Regional Euro-Currency Integration
And Economic Growth in Sub-Saharan Africa

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Abstract

This paper examines the effects of regional euro-currency integration on economic growth in Sub-Saharan Africa. As having a currency pegged to the euro is likely to be a source of shocks to output, we parameterize a Solow growth model by specifying the level of technology as a function of membership in a regional euro-currency union. Parameter estimates from panel data indicate that for the two regional euro-currency unions we consider, membership in the Central African Economic and Monetary Community (CEMAC) had a positive effect on economic growth over the 1999 - 2007 time period. Our results suggest that the regional integration of currency has benefits in Sub-Saharan Africa, and the expansion of existing, or creation of additional euro-currency unions similar to CEMAC could improve living standards.

JEL Classification: *C23, F15, N17, O11, O47*

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I. Introduction

For 14 countries in Sub-Saharan Africa, January 1, 1999 represented a new macroeconomic policy regime. On that date, the euro, by becoming the common currency for 11 countries in Europe, the 14 countries in Sub-Saharan Africa that constitute the CFA Franc Zone (CFAZ) are economies subject to business cycle dynamics associated with the euro, as the franc—the currency of these countries—is pegged to the euro.[1] To the extent that a common currency regime reduces exchange rate risk among members of a currency union , it can also reduce the effects of monetary policy shocks (Boivin, Giannoni and Mojon, 2008). As such, relative to non-CFAZ countries in Sub-Saharan Africa, the growth path of CFAZ countries could be superior. This is especially true if a source of output volatility for an economy are technology shocks that are transmitted through currency union membership.

The implementation of the euro as a single currency for Europe represented a major innovation in the international monetary system in terms its impact on the economies of the world. The introduction of the euro created a single market to facilitate and ensure the free movement of goods, services, people, and capital within the trading bloc of the European Union. With member countries, monetary policy sovereignty has been lost as the European Central Bank is solely responsible for monetary policy. The euro has created and reinforced trade within Europe by facilitating trade flows, and capital mobility via the provision of credit facilities and single currency access within the region. Sub-Saharan African economies have a large stake in the euro. It is estimated that 40 to 50 percent of Sub-Saharan African trade with Europe is linked to countries that are members of CFAZ (Irving, 1999). For member countries, approximately 65 percent of their external reserves are held in currency pegged to the euro (Irving, 1999). Such an arrangement reduces the monetary policy

[1]CFAZ consists of two separate regional currency and economic unions in Sub-Saharan Africa: 1. West African Economic and Monetary Union (WAEMU) and 2. Central African Economic and Monetary Community (CEMAC). WAEMU) includes the countries of Benin, Burkina Faso, Cote d' Ivoire, Mali, Niger, Senegal and Togo. CEMAC includes the countries of Cameron, Central African Republic, Chad, Congo Republic of, Equatorial Guinea, and Gabon.

sovereignty of CFAZ members and exposes them to potential monetary shocks originating in the European Union. Even though WAEMU is part of CFAZ, WAEMU countries constitute half of the countries that belong to the Economic Community for West African States (ECOWAS).[2] As such, WAEMU countries are regionally integrated with respect to currency and trade.

In this paper, we examine the effects of regional euro-currency integration on economic growth in Sub-Saharan African over the 1999 - 2007 time period. We parameterize a simple Solow growth model (Solow, 1956) to accommodate the effects of euro-currency integration by embedding it in the level of technology. Such an approach is sensible as the canonical Solow growth model leaves the level of technology unspecified, allowing it to be a source of unobserved differences in technology as at least a partial explanation of cross-country differences in growth. As the benefits of a common currency include lower transaction costs for buyers and sellers, an lower exchange rate risk, existing technology can become more productive engendering more output. The lower exchange rate risk for countries that share a common currency can also engender positive shocks to output. In this context, the level of technology in a Solow model can be specified as a function of membership in a currency union, and capture output dynamics similar to those of real business cycle models (McCallum, 1989).

