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Foreign Aid and the Real Exchange Rate in the West African Economic and Monetary Union

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Abstract

The aim of this paper is to re-examine the relationship between foreign aid and the real exchange rate, using the recent econometric methods developed for non-stationary dynamic panels and an estimator that imposes a weaker homogeneity assumption on the slope coefficients. The investigation shows that foreign aid led to an appreciation of the real exchange over the period 1975-2005. In addition, the paper finds that other variables, such as labour productivity (a proxy for Balassa-Samuelson effect), terms of trade improvement, and government consumption of non-tradable goods are also associated with an appreciation of the real exchange rate. To avoid an appreciation of the real exchange rate and a decline in competiveness, we recommend that WAEMU countries use foreign exchange from aid inflows to import capital goods, which will not only lead to export expansion, but also to faster economic growth.

JEL Classification: C23, F31, F35
Key words: Foreign Aid; Real Exchange Rate; Pooled Mean Group

I. Introduction

Many disagreements over foreign aid are rooted in different views about how it affects some important economic variables – growth, investment, savings, and the real exchange rate. In a classic paper, Van Wijnbergen (1986) demonstrated that foreign aid inflows (through foreign exchange) can cause the price of non-tradable goods to rise relative to tradable goods, leading to an appreciation of the real exchange rate – a phenomenon known in the literature as the "Dutch Disease". This relationship is certainly an important one, at least from a policy perspective, since an appreciation of the real exchange rate will ultimately lead to a loss of export competitiveness, which can hurt export orientation, and even reduce growth. This effect is even worse if the export sector can benefit from both static and dynamic gains from trade; for example, increased specialization and learning-by-doing[2].

Yet there is another notion that foreign aid inflows will not lead to an appreciation of the real exchange rate when spent on traded goods - imported investment goods and on factors that are limited in supply (Berg *et al.*, 2005). In this case, the import of capital goods will permit greater domestic investment, which then, can lead to export expansion[3] (and increased competiveness) and growth. Thus, the long-term impact of foreign aid inflows on the real exchange rate can only be determined empirically.

However, while the past two to three decades have witnessed an outpouring of empirical research on the impact of foreign aid on growth[4], only a few studies have examined the link between foreign aid and the real exchange rate. To the

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[2] There is a general consensus that trade (in particular export) is the engine of growth, so a decline in export will have a negative implication for growth.

[3] The Aid-tying argument which requires a recipient country to use the aid money to import capital goods from the donors (Morrissey and White, 1996), though no longer popular, can also mitigate the effect of foreign aid on the real exchange rate.

[4] The main studies include Burnside and Dollar (2000), Levy, (1988), Boone (1994), Hansen and Tarp (2001), Clemens, Radelet, and Bhavnani (2004), Rajan and Subramanian (2005), Islam (2005), and Roodman (2007).

extent that we are aware, the studies on foreign aid and the real exchange rate for Sub-Saharan Africa (SSA) include: Nyoni (1998) for Tanzania; Sackey (2001) for Ghana; Adenauer and Vagassky (1998) for four countries in West Africa; Elbadawi (1999) for 62 developing countries, including 28 from Africa; and Ouattara and Strobl (2004) for the 13 CFA zone countries. But results from many of these studies produce conflicting conclusions. More generally, Edwards (1994) examines the effect of capital inflows on the real exchange rate, but does not consider the foreign aid-real exchange rate nexus. Recently, Rajan and Subramanian (2009) find that foreign aid inflows led to an appreciation of the real exchange rate in a large sample of developing countries, but it is far from the last word on this subject. Our paper differs by using an extended data set, sample of economies with similar characteristics, and a different estimation technique.

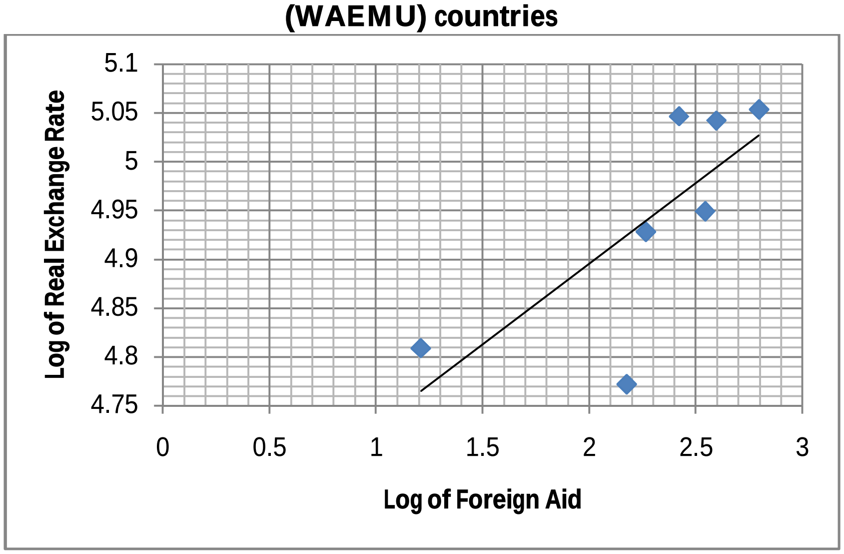
While time series studies (for example, Nyoni, 1998 and Sackey, 2001) find that increases in foreign aid inflows are in the long-run associated with a depreciation of the real exchange rate, existing panel data studies on this subject produce mixed results. For example, Adenauer and Vagassky (1998) apply the Generalized Least Squares (GLS) method to four CFA countries and find that foreign aid inflows led to an appreciation of the real exchange rate. In Elbadawi (1999), an equation in which a one period lag of the real exchange rate appeared on the right hand side was estimated using the fixed effects methodology. In line with the Dutch disease prediction, Elbadawi reported a strong evidence of real exchange rate appreciation. However, the traditional fixed effects technique applied by Elbadawi is not suitable when the lag of the dependent variable appears on the right hand side of the estimating equation. One reason for this is that fixed effects assume exogeneity of the independent variables, implying that the coefficients may not be consistently estimated.[5]

[5] The presence of the fixed effects leads to a correlation between the lagged real exchange rate and the residual, which biases the results. In this case, the coefficient of the lagged variable is

Ouattara and Strobl (2004) apply the first difference-generalized methods of moments (GMM) estimator on a sample of 13 CFA countries and find that foreign aid inflows are associated with a depreciation of the real effective exchange rate. While this estimator can address the problem of endogeneity, it is not completely immune to other estimation problems. For example, by using the first difference of the variables, the information relevant for a long –run stable relationship between the real exchange rate and its determinants is lost. Thus, a long run stable relationship cannot be inferred from this methodology, and the coefficients cannot be interpreted as long-run values.

