Convergence and the Short-Run Dynamics of Prices in a Currency Union*

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Abstract

This paper investigates the formation of prices for members of a currency union in west and central Africa, the Communauté Financière Africaine (CFA) Zone. We use fully-modified estimation methods for nonstationary time series and panel data to establish the degree of long-run transmission between the French price level (the anchor of the CFA currency union) and the price levels of Zone members. We also investigate the short-run dynamics of Zone inflation rates by estimating a multivariate error-correction model. We find a significant long-run linkage between French and Zone price levels that for most members is not significantly different from unity. Moreover, Zone inflation rates respond significantly to the divergence of prices from their long-run equilibrium level, but the speed of adjustment is relatively slow.

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1 Introduction

The debate on the merits of alternative exchange rate regimes and their impact on the macroeconomic performance of developing countries has a long history. One of the purported advantages of a fixed exchange rate regime is that it imposes monetary discipline by tying a small open economy's price level to that of a large (industrial) economy that serves as the anchor for the system. Anchoring is thought to be especially important for countries embarking on disinflation programs after prolonged periods of sustained inflation. One important factor in the success of such programs is the credibility of the fixed exchange regime: frequent parity realignments negate the discipline benefits imposed by the fixed peg.

The experience of a group of west and central African economies provides an example of a fixed exchange rate system that has avoided the frequent devaluations that have plagued fixed exchange regimes elsewhere. The members of the Communauté Financière Africaine (CFA) Zone have maintained a fixed peg vis-à-vis the French franc for the past fifty years. Their currency (CFA franc) underwent only one nominal devaluation in January 1994.[1] Central to evaluating the benefits of a fixed exchange regime, such as that maintained by the members of the CFA Zone (the Zone hereafter), is the degree to which price changes at the 'centre' of the union (France) are transmitted to the members of the Zone and the convergence of the price levels of member countries to that of the center. These issues are important to our understanding of the price stability benefits of fixed exchange regimes.

[1] The CFA Zone comprises of 14 countries in West and Central Africa that belong to two monetary unions with separate central banks: the Banque Centrale des Etats de l'Afrique de l'Ouest (BCEAO) in West Africa and the Banque des Etats de l'Afrique Centrale (BEAC) in Central Africa. The two central banks issue a separate currency but the two currencies are equivalent in value and are collectively known as the CFA franc. During the past 50 years, Zone members have formed a currency union with France. France guarantees the convertibility of the CFA franc, hence no active black market for foreign exchange has developed in these countries. In addition, the Zone shares several other specific features unique to developing countries, discussion of which is beyond the scope of this paper (see Boughton (1991), for additional details and Clément (1996), for the 1994 devaluation).

Moreover, they are of special significance to the African economies, in general, and the CFA Zone, in particular. As several European currencies (including the French franc) are replaced by the euro, a debate is taking place on the future CFA franc peg to the euro and whether other African monetary unions ought to follow by adopting the euro as an anchor. Despite their significance, the empirical literature has paid relatively little attention to these issues.

Related to these issues is the hypothesis of purchasing power parity (PPP). In a system where the nominal exchange rate has not been allowed to fluctuate to adjust to divergences in the path of domestic and foreign prices, the degree to which foreign price changes are transmitted to those of the member countries is central to the PPP hypothesis. The literature on PPP has witnessed a tremendous resurgence in recent years; general surveys can be found in Froot and Rogoff (1995) and Rogoff (1996).[2] While the PPP literature has seen an exponential growth, the issue of the transmission of foreign price shocks for countries pursuing fixed exchange rates has received little attention.

The main objective of this paper is to investigate the link between prices at the center of the CFA union and the price levels of Zone members. We accomplish our objective by estimating a long-run cointegrating relationship between domestic (Zone) prices, French prices and the domestic monetary aggregate using fully-modified estimation methods for nonstationary time series and panel data. Our results reveal a significant long-run cointegrating relationship between French and Zone prices. The long-run linkage between the French and Zone price levels is not significantly different from one for five of the eight Zone

[2]The vast majority of the contributions to the PPP literature concern the developed economies. Some recent studies, however, use panel data from a large number of countries including both developed as well as developing economies (e.g. Frankel and Rose (1996) and Oh (1996)). Cheung and Lai (2000) test the mean reverting dynamics of real exchange rates for a large cross section of countries which are subsequently grouped according to geographic region, level of economic development, and exchange rate arrangement. There are comparatively few studies on developing countries alone. McNown and Wallace (1989) test PPP for four economies that have experienced high inflation using the Engle/Granger two-step procedure. Phylaktis and Kassimatis (1994) investigate the existence of unit roots in parallel-market real exchange rates for eight Pacific Basin economies.

members in our sample. In addition, a domestic monetary aggregate enters the long-run relationship significantly for six of the countries. The short-run dynamics of Zone inflation rates are studied by estimating a multivariate error-correction model based on the fully-modified least squares estimates of the cointegrating vector. Systems estimation enables one to test cross-equation restrictions for Zone members, especially the restriction concerning the speed of adjustment of prices to their long-run equilibrium. We find that there is significant convergence to the long-run price level but for the majority of Zone members the speed of adjustment is rather slow.

The paper is organized as follows. The following section provides a brief review of the relevant literature and a description of the methodology we use to test the main hypotheses of interest. Section 3 discusses our results. The final section offers a brief conclusion.

2 Methodology

The focus of our study is the link between the French price level and prices of Zone members, an issue central to the working of a currency union such as the CFA Zone. There are many excellent discussions on the merits and drawbacks of fixed exchange rate regimes including Obstfeld and Rogoff (1995). The extent, however, to which foreign price shocks are transmitted to domestic price levels for countries pursuing fixed exchange regimes has, on the whole, been neglected by the empirical literature. Edwards (1993) finds that countries that maintained a fixed exchange rate in 1980 tended to experience a lower inflation rate during the subsequent decade. Edwards and Losada (1994) investigate whether the fixed peg maintained by two Central American nations (Guatemala and Honduras) vis-à-vis the U.S. dollar for nearly seven decades imposed macroeconomic discipline by 'tying' the domestic inflation rate to that of the United States. They are unable to find evidence supporting this conjecture. Fielding and Bleaney (2000) empirically test two mechanisms by which fixed exchange rates reduce inflation in developing economies. These are the monetary discipline imposed by the requirement to maintain the fixed peg and the balance of payments effect: excess money balances build up as a balance of payments deficit rather than higher prices. They find partial evidence favorable to both effects. They, however, do not model the short-run price dynamics of fixed exchange rate regimes but use cross section data for 80 developing economies during 1980-89.