Our inquiry is related to at least two broad strands of literature,and has implications for policy. As we extend the canonical Solow model of economic growth to incorporate the effects of common currency unions, our paper adds to our understanding of whether or not differences in cross-country living standards can be explained by differences in capital investment, population growth, and observed or unobserved differences in technology.[3] Our analysis also contributes to the literature that attempts to explain the causes and consequences of economic growth in Sub-Saharan Africa.[4] Our particular inquiry aims to determine if one of the causes of economic growth in Sub-Saharan Africa is regional euro-currency integration. If membership in a euro-currency region causes economic growth, it would have policy consequences, suggesting that the establishment of more, or the expansion of

[2]See Elu (1998) for an overview of ECOWAS.

[3]See for example Mankiw, Romer and Weil, 1992; Islam, 1995; Okada,2006.

[4]See for example Easterly and Levine (1997); Sachs and Warner (1997); Temple (1998b); Azam, Fosu, and Ndung"u (2002); Papp, Franses, and van Dijk (2005); and Gyimah-Brempong and Corley (2005).

existing regional euro-currency unions could improve living standards in Sub-Saharan Africa.

The remainder of this paper is organized as follows. In the second section, we motivate the use of the canonical Solow growth model as a specification for explaining the effects of membership in a euro-currency union on economic growth. We parameterize the effects of euro-currency membership in the technology specification. We report parameter estimates in the third section. As our measure of currency union membership is dichotomous and does not vary for member countries over the time period under consideration, our primary estimates are based on the Instrumental Variables estimator of Hausman and Taylor (1981), which allows us to control for unobserved heterogeneity, and the nature of our dichotomous measure of membership in a euro-currency union. We find that for Sub-Saharan African countries, membership in euro-currency union does matter for economic growth—and has beneficial effects at least for one particular euro-currency union. The last section concludes.

II. Theory and Empirical Methodology

To estimate the effects of regional euro-currency integration on economic growth in Sub-Saharan Africa, we assume that for a typical country, output evolves as in the standard Solow growth model (Solow, 1956). While there is evidence that the canonical Solow specification is inadequate for explaining cross-country differences in growth (Tsangarides, 2001), there is also evidence that its explanatory power is adequate. Mankiw, Romer and Weil (1992) report results suggesting that international differences in per capita are best understood with a Solow model augmented with human capital. Temple (1998a) for example, finds that the Solow growth model explains cross-country differences in growth well when the countries under consideration are similar. Durlauf (2001) reports similar results when country-specific heterogeneity is accounted for. In the context of Sub-Saharan Africa, Price (2003) finds that the Solow growth specification can accommodate the effects of colonial heritage on steady-state growth paths.

We adopt as in Mankiw, Romer and Weil (1992) a simple Solow growth model. It it is assumed that at time t, output for a particular country is determined by a Cobb-Douglas production function with physical capital, labor, and technical efficiency as arguments:

$$Q_t = K(t)^{\alpha}(A(t)L(t))^{1-\alpha}$$

where K(t) is physical capital, L(t) is labor supply, and A(t) is technology. It is assumed that the share of investment in physical capital is constant at s_k, and has a depreciation rate of δ. L(t) and A(t) are assumed to grow exogenously according to:

$$L(t) = L(0)e^{nt}$$

$$A(t) = A(0)e^{gt}$$

Instead of having the level of technology A(t) completely unspecified, we posit that it is a function of membership in a regional euro-currency union. At a micro-level a common medium of exchange reduces transaction costs (Alesina and Barro, 2002) for buyers and sellers across countries sharing the currency and as such, increase the value of markets within a country. At a macro-level, between countries a common currency can reduce and/or eliminate foreign exchange risk. One plausible macroeconomic effect is an increase in the value of markets for import and exports, which would increase output and trade between countries in a currency union (Geda and Kebret, 2007).