Figure 1 shows a positive association between foreign aid and movements in real exchange rates, indicating that the Dutch disease effect seems to be at work for our sample of 7 WAEMU countries. But whether foreign aid leads to an appreciation of the real exchange rate is a matter for empirical examination.

Figure 1: Average log of foreign aid and real exchange rate for a sample of 7 CFA (WAEMU) countries



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negatively biased. As discussed in Baltagi, Griffin, and Xiong (2000), FE model is subject to a simultaneous equation bias from the endogeneity between the error term and the lagged dependent variable.

The aim of this paper is to re-examine the relationship between foreign aid and the real exchange rate using the Pooled Mean Group estimator (PMG) estimator and the recent econometric methods developed by Im, Pesaran and Shin [henceforth IPS] (2003), Hadri (2000) and Pedroni (1999, 2004) for non-stationary dynamic panels. These methods, except the PMG estimator, have successfully been applied by Abdih and Tsangarides (2006) and Roudet *et al.* (2007) in estimations involving equilibrium real exchange rates. Similarly, Drine and Rault (2003) have used this approach to examine the relationship between productivity and the real exchange rate for Middle East and North Africa (MENA) countries. However, there are no studies using these techniques to examine the foreign aid-real exchange rate relationship for WAEMU.

Again, from a policy view point, it is important to examine whether foreign aid hurts the competiveness of the WAEMU given that these countries receive substantial amounts of foreign exchange in foreign aid as opposed to other inflows. This reasoning is in line with Corden and Neary (1982) who argue that different capital flows can have different effects on the real exchange rate, depending, in part, on the reversibility of capital inflows.

In line with the Dutch Disease prediction, we find that foreign aid led to an appreciation of the real exchange rate between 1975 and 2005. Similarly, other factors, such as, improved terms of trade and government consumption of non-tradable goods, are associated with an appreciation of the real exchange rate. The Balassa-Samuelson effect is also evident in the data.

The remainder of the paper is organised as follows: Section II presents the theoretical foundation. Section III lays out the econometric model and discusses the data while the estimation results are discussed in section IV. Finally, section V concludes.

II. Theoretical Foundation

The Salter-Swan framework[6] is the main building block of the analysis of the relationship between foreign aid and the real exchange rate. The model considers an open economy that is split into traded goods and non-traded goods sectors. Here, traded goods comprise exportable and importable goods whose prices are determined on the world market. In line with the small country assumption of orthodox trade theory, the model assumes that the demand for traded goods is a perfectly elastic - countries are price takers. In contrast, non-traded goods do not enter the world market; hence their price is determined by the forces of demand and supply in the domestic economy.

Van Wijnbergen (1986), building on this framework and on the work of Corden and Neary (1982), showed how foreign aid can lead to an appreciation of the real exchange rate. In this context, foreign aid (through foreign exchange) leads to an increase in domestic income. This additional income then leads to an increase in the demand for, and expenditure on, both traded and non-traded goods – *aid-induced spending* effect. As demand increases, the price of non-traded goods, which is domestically determined, also increases. Given that the demand for traded goods is assumed to be perfectly elastic – price taker assumption, its price will remain unchanged. Thus, the real exchange rate, defined in this set-up as the ratio of price of non-traded goods to price of traded goods, will appreciate.

In addition to the aid-induced spending effect, the increase in domestic income produces a *resource switching* effect, which reduces competiveness and shrinks the traded goods sector. Here, as the non-traded goods sector becomes more

[6] Salter (1959) divides total production and expenditure into traded and non-traded goods. Traded goods are those with prices determined on world markets. They consist of exportable goods, of which the deficiency between consumption is exported; and importable goods, of which the deficiency between consumption and home production is imported. Non-traded goods are those goods which do not enter into world trade; their prices are determined solely by internal costs and demand.

attractive due to a price and wage rise, resources move from the traded goods sector to the non-traded goods sector and cause a decline in exports.

III. Econometric Methodology and Data

In this section we discuss the main econometric techniques and the variables used in the study. So far in this paper, we have explained the mechanisms through which aid can affect the real exchange rate, as well as reviewed the empirical literature. We now move on to a consideration of the main estimator, and later to the cointegration and unit root techniques.

A. Model Specification

This section presents the Pooled Mean Group estimator developed by Peseran *et al*. (1999). While the traditional fixed and random effects estimators can be used for panel data estimations, they are not suitable for dynamic specifications. One reason for this that we mentioned earlier is that the coefficients of the independent variables will not be consistently estimated. Moreover, these estimators are based on strict homogeneity assumptions, which studies have shown not to be frequently satisfied. In view of these arguments, it is important to employ an estimator that imposes weaker homogeneity assumptions. Thus, the PMG estimator becomes the natural choice for this analysis.