Of more direct relevance to our study is Honohan (1992) who estimates univariate price adjustment equations where Zone inflation rates depend only on the French inflation rate, the (lagged) gap between Zone and French price levels and deterministic time components. He finds a relatively high degree of contemporaneous pass-through between the French and Zone inflation rates and a small catch up coefficient in closing the gap between French and Zone price levels. This framework, however, is not well suited to address the issue at hand; it assumes that there is a one-to-one long-run relationship (cointegrating vector) between the French and Zone price levels and does not adequately model the dynamics of Zone inflation rates. Moreover, single-equation estimation methods ignore contemporaneous correlation among members of the CFA currency union and do not permit cross-country hypotheses to be directly tested. These issues are tackled in the remainder of this paper. Fully-modified estimation methods (Phillips and Hansen 1990) are used to investigate the long-run cointegrating relationship between the price level of Zone members, the domestic money stock and the French price level. In addition the short-run dynamics of Zone inflation rates are modeled via a multivariate error-correction model.

The proposed long-run model of domestic prices for each member of the CFA currency union is:

$$p_t = \gamma_0 + \gamma_1 \, p_t^* + \gamma_2 \, m_t + \delta \, d_t + \varepsilon_t \tag{1}$$

where p_t is the log of the domestic (CFA Zone) price level, p_t^* is the log of the foreign price level (approximated here by the French price level), m_t is the log of the domestic monetary aggregate and d_t is a vector of deterministic time dependent variables. Equation (1) can be thought of as a reduced-form equation from a traded/nontraded goods monetary model of a small open economy under fixed exchange rates [see Ch. 7, Dornbusch (1980) for a description].[3]

The inclusion of the domestic monetary aggregate in (1) may seem surprising given the well-known proposition that under a system of fixed exchange rates and capital mobility (characteristics of the CFA Zone) the central bank loses control over the domestic money supply (see Obstfeld and Rogoff (1995)). In practice, central banks may have considerable leeway in setting monetary targets even under rigidly pegged exchange rate systems. Related to the question of an independent monetary policy is the issue of whether a meaningful distinction can be made between the monetary aggregate for each Zone member, given the existence of a single currency for the Zone. Honohan (1991) and Savvides (1998) have explained that the two CFA regional central banks engage in a monetary programming exercise that results in separate monetary targets for each member of the Zone. The latter study has also shown that the monetary authorities of the three most important Zone economies (Cameroon, Côte d'Ivoire and Gabon) have been able to pursue an independent monetary policy. Each of the CFA central banks pools the foreign exchange reserves of its members and these reserves (except for balances required for operational purposes) are maintained in an Operations Account (*compte d'opérations*) with the French Treasury. A number of commentators have pointed out that until the early 1990s, access to the French overdraft in the *compte d'opérations* was automatic, thus effectively relaxing the constraint on monetary growth imposed by a fixed exchange rate regime.

Two dummy variables are included in d_t to permit changes in the long-run relationship stemming from the 1994 nominal devaluation of the CFA franc. The first ($D1$) takes the value of 1 for 1994:1 and beyond while the second (DB) is coded 1 for 1994:1 and zero for all

[3] Fielding and Bleaney (2000) derive an equation similar to (1) for the nontraded-exportable-importable model. In addition to the money stock and foreign prices, the terms of trade are a determinant of domestic prices. Since the terms-of-trade effect is consistently statistically insignificant, they do not consider it in their discussion.

other observations. Including these two variables is tantamount to the innovational-outlier model of Perron (1990). Also included in d_t is a time trend, D_T. The maintained hypothesis in this study is that (1) represents the long-run relationship between domestic prices, French prices, and the domestic money supply. Furthermore, by including the dummy variables we have allowed for the trend to experience permanent as well as an event shock due to the CFA franc devaluation in 1994.[4]

Although there is no statistical distinction between dependent and independent variables in a cointegrated system, in this case an economic argument can be made for normalizing the cointegrating vector using the domestic price level. In view of the relative size difference between the French and Zone economies, few would argue that changes in CFA prices would materially change French prices and theory suggests that control of the money supply affects domestic prices. Borland and Ouliaris (1994) make a similar argument in their model of Australian union membership. They interpret the cointegrating relationship in their model as a structural equation in a system of simultaneous equations. If there is more than one cointegrating vector, there needs to be at least two predetermined variables excluded from (1) in order for it to be identified. The results of the Johansen tests, to be discussed shortly, suggest the existence of only one cointegrating vector in every country except Togo, which has none. Interpreted as a structural equation, a means of consistently estimating the model's parameters, where there is correlation between the explanatory variables and the true residuals, should be used.

Among the many procedures by which one could estimate the parameters of this model, the fully modified least squares (FMOLS) es-

[4] In modeling prices in (1), our intention is not to capture the various structural determinants of price fluctuations for Zone members. Rather, our intent is to estimate as parsimonious a model as possible that addresses the main issues of interest, the link between French and Zone price levels. Moreover, given our interest in the time series properties of prices, it is necessary to use data of higher frequency than annual. As Honohan (1992) points out, higher frequency non-monetary data for the CFA Zone are either nonexistent or of very poor quality.

timator (Phillips and Hansen 1990) is particularly attractive. The FMOLS estimator essentially corrects the variables in (1) for system endogeneity due to cointegration as well as for serial correlation caused by the unit roots in p^* and m. FMOLS yields t-ratios that are asymptotically normally distributed and is not dependent on the correct choice of lag length of the underlying vector autoregression. According to Borland and Ouliaris (1994), the FMOLS estimator permits inference based on normal distribution theory "by means of a nonparametric correction to the data that effectively purges any long-run dependence between the true residuals of the cointegrating regression and the innovations of the explanatory variables." The FMOLS estimator for the model in (1) that includes exogenous deterministic components is explained in the Appendix which also provides additional information on the statistical properties of the estimator. As will be explained in the following section, equation (1) is estimated for each member of the CFA Zone separately as well as for all the CFA countries as a panel. Details on computation of the FMOLS panel estimator are also provided the Appendix.

