INFLATION ANALYSIS IN THE CENTRAL AMERICAN MONETARY COUNCIL

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ABSTRACT

Though not working towards an imminent transition to a monetary or currency union, the Central American Monetary Council (or CMCA, from Spanish *Consejo Monetario Centroamericano*) serves as an institution promoting economic and financial stability among five Central American countries (Costa Rica, El Salvador, Guatemala, Honduras and Nicaragua) and the Dominican Republic. Econometric studies conducted by researchers from CMCA have mostly focused on studying inflation levels of these countries, making use of econometric tools such as VECM and cointegration. We expand the study of inflation stability in the member countries of the CMCA by adopting a long memory and fractionally integrated approach and implementing cointegration methods that have not yet been used in the context of the Central American Monetary Council. Our results first show that all the series of prices are nonstationary, with orders of integration equal to or higher than 1, implying high levels of persistence. Looking at long run equilibrium relationships among the countries, we only found strong evidence of cointegration in the case of Honduras with El Salvador. All the other vis a vis relationships seem to diverge in the long run. Policy implications of the results obtained are also derived in the paper.

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1. Introduction

The Central American Monetary Council (or CMCA, from Spanish *Consejo Monetario Centroamericano*) attempts to provide economic and financial stability to five Central American countries (Costa Rica, El Salvador, Guatemala, Honduras and Nicaragua) and the Dominican Republic. Measuring and controlling inflation levels constitutes a very important task for SECMCA (from the Spanish *Secretaría Ejecutiva del Consejo Monetario Centroamericano*), which acts as the research branch of the CMCA.

If there are differences in the rate at which inflation returns to its baseline following a shock, policy makers in SECMCA will be confronted with the design of a monetary policy for diverse or even conflicting economic environments. This is the reason why policy aimed at stimulating growth may not influence price stability in one of the countries in the region but might have the opposite effect in another with further knock-on effects. Very often the design of monetary policy assumes that inflation series are stationary, in such a way that if there is low persistence in inflation among all member countries then inflation levels will tend to move close to some average value within a year or two. However if there is varying degrees of persistence, the more asymmetric the shocks are then the greater the risks to the stability of the CMCA could be. Knowing whether inflation rates react in a similar manner to shocks is crucial for the design of a successful common monetary policy strategy.

We conduct the following study using a long memory modeling framework, based on fractional integration, with the aim of analyzing the level of persistence in the inflation levels of the countries belonging to the CMCA. Our results show that all the series of prices are nonstationary, with orders of integration equal to or higher than 1 and thus implying high levels of persistence even in the inflation rates. Looking at long run equilibrium relationships among the countries, we only found strong evidence of cointegration relationship in the case of Honduras with El Salvador, and partially in the cases of Costa Rica with the Dominican Republic, El Salvador and Guatemala. All the other vis a vis relationships seem to diverge in the long run. The rest of the paper is organized as follows: Section 2 briefly describes the history of the CMCA. Section 3 deals with the literature review. Section 4 presents the data and the methodology used in the paper. Section 5 is devoted to the empirical results, while Section 6 concludes the paper.

2. Brief history of the CMCA

The history of the Central American Monetary Council can be summarized as an outstanding integrationist effort made by the Central Banks of its member countries (Costa Rica, Dominican Republic, El Salvador, Guatemala, Honduras and Nicaragua). As part of a clear movement towards more integration in the region during the 1950's, the central banks of these countries decided to hold informal meetings and keep negotiations with the ultimate aim of achieving a general consensus on monetary integration.

Between 1951 and 1957 several bilateral agreements among these Central American countries were signed, constituting thus the basis for the creation of a new system of Central Banks in Central America with the initial goal of achieving monetary integration. The first step was to create a mechanism of multilateral payment compensations, which was established under the Central American Compensation Chamber Agreement, signed in July 1961. This was followed by the Central American Monetary Union Establishment agreement in February 1964, which lead to the formation of the Central American Monetary Council. Later on the Central American Monetary Establishment Fund was established in 1969 with the aim of establishing an equilibrium in the balance of payments between the member countries that could affect their corresponding exchange rates stability.