To the extent that a common currency reduces the likelihood and/or effects of monetary policy shocks (Boivin, Gainnoni and Mojon, 2008) among currency union members, output can also be less volatile if business cycle dynamics are a function of technology shocks as suggested by real business cycle theories (McCallum, 1989). In this context, it is theoretically sensible to specify the level of technology as a function of whether or not a country shares a common currency with other countries. Given these plausible theoretical reasons for the importance of currency union membership on a country's output we hypothesize that:

$$A(t) = A(0)e^{gt+\beta RCU_i}$$

where RCU$_i$ is a dichotomous variable indicating whether or not a country belongs to a regional euro-currency union, and β is its impact on initial technology.

Let q_{oit} and q^*_{it} be actual/initial and steady-state income per worker (e.g. $q_{it} = Q_{it}/L_{it}$) for country i at time t respectively. Approximating about the steady-state, where d$lnq(it)$/dt = $\lambda[ln(q^*_{it})$ - $ln(q_{oit})]$, and given the steady state values of the stock of physical capital per worker, an explicit regression specification for economic growth that can accommodate the effects of membership in a regional euro-currency union is:

$$\begin{aligned} ln(q^*_{it}) - \quad ln(q_{oit}) \quad &= \quad \theta a - \theta lnq_{oit} - \theta\left(\frac{\alpha}{1-\alpha}\right)ln(n_{it}+g+\delta) + \theta\left(\frac{\alpha}{1-\alpha}\right)ln(s_{kit}) \\ &+ \quad \theta\beta\ RCU_{it} + \varepsilon_{it} \end{aligned} \qquad (1)$$

where $lnA(0) = a = gt$, s_{kit} is the share of income invested in physical capital, θ = 1 - exp(-λt), λ = (n$_{it}$ + g + δ)(1 - α) is the rate at which a country converges to its steady state income per capita, and ε_{it} is a stochastic error term.[5]

The specification in equation (1) is the empirical Solow growth model we are interested in, and in particular $\theta\beta$—the parameter that measures the impact of currency union membership on economic growth for Sub-Saharan African countries. If, following Mankiw, Romer and Weil (1992) it is assumed that s_{it} and n_{it} are independent of ε_{it}, and where $E(\varepsilon_{it} \mid \varepsilon_{it-r}) = 0$, for lag r over time, ordinary least

[5]The specification in (1) results from the noting that in the steady state, the equilibrium stock of physical capital per effective worker ($k_{it} = K_{it}/A_{it}L_{it}$) is:

$$k^*_{it} = \left[\frac{s_{kit}}{n_{it}+g+\delta}\right]^{\frac{\alpha}{1-\alpha}}$$

Substituting the steady-state value of k into the intensive form of the production function (Q$_{it}$/A$_{it}$L$_{it}$ = (K$_{it}$/A$_{it}$L$_{it}$)$^{\alpha}$), taking logs and simplifying yields the specification in (1).

squares (OLS) estimation of the specification in (1) is appropriate. These identifying assumptions enable unbiased estimation of the growth model parameters, and permit an assessment of the extent to which it explains cross-country differences in growth rates.

To the extent that ε_{it} represents country-specific shifts in technology, it is rather implausible that it is uncorrelated with s_{kit} and n_{it} (Islam, 1995).[6] Neglecting this correlation if it exists leads to biased parameter estimates of the Solow growth model. Our estimation approach proceeds under the assumption that heterogeneity in technology possibly exists. For panel data, it is standard to account for heterogeneity with a fixed effect estimator or with a dynamic panel estimator—based on Instrumental Variables (Arellano and Bond, 1991), or Generalized Methods of Moments (Arellanno and Bover, 1995). Our choice is the Instrumental Variables random effect panel data estimator of Hausman and Taylor (1981) as we measure euro-currency union membership for Sub-Saharan African countries as a unit effect dummy variable. In fixed effect and dynamic panel data estimators, unit effect dummies and time invariant exogenous are collinear, rendering them unable to estimate the effects of exogenous variables measured dichotomously. The Hausman-Taylor random effect panel data estimator permits the use of unit effect dummies, permitting unbiased estimation of their parameters, and in contrast to traditional random effect estimators, allows a correlation between the regressor and the unobserved random effect.