A.1. Pooled Mean Group Estimator (PMG)

The PMG follows an autoregressive distributed lag process (ARDL) of order (*p q*) expressed as:

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}^{'} x_{i,t-j} + \alpha_i + \varepsilon_{it} \qquad (1)$$

where y_{it} the is real effective exchange rate for country *i* in time *t;* x_{it} is the vector of explanatory variables for country *i* in time *t;* α_i represents country-specific effects; the coefficients of the lagged dependent variables, λ_{ij}, are scalars; and δ_{ij} are coefficient vectors. Re-parameterizing (1) gives the error correction equation:

$$\Delta y_{it} = \phi_i y_{it-1} + \beta_i^{'} x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta x_{it-j} + \alpha_i + \varepsilon_{it} \qquad (2)$$

i = 1, 2, …, *N*, and *t* = 1, 2, …, *T,* where ϕ_i = - $(1 - \sum_{j=1}^{q_i} \lambda_{ij})$, $\beta_i = \sum_{j=0}^{q_i} \delta_{ij}$,

j = 1, 2, …, *p*-1,

and

$$\delta_{ij}^{*} = -\sum_{m=j+1}^{q} \delta_{im}, \quad j = 1, 2, \ldots, q\text{-}1,$$

Two assumptions of this estimator are important: First is that the ARDL (*p q*) model (1) is stable. This assumption ensures that $\phi_i < 0$, and that y_{it} and x_{it} are related (cointegrated). The parameter relates to the adjustment path to the long-run value of the real exchange rate after some shock. Satisfying the condition, $\phi_i < 0$, would reinforce any evidence of a long-run stable relationship between y_{it} and x_{it}, when established through a formal cointegration test. Second is that the long-run effects are the same across countries. That is:

$$\beta_i = \beta, \; i = 1, 2, \ldots, N$$

We shall verify the validity of this assumption by applying the Hausman statistic. In addition to conducting this formal test, several arguments can validate this assumption. Since member countries use the same currency [*CFAF*], some homogeneity in the relationship between the real exchange rate and its main determinants can be expected. More importantly, the absence of forward markets supports the long-run homogeneity of the slope coefficients. On the other hand, one can expect some short-run differences in the size and effects of some shocks to the economy. This paper is not necessarily concerned with the short-run movements in the real exchange rates, instead, with the long-run coefficients, β s, and the speed of adjustment parameter - ϕ_i .

Following the lead by Peseran *et al.*, this paper also applies the dynamic fixed effects (DFE) estimator. Abdih and Tsangarides (2006) and Roudet *et al.* (2007) do not use this estimator. In principle, the DFE serves as a robustness check on the PMG results. Again, the Hausman statistic will be used to determine whether any simultacneity bias is present in the DFE estimates.

Another estimator that is similar to the PMG estimator is the Mean Group [MG] estimator. The MG estimator estimates the model for each country separately before averaging over the coefficients. However, the efficiency of the MG estimator is affected by the inclusion of a lagged dependent variable, especially when *T* is small. Also, the MG is inefficient in the presence of homogenous coefficients, hence not compatible with assumption 2 above.

B. Unit Root and Cointegration Tests

Evidence from recent studies is often quoted to support the view that most macroeconomic variables, especially real exchange rates, are non-stationary (see, for example, Maeso-Fernandez, 2006). This, therefore, necessitates testing the series for unit root. Unless non-stationary variables are cointegrated, any regression based on them will yield spurious results.

In this paper, we shall employ the IPS *t-bar* test[7] for unit-root. The IPS test is developed for dynamic panels and takes account of serial correlation in the series[8]. We augment the IPS test with a residual-based Lagrange Multiplier [LM] test, developed by Hadri (2000). The LM test is derived from the Kwiatkowski *et al.* (1992) statistic which tests the hypothesis that the time series for each cross-sectional unit is stationary around a deterministic level or a deterministic trend.

In the traditional time series literature, a set of variables that are individually integrated (of the same order) are cointegrated if some linear combination of the series can be described as stationary. Many, perhaps most economists, (e.g. Engle and Granger, 1987; Johansen, 1988, 1995; Pesaran, 1997) working in the field of econometrics have expressed the view that cointegration analysis is concerned with long-run behaviour. In this case, since our main interest is in the medium to long-term determinants of the real exchange rate, the use of cointegration techniques is therefore appropriate.

For the cointegartion test, we shall apply the panel residual-based statistics developed by Pedroni (1997, 1999). These tests are primarily used for testing long-run relationships in dynamic panels within a multivariate framework. There are two categories of the Pedroni tests, namely, the within-dimension based tests and between-dimension based tests (see the appendix for further discussion)[9].

C. Data and Variables

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[7] For more insights, see Levin, Lin and Chu (2002); Breitung (2000); Maddala and Wu (1999); Choi (2001).

[8] A Monte Carlo experiment by IPS justifies this choice, since it was shown that *t-bar* is powerful even when the value of N is less than 5.

[9] There are alternative tests by Kao (1999) and McCoskey and Kao (1998). Kao's test is not applicable in this context because it does not allow for multiple exogenous variables in the cointegrating equation.

This paper uses annual data from 1975 to 2005, obtained from the *IMF International Financial Statistics, World Economic Outlook, World Development Indicators and the Organisation for Economic Cooperation and Development (OECD) database*. The choice of variables is motivated by the real exchange rate and the Dutch disease theories, the empirical literature, data availability considerations and the structure of WAEMU economies. The summary statistics and graphs of foreign aid and the real exchange rate are presented in the appendix.

C.1. Real Exchange Rate

In the preceding theoretical discussions, the real exchange rate was defined as the ratio of non-tradable to tradable goods prices (PNT to PT). But the problem associated with this definition is that the national accounts do not make any distinction between tradable and non-tradable goods. The application of this concept may therefore pose some problems if one was to estimate a real exchange rate equation. For this reason, we re-define the real exchange rate in effective (multilateral) terms, which for country *i*, (E_i) is expressed as:

$$REER_{it} = \prod_{j=1}^{n} \left(\frac{P_{it} S_{ijt}}{P^*_{jt}} \right)^{\varpi_{ij}} \tag{3}$$

In this equation, $REER_{it}$ is real effective exchange rate for country *i* at time *t*, S_{ij} is the nominal exchange rate defined as the foreign price of domestic currency; P_i is the domestic price level in country *i*; P_j^* is the foreign price level in country *j*; ϖ_{ij} is the trade weight of country *j* in country *i*'s effective exchange rate index. An increase in $REER$ implies currency appreciation or loss in competitiveness. While the domestic and foreign price levels can be measured in various ways depending on which definition of the real exchange rate one is interested in, this study uses the consumer price index (CPI).