These three agreements were united in 1974 under the Central American Monetary Agreement, which was later modified in 1999 in order to include some of the integrationist achievements that took place during the 1990s. Among these we shall point out the Tegucigalpa Protocol in December 1991, which lead to the foundation of the Central American Integration System; and the Guatemala Protocol in October 1993, which substituted the Central American Economic Integration General Treaty that had originally been signed in 1960. The Central American Monetary Agreement is still today the main pillar of the monetary and financial integration in Central America. During its 50 years of existence the Central American Monetary Council has held multiple meetings, all of them with the aim of improving and fostering economic integration among its country members.

Despite not having as ultimate goal the adoption of a new common currency, the CMCA attempts to achieve economic and financial stability in the region, in order to promote the integration and mutual collaboration of its member countries. Inflation levels are crucial in determining the stability of an economy, and targeting it is therefore one of the most important activities at CMCA. This is the reason why we have decided to carry out the following study, including several econometric techniques which up to what we know have not yet been used in analyzing the six countries of the CMCA.

3. Literature review

Fractional integration or $I(d)$ models have been widely used for modelling inflation in developed countries. For instance, regarding inflation persistence, Backus and Zin (1993) found a fractional degree of integration in US monthly data, arguing that aggregation across agents with heterogeneous beliefs results in long memory of the inflation rate. Specifically on US inflation persistence, Cogley and Sargent (2002) find that inflation persistence in the United States rose in the 1970s and declined from 1980s to the present. Brainard and Perry (2000) and Taylor (2000) found similar results. Pivetta and Reis (2007), estimating a Bayesian non-linear model, found that inflation persistence did not change over the past three decades in the United States. Hassler (1993) and Delgado and Robinson (1994) provided strong evidence of long memory or $I(d)$ behaviour in the Swiss and Spanish inflation rates, respectively.

There is however less evidence of $I(d)$ behaviour in the inflation rates for developing countries (Kallon, 1994; Moriyama and Naseer, 2009). Almost all existing studies assume integer degrees of differentiation, testing stationarity/nonstationarity with unit root tests and cointegration techniques. Masale (1993) tested for stationarity of the inflation rate in Botswana and found evidence of nonstationarity and cointegration with South African prices. Gaomab (1998) used an error correction model based on cointegration for the inflation rate in Namibia. Moriyama and Naseer (2009) forecasted the 3-month average inflation rate in Sudan with data from January 2000 to October 2008 using an ARMA(4,5) model. Chhibber et al. (1989) estimated a multivariate model for Zimbabwe, including both monetary factors and economic fundamentals as determinants of inflation. A similar model for Ghana was employed by Chhibber and Shafik (1990) and Sowa and Kwakye (1991).

Econometric studies conducted by the SECMCA have mostly focused on studying inflation levels of these countries, making use of econometric tools such as VECM and cointegration. Publically available on the CMCA website (http://www.secmca.org/), one can find interesting studies and analysis on several macroeconomic aspects which are of big importance to the Monetary Council. Iraheta, Medina and Blanco (2007) provide empirical evidence of inflation transmissions among CMCA countries by making use of structural VAR models. Iraheta and Blanco (2007) also present a macroeconomic model for analyzing and forecasting the effects of external shocks to the collective economies of Central America and the Dominican Republic. Granados (2002) presented an approach based on a VAR estimation about how the future of the Central American Monetary System could be under an optimum currency area and the new role for the Central American Monetary Council which was established to improve monetary policy coordination in Central America. Galindo and Moreno-Brid (2007) provide an outstanding overview of the diverse econometric techniques which are being carried out not only at the CMCA but also at each Central Bank of the country members. With our study based on fractional integration we provide an analysis of the inflation persistence levels of the countries belonging to the CMCA, with the aim of determining the degree of homogeneity in which the member countries behave from an economic perspective.

4. Data and methodology

We use monthly data from January 1993 up to December 2013 corresponding to CPI levels of the six countries that belong to CMCA, having thus series of 252 data values. We obtained them from the official CMCA statistical data base called SIMAFIR (http://www.secmca.org/simafir.html).