III. Data and Results

The data consist of a panel of 48 Sub-Saharan African countries over the period 1999 - 2007. We choose this time period as it allows us to capture the effects of the establishment of a euro-currency zone for a subset of the countries in our sample. The data are from two sources. Our measures of $ln(q_{it})$ - $ln(q_{it-1})$, n_{it}, and s_{kit} are from *World Development Indicators 2007* compiled by the International Bank for Reconstruction and Development (World Bank). Gross domestic product

[6] For additional evidence of heterogeneity among countries and its econometric implications, see the results of Phillips and Sul (2007), Okada (2006), Durlauf, Kourtellos, and Minkin (2001), and Lee, Pesaran, and Smith (1998).

per worker (q_{it}) is measured as the ratio of gross domestic in constant year 2000 U.S dollars to the size of the labor force. We assume, following Mankiw, Romer, and Weil (1992) that $g + \delta = .05$, and use a country's population growth rate as a measure of n_{it}. Our measure of s_{kit} is the ratio of a country's gross capital formation in constant year 2000 U.S. dollars to its gross domestic product. Our measure of belonging to a regional euro-currency zone is dichotomous, and based on a country's membership in the Central African Franc Zone as identified by Irving (1999). We create dummy variables for a country's membership in the: (1) Central African Franc Zone (CFAZ), (2) West African Economic and Monetary Union (WAEMU), and (3) Economic and Monetary Union of Central Africa (CEMAC).

Table 1 reports the mean and standard deviation of the covariates for the entire sample and relevant subsample of Sub-Saharan African countries for our 3 measures of regional euro-currency uniom membership over the 1999 - 2007 time period. The differences in the mean for this variable suggest that being a member of a euro-currency union has an effect on economic growth. Relative to all Sub-Saharan African countries, CFAZ countries had a growth rate that was approximately 22 percent higher. Nonetheless it appears that any growth effects associated with membership in a euro-currency union is associated with CEMAC membership. Relative to all Sub-Saharan African countries, WAEMU countries had a growth rate that was approximately 27 percent lower. In contrast, CEMAC countries had a relative growth rate that was approximately 78 percent higher.

The differences in growth rates by membership in euro-currency union in Table 1 suggest that membership matters for growth. Table 2 reports pooled ordinary least squares (OLS) parameter estimates for the specification in (1). We report on four specifications of the dummy for membership in a euro-currency union: 1) Being a member of a euro-currency union—CFAZ, 2) Being a member of WAEMU only, 3) Being a member of CEMAC only, and 4) Being a member of WEAMU or CEMAC. In each case, we allow for clustering within currency unions as there could be correlated effects—within group dependence—associated with membership. As such, the estimated standard errors are adjusted for clustering within the relevant euro-currency union.[7]

[7]Adjusting standard errors for clustering proceeds simply by modifying the robust sandwich estimator of the

The pooled OLS clustered parameter estimates in Table 2 support the implications of the pattern of growth rates in the sample reported in Table 1. Across all 4 specifications the dummies for membership in a euro-currency union are significant. Being a member of CFAZ and being a member of CEMAC have had a positive effect on growth rates over the 1999 - 2007 time period. In contrast, WAEMU membership had a negative effect on economic growth in Sub-Saharan Africa. As for the adequacy of the Solow growth model, there are several reasons to be skeptical. The relatively low R^2's imply that the specification explains very little of the variation in growth rates among Sub-Saharan African countries during the time period under consideration. There is also no support for the idea of convergence to the steady-state, given the sign and lack of significance on $ln(q_{oit-1})$.[8] The signs on $ln(n_{it}+g+\delta)$ are positive, and lack significance, suggesting that contrary to the prediction of the Solow model, higher population growth causes higher growth.