Residuals generated from each cointegrating relation (estimates of ε_t) represent deviations from long-run equilibrium and can be used to estimate the short-run dynamics of CFA zone inflation using the following error-correction model:

$$\begin{aligned}\Delta p_{it} \quad = \quad & \delta_i - \lambda_i ecm_{it-1} + \sum_{j=1}^{k_i} \pi_{ij} \Delta p_{it-j} + \sum_{j=0}^{k_i} \mu_{ij} \Delta m_{it-j} \\ & + \sum_{j=0}^{k_i} \pi^*_{ij} \Delta p^*_{it-j} + \delta_{1i} D1_i + \delta_{Bi} DB_i + \delta_{Ti} D_T + \nu_{it} \qquad (2)\end{aligned}$$

where ecm_{it-1} is the residual from estimating (1) using FMOLS, $i = 1, ..., 8$ is the index indicating the CFA country, ν_{it} $iid(0, \sigma_i^2)$, and $E(\nu_{it}\nu_{js}) = \sigma_{ij}$ for $s = t$ and 0 otherwise. The model in (2) is estimated via (unbalanced) seemingly unrelated regressions (SUR). The lag length for each country, k_i, is chosen by minimizing the AIC criterion from least squares estimation for each country. The maximum lag length considered for model selection is 8.[5] Since the number of observations across countries is different, the resulting panel is unbalanced and this has been accounted for in the estimation of the SUR.[6] The systems-of-equations approach also allows one to test cross-equation restrictions on the estimated parameters of interest. Of special interest is the estimates of the error correction parameters in the model, λ_i. This parameter is important because it measures the speed of adjustment of the system to disequilibrating shocks. The coefficients are expected to be negative and larger (absolute) values indicate faster adjustment to economic shocks.

Another possible approach is to estimate the cointegrating vector(s) in (1) using recently developed methods for cointegrated panel data. The coefficients on p^* and m in the panel approach measure the *long-run average* relationship among the variables across the Zone. This could be useful if one views the CFA Zone members as a regional block. Recent contributions by Phillips and Moon (1999) permit such a treatment while allowing for different individual country effects among members of the panel. As shown below, estimates of the short-run dynamics based on error correction terms generated from panel regressions are quite similar to those from the disaggregated methods.

3 Empirical Results

The data are measured quarterly for each of the eight members of the CFA Zone: Burkina Faso, Cameroon, Congo, Côte d'Ivoire, Gabon, Niger, Senegal and Togo. These are the Zone members for which continuous data on the necessary variables exist. These countries include the most important members of the Zone which, during 1990-97, ac-

[5] AIC is used specifically because it tends to overparameterize a model. Since our goal is to test hypotheses based on the error correction model, it is important not to under fit the model. Omission of relevant lags invalidates subsequent hypothesis tests whereas overfitting merely results in loss of statistical power. Another SUR model was estimated using lags chosen by the more aggressive SBC criterion. As expected the lags selected by the SBC were much shorter, but its use resulted in no substantive changes to the results.

[6] Unbalanced SUR was programmed by the authors using GAUSS 3.5.

counted for 87 percent of Zone GDP and 91 percent of Zone exports. It should be noted that there are substantial gaps in the monthly data for some of the necessary variables rendering the use of such data infeasible. Price levels are measured by the CPI and the domestic monetary aggregate is the M1 stock. Given that the monetary aggregate is measured at the end of the quarter while prices are calculated as period averages, our measure of the monetary aggregate is a two-period moving average of M1. The data cover the period 1973:1-1998:3, except for Cameroon (1973:1-1998:1) and Congo (1973:1-1997:1) due to the lack of recent CPI data for these two countries. All data are from the International Financial Statistics (IFS) CD-ROM of the IMF.

Figures 1 and 2 portray the ratio of each CFA country's CPI to that of France. For most Zone members the visual evidence points to a long-run co-movement between the two price levels, though substantial short run variations are also evident. Also evident is the jump coincident with the 1994 devaluation. In the remainder of this paper we explore systematically and statistically both the long-run and short-run relationship between Zone and French price levels.

3.1 Long-run Empirical Model

In Table 1 the FMOLS results from estimating (1) for each country in our sample are presented. The FMOLS parameter estimates are computed using a Gauss-Weierstrauss kernel with a bandwidth of 8.[7] Several other choices of bandwidth and nonparametric kernel were combined with no substantive effect on our main results.[8] Also reported

[7]The bandwidth parameter in the time series kernel is the number of autocovariances over which the sample autocovariances are smoothed. As a practical matter it should be large enough to capture most of the nonzero sample autocovariances but small enough not to dilute the substantially nonzero autocovariances with a large number of near zero terms. In this paper, a bandwidth of 8 represents two years of data and is the same number used as a maximum lag in the error correction model selection process. Phillips (1995) gives some general guidelines as to the rate at which the bandwidth parameter is allowed to expand with sample size. For information on the Gauss-Weierstrauss kernel see Ouliaris and Phillips (1994), which was used for computations, or Brillinger (1981).