The methodology that we employ in this paper to test the persistence of inflation rates is based on the concepts of fractional integration and cointegration. We first need to introduce some definitions. A covariance stationary process $\{x_t, t = 0, \pm 1, ...\}$ is integrated of order 0 (and denoted by I(0)) if the infinite sum of the autocovariances $\gamma_u = E[(x_t - Ex_t)(x_{t+u} - Ex_t)]$ is finite, i.e.,

$$\lim_{T\to\infty} \sum_{j=-T}^{T} \left|\gamma_j\right| < \infty.$$

It can also be defined in the frequency domain. Supposing that x_t has an absolutely continuous spectral distribution function, implying that it has a spectral density function, denoted by $f(\lambda)$, and defined as

$$f(\lambda) = \frac{1}{2\pi} \sum_{j=-\infty}^{\infty} \gamma_j \cos\lambda\, j, \qquad -\pi < \lambda \leq \pi.$$

Then, x_t is I(0) if the spectral density function is positive and finite, i.e.,

$$0 < f(\lambda) < \infty, \qquad \lambda \in [0, \pi].$$

A process is fractionally integrated or integrated of order d ($x_t \approx I(d)$) if

$$(1 - L)^d x_t = u_t, \qquad t = 0, \pm 1, \ldots, \tag{1}$$

where d can be any real value (and thus including fractional values), L is the lag-operator ($Lx_t = x_{t-1}$) and u_t is I(0) as defined above.

Given the parameterization in (1), the fractional differencing parameter d plays a crucial role. If d = 0, $x_t = u_t$, x_t is said to be “short memory” or I(0), and if the observations are autocorrelated they are of a “weak” form, in the sense that the values in the autocorrelations are decaying exponentially; if d > 0, x_t is said to be “long memory”, so named because of the strong association between observations far distant in time. If d belongs to the interval (0, 0.5) x_t is still covariance stationary, while $d \geq 0.5$ implies nonstationarity. Finally, if d < 1, the series is mean reverting in the sense that the effect of the shocks disappears in the long run, contrary to what happens if $d \geq 1$, with shocks persisting forever. In this study, we find t hat most of the log CPI series present orders of integration which are above 1, implying the permanency of the shocks. Additionally, the fact that the analysis is conducted in logarithms indicates that the first differences (i.e., the inflation rates) present orders of integration which are above 0 and thus showing long memory behaviour.

Several methods exist for estimating and testing the fractional differencing parameter d. Some are parametric while others are semiparametric and can be specified in the time or in the frequency domain. In our study we use a parametric Whittle estimation approach (Dahlhaus, 1989) along with a testing procedure (Robinson, 1994), which is based on the Lagrange Multiplier (LM) principle and uses the Whittle function in the frequency domain. In addition, we also perform several semiparametric methods (Robinson, 1995a,b; Velasco, 1999a,b; Abadir et al., 2007).

The natural extension of fractional integration to the multivariate case is the concept of fractional cointegration. Engle and Granger (1987) suggested that, if two processes x_t and y_t are both I(d), then it is generally true that for a certain scalar $a \neq 0$, a linear combination $w_t = y_t - ax_t$ will also be I(d) or I(d-b)) with $b > 0$.[1] This is the concept of cointegration, which they adapted from Granger (1981) and Granger and Weiss (1983). Given two real numbers d, b, the components of the vector z_t are said to be cointegrated of order d, b, denoted $z_t \sim CI(d, b)$ if:

(i) all the components of z_t are I(d),

(ii) there exists a vector $\alpha \neq 0$ such that $s_t = \alpha' z_t \sim I(\gamma) = I(d - b)$, $b > 0$.

Here, α and s_t are called the cointegrating vector and error respectively. Note that by allowing fractional degrees of differentiation, we allow a greater degree of flexibility in representing equilibrium relationships between economic variables than the traditional use of integer differentiation.

We conduct the following strategy: We first estimate individually the orders of integration of the series using, in addition to the previous methods, the log-periodogram-type of estimator as devised by Robinson (1995b), Kim and Phillips

[1] Classical cointegration as widely employed in the literature occurs then if d = 1 and b = 1.

(1999, 2006), Velasco (1999b) and others. This method is a generalization of the one proposed earlier by Geweke and Porter-Hudak (GPH, 1983), and is defined as:

$$\hat{d}(l) = \sum_{j=l+1}^{m} \left(a_j - \bar{a}\right) \log I(\lambda_j) / S_l, \tag{2}$$

where

$$a_j = -log\left(4\,sin^2\left(\frac{\lambda_j}{2}\right)\right), \quad \bar{a} = \frac{1}{m-l}\sum_{j=1}^{p} a_j u_{t-j} + \varepsilon_t, \quad S_l = \sum_{j=l+1}^{m}\left(a_j - \bar{a}\right)^2, \quad \lambda_j = \frac{2\pi\, j}{T},$$

and $0 \leq l < m < T$.