The Solow growth model also implies that the coefficients on $ln(s_{kit})$ and $ln(n_{it}+g+\delta)$ are identical. Table 3 reports pooled OLS clustered parameter estimates under this restriction. Across the specifications there is no evidence that this restriction is valid, which further undermines the explanatory power of our simple Solow growth specification. However, even under the restriction the euro-currency membership dummies maintain their pattern of significance suggesting that mem-

covariance matrix (Huber, 1967) for clustering as in Zorn(2006). Let V be the information matrix, and u_i an empirical and consistent estimate of the true population residuals, then for sample size N, a robust estimate of the covariance matrix (V_R) :

$$V_R = V\sum_{i=1}^{N}[u_i'u_i]V$$

If there are a total of C clusters, each with n_i observations, the robust estimate of the covariance matrix adjusted for clustering (V_C) is:

$$V_C = V\sum_{j=1}^{N}[(\sum_{i=1}^{n}u_{ij})'(\sum_{i=1}^{n}u_{ij})]V$$

[8] Evidence for convergence to the steady-state per capita gdp requires that the coefficient on initial gdp per capita—or in our case, the previous period gdp per capita—be negative and significant (Mankiw, Romer, and Weil; 1992).

bership in a euro-currency union matters for economic growth in Sub-Saharan Africa. While the estimated parameters, fit diagnostics, and restrictions appear incompatible with the Solow model, it is conceivable that the OLS parameter estimates are biased. This is particularly the case if there is cross-country unobserved heterogeneity in technology that is correlated with $ln(s_{kit})$ and $ln(n_{it} + g + \delta)$. The direction of this bias for each parameter could be such that the true parameter values are of a magnitude, sign and significance compatible with the Solow growth model.

Of course, the pooled OLS parameter estimates in Tables 2 - 3 ignore any unobserved heterogeneity in the error term—or country-specific differences in technology—which leads to biased estimates on the effects of $ln(s_{kit})$, $ln(n_{it} + g + \delta)$, and RCU_{it}. This is especially likely as shocks evolve over time, and our parameter estimates of the Solow growth model are based on panel data. Given the possibility of unobserved country-specific differences in technology, Table 4 reports Hausman-Taylor random effect parameter estimates of the Solow growth model. Our choice of this estimator, relative to the dynamic panel estimators based on IV—Instrumental Variables (Arellano and Bond, 1991), or GMM—Generalized Method of Moments (Arellano and Bower, 1995), is based on how we measure membership in a regional euro-currency. In our data, membership in a euro-currency union does not vary over the sample time period. As such, their impact in dynamic fixed effect panel data estimators based on IV or GMM are differenced out to equal zero, and their impact on the dependent variable cannot be estimated.[9] Viewing the unobserved heterogeneity as a random effect is also compatible

[9]Greene (2003) provides a useful overview of the analytics of dynamic panel data estimators. To see the virtue and necessity of the Hausman-Taylor estimator, consider a dynamic panel formulation of our Solow growth model:

$$\mathrm{q}_{it} - q_{it-1} = \delta^{'}\mathbf{w}_{it} + \mu_i + \varepsilon_{it}$$

$$\mathbf{w}_{it} = [q_{it-1}, \mathbf{x}_{1it}^{'}, \mathbf{x}_{2it}^{'}, \mathbf{z}_{1i}^{'}, \mathbf{z}_{2i}^{'}]^{'}$$

where $\delta^{'}$ is a vector of coefficients, μ_i is an unobserved country-specific error term—the random effect, ε_{it} is an idiosyncratic error term, $\mathbf{x}_{1it}$ is a vector of time varying variables uncorrelated with μ_i, $\mathbf{x}_{2it}$ is a vector of time varying variables correlated with μ_i, $\mathbf{z}_{1i}$ is a vector of time invariant variables uncorrelated with μ_i, and $\mathbf{z}_{2i}$ is a vector of time varying variables correlated with μ_i. To identify $\delta^{'}$ given unobserved heterogeneity (μ_i), the key assumptions are $E[\mu_i \mid \mathbf{x}_{1it}, \mathbf{z}_{1i}] = 0$ and $E[\mu_i \mid \mathbf{x}_{2it}, \mathbf{z}_{2i}] \neq 0$. Hausman and Taylor (1981) show that the group mean deviations of with how we view the possible source of country-specific differences in output—shocks in production as a result of exchange rate dynamics that induce output changes. In this context, modeling the unobserved effect as random is similar to real business cycle models that view technology shocks as random variables (Dore, 1993).