C.2. Foreign aid

Foreign aid is defined as Official Development Assistance (ODA). ODA are pure grants and loans provided on concessional terms (with a grant element of at least 25 percent on loans) which are aimed at promoting economic development and welfare of recipients. It is normalised in percent of GDP.

C.3. Labour Productivity

Balassa (1964) and Samuelson (1964), working independently reached the same conclusion that the real exchange rate appreciates in countries experiencing rapid economic growth (technological progress). This is because labour productivity, which historically, has been a feature of traded rather than non-traded goods sector is higher in, developed than developing countries. Furthermore, productivity is assumed to be the same in the non-traded sector for both countries, while wages are the same in the traded and non-traded sectors within each economy and is positively related to productivity. Because increases in productivity induce a wage rise, prices of goods tend to increase as a result.

In effect, the main point that arises from the Balassa-Samuelson story is that rich countries tend to have overall high price indices, and poor countries low price indices, when aggregate baskets of traded and non-traded goods are converted into a common currency. Following Abdih and Tsangarides (2006), Roudet *et al.*, (2007) and Li (2004), we take *real per capita GDP relative to the main trade partners* as a proxy for the Balassa-Samuelson (a measure of productivity) effect on the real exchange rate.

C.4. Terms of Trade

Most African countries export primary commodities whose prices are determined in the world commodity markets and are also subject to erratic shocks. In this case, the real exchange rate will be affected by movements in the relative prices of exports and imports. Terms of trade (TOT) is defined as the *ratio of export to import prices* and can affect the real exchange rate through both the income and substitution channels (Edwards, 1989).

The income effect is when the real exchange rate appreciates or depreciates as a result of a rise or fall in the relative price of exports. The rise or fall in the relative price of exports leads to a rise or fall in real income of the economy, hence, a rise or fall in demand for, and the relative price of, domestic goods (non-tradables). The substitution effect of TOT improvement or worsening is not well understood. Assuming non-tradables and tradables are substitutes, an improvement in TOT will cause the price of non-tradables to increase relative to importables and decline relative to exportables. With these opposite effects, the overall change in the relative price of non-tradables to the tradables is difficult to disentangle. However, we will expect an improvement in terms of trade to cause an appreciation of the real effective exchange rate.

C.5. Government Consumption

Changes in government expenditure can also affect the real effective exchange rate through the domestic price level. An increase in government spending leads to a rise in the demand for, and price of, domestic goods, thus, causing an appreciation of the real exchange rate. This is the substitution effect. It is also plausible to contend that an increase in government spending will be financed through higher taxes. This will lead to a fall in disposable income, and subsequently, a decrease in demand for domestic goods. This represents the income effect of an increase in government spending. On this basis, the effect of government spending on the real exchange rate will depend on whether the substitution or income effect dominates (Edwards, 1989)[10]. In line with other studies, government consumption expenditure is expressed as a percent of GDP.

IV. Estimation Results

The first set of regressions investigates the variables for unit root and cointegration. The second set examines the presence or otherwise of the Dutch disease phenomenon. In addition, it gives some insights into the relationship between the real exchange rate and other exogenous variables, and sheds some light on the stability of our core findings.

A. Unit Root and Cointegration

The results of the IPS and Hadri tests are summarised in Table 1, and show that all the variables are non-stationary. At the country level, however, the results are somewhat mixed. The ADF and KPSS tests (see Table A.1 in the appendix) show that government consumption is stationary for Benin, Burkina Faso and Togo. Since our study is based on panel data, and both panel unit root techniques show that government consumption has a unit root, we accept that government consumption is non-stationary. Perhaps, what is striking here is the presence of unit root in the real exchange rate which confirms to the widely held view in the literature.

Also, additional panel unit root tests conducted on the first difference of the variables confirm that the series are indeed I (1) variables, that is, integrated of order one (see Table A.2 in the appendix). As a first step, these results suggest that a meaningful relationship between the real effective exchange rate and its main determinants could exist.

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[10] Alternatively, the effect of government consumption on the real effective exchange rate will depend on whether consumption is biased towards tradables or non-tradable goods.

Table 1: Panel Unit Root Tests

Panel: WAEMU	Series	t-bar statistic	Hadri
	reer	-1.084	8.218[0.000]
	prod	-1.646	5.102[0.000]
	tot	-2.123	3.920[0.000]
	govt	-2.427	0.767[0.009]
	oda	-2.210	5.333[0.000]

For t-bar test, Ho: Unit root; Hadri, Ho: Stationarity. [ ] are P-values. For 1%, 5% and 10% significance levels; the *t-bar* critical values are -2.93, -2.69 and -2.57, respectively.

In view of the non-stationarity of the variables, it is important to employ the panel cointegration test technique proposed by Pedroni (1997, 1999). The results summarised in Table 2 show that two out of the four within-dimension based tests are significant at the 5 percent level, while the other two are significant at 10 percent. On the other hand, all the three between-dimension based tests are statistically significant at the 5 percent level. Pedroni (1997, 2004) show that all the 7 statistics (see appendix) produce reliable estimates when T is as large as 100, while the Group ADF-Statistic, Panel V-Statistic and the Panel Rho-Statistic produce more stable estimates for smaller samples. On the basis of these results, we conclude that a long-run stable relationship exists between the real effective exchange rate and its main determinants for WAEMU.

With this preliminary insight, we now proceed to estimate the long-run REER equation for two different specifications. First, we estimate the REER without the aid term. Second, we augment the first equation by including the aid term. The latter serves as our main specification.

Table 2: Panel Cointegration Tests

Series: $reer$, $prod$, tot , $govt$, oda Ho: No Cointegration	***Pedroni Residual Tests*** ***Test statistic***	 ***p-value***
Within-dimension tests		
Panel v-Statistic	-2.1257	0.0417
Panel rho-Statistic	3.4042	0.0012
Panel PP-Statistic	1.8016	0.0787
Panel ADF-Statistic	1.9952	0.0545
Between-dimension tests		
Group rho-Statistic	4.1988	0.0001
Group PP-Statistic	2.1375	0.0406
Group ADF-Statistic	3.0521	0.0038

Model includes deterministic intercept and trend.