[8]This is in contrast to results one obtains with Johansen's (1988) MLE. With the MLE the estimates of the cointegrating vectors can vary substantially depending on

in Table 1 are the Z_α and Z_t statistics for the null hypothesis of no cointegration. The results based on Z_α indicate that the null of no cointegration can be rejected for six of the eight members of the Zone (the rejection is at the 5 percent level except for Niger at the 10 percent level). The two exceptions are the Côte d'Ivoire and Togo.[9] According to the Z_t test the no cointegration null hypothesis cannot be rejected at the 10% level for Togo, Niger, and Côte d'Ivoire.[10]

We explored further the null hypothesis of no cointegrating relationship for the model in (1) using the Johansen (1988) ML test. Table 2 presents results for the Johansen likelihood ratio test based on the maximal eigenvalue of the stochastic matrix. Lag length was chosen on the basis of the SBC criterion. The null can be rejected in all cases except Togo. On the other hand the hypothesis that there is one cointegrating vector cannot be rejected for any of the eight countries. Given the somewhat contrasting results of our cointegration tests, we decided to include all eight Zone members in the estimation of the multivariate error correction model in (2).

For the six countries for which no cointegration is rejected, the long-run link between French and Zone price levels is significantly different from zero for all except Cameroon (see Table 1). Indeed the transmission coefficient is greater than unity in all five cases (though not statistically significantly). A long run transmission coefficient greater than unity is consistent with real overvaluation of the CFA franc. Many commentators have reported on the overvaluation of the CFA franc during the 1980s and 90s. For example, Devarajan (1997) calculates the degree by which the real CFA exchange rate differs from its equilibrium

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the user's choice of lag length. On the other hand, the Johansen procedure is an excellent means of testing for the number of cointegrating vectors.

[9] In our study, the bandwidth and kernel choice did not have an important impact on the estimates of the error correction models; however, these choices did affect the outcomes of the Z tests. For instance, using a Parzen window with bandwidth 8 or 5, Togo appears to be cointegrated while Cameroon, Senegal, Niger and Congo do not. Given the uncertainty associated with the Z tests, Johansen tests were also performed to determine whether the variables in the hypothesized long run relationship are cointegrated.

[10] Phillips and Ouliaris (1990) show that in many situations Z_α is more powerful than Z_t.

value (computed on the basis of three-sector Salter-Swan model) and finds that, in 1993, eleven of the Zone members experienced overvalued exchange rates (in the case of Congo and Gabon the real overvaluation was in excess of 50 percent).

Table 1 also indicates that for five of the eight Zone members the coefficient of the monetary aggregate is significantly different from zero at the 5 percent level (and for Burkina Faso at the 10 percent level). In these countries monetary policy has exerted effective control on domestic prices. The long run monetary price elasticity, however, is quite small. Moreover, the monetary neutrality hypothesis can be rejected in all cases. The bottom of Table 1 reports p-values associated with Wald tests performed for several joint hypotheses. The joint hypothesis of unit long run transmission between French and Zone prices ($\gamma_1 = 1$) and monetary neutrality ($\gamma_2 = 1$) can be rejected in all cases. Other hypotheses concerning combinations of unit transmission and monetary neutrality can also be rejected (the exception is the joint hypothesis of unit transmission and insignificant monetary transmission for Cameroon and Congo at the 5 percent level). On the other hand, the hypothesis that the combined transmission of French prices and the monetary aggregate is significantly different from unity cannot be rejected at the 5% level for six of the eight countries. This result is consistent with the hypothesis that long run movements in Zone prices can be explained by foreign price shocks and domestic monetary developments.

3.2 Short-run Price Dynamics

We investigate the short-run dynamics of inflation rates in Zone members by estimating the multivariate error-correction model in (2). The results for the error correction parameters from the unrestricted SUR estimation of (2) appear in Table 3. The middle panel in the table presents the Wald statistics for testing several cross-equation restrictions on coefficients. The hypothesis that the coefficients associated with the regime change dummy variables ($D1$ and DB) are equal across countries is rejected (5 percent level). On the other hand, the hypothesis that the coefficients on the time trend (D_T) are equal cannot be rejected. The null hypothesis that the error correction coefficients are equal across countries is rejected. Based on the pretests, the multivariate SUR model is re-estimated imposing the cross-equation restriction on the time trend. The results appear in the bottom panel of Table 3. After imposing the restriction of a common trend, there is very little change in the size or significance of the coefficients. The biggest differences are for Senegal and Cameroon; the estimated adjustment speed for Senegal falls by about 2.5% while that of Cameroon increases by less than 2%.[11]

Several conclusion emerge from Table 3. First, all of the estimated error correction coefficients are negative and significant. The negative sign indicates that prices converge to their long run equilibrium. The significance of the error correction coefficients is additional evidence of a cointegrating relationship among the variables in equation (1). Second, there is wide variation in adjustment speeds across countries but, on the whole, they are relatively low in magnitude. The estimates indicate that between 11% and 45% of the quarter-to quarter change in prices can be attributed to divergence of prices from their long run equilibrium level. Finally, the half life of deviations of prices from their long run equilibrium varies from 1.2 to 5.6 quarters. Our estimates of the speed of adjustment are generally higher than those of Honohan (1992).[12] It must be recalled, however, that he assumes (without testing) a one-to-one relation between French and domestic prices, and uses only deterministic time components and a single-equation method in modeling the short run dynamics.

Because SUR can transmit any model mispecification in one equation to all others, Côte d'Ivoire and Togo, whose long-run models were

[11]The complete set of SUR results is available on request. For the restricted model, the hypothesis of cross-equation equality on the coefficients of the dummy variables ($D1$ and DB) and the error correction term could not be rejected.

[12]Honohan's (1992) estimates range (in absolute value) from 0.075 (Côte d'Ivoire) to 0.348 (Congo) and a half life that varies from 2 to 9 quarters. With the exception of Cameroon and Senegal, all our estimates of the speed of adjustment are higher.

possibly spurious (based on the Z tests in Table 1), were dropped and the system was re-estimated. The SUR results appear in Table 4. There is very little change in the coefficients on the error correction term. In this instance, however, the common trend hypothesis is rejected.