The Hausman-Taylor parameter estimates reported in Table 4 assume that unobserved country-specific heterogeneity is random, and correlated with $ln(s_{kit})$ and $ln(n_{it} + g + \delta)$.[10] The standard errors are adjusted for clustering within euro-currency unions, and are bootstrapped.[11] Similar to the pooled OLS parameter estimates, the Hausman-Taylor parameter estimates in Table 4 support the implications of the pattern of growth rates in the sample reported in Table 1. Membership in CFAZ and CEMAC have a positive effect on growth, and WAEMU membership has a negative effect on growth, but insignificant when membership in WAEMU and CEMAC are both in the specification. Relative to the pooled OLS parameter estimates, the Hausman-Taylor parameter estimates are more favorable toward the Solow growth model's adequacy. The negative and significant sign on $ln(q_{oit-1})$ supports the idea of convergence to a steady-state, the sign on $ln(n_{it}+g+\delta)$ is negative, and significant

the vectors $\mathbf{x}$ and $\mathbf{z}$ can be used as instruments to obtain a random effect estimate of δ', and some of the regressors are allowed to be correlated with the unobserved random effect. This estimator allows the inclusion of time invariant dichotomous variables, and they are not dropped if used as instruments since the IVs are transformations based on deviations from group means. This in contrast to the dynamic panel estimators proposed by Arellano and Bond (1991) and Arellano and Bower (1995). These estimators use as instruments first differences of the regressors, which drop dichotomous variables included in the regression specification.

[10]In particular, our Hausman-Taylor parameter estimates assume the following: 1) $ln(q_{oit-1})$ is time varying, exogenous and uncorrelated with the random effect, 2) $ln(n_{it} + g + \delta)$ and $ln(s_{kit})$ are time varying, endogenous and correlated with the random effect, and 3) RCU_i is time invariant, exogenous, and uncorrelated with the random effect. The first assumption char in technology that condition $ln(n_{it}+g+\delta)$ and $ln(s_{kit})$. The third assumption characterizes the random nature of real business cycle type shocks to output that could occur as a result of membership in a regional euro-currency union.

[11]The properties of robust clustered standard errors in fixed/random effect models assumes that the number of clusters goes to infinity. When the number of clusters is small—as in our case—Cameron, Gelbach and Miller (2007) show that bootstrapped standard errors have superior asymptotic properties, and lead to improved inference.

in one instance. As a measure of goodness-of-fit, Table 4 reports the correlation (ρ) between the actual and estimated value of the dependent variable. Given the value of ρ, our specification of the Solow growth model explains approximately 71 percent of the variation in growth rates for Sub-Saharan Africa. Undermining the adequacy of the Solow growth model is the lack of significance on the estimated parameters for $ln(s_{kit})$.

Table 5 reports Hausman-Taylor clustered parameter estimates under the restriction that the coefficents on $ln(s_{kit})$ and $ln(n_{it} + g + \delta)$ are identical. In none of the specifications can this restriction be rejected. Moreover, the signs on $ln(q_{oit-1})$ are significant and consistent with convergence, and the value of indicates that the specification accounts for a substantial amount of the variation in growth rates. These results are highly favorable for the adequacy of the Solow growth model in explaining cross-country difference in growth among Sub-Saharan African countries. The pattern of significance for the euro-currency dummies change remain as they were for the parameter estimates in Table 4. As our parameterization of the Solow model views a country's technology as a function of euro-currency union membership, the Hausman-Taylor parameter estimates in Table 4 provide favorable evidence that euro-currency union membership matters for economic growth in Sub-Saharan Africa. As the currency-union dummies control for their direct effects on technology, the unobserved country-specific random differences in technology could represent shocks, some of which could originate in the foreign exchange market, or or from policy responses to monetary surprises—all of which shift the production function engendering output dynamics that affect growth rates.