B. Does Foreign Aid Cause an Appreciation of the Real Exchange Rate?

The results presented in Table 3 show that the coefficient on the foreign aid variable is significantly positive for both estimators[11]. Thus, the "Dutch Disease" prediction is confirmed for the WAEMU.

Adenauer and Vagassky (1998), Elbadawi (1999), and Rajan and Subramanian (2009) all find evidence of the 'Dutch disease' phenomenon. Therefore, the finding that foreign aid led to an appreciation of the real effective exchange rate is in line with this evidence. In contrast, Ouattara and Strobl (2004) find that foreign aid is associated with a depreciation of the real exchange rate.

The empirical evidence for the other variables is broadly in line with their theoretical predictions. As for the adjustment of the real effective exchange rate to the equilibrium, no adjustment will take place if, ϕ, the adjustment parameter

[11] To avoid the problem of cross-sectional dependence in the panel, the standard errors of the DFE estimates are corrected for cross-sectional error variances.

is *zero*. However, the significantly negative coefficient of the adjustment parameter suggests that the real effective exchange rate reverts to its long-run position after any shock or disturbance. This strengthens the evidence of a cointegrating relationship (long-run relationship), which was earlier established using the Pedroni residual-based tests. More importantly, this implies that the stability assumption of the model, that is, $\phi < 0$ is fully satisfied.

To check whether these results are sensitive to the lag structure of the variables, we re-assessed the results in Table 3 using a lag length of up to 2 for the dependent variable. The results presented in Table 4 suggest that the signs are robust to the lag length and also to the choice of econometric approach.

Furthermore, in all the PMG specifications the Hausman statistics show that the long-run homogeneity assumption imposed on the coefficients is adequate and fully satisfied. On this basis, we conclude that these countries are not too heterogeneous to be pooled, and that the PMG is preferred to the MG (see Table A.5 in the appendix for MG estimates).

As Baltagi *et al*. (2000) discussed, Fixed Effects models are subject to a simultaneous equation bias from the endogeneity between the error term and the lagged dependent variable. Here, the Hausman statistics under the DFE in Tables 3 and 4 indicate that the simultaneous equations bias is not present for these data, implying that the DFE is also preferred to the MG. Like the PMG estimator, the DFE estimator assumes that the long-run estimates are the same across countries.

Table 3: Panel regression of Dutch disease effect of Aid using annual data (1975 -2005)

Dependent Variable: log of real effective exchange rate (***LREER***)				
	PMG		**DFE**	
Variables	**1**	**2**	**3**	**4**
prod	0.639***	0.200*	0.284*	0.242***
	(0.157)	(0.107)	(0.157)	(0.062)
tot	0.269**	0.256***	0.108	0.252***
	(0.127)	(0.102)	(0.164)	(0.078)
govt	0.269**	0.268***	0.087	0.191**
	(0.109)	(0.090)	(0.127)	(0.075)
oda		0.319***		0.328***
		(0.082)		(0.068)
Adjust. (ϕ)	-0.298**	-0.389***	-0.277***	-0.362***
	(0.136)	(0.112)	(0.079)	(0.085)
	Diagnostics			
Log likelihood	163.131	173.429		
Hausman-test	5.25	7.18	0.01	0.06
	[0.386]	[0.304]	[1.000]	[0.999]
Observations	217	217	217	217

The model is estimated for both PMG and DFE using ARDL (1, 1, 1, 1, 1) specification. Note: Robust Standard errors are in parentheses ( ). Numbers in brackets [ ] indicate p –values. * indicates that a coefficient is significant at 10 percent level; ** indicates 5 percent significance level; *** indicates significance at 1 percent level.

However, we suspect that the large devaluation of the CFAF in 1994 will have some implications for the real exchange rate and the competiveness of WAEMU countries. Therefore, the results presented in Tables 3 and 4 may suffer from omitted variable bias. To deal with this issue, we included a shift dummy in the main specification.

Table 4: Panel regression of Dutch disease effect of Aid using annual data (1975 -2005)

Dependent Variable: log of real effective exchange rate (*LREER*)				
	PMG		**DFE**	
Variables	**1**	**2**	**3**	**4**
prod	0.808***	0.250***	0.386***	0.309***
	(0.148)	(0.104)	(0.169)	(0.094)
tot	0.617***	0.287***	0.164	0.280***
	(0.141)	(0.099)	(0.158)	(0.078)
govt	0.023	0.240***	0.037	0.105
	(0.104)	(0.085)	(0.127)	(0.105)
oda		0.287**		0.236***
		(0.076)		(0.066)
Adjust. (ϕ)	-0.275***	-0.333***	-0.266***	-0.315***
	(0.101)	(0.099)	(0.060)	(0.063)
	Diagnostics			
Log likelihood	228.278	235.997		
Hausman-test	2.12	4.85	5.30	7.14
	[0.832]	[0.564]	[0.623]	[0.387]
Observations	217	217	217	217

The model is estimated for both PMG and DFE using ARDL (2, 1, 1, 1, 1) specification. Note: Robust Standard errors are in parentheses ( ). Numbers in brackets [ ] indicate p –values. * indicates that a coefficient is significant at 10 percent level; ** indicates 5 percent significance level; *** indicates significance at 1 percent level.

C. The 1994 Nominal Devaluation

Between 1986 and 1993, CFA zone, including WAEMU countries, experienced a gradual appreciation of its currency. The appreciation of the French franc, coupled with a series of commodity price shocks forced the economies to devalue the CFA franc in January 1994. Before the devaluation, however, the CFA franc maintained a fixed parity with the French franc. During this period, the France pursued the 'franc fort' strategy aimed at stabilizing the French economy after an expansionary policy adopted by the socialist government in

1981-83. In addition, France attempted to gain some monetary leadership in Europe along with Germany, preparatory to the European Monetary Union. According to Blanchard and Muet (1993), this ‘franc fort’ strategy pursued over this period was largely responsible for the appreciation of the CFA. Given the scale of the devaluation (50 percent), it is important to control the effect of this policy shift in our estimations.