3.3 Panel Estimation of the Long-run and Short-run Dynamics

Given the close economic links between the members of the CFA Zone, a natural extension of our methodology is to investigate the long-run relationship in (1) in a panel setting. The corresponding panel specification is:

$$p_{it} = \gamma_{i0} + \gamma_1 \, p^*_{it} + \gamma_2 \, m_{it} + \delta_i \, d_{it} + \varepsilon_{it} \tag{3}$$

This is similar to (1) except that the intercept and time deterministic components are now allowed to vary across countries while γ_1 and γ_2 do not. In a cointegrated panel, these two coefficients can be thought of as describing the *average* long-run relationship among p, p^*, and m.

First, we test the hypothesis of a cointegrating relationship in (3) using a panel cointegrating test proposed by McCoskey and Kao (1999). For the test, each cross section regression is estimated individually and the ADF statistic is computed. The average ADF statistic, $\bar{t}_{ADF}$, has asymptotic distribution

$$\sqrt{N}(\bar{t}_{ADF} - \mu_{ADF}) \Longrightarrow N(0, \sigma^2_{ADF}). \tag{4}$$

McCoskey and Kao (1999) compute μ_{ADF} and $\sqrt{\sigma^2_{ADF}/N}$ using Monte Carlo simulations for various numbers of regressors in the ADF regressions. In our sample $\bar{t}_{ADF} = -3.665$. Using the simulated mean and standard deviation from McCoskey and Kao (1999) of -2.468 and 0.8, respectively, the computed test statistic is -1.495 which is statistically less than zero at the 10% level. The hypothesis of no cointegrating relationship between domestic prices, French prices and the domestic monetary aggregate can be rejected for the CFA panel.

Next the cointegrating relationship in (3) is estimated using fully modified panel methods. Details on the estimation are provided in the Appendix. First, we estimate a model that includes heterogeneous deterministic effects; the results are columns (1) and (2) of Table 5. The table also presents results of hypotheses tests on the equality of the heterogenous effects across members of the panel. The cross-equation restriction on the coefficient of DB cannot be rejected. Therefore, this restriction is imposed and the restricted results are presented in columns (3) and (4). There is little difference between the unrestricted and restricted estimates. It is evident from Table 5 that there exists a significant long-run relationship between both the French price level and the domestic monetary aggregate, on the one hand, and domestic prices, on the other. Moreover, the long-run transmission coefficient between French and domestic prices is not significantly different from one. This result provides support for the discipline-benefits argument for currency unions. On the other hand, the monetary elasticity is significantly different from zero, indicating that the monetary authorities in these countries had substantial leeway in setting domestic prices and that the monetary irrelevance hypothesis for fixed exchange regimes is not supported by the CFA Zone experience.

Finally, we estimate the multivariate model in (2) using the estimated residuals from the panel FMOLS estimates of (3) to construct error correction terms. Because the cointegrating vector represents the long-run *average* across countries, (3) can be thought of as a regional equilibrium relationship that allows for certain heterogenous differences across countries. The corresponding interpretation of the coefficients on the error correction terms is changed in a subtle way. The ecm_{t-1} terms are now equilibrium displacements from the long-run country averages and the λ_i measure the speed at which each member of the system returns to regional equilibrium. The results are in Table 6. All our estimates for the adjustment speed are negative and significant but fall in a narrower range than those of Table 3. Our estimates of the half life of price deviations from their equilibrium value range from 2.0 to 3.9 quarters. Table also includes results of hypotheses tests of several cross-equation restrictions. In particular, the hypothesis that the speed at which members return to regional equilibrium is equal across countries cannot be rejected. Therefore, the model is reestimated imposing this restriction. The restricted SUR estimates reveal a common average speed of adjustment across the CFA Zone of 0.219, corresponding to a half life of 2.8 quarters. Taken together, our panel FMOLS and SUR results indicate that while domestic prices are anchored by French prices in the long run, there may be important deviations from equilibrium in the short run that may take up to one year to be (half-way) eliminated.

4 Conclusion

The main objective of this paper has been to investigate the link between prices of members of the CFA Zone and the French price level, the country to which these countries have pegged their currency for the past fifty years. The transmission of foreign price shocks was examined by estimating a long-run cointegrating relationship between domestic (Zone) prices, French prices and the domestic monetary aggregate using fully-modified estimation methods for nonstationary time series. The results indicate that there is a significant long-run cointegrating relationship between French and Zone prices for at least six of the Zone members and that the relationship between prices is close to one-to-one for five of these countries. Further, there is some evidence that domestic monetary policy also plays a role in determining the price level of CFA Zone members in the long-run. In terms of the short-run dynamics, Zone members adjust to shocks rather slowly, though our estimates suggest that the speed of adjustment is faster than previously thought. Furthermore, SUR estimation of the adjustment speeds indicates significant differences across members of the Zone.

Interestingly, when the regional (average) long-run relationship is estimated for Zone members using recently developed panel cointegrated estimation procedures the dissimilarity of adjustment speeds diminishes. In fact, the hypothesis that as a region the adjustment speeds are equivalent across countries cannot be rejected.

Table 1: Fully modified OLS estimates of the cointegrating relationship for each country. The t-ratios are in parentheses below the parameter estimates. The Z_α and Z_t statistics for the null hypothesis of no cointegration are also presented. The 5% and 10% critical values for Z_α are -27.1 and -24.8, respectively; the 5% and 10% critical values for Z_t are -3.80 and -3.65, respectively. A Gauss-Weierstrauss kernel with a bandwidth of 8 is used in the FMOLS estimator and for the Z statistics.