Next we test the homogeneity of the orders of integration in the bivariate systems (i.e., H_o: $d_x = d_y$), where d_x and d_y are now the orders of integration of the two individual series, by using an adaptation of Robinson and Yajima (2002) statistic $\hat{T}_{xy}$ to log-periodogram estimation. The statistic is:

$$\hat{T}_{xy} = \frac{m^{1/2}\left(\hat{d}_x - \hat{d}_y\right)}{\left(\frac{1}{2}\left(1 - \hat{G}_{xy} / (\hat{G}_{xx}\hat{G}_{yy})\right)\right)^{1/2} + h(n)} \tag{3}$$

where $h(n) > 0$ and $\hat{G}_{xy}$ is the $(xy)^{th}$ element of

$$\hat{\mathrm{G}} = \frac{1}{\mathrm{m}}\sum_{\mathrm{j=1}}^{\mathrm{m}} \mathrm{Re}\left[\hat{\Lambda}(\lambda_{\mathrm{j}})^{-1}\mathrm{I}(\lambda_{\mathrm{j}})\hat{\Lambda}(\lambda_{\mathrm{j}})^{-1*}\right], \quad \hat{\Lambda}(\lambda_j) = diag\left\{e^{i\pi \hat{d}_x/2}\lambda^{-\hat{d}_x}, e^{i\pi \hat{d}_y/2}\lambda^{-\hat{d}_y}\right\},$$

with a standard normal limit distribution (see Gil-Alana and Hualde (2009) for evidence on the finite sample performance of this procedure). Finally we perform the Hausman test for no cointegration of Marinucci and Robinson (2001) comparing the estimate $\hat{d}_x$ of d_x with the more efficient bivariate one of Robinson (1995a), which uses the information that $d_x = d_y = d_*$. Marinucci and Robinson (2001) show that

$$H_{im} = 8m\left(\hat{d}_* - \hat{d}_i\right)^2 \rightarrow_d \chi_1^2 \quad as \quad \frac{1}{m} + \frac{m}{T} \rightarrow 0, \tag{4}$$

with i = x, y, and where m < [T/2] is again a bandwidth parameter, analogous to that introduced earlier; $\hat{d}_i$ are univariate estimates of the parent series, and $\hat{d}_*$ is a restricted estimate obtained in the bivariate context under the assumption that $d_x = d_y$. In particular,

$$\hat{d}_* = -\frac{\sum_{j=1}^{s} 1_2' \hat{\Omega}^{-1} Y_j v_j}{2 1_2' \hat{\Omega}^{-1} 1_2 \sum_{j=1}^{s} v_j^2}, \qquad (5)$$

where 1_2 indicates a (2x1) vector of 1s, and with $Y_j = [\log I_{xx}(\lambda_j), \log I_{yy}(\lambda_j)]^T$, and $v_j = \log j - \frac{1}{s}\sum_{j=1}^{s} \log j.$ The limiting distribution above is presented heuristically, but the authors argue that it seems sufficiently convincing for the test to warrant serious consideration.

5. Empirical results

We start the empirical analysis by estimating the fractional differencing parameter d in a model given by the following form:

$$y_t = \alpha + \beta t + x_t; \quad (1 - L)^d x_t = u_t, \quad t = 1, 2, ..., \qquad (5)$$

where y_t is the observed time series for each country (logged CPI), α and β are the unknown coefficients corresponding to an intercept and a linear trend, and the resulting errors, x_t, are supposed to be I(d). First we employ a parametric approach, and thus, we need to specify a functional form for the d-differenced process. We considered four different cases corresponding to white noise disturbances, Bloomfield-type and seasonal and non-seasonal (monthly) AR. The model of Bloomfield (1973) is a non-parametric approach that produces autocorrelations decaying exponentially as in the AR(MA) case and that accommodates extremely well in the context of fractional integration.[2] Finally, we use monthly AR(1) disturbances based on the monthly nature of the series examined.