The results presented in Table 5 show that adding a shift dummy to the equations increased the significance of the estimates considerably. Perhaps, the most striking feature of this result is that the effect of the devaluation on the real exchange rate was fairly large. However, with the inclusion of a shift dummy, both estimators now suggest a slower speed of convergence to the equilibrium - of around 20 percent per year.

All told, a number of conclusions can be drawn from the totality of these results. First, there exists a long run stable relationship between real exchange rate and its main determinants. Second, the long run homogeneity assumption holds even after accounting for the nominal devaluation of the exchange rate. Third, once account is taken of other factors, foreign aid inflows led to an appreciation of the real exchange rate. Fourth, the quality of the results is not affected by the lag structure of the model.

Table 5: Panel Regression of Dutch Disease Effect of Aid Using Annual Data (1975-2005): Devaluation Dummy Included

Dependent Variable: real effective exchange rate (*REER*)				
	PMG	**DFE**	**PMG**	**DFE**
Variables	**1**	**2**	**3**	**4**
prod	0.336***	0.179***	0.261***	0.207**
	(0.120)	(0.069)	(0.093)	(0.087)
tot	0.280***	0.247***	0.359***	0.249***
	(0.104)	(0.084)	(0.093)	(0.082)
govt	0.321***	0.191***	0.158**	0.217**
	(0.093)	(0.026)	(0.078)	(0.093)
oda	0.113***	0.234***	0.161***	0.184***
	(0.041)	(0.058)	(0.044)	(0.060)
dev. dummy	-0.111**	-0.227***	-0.132**	-0.236***
	(0.055)	(0.062)	(0.058)	(0.059)
Adjust. (ϕ)	-0.238***	-0.285***	-0.224***	-0.270***
	(0.053)	(0.058)	(0.043)	(0.049)
	Diagnostics			
Log likelihood	273.826		221.272	
Hausman-test	1.131	0.44	2.85	0.62
	[0.932]	[0.996]	[0.764]	[0.984]
Observations	217	217	217	217
Specification	*ARDL (1, 1, 1, 1, 1)*		*ARDL (2, 1, 1, 1, 1)*	

Robust Standard errors are in parentheses ( ). Numbers in brackets [ ] indicate p – values. * indicates that a coefficient is significant at 10 percent level; ** indicates 5 percent significance level; *** indicates significance at 1 percent level.

V. Summary and Conclusion

This paper has systematically re-examined the relationship between foreign aid and the real effective exchange rate using a multivariate econometric approach that relied on a sample of 7 countries in West African the Economic and Monetary Union. Even so, we believe this paper has a few limitations. To the extent that the paper uses a relatively small sample size, the results need to be interpreted with caution. Again, since direct comparison cannot be made on studies that apply different specifications, estimation techniques and sample size, we recognise this as one of the limitations of the study.

Before estimating the real exchange rate equation, the paper examined the variables for stationarity and cointegration, using both the traditional time series and dynamic panel data techniques. The panel unit root tests showed that the variables are non-stationary while the panel cointegration tests provided a strong support for cointegration. The presence of cointegration, re-enforced by the model's stability, permitted a long-run interpretation of the estimates of the regressions.

In line with the theoretical prediction of the Dutch Disease model, both estimators show that foreign aid led to an appreciation of the real exchange rate during the 1975-2005 period. Also, other factors, for example, labour productivity, terms of trade and government consumption of non-tradable goods were associated with an appreciation of the real exchange rate.

A key policy message from this study is that foreign aid should be targeted in such a way that it is not spent on non-traded goods. This way, the real appreciation effect of foreign aid inflows which can reduce the region's competitiveness (and growth) can be easily avoided. Again, measures should be taken to use foreign aid inflows to import investment goods. Such measures can help to quicken the rate of industrialization and export expansion in West Africa and WAEMU, in particular. In addition, donors should consider the absorptive capacity of recipients before scaling-up aid.

APPENDIX

Table A.1 Descriptive Statistics

Variable	Mean	Standard Deviation	Observations
Real Effective Exchange Rate	146.13	43.66	217
Foreign aid (% of GDP)	11.51	5.34	217
Terms of Trade	125.34	40.44	217
Government consumption (% of GDP)	14.75	4.72	217
Real per capita GDP (in relation to trade partners	1.46	0.64	217

Table A.2: Correlation Statistics

	REER	PROD	TOT	GOVT	ODA
REER	1.00	0.55	0.37	0.47	0.13
PROD	0.55	1.00	0.29	0.23	-0.24
TOT	0.37	0.29	1.00	0.04	0.36
GOVT	0.47	0.23	0.04	1.00	0.21
ODA	0.13	-0.24	0.36	0.21	1.00

Table A.3: Unit Roots Tests for Individual Countries

Country	Series	ADF	KPSS
Benin	reer	-1.362	0.169**
	oda	-1.564	0.152**
	prod	-0.722	0.159**
	tot	-2.633	0.090
	govt	-4.937***	0.099
Burkina Faso	reer	-2.696	0.620**
	oda	-2.304	0.351*
	prod	-2.483	0.186**
	tot	-0.853	0.172**
	govt	-5.742***	0.082
Côte d'Ivoire	reer	-2.490	0.083
	oda	-1.423	0.151**
	prod	-1.722	0.130*
	tot	-2.735	0.135*
	govt	-2.987	0.518**
Mali	reer	-2.789	0.136*
	oda	-2.247	0.162**
	prod	-1.986	0.178**
	tot	-1.977	0.149**
	govt	-2.270	0.133*
Niger	reer	-2.414	0.651**
	oda	-2.421	0.119*
	prod	-1.859	0.115
	tot	-3.151	0.142*
	govt	-2.390	0.096
Senegal	open	-3.003	0.122*
	reer	-1.829	0.095
	oda	-3.327*	0.173**
	prod	-0.701	0.151**
	tot	-2.561	0.083
	govt	-1.548	0.103
Togo	reer	-3.226*	0.066
	oda	-2.181	0.185**
	prod	-2.048	0.111
	tot	-2.422	0.129*
	govt	-5.465***	0.097

For ADF, H0: Unit root; KPSS, Ho: Stationarity. ***, **, and * is 1%, 5%, 10 % significance level, respectively. For 1 %, 5 % and 10 % significance levels, the ADF critical values are -4.30, -3.57 and -3.22, respectively; KPSS critical values are 0.216, 0.46 and 0.119, respectively.