Variable	Burkina Faso	Cameroon	Congo	Gabon	Côte d'Ivoire	Niger	Senegal	Togo
p^* (γ_1)	1.152	0.278	1.091	1.162	0.852	1.220	1.045	0.458
	(8.876)	(0.812)	(12.532)	(16.136)	(4.205)	(2.973)	(5.888)	(3.004)
m (γ_2)	0.171	0.275	-0.037	0.279	0.177	0.314	0.143	0.340
	(1.709)	(2.343)	(-1.006)	(10.155)	(2.241)	(2.201)	(1.449)	(7.131)
Constant	-0.689	1.089	-0.134	-1.976	-0.235	-1.155	-0.613	1.081
	(-2.178)	(1.534)	(-0.685)	(-11.407)	(-0.435)	(-1.136)	(-1.618)	(2.712)
$D1$	0.344	0.252	0.531	0.230	0.296	0.558	0.284	0.311
	(5.971)	(4.115)	(27.001)	(7.06)	(3.526)	(5.103)	(4.155)	(5.223)
DB	-0.068	0.150	-0.008	0.155	-0.116	-0.049	0.153	-0.057
	(-0.634)	(1.110)	(-0.220)	(2.679)	(-0.667)	(-0.245)	(1.201)	(-0.550)
D_T	-0.011	0.008	-0.001	-0.009	0.000	-0.018	-0.005	0.000
	(-5.432)	(2.249)	(-1.075)	(-8.610)	(0.145)	(-3.542)	(-2.437)	(-0.155)
Z_α	-36.597	-29.509	-32.391	-34.474	-16.631	-25.997	-36.629	-18.313
Z_t	-4.446	-4.030	-5.242	-4.562	-2.915	-3.296	-4.435	-3.279
$\gamma_1 = 1, \gamma_2 = 0$	6.983	2.772	0.551	145.310	2.999	19.091	3.559	34.953
p-value	0.001	0.068	0.578	0.000	0.055	0.000	0.032	0.000
$\gamma_1 = 0, \gamma_2 = 1$	43.540	120.030	1016.800	352.210	60.652	16.343	39.109	216.340
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
$\gamma_1 = 1, \gamma_2 = 1$	55.397	273.450	2509.100	628.840	105.370	49.347	93.999	471.960
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
$\gamma_1 + \gamma_2 = 1$	12.078	3.601	0.962	62.635	0.031	3.423	2.556	3.028
p-value	0.001	0.061	0.329	0.000	0.860	0.067	0.113	0.085

Table 2: Johansen Maximum Likelihood Tests for a Cointegrating Relationship. λ_{max} is the cointegration likelihood-ratio test based on the maximal eigenvalue of the stochastic matrix, respectively, and r is the number of cointegrating vectors. The 5 per cent critical values for $r = 0$ and for $r \leq 1$ are 21.12 and 14.88, respectively. Significance at the 5 per cent level is indicated by **.

Country	Null	λ_{max}
Burkina Faso	$r = 0$	26.75**
	$r \leq 1$	12.21
Cameroon	$r = 0$	29.99**
	$r \leq 1$	11.68
Congo	$r = 0$	41.68**
	$r \leq 1$	11.63
Côte d'Ivoire	$r = 0$	33.31**
	$r \leq 1$	8.36
Gabon	$r = 0$	27.19**
	$r \leq 1$	8.94
Niger	$r = 0$	22.75**
	$r \leq 1$	4.75
Senegal	$r = 0$	31.80**
	$r \leq 1$	11.69
Togo	$r = 0$	18.88
	$r \leq 1$	12.25

Table 3: Estimates of the short-run transmission speeds from unrestricted and restricted SUR estimation of the error correction model and relevant hypothesis tests.

Unrestricted SUR

Country	Variable	Estimate	t-ratio
Burkina Faso	ecm_{t-1}	-0.3754	-5.7321
Cameroon	ecm_{t-1}	-0.1171	-2.8187
Congo	ecm_{t-1}	-0.4507	-6.7565
Gabon	ecm_{t-1}	-0.2458	-5.4095
Côte d'Ivoire	ecm_{t-1}	-0.2576	-5.1192
Niger	ecm_{t-1}	-0.1515	-3.0841
Senegal	ecm_{t-1}	-0.1997	-4.9659
Togo	ecm_{t-1}	-0.3546	-5.3038

Variable	Null Hypothesis	Statistic	p-value
$D1$	δ_{1i} equal for all i	4.8661	$< .0001$
DB	δ_{Bi} equal for all i	11.500	$< .0001$
D_T	δ_{Ti} equal for all i	0.994	0.434
ecm_{t-1}	λ_i equal for all i	4.308	$< .0002$

Restricted SUR

Country	Variable	Estimate	t-ratio
Burkina Faso	ecm_{t-1}	-0.3807	-5.8814
Cameroon	ecm_{t-1}	-0.1339	-3.3697
Congo	ecm_{t-1}	-0.4490	-6.7328
Gabon	ecm_{t-1}	-0.2380	-5.3432
Côte d'Ivoire	ecm_{t-1}	-0.2576	-5.6059
Niger	ecm_{t-1}	-0.1467	-3.0329
Senegal	ecm_{t-1}	-0.1668	-4.9199
Togo	ecm_{t-1}	-0.3571	-5.3451

Table 4: Estimates of the short-run transmission speeds from Unrestricted SUR estimation of the error correction model and relevant hypothesis tests. Côte d'Ivoire and Togo are omitted from the model.

Country	Variable	Estimate	t-ratio
Burkina Faso	ecm_{t-1}	-0.3818	-5.7160
Cameroon	ecm_{t-1}	-0.1203	-2.8800
Congo	ecm_{t-1}	-0.4365	-6.5227
Gabon	ecm_{t-1}	-0.2481	-5.4433
Côte d'Ivoire	ecm_{t-1}	-	-
Niger	ecm_{t-1}	-0.1552	-3.1335
Senegal	ecm_{t-1}	-0.2007	-4.9887
Togo	ecm_{t-1}	-	-

Variable	Null Hypothesis	Statistic	p-value
$D1$	δ_{1i} equal for all i	8780	$< .0001$
DB	δ_{Bi} equal for all i	18488	$< .0001$
D_T	δ_{Ti} equal for all i	1700	$< .0001$
ecm_{t-1}	λ_i equal for all i	6723	$< .0001$

Table 5: Fully modified panel estimates of the cointegrating relationship. Heterogeneous deterministic effects are estimated for each country. The null hypothesis that the DB parameter is equal across countries is not rejected (5% or 10% levels) and this restriction is imposed, yielding columns (3)-(4).