[Insert Tables 1 - 4 about here]

We display the estimated values of d under three different specifications assuming 1): no deterministic terms (i.e., $\alpha = \beta =$ in equation (5)), 2): an intercept (α unknown and $\beta = 0$), and c): an intercept with a linear time trend (α and β unknown). We observe that for most of the cases, the estimated values of d are above 1. Only for Guatemala and Nicaragua, in the cases of AR(1) and Bloomfield disturbances, the unit root null hypothesis (d = 1) cannot be rejected. In all the other cases, this hypothesis is decisively rejected in favour of higher degrees of integration (i.e., d > 1). This implies that shocks affecting the series will be clearly non-mean reverting and strong policy measures should be adopted to recover in the series the original trends (levels). Another consequence of these results is that the inflation rates will be long memory, with estimated values of d above 0 except in the cases of Guatemala and Nicaragua.

[Insert Table 5 about here]

We next use a semiparametric method proposed by Robinson (1995) and modified later by Abadir et al. (2007) among many others. This is a "local" Whittle estimator in the frequency domain, using a band of frequencies that degenerates to zero. Evidence of unit roots is obtained, for Guatemala and especially for Nicaragua. Also, in some minor cases for the Dominican Republic. Thus, these results are completely in line with those reported above and based on a parametric model.

[Insert Table 6 about here]

Next, we investigate the homogeneity condition in a pairwise representation through the Robinson and Yajima (2002) approach as presented in the previous section.

[2] See, Gil-Alana (2004).

The results (though not reported) are summarized in Table 6, indicating the cases where the hypothesis of $d_x = d_y$ cannot be rejected. It is observed that there are only two cases where this hypothesis is rejected, corresponding to Nicaragua with the Dominican Republic and Honduras. In all the other bivariate representations of the series, the hypothesis of equal orders of integration cannot be rejected at standard statistical levels.

[Insert Table 7 about here]

Noting that the series seem to be clearly nonstationary, the next step we conducted was to perform the OLS regression of one of the series over another. The fact that the two individual series are I(1) validates the use of standard OLS methods under the standard setting of cointegration (Phillips and Durlauf, 1986). In a fractional setting, things are more complicated and the properties depend on the specific orders of integration of the parent series and that of the cointegrating regression (Gil-Alana and Hualde, 2009).[3] Table 7 displays the estimated coefficients in the potentially cointegrated relationships. All the estimated coefficients are statistically significant at conventional statistical levels.

[Insert Tables 8 and 9 about here]

Next, in Table 8 we display the estimated coefficients of d in the estimated residuals under the assumption of white noise errors. It is observed that only for the relationship between Costa Rica and Salvador, and Honduras and El Salvador, the estimated values of d are smaller than 1 and the unit root null cannot be rejected. For the relationship between Guatemala and Nicaragua, the estimated d is slightly above 1 but the unit root cannot be rejected either. In all the other cases, this hypothesis is rejected in favour of d > 1.

[3] Alternative methods for the estimation of the cointegrating parameters were also employed including a Narrow Band Least Squared (NBLS) estimator as proposed in Robinson (1994b) and a Fully Modified NBLS as in Nielsen and Frederiksen (2011).

Using the Whittle semiparametric method (Robinson, 1995a), which is also valid in the context of cointegration, the estimated values of d are reported in Table 9. We only observe few cases where the unit root null cannot be rejected. They are the relationships of Costa Rica with El Salvador and Guatemala, Honduras with El Salvador, and Guatemala with Nicaragua. Thus, these results are completely in line with those based on the parametric approach in Table 8.

[Insert Table 10 about here]

Table 10 displays the results of testing the null hypothesis of no cointegration against the alternative of fractional cointegration through the Hausman test of Marinucci and Robinson (2001). For Costa Rica we found some partial evidence of cointegration with the Dominican Republic, El Salvador and Guatemala. The orders of integration of the estimated residuals in these cointegrating relationships are respectively 1.303, 1.237 and 1.077, and the rejection of the null is a consequence of the large order of integration obtained for Costa Rica, which is in some cases even above 1.50. Apart from these partial cases, the only clear evidence of cointegration occurs between Honduras and El Salvador, with an estimated order of integration of 1.082 for the cointegrating regression, much lower than the values obtained for the individual series, which are higher than 1.3 in the two series.