Table A.4 Panel Unit Root Tests Using Various Techniques: First Difference

Variables	*Ho: Unit root*					*Ha: Stationarity*
	LLC	**Breitung**	**IPS**	**Maddala-Wu**	**Choi**	**Hadri**
$\Delta reer$	-4.730 [0.000]	-5.534 [0.000]	-5.218 [0.000]	20.897 [0.000]	176.474 [0.000]	0.377 [0.354]
Δoda	-4.932 [0.000]	-5.364 [0.000]	-4.574 [0.000]	18.421 [0.000]	18.421 [0.000]	0.250 [0.401]
$\Delta prod$	-1.837 [0.033]	-3.517 [0.000]	-1.823 [0.034]	6.486 [0.039]	17.448 [0.000]	0.715 [0.237]
Δtot	-1.790 [0.037]	-3.853 [0.000]	-4.459 [0.000]	17.420 [0.000]	18.442 [0.000]	1.145 [0.126]
$\Delta govt$	-3.612 [0.000]	-3.872 [0.000]	-2.591 [0.005]	9.455 [0.009]	8.786 [0.012]	0.168 [0.433]

Model includes individual effects and individual linear trends. P-values are in brackets.

Table A.5: Mean Group (MG) Estimates

	Equations					
	1	2	3	4	5	6
prod	0.088	0.441	0.161	0.115	0.917	0.426
	(0.742)	(0.505)	(0.238)	(1.038)	(0.581)	(0.400)
tot	0.593	0.650**	0.288	0.488	0.622*	0.274
	(0.427)	(0.312)	(0.157)	(0.500)	(0.346)	(0.201)
govt	0.055	0.127	-0.118	0.020	0.191	-0.080
	(0.269)	(0.273)	(0.127)	(0.331)	(0.316)	(0.179)
oda		-0.053	-0.058		0.031	0.013
		(0.199)	(0.068)		(0.213)	(0.091)
dev.			-0.426***			-0.444***
dummy			(0.037)			(0.052)
Adjust.	-0.338***	-0.388***	-0.694***	-0.300***	-0.347***	-0.615***
(ϕ)	(0.035)	(0.027)	(0.063)	(0.042)	(0.036)	(0.061)
ARDL	*(1, 1, 1)*	*(1, 1, 1, 1)*	*(1,1,1,1,1)*	*(2, 1, 1)*	*(2, 1, 1, 1)*	*(2,1,1,1,1)*

A.6: Unit Root and Cointegration Tests

I. Unit Root Tests

The Im et al. (2003) test (IPS t-bar test) is based on the null hypothesis of a unit roots and follows the augmented Dickey-Fuller approach which relies on the following equation:

$$\Delta y_{it} = \rho_i y_{t-1} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta y_{it-j} + \alpha_i + \gamma_i t + \varepsilon_{it} \qquad \text{for t=1… T; i = 1 …N}$$

On the other hand, the test proposed by Hadri (2000) is based on the null of stationarity and the following regression:

$$y_{it} = \alpha_i + \gamma_i t + \sum_{t=1}^{T} \mu_{it} + \varepsilon_{it}$$

II. Cointegration

The Pedroni's Cointegration tests are based on the following regression equation[12]:

$$y_{it} = \alpha_i + \delta_i t + \beta_{1i} x_{1it} + \ldots + \beta_{ki} x_{kit} + \varepsilon_{it}$$

for $t = 1, \ldots, T \; ; \; i = 1, \ldots, N$

[12] Some studies that have applied Pedroni's test include; Maeso-Fernandez *et al.* (2006), Abdih and Tsangarides (2006) and Roudet et al. (2007), Drine and Rault (2003).

where T is the number of observations over time; N is number of countries in the panel; k is the number of regressors. The slope coefficients $\beta_{1i},...,\beta_{ki}$ are allowed to vary across countries, while $\alpha_i \, and \, \delta_i t$ are country-specific intercepts and deterministic trends, respectively. These tests are based on the absence of cross-sectional correlation and are constructed from the cointegrating residuals in the above equation.

The first category of these statistics comprises four within dimension based tests which have an alternative hypothesis of common autoregressive [AR] coefficients. These tests pool the AR coefficients across different sections of the panel for the unit-root tests on the residuals. Practically, the tests are implemented by calculating the average test statistics for cointegration in the times series framework across the different sections. These statistics include:

$$Z_\rho^w = \left(\sum_{i=1}^{N} \sum_{t=1}^{T} L_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} L_{11i}^{-2} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right) : \quad \text{Panel Rho}$$

$$Z_t^w = \left(\tilde{s}_{NT}^{*2} \sum_{i=1}^{N} \sum_{t-1}^{T} L_{11i}^{-2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} L_{11i}^{-2} \left(\hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*} \right) : \text{Panel ADF}$$

$$Z_{pp}^w = \left(\tilde{\sigma}^2 \sum_{i=1}^{N} \sum_{t=1}^{T} L_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} L_{11i}^{-2} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right) : \text{Panel PP}$$

$$Z_v^w = \left(\sum_{i=1}^{N} \sum_{t=1}^{T} L_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1} : \qquad \text{Panel V}$$

The second category includes three tests that are based on between dimension effects, with an alternative hypothesis of individual autoregressive coefficients. This involves averaging the AR coefficients for each of the panel for unit-root test on the residuals.