		Unrestricted		Restricted	
		(1)	(2)	(3)	(4)
Country	Parameter	Estimate	t-ratio	Estimate	t-ratio
	p^* (γ_1)	0.9734	12.83	0.9791	12.99
	m (γ_2)	0.1851	6.16	0.1822	6.12
Burkina Faso	Constant	-0.1223	-0.61	-0.1341	-0.67
	$D1$	0.2745	4.95	0.2741	5.01
	DB	0.0189	0.13	0.0560	1.07
	D_T	-0.0084	-7.89	-0.0084	-7.91
Cameroon	Constant	-0.8924	-5.06	-0.8988	-5.11
	$D1$	0.4231	8.09	0.4265	8.28
	DB	0.1171	0.82	0.0560	1.07
	D_T	-0.0021	-1.98	-0.0022	-2.01
Congo	Constant	-0.4162	-2.14	-0.4267	-2.20
	$D1$	0.5873	10.68	0.5886	10.91
	DB	0.0646	0.45	0.0560	1.07
	D_T	-0.0043	-3.86	-0.0043	-3.89
Gabon	Constant	-0.8431	-5.08	-0.8470	-5.10
	$D1$	0.1759	3.14	0.1794	3.25
	DB	0.0790	0.55	0.0560	1.07
	D_T	-0.0042	-3.49	-0.0043	-3.55
Côte d'Ivoire	Constant	-0.6736	-3.68	-0.6817	-3.73
	$D1$	0.3335	6.27	0.3373	6.43
	DB	0.0964	0.67	0.0560	1.07
	D_T	-0.0021	-1.85	-0.0022	-1.90
Niger	Constant	0.0600	0.31	0.0481	0.25
	$D1$	0.3953	7.67	0.3856	7.57

		Unrestricted		Restricted	
		(1)	(2)	(3)	(4)
Country	Parameter	Estimate	t-ratio	Estimate	t-ratio
	DB	-0.1083	-0.76	0.0560	1.07
	D_T	-0.0104	-9.07	-0.0104	-9.09
Senegal	Constant	-0.5450	-3.08	-0.5517	-3.12
	$D1$	0.2856	5.27	0.2895	5.41
	DB	0.0926	0.65	0.0560	1.07
	D_T	-0.0049	-4.09	-0.0049	-4.16
Togo	Constant	-0.1425	-0.74	-0.1523	-0.79
	$D1$	0.4957	9.11	0.5000	9.35
	DB	0.0982	0.68	0.0560	1.07
	D_T	-0.0074	-6.29	-0.0075	-6.36

Variable	Null Hypothesis	Statistic	p-value
$D1$	δ_{1i} equal for all i	6.4775	$< .0001$
DB	δ_{Bi} equal for all i	0.26268	0.96804
D_T	δ_{Ti} equal for all i	22.347	$< .0001$

Table 6: Estimates of the short-run transmission speeds from unrestricted and restricted SUR estimation of the error correction model and relevant hypothesis tests. The error correction terms are generated using the FM panel estimator. The null hypothesis that λ_i is equal for all i is rejected and the restrictions imposed. A Gauss-Weierstrauss kernel with bandwidth 8 is used to estimate the long-run relationship.

Estimator	Country	Coefficient	Estimate	t-Ratio
Unrestricted	Burkina Faso	ecm_{t-1}	-0.2724	-4.5666
	Cameroon	ecm_{t-1}	-0.2295	-6.7678
	Congo	ecm_{t-1}	-0.2919	-5.7284
	Gabon	ecm_{t-1}	-0.2126	-4.9875
	Côte d'Ivoire	ecm_{t-1}	-0.1696	-4.1083
	Niger	ecm_{t-1}	-0.2202	-5.7087
	Senegal	ecm_{t-1}	-0.2410	-5.5852
	Togo	ecm_{t-1}	-0.1638	-3.7219
Restricted	All	ecm_{t-1}	-0.2190	-13.8909

Variable	Null Hypothesis	Statistic	p-value
$D1$	δ_{1i} equal for all i	2.7687	0.0077
DB	δ_{Bi} equal for all i	11.443	$< .0001$
D_T	δ_{Ti} equal for all i	0.0077932	0.0515
ecm_{t-1}	λ_i equal for all i	0.88099	0.52116

A Appendix

This appendix discusses the fully modified least squares estimator and its computation when the model contains deterministic regressors. First, the basic model is presented and then it is modified to allow for intercepts, time trends and/or dummy variables. Then, the statistical properties of the FMOLS estimator and test statistics are summarized. Finally, the implementation of the FMOLS panel estimator is discussed.

A.1 The Fully Modified OLS Estimator

Let the data generation process for y_t be described by the following cointegrating system:

$$y_t = \beta' x_t + u_{1t} = \gamma' z_t + u_{1t} \tag{5}$$

$$Dx_t = u_{2t} \tag{6}$$

where D is the first difference operator, y_t is a scalar, x_t is a m vector of $I(1)$ stochastic processes, and the model is reparameterized using γ and z_t to permit easy generalization below. Let

$$u_t = \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix} \tag{7}$$

which has finite covariance matrix $\Sigma > 0$ and continuous spectral density $f_{uu}(\lambda)$ with long-run covariance

$$\Omega = 2\pi f_{uu}(0) = \Omega = \begin{pmatrix} \omega_{11} & \omega_{21}' \\ \omega_{21} & \Omega_{22} \end{pmatrix} > 0. \tag{8}$$

The fully modified OLS estimator is a single equation method based on the use of OLS on (5) where a semi-parametric correction for serial correlation and endogeneity have been made. The method is developed in Phillips and Hansen (1990) and the discussion of it here is largely due to Phillips (1995).