IV. Conclusion

Viewing euro-currency membership as a determinant of a country's level of technology that can engender output dynamics, this paper examined the impact of regional euro-currency union membership on economic growth in Sub-Saharan Africa. Parameter estimates from a Solow growth model specification that accounts for unobserved country-specific differences in technology show that membership in a regional euro-currency union does matter for economic growth in Sub-Saharan Africa.

Relative to Sub-Saharan African countries in general, countries in the Central African Economic and Monetary (CEMAC) Union experienced higher growth rates over the 1999 - 2007 time period. Our results suggest that this effect is causal, given our parameterization of euro-currency union membership in the Solow growth model. In contrast, we found that countries in the West African Economic and Monetary Union (WAEMU) experienced relatively lower economic growth—suggesting that regional euro-currency integration per se is not necessarily beneficial.

With respect to the canonical Solow growth model, our Hausman-Taylor restricted parameter estimates suggest as a specification, it is adequate for explaining cross-country differences in economic growth among Sub-Saharan African countries. This is indeed promising, as it suggests that the steady-state standard of living can be derived from a theory that is parsimonious with respect to variables that can inform policy. As a large part of the literature on growth features cross-country growth specifications with no obvious theoretical justifications—so called "Barro-regressions"'(Barro, 1991)—our results are perhaps more useful for policymakers interested in improving living standards in Sub-Saharan Africa. That our results are consistent with the Solow growth model suggests that policymakers need only focus on a few policy instruments: capital investment, population growth—and the determinants of technology such as regional euro-currency unions.

Our results suggest that regional euro-currency unions such as CEMAC are able to induce favorable output dynamics, perhaps as a result of mitigating exchange rate risks that reduce the likelihood and severity of shocks to output. As such, there are benefits to regional currency integration in Sub-Saharan Africa. However, euro-currency unions like WAEMU appear to induce output dynamics that lower living standards. Of course WAEMU countries are also member of ECOWAS, a regional economic union that effects more broad economic relationships between members. As such, it could be the case that regional integration that includes both currency and formal trade, employment, and regulatory integration that are impacted through ECOWAS induce output dynamics not favorable for economic growth.

Table 1

Mean and Standard Deviation of Covariates

Sample:	**All Countries**	**CFAZ Countries**	**WAEMU Countries**	**CEMAC Countries**
Variable				
Growth Rate: 1999 - 2007:	.388	.472	.284	.692
(GDP Per Worker)	(.398)	(.599)	(.167)	(.815)
Population Growth	.025	.026	.029	.024
	(.009)	(.006)	(.005)	(.006)
Gross Capital Formation:	.218	.233	.194	.281
As a Percentage of GDP	(.108)	(.142)	(.056)	(.193)

NOTES: Standard deviation reported in parentheses. Due to missing observations on countries, the number of observations for each covariate varies. CFAZ = Central African Franc Zone, WAEMU = West African Economic and Monetary Union, CEMAC = Economic and Monetary Union of Central Africa.

Table 2

Pooled Ordinary Least Squares Clustered Parameter Estimates

Dependent variable: $ln(q_{it})$ - $ln(q_{it-1})$ (1999 - 2007)

	(1)	(2)	(3)	(4)
$ln(q_{oit-1})$	0.00445**	0.00523**	0.00291	0.00312
	(0.00150)	(0.00123)	(0.00188)	(0.00310)
$ln(n_{it} + g + \delta)$	0.0306**	0.0463**	0.0260	0.0317**
	(0.0114)	(0.00729)	(0.0234)	(0.0175)
$ln(s_{kit})$	0.0416**	0.0409**	0.0413**	0.0410**
	(0.0200)	(0.00282)	(0.0166)	(0.0193)
CFAZ	0.00523**			
	(0.00145)			
WAEMU		-0.00895***		-0.00568**
		(0.000222)		(0.00146)
CEMAC			0.0196**	0.0185**
			(0.00153)	(0.00138)
Constant	-0.0115	-0.0463*	0.00661	-0.00558
	(0.0208)	(0.0277)	(0.00610)	(0.0188)
N	252	252	252	252
R^2	0.168	0.170	0.180	0.182
adj. R^2	0.154	0.156	0.167	0.165