6. Concluding comments

In this study we have tried to expand the analysis of inflation stability in the member countries of the CMCA by using a long memory and fractionally integrated approach, and implementing cointegration methods that have not yet been used in the study of the Central American Monetary Council. We have done so with the aim of trying to detect the level of persistence in the inflationary rates of these Central American countries, and also in order to detect any possible case of cointegration between them, which could be a sign of homogeneity within the region.

Our results first show that all the series of prices are nonstationary, with orders of integration equal to or higher than 1 in all cases. Only for Guatemala and Nicaragua, in the cases of AR(1) and Bloomfield disturbances, the unit root null hypothesis (d = 1) cannot be rejected. In all the other cases, this hypothesis is decisively rejected in favor of higher degrees of integration. Similar results are obtained when using a semiparametric approach. These results will imply that shocks to the inflationary rates of the CMCA countries will have very long lasting effects, not disappearing by themselves in the long run. We believe this should be an important aspect to be taken into account by the CMCA policy makers.

Looking at the long run equilibrium relationships among the countries, we only found strong evidence of a cointegration relationship in the case of Honduras with El Salvador. All the other vis a vis relationships seem to diverge in the long run. Although this could be seen as an obstacle towards more and better integration within the region, we must not forget that the ultimate goal of the CMCA is not to gain a currency or monetary union in the near future. Its main objective remains to serve as an institution promoting economic and financial stability in the region. Thus we believe that the results we have obtained, though it might be a bad indication for a close monetary union of countries, the fact that shocks affecting a certain country might not have a direct effect on any of the neighboring countries, as suggested by the lack of cointegration relationship between almost all of them, can be seen as a positive sign, a fact that also needs to be taken into account. Nevertheless, further research should be conducted to verify these results.

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Table 1: Estimates of d for the case of white noise disturbances

Country	No regressors	An intercept	A linear time trend
COSTA RICA	0.97 (0.90, 1.07)	**1.40 (1.32, 1.51)**	1.33 (1.26, 1.45)
DOMINICAN REP.	0.97 (0.90, 1.07)	**1.45 (1.36, 1.57)**	1.45 (1.36, 1.57)
HONDURAS	0.99 (0.91, 1.08)	**1.42 (1.36, 1.49)**	1.34 (1.29, 1.41)
EL SALVADOR	0.98 (0.91, 1.08)	**1.07 (1.01, 1.17)**	1.07 (1.01, 1.15)
GUATEMALA	0.98 (0.90, 1.08)	**1.32 (1.23, 1.45)**	1.26 (1.16, 1.39)
NICARAGUA	0.99 (0.91, 1.09)	**1.15 (1.02, 1.33)**	1.14 (1.02, 1.30)

In bold the significant models according to the deterministic terms.

Table 2: Estimates of d for the case of AR(1) disturbances

Country	No regressors	An intercept	A linear time trend
COSTA RICA	1.37 (1.24, 1.54)	**1.28 (1.18, 1.39)**	1.19 (1.11, 1.28)
DOMINICAN REP.	1.36 (1.23, 1.53)	**1.25 (1.03, 1.43)**	1.25 (1.04, 1.43)
HONDURAS	1.38 (1.25, 1.55)	**1.47 (1.39, 1.57)**	1.37 (1.29, 1.46)
EL SALVADOR	1.38 (1.25, 1.54)	**1.15 (1.00, 1.28)**	1.11 (1.01, 1.25)
GUATEMALA	1.38 (1.24, 1.55)	**1.05 (0.98, 1.29)**	1.01 (0.86, 1.16)
NICARAGUA	1.39 (1.25, 1.57)	**0.88 (0.81, 1.02)**	0.86 (0.68, 1.01)

In bold: Statistical evidence of mean reversion at the 5% level.