$$Z_{\rho}^{B} = \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{t=1}^{T} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_{i} \right): \quad \text{Group Rho}$$

$$Z_{t}^{B} = \sum_{i=1}^{N} \left(\hat{\sigma}_{i}^{2} \sum_{t=1}^{T} \hat{e}_{it-1}^{2} \right)^{-1} \sum_{t=1}^{T} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_{i} \right): \text{ Group ADF}$$

$$Z_{pp}^{B} = \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{s}^{*2} \hat{e}_{it-1}^{*2} \right)^{-1} \sum_{t=1}^{T} \left(\hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*} \right): \quad \text{Group PP}$$

with,

$$\hat{\lambda} = \frac{1}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^{t} \hat{\mu}_{it} \hat{\mu}_{it-s} ,$$

$$\hat{s}_i^2 = \frac{1}{T} \sum_{t=s+1}^{t} \hat{\mu}_{it}^2 , \quad \hat{\sigma}^2 = s_i^2 + 2\hat{\lambda}_i ,$$

$$\tilde{\sigma}_i = \hat{s}_i^2 + 2\hat{\lambda}_i , \quad \tilde{\sigma}_{NT}^2 = \frac{1}{T} \sum_{i=1}^{N} \hat{L}_{11i}^{-2} \hat{\sigma}_i^2 ,$$

$$\hat{s}_i^{*2} = \frac{1}{T} \sum_{t=s+1}^{t} \hat{\mu}_{it}^{*2} , \quad \hat{s}_{NT}^{*2} = \frac{1}{T} \sum_{t=s+1}^{t} \hat{s}_{it}^{*2} ,$$

$$\hat{L}_{11i}^2 = \sum_{t=1}^{T} \hat{\eta}_{it}^2 + \frac{2}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^{t} \hat{\eta}_{it} \hat{\eta}_{it-s} ,$$

$$\hat{e}_{it} = \hat{\rho} \hat{e}_{it-1} + \hat{\mu}_{it} = \hat{\rho} \hat{e}_{it-1} + \sum_{k=1}^{K_i} \hat{\gamma}_{ik} \Delta \hat{e}_{it-k} + \hat{\mu}_{it} ,$$

$$\Delta y_{it} = \sum_{m=1}^{M} \hat{b}_{mi} \Delta X_{mit} + \hat{\eta}_{it} ,$$

In the above representation, L_i represents the i^{th} component of the cholesky decomposition of the residual variance-covaraince matrix, $\hat{\lambda}$ and $\tilde{\sigma}_{NT}^2$ are autocorrelation parameters, σ_i and s_i^2 are the contemporaneous and long-run individual variances.

Figure A.1: Real Exchange Rate Movements for 7 WAEMU countries, 1975-2005

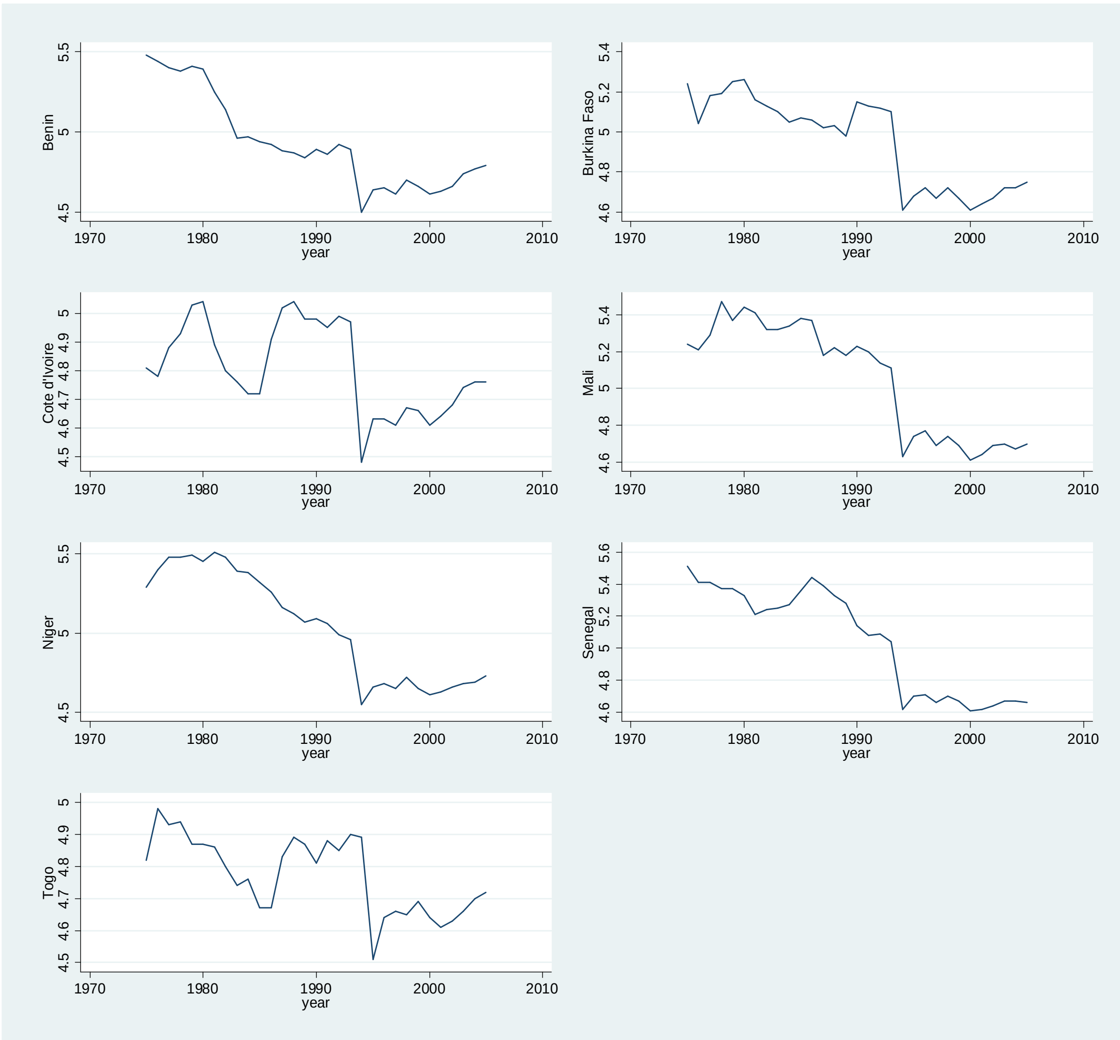

Figure A.2: Foreign Aid Inflows to 7 WAEMU countries, 1975-2005

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