible that  $F_t$  and  $e_{it}$  are integrated of different orders.

The Moon and Perron test's procedure include three steps. First, estimate the pooled estimator  $\hat{\rho}_i$  and derive the estimated residuals  $\hat{\mu}_{i,t}$ . Second, apply the method of principal component to the residuals  $\hat{\mu}_{i,t}$  to get the estimates of common factors and factor loadings parameters.<sup>19</sup> Finally, construct two modified t-statistics,  $t_a^*$  and  $t_b^*$ , for the unit root test which have a standard normal distribution under the null hypothesis.

*Pesaran (2007) unit root test*

Pesaran (2007) also uses a one-factor model to dealing with the problem of cross-section dependence as in line with Phillips and Sul (2003). However, instead of basing the unit root tests on deviations from the estimated factors, Pesaran (2007) augments the standard augmented Dickey-Fuller regressions with the cross-section averages of lagged levels and first-differences of the individual series. When the residuals are not serially correlated, the Pesaran unit root test for inflation is based on the following cross-sectionally augmented Dickey-Fuller (CADF) regression:

$$\Delta y_{i,t} = c_i + \rho_i y_{i,t-1} + a_i \bar{y}_{t-1} + b_i \Delta \bar{y}_t + e_{i,t}.$$

where  $\bar{y}_t = \frac{1}{N} \sum_{i=1}^N y_{i,t}$ . The null hypothesis of this test is expressed as  $H_0: \rho_i = 0$ , for all  $i$ , against the possibly heterogeneous alternative  $H_1: \rho_i < 0$  for  $i=1, \dots, N_1$ , and  $\rho_i = 0$  for  $i=N_1 + 1, \dots, N$ , with  $0 < N_1 \leq N$ .

The Pesaran test statistic is then constructed as a cross-sectionally augmented version of the IPS test (hence denoted CIPS) by taking the mean of the  $t$ -ratio of the OLS estimate  $\hat{\rho}_i$ . A truncated version of the test is also considered where the individual CADF statistics are suitably truncated to avoid undue influences of extreme outcomes that could arise when  $T$  is small (see details in Pesaran, 2007). The simulated critical values of these tests are proposed in Pesaran (2007).

The CIPS testing procedure also readily extends to the situation in which the individual-specific error terms are serially correlated. In such a case, Pesaran (2007) modifies the above cross-sectionally augmented Dickey-Fuller (CADF) regression as follows:

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<sup>19</sup> The number of factors  $r$  in the Moon and Perron test is selected by the same criteria used in Bai and Ng (2004).

$$\Delta y_{i,t} = c_i + \rho_i y_{i,t-1} + a_i \bar{y}_{t-1} + \sum_{j=0}^k b_{i,j} \Delta \bar{y}_{t-j} + \sum_{j=1}^k \delta_{i,j} \Delta y_{i,t-j} + e_{i,t}.$$

*Choi (2006) unit root test*

Differing from the three tests above, which use factor structures to model the cross-section dependence, Choi (2006) proposes a different approach by utilizing error-component models. In our context, the Choi unit root test is based on the two-way error-component model as follows:

$$y_{i,t} = \mu_i + \lambda_t + v_{i,t}$$

$$v_{i,t} = \sum_{l=1}^{p_i} \rho_{il} v_{i,t-l} + e_{i,t}$$

where  $\beta_0$  is the common mean for all  $i$ ,  $\mu_i$  is the unobservable individual effect,  $\lambda_t$  is the unobservable time effect and  $v_{i,t}$  is the remaining random component which is modeled by the autoregressive process of the order  $p_i$ . The null hypothesis is described as  $H_0: \sum_{l=1}^{p_i} \rho_{il} = 1$ , for all  $i$ , against the possibly heterogeneous alternative  $H_1: \sum_{l=1}^{p_i} \rho_{il} < 1$  for some  $i$ .

In order to test this hypothesis, it is required to eliminate the constant term and all error components except  $v_{i,t}$  from the observed panel data  $\tilde{\pi}_{i,t}$ . This is done by using the Elliot, Rothenberg, and Stock GLS-based detrending and the conventional cross-sectional demeaning for panel data. Let us denote the transformed variables  $z_{i,t}$ , which is independent across  $i$ . Choi then applies the augmented Dickey-Fuller test to each  $z_{i,t}$  and calculates the corresponding p-values.<sup>20</sup> Based on these p-values, Choi proposes three different Fisher type statistics  $P_m$ ,  $Z$  and  $L^*$  which are proved to have a standard normal distribution.

*Chang (2002) unit root test*

Chang (2002) introduces another strategy to deal with the cross-section dependence in panel-based unit root tests. The proposed test is based on nonlinear IV estimation of the

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<sup>20</sup> This test is known as the Dickey-Fuller-GLS test.

augmented Dickey-Fuller type regression for each cross-sectional unit. To apply this unit root test, we consider the following regression:

$$y_{i,t}^{\mu} = \rho_i y_{i,t-1}^{\mu} + \sum_{k=1}^{k_i} \alpha_{i,k} \Delta y_{i,t-k}^{\mu} + e_{i,t}$$

where  $y_{i,t}^{\mu}$  are the adaptive demeaning of  $y_{i,t}$  and  $e_{i,t}$  are the regression errors.<sup>21</sup> The null hypothesis is  $H_0: \rho_i = 1$ , for all  $i$ , against the alternative  $H_1: \rho_i < 0$  for some  $i$ . To test this hypothesis, we first consider the IV estimation of the above equation. The instrument of  $y_{i,t}^{\mu}$  is generated by a nonlinear function  $F(y_{i,t-1}^{\mu})$ , which is called the *instrument generating function (IGF)*. The *IGF* is required to be regularly integrable and satisfy  $\int_{-\infty}^{\infty} xF(x)dx \neq 0$ . In this paper, we consider the function of  $IGF(x) = x \exp(-c_i|x|)$  where  $c_i = 3T^{-\frac{1}{2}}s^{-1}(\Delta y_{i,t}^{\mu})$  where  $s(\Delta y_{i,t}^{\mu})$  is the sample standard deviation of  $\Delta y_{i,t}^{\mu}$ .

Based on the IV estimates of  $\hat{\rho}_i$ , Chang constructs the individual IV  $t$ -ratio statistic and then average them to test for the joint unit root hypothesis  $H_0: \rho_i = 1$  for all  $i$ . In a balanced panel, the average IV  $t$ -ratio statistic, denoted  $t_{IV}^*$ , is proved to have a limit standard normal distribution.

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<sup>21</sup> The adaptive demeaning is described in Section 5 of Chang (2002).