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 3

Pooled Ordinary Least Squares Restricted Clustered Parameter Estimates

Dependent variable: $ln(q_{it})$ - $ln(q_{it-1})$ (1999 - 2007)

	(1)	(2)	(3)	(4)
$ln(q_{oit-1})$	0.00520***	0.00489***	0.00397	0.00375
	(0.000996)	(0.000583)	(0.00251)	(0.00305)
$ln(s_{kit}) - ln(n_{it} + g + \delta)$	0.0408**	0.0414****	0.0401**	0.0403**
	(0.0189)	(0.00202)	(0.0166)	(0.0187)
CFAZ	0.00468**			
	(0.00181)			
WAEMU		-0.00871***		-0.00616***
		(0.000245)		(0.00156)
CEMAC			0.0191***	0.0180***
			(0.00107)	(0.00134)
Constant	-0.0384***	-0.0336***	-0.0307**	-0.0281*
	(0.000653)	(0.00341)	(0.0118)	(0.0151)
N	252	252	252	252
R^2	0.168	0.169	0.180	0.181
adj. R^2	0.158	0.159	0.170	0.168

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 4

Hausman-Taylor Random Effect Clustered Parameter Estimates

Dependent variable: $ln(q_{it})$ - $ln(q_{it-1})$ (1999 - 2007)

	(1)	(2)	(3)	(4)
$ln(q_{oit-1})$	-0.109**	-0.108**	-0.109**	-0.109**
	(0.0542)	(0.0490)	(0.0513)	(0.0526)
$ln(n_{it} + g + \delta)$	-0.0833*	-0.0816	-0.0838	-0.0832
	(0.0455)	(0.0665)	(0.0564)	(0.0744)
$ln(s_{kit})$	0.0228	0.0229	0.0228	0.0228
	(0.0168)	(0.0164)	(0.0148)	(0.0161)
CFAZ	0.0335***			
	(0.00203)			
WAEMU		-0.0433***		-0.0264
		(0.00168)		(0.0395)
CEMAC			0.110***	0.105***
			(0.00161)	(0.00222)
Constant	0.971**	0.983**	0.968**	0.971**
	(0.436)	(0.426)	(0.420)	(0.484)
N	252	252	252	252
ρ	.708	.708	.708	.708

Bootstrap standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

ρ is the correlation between $ln(q_{it})$ - $ln(q_{it-1})$ and its predicted value.

Table 5

Hausman-Taylor Random Effect Restricted Clustered Parameter Estimates

Dependent variable: $ln(q_{it})$ - $ln(q_{it-1})$ (1999 - 2007)

	(1)	(2)	(3)	(4)
$ln(q_{oit-1})$	-0.103*	-0.103*	-0.103*	-0.103*
	(0.0564)	(0.0495)	(0.0529)	(0.0528)
$ln(s_{kit}) - ln(n_{it} + g + \delta)$	0.0188	0.0189	0.0187	0.0188
	(0.0167)	(0.0166)	(0.0147)	(0.0168)
CFAZ	0.0281***			
	(0.00202)			
WAEMU		-0.0487***		-0.0326
		(0.00157)		(0.0425)
CEMAC			0.107***	0.101***
			(0.00159)	(0.00214)
Constant	0.724*	0.739*	0.719*	0.725*
	(0.424)	(0.369)	(0.390)	(0.398)
N	252	252	252	252
ρ	.704	.704	.704	.704

Bootstrap standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

ρ is the correlation between $ln(q_{it})$ - $ln(q_{it-1})$ and its predicted value.

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