Table 3: Estimates of d for the case of Bloomfield (1973) disturbances

Country	No regressors	An intercept	A linear time trend
COSTA RICA	0.95 (0.83, 1.12)	**1.23 (1.14, 1.37)**	1.16 (1.08, 1.26)
DOMINICAN REP.	0.95 (0.80, 1.12)	**1.24 (1.11, 1.41)**	1.24 (1.11, 1.40)
HONDURAS	0.97 (0.84, 1.14)	**1.49 (1.40, 1.62)**	1.37 (1.29, 1.49)
EL SALVADOR	0.97 (0.84, 1.14)	**1.14 (1.02, 1.31)**	1.12 (1.01, 1.27)
GUATEMALA	0.97 (0.83, 1.13)	**1.09 (0.80, 1.25)**	1.04 (0.93, 1.17)
NICARAGUA	0.94 (0.83, 1.11)	**0.90 (0.84, 1.01)**	0.88 (0.78, 1.00)

In bold: Statistical evidence of mean reversion at the 5% level.

Table 4: Estimates of d for the case of seasonal AR(1) disturbances

Country	No regressors	An intercept	A linear time trend
COSTA RICA	0.97 (0.88, 1.08)	**1.39 (1.30, 1.50)**	1.33 (1.24, 1.44)
DOMINICAN REP.	0.97 (0.89, 1.07)	**1.46 (1.37, 1.58)**	1.45 (1.36, 1.58)
HONDURAS	0.98 (0.89, 1.08)	**1.40 (1.34, 1.49)**	1.35 (1.28, 1.44)
EL SALVADOR	0.98 (0.89, 1.08)	**1.08 (1.00, 1.17)**	1.07 (1.00, 1.14)
GUATEMALA	0.97 (0.88, 1.08)	**1.32 (1.23, 1.44)**	1.25 (1.16, 1.38)
NICARAGUA	0.99 (0.90, 1.09)	**1.16 (1.01, 1.33)**	1.15 (1.02, 1.32)

In bold: Statistical evidence of mean reversion at the 5% level.

Table 5: Estimates of d based on a semiparametric method

	CR	DOM	HOND	ELS	GUAT	NIC	Lower	Upper
10	1.500	**1.187**	1.500	1.295	**1.050**	**1.015**	0.739	1.260
11	1.500	**1.204**	1.500	1.342	**1.134**	**0.988**	0.752	1.247
12	1.500	**1.191**	1.500	1.336	**1.172**	**1.041**	0.762	1.237
13	1.500	1.252	1.500	1.230	**1.147**	**1.102**	0.771	1.228
14	1.500	1.277	1.500	**1.186**	**1.135**	**1.166**	0.780	1.219
15	1.500	1.232	1.500	1.237	**1.144**	**1.082**	0.787	1.212
16	1.500	1.264	1.500	1.292	**1.190**	**1.070**	0.794	1.205
17	1.480	1.316	1.500	1.304	1.210	**1.081**	0.800	1.199
18	1.438	1.361	1.500	1.344	1.257	**1.101**	0.806	1.193
19	1.447	1.341	1.500	1.348	1.289	**1.070**	0.811	1.188
20	1.459	1.286	1.487	1.280	1.237	**1.085**	0.816	1.184

In bold, evidence of unit roots at the 5% level.

Table 6: Testing the homogeneity condition in the orders of integration

	DOM. REP.	HOND.	EL SALV.	GUATEMALA	NICARAGUA
COSTA RICA	**V**	**V**	**V**	**V**	**V**
DOM. REP.	XXXXX	**V**	**V**	**V**	**NO**
HONDURAS	XXXXX	XXXXX	**V**	**V**	**NO**
EÑ SALV.	XXXXX	XXXXX	XXXXX	**V**	**V**
GUATEMALA	XXXXX	XXXXX	XXXXX	XXXXX	**V**

Table 7: OLS estimates on the cointegrating regressions

OLS regression	α	β
Costa Rica / Dom Rep.	0.417 (8.05)	0.981 (74.37)
Costa Rica / Honduras	-0.768 (-16.39)	1.041 (107.22)
Costa Rica / El Salvador	-8.277 (-88.26)	2.853 (133.41)
Costa Rica / Guatemala	-2.101 (-102.12)	1.531 (308.66)
Costa Rica / Nicaragua	-0.916 (-27.04)	1.186 (154.65)
Dom. Rep. / Honduras	-0.875 (-8.398)	0.991 (45.94)
Dom. Rep. / El Salvador	-8.225 (-39.43)	2.7636 (58.09)
Dom. Rep. / Guatemala	-2.313 (-29.91)	1.499 (80.40)
Dom. Rep. / Nicaragua	-1.136 (-15.28)	1.158 (67.93)
Honduras / El Salvador	-6.982 (-53.18)	2.688 (89.78)
Honduras / Guatemala	-1.125 (-16.51)	1.433 (87.24)
Honduras / Nicaragua	-0.004 (-0.06)	1.108 (73.52)
El Salvador / Guatemala	2.191 (134.56)	0.530 (134.97)
El Salvador / Nicaragua	2.602 (155.32)	0.410 (106.87)
Guatemala / Nicaragua	0.771 (43.93)	0.775 (192.37)