The long-run covariance matrix, Ω, can be decomposed as

$$\Omega = \Sigma + \Lambda + \Lambda' \tag{9}$$

where $\Sigma = E(u_0 u_0')$, and $\Lambda = \sum_{k=1}^{\infty} E(u_0 u_k')$. Define

$$\Delta = \Sigma + \Lambda \tag{10}$$

The FMOLS estimator of γ is

$$\gamma^+ = (Z'Z)^{-1}(Z'y^+ - T\hat{\delta}^+) \tag{11}$$

with

$$y_t^+ = y_t - \hat{\omega}_{21}' \hat{\Omega}_{22}^{-1} Dx_t \tag{12}$$

and

$$\hat{\delta}^+ = \hat{\Delta} \begin{bmatrix} 1 \\ -\hat{\Omega}_{22}^{-1}\hat{\omega}_{21} \end{bmatrix} \tag{13}$$

The estimator requires consistent estimation of Σ and Λ. The latter is obtained using a time series kernel estimators based on preliminary OLS estimates of the cointegrating relationship. Specifics of how these computations were performed in the current context are presented below.

A.2 Adding deterministic components

First, the model estimated in this paper contains deterministic time components that must be accounted for. Therefore, (5) is generalized:

$$y_t = \delta_0 + \delta_1' d_t + \beta' x_t + u_{1t} = \gamma_0' k_t + \beta' x_t + u_{1t} = \gamma' z_t + u_{1t} \tag{14}$$

where $\gamma_0' = (\delta_0\ \delta_1')$, $\gamma' = (\gamma_0'\ \beta')$,$d_t$ is a sx1 vector of independent time variables, and $k_t' = (1\ d_t')$.

1. The first step is to estimate (14) using OLS. The residuals from

this equation, $\hat{u}_1$, are obtained.

2. The influence of deterministic time components are purged from the model. To do this, the $I(1)$ variables in x_t are regressed on the deterministic portion of the model, k, and the predicted values are subtracted from x_t. In vector notation:

$$\hat{x}^k = x - k(k'k)^{-1}k'x = x - \hat{x} \tag{15}$$

and

$$\hat{u}_2 = D\hat{x}^k \tag{16}$$

Again, D is the first difference operator.

3. The Tx$(m+1)$ matrix of system residuals is then formed in order to estimate the long-run covariances. That is:

$$\hat{E} = \left[\begin{array}{cc} \hat{u}_1 & \hat{u}_2 \end{array} \right] \tag{17}$$

4. Using $\hat{E}$, the finite covariance matrix, $\hat{\Sigma}$, is easily computed and $\hat{\Lambda}$ is obtained using a suitable lag kernel.

$$\hat{\Sigma} = \hat{E}'\hat{E}/T \tag{18}$$

and

$$\hat{\Lambda} = \sum_{j=0}^{T-1} \omega(\frac{j}{K})\hat{\Gamma}(j) \tag{19}$$

where

$$\hat{\Gamma}(j) = \frac{1}{T}\sum_t \hat{E}_t'\hat{E}_{t+j} \tag{20}$$

and $\omega(j)$ is the lag kernel. In this paper the Gauss-Weierstrauss lag kernel was used with bandwidth, K=8. In this estimator K

is the number of autocovariance terms used to compute the spectrum at frequency zero when estimating the long-run covariance of u_t.

5. Using $\hat{\Sigma}$ and $\hat{\Lambda}$, form the matrices and $\hat{\Omega} = \hat{\Sigma} + \hat{\Lambda} + \hat{\Lambda}'$ and $\hat{\Delta} = \hat{\Sigma} + \hat{\Lambda}$ and partition $\hat{\Omega}$ according to u_t.

6. Transform y_t using equation (12) after replacing Dx_t with $\hat{u}_2 = D\hat{x}^k$ from equation (16).

7. Finish the transformation using equation (13) that has been augmented with $s+1$ rows of zeros corresponding to the constant and deterministic components of Z.

8. Using the transformed data, compute the fully modified OLS estimator (11).

A.3 Some properties of FMOLS

Phillips (1995) examines the asymptotic behavior of the FMOLS estimator in models with full rank I(1) regressors (no cointegration among the x_t), models with I(1) and I(0) regressors, models with unit roots, and models with only I(0) regressors. Some of his conclusions are:

1. FMOLS is applicable in model with either full rank or cointegrated I(1) regressors. If the variables in x_t are stationary, then the FMOLS estimator has the same Gaussian limit distribution as OLS.

2. The FM estimates of the nonstationary components are also optimal in the sense that they are equivalent to maximum likelihood estimation of the cointegrating relationship.

3. The fully modified estimators have normal and mixed normal limiting distributions (depending on whether regressors do or do not have unit roots) and consequently Wald tests can be constructed and used in practice. When the null hypothesis involves both I(1) and I(0) variables the limiting distribution of the Wald statistic

is a linear combination of chi-squares. If q is the total number of restrictions then the χ^2_q critical values can be used as upper bounds for nominal α sized tests.

A.4 Panel FMOLS estimator

Phillips and Moon (1999) have surveyed some recent developments in the analysis of nonstationary panel data. Following their analysis, a cointegrated panel regression was estimated for the CFA countries in our sample. The primary difference in the construction of the FMOLS estimator is in the computation of the long-run matrices $\hat{\Omega}$ and $\hat{\Delta}$. In the panel estimator these are defined by taking the average over the countries in the sample. Hence,

$$\hat{\Omega} = \frac{1}{n}\sum_{i=1}^{n}\hat{\Omega}_i \tag{21}$$

and

$$\hat{\Delta} = \frac{1}{n}\sum_{i=1}^{n}\hat{\Delta}_i \tag{22}$$

where $\hat{\Omega}_i$ and $\hat{\Delta}_i$ are estimated for each of the $i = 1, 2, \ldots, 8$ countries as in step 5 above. The dependent variable for each country is transformed using these averages, adjusted for serial correlation using $nT\delta^+$ rather than $T\delta^+$ as in (11) and the FMOLS estimator is computed as before. In this way, one is able to include a large number of deterministic components and/or individual effects with relative ease. The resulting estimator is consistent and asymptotically normally distributed.

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