Table 8: Estimated values of d on the residuals from the cointegrating regression

	No regressors	An intercept	A linear time trend
Costa Rica / Dom Rep.	1.14 (1.06, 1.23)	**1.40 (1.31, 1.51)**	1.40 (1.31, 1.50)
Costa Rica / Honduras	1.02 (0.95, 1.12)	**1.19 (1.10, 1.28)**	1.18 (1.10, 1.26)
Costa Rica / El Salvador	0.97 (0.88, 1.09)	**0.99 (0.92, 1.08)**	0.99 (0.92, 1.08)
Costa Rica / Guatemala	1.18 (1.07, 1.33)	**1.19 (1.08, 1.33)**	1.19 (1.08, 1.33)
Costa Rica / Nicaragua	1.23 (1.11, 1.40)	**1.12 (1.02, 1.29)**	1.13 (1.02, 1.30)
Dom. Rep. / Honduras	1.10 (1.03, 1.17)	**1.41 (1.34, 1.51)**	1.40 (1.33, 1.50)
Dom. Rep. / El Salvador	1.07 (0.98, 1.17)	**1.18 (1.11, 1.26)**	1.17 (1.11, 1.24)
Dom. Rep. / Guatemala	1.16 (1.08, 1.26)	**1.40 (1.31, 1.51)**	1.39 (1.31, 1.50)
Dom. Rep. / Nicaragua	1.12 (1.06, 1.20)	**1.33 (1.24, 1.45)**	1.33 (1.24, 1.44)
Honduras / El Salvador	0.96 (0.90, 1.04)	**0.96 (0.88, 1.04)**	0.96 (0.89, 1.04)
Honduras / Guatemala	1.08 (1.01, 1.17)	**1.25 (1.14, 1.34)**	1.24 (1.15, 1.34)
Honduras / Nicaragua	1.18 (1.07, 1.31)	**1.21 (1.13, 1.34)**	1.21 (1.13, 1.34)
El Salvador / Guatemala	1.08 (1.01, 1.16)	**1.25 (1.18, 1.34)**	1.24 (1.18, 1.34)
El Salvador / Nicaragua	1.18 (1.08, 1.32)	**1.21 (1.13, 1.34)**	1.21 (1.13, 1.33)
Guatemala / Nicaragua	1.17 (1.05, 1.36)	**1.06 (0.95, 1.20)**	1.06 (0.95, 1.21)

Table 9: Estimated values of d on the residuals of the cointegrating regression with a semiparametric Whittle method

OLS regression	m = 15	m = 16
Costa Rica / Dom Rep.	1.305	1.303
Costa Rica / Honduras	1.500	1.497
Costa Rica / El Salvador	**1.093***	**1.133***
Costa Rica / Guatemala	**1.088***	**1.077***
Costa Rica / Nicaragua	1.366	1.406
Dom. Rep. / Honduras	1.396	1.418
Dom. Rep. / El Salvador	1.370	1.416
Dom. Rep. / Guatemala	1.278	1.303
Dom. Rep. / Nicaragua	1.385	1.352
Honduras / El Salvador	**1.082***	**1.127***
Honduras / Guatemala	1.391	1.424
Honduras / Nicaragua	1.500	1.500
El Salvador / Guatemala	**1.119***	**1.166***
El Salvador / Nicaragua	1.221	1.239
Guatemala / Nicaragua	1.238	**1.164***

*: Evidence of unit root in the estimated residuals.

Table 10: Testing the null of no cointegration against the alternative of fractional cointegration

	DOM. REP.	HOND.	EL SALV.	GUATEMALA	NICARAGUA
COSTA RICA	4.928* 0.193 1.303	0.002 0.001 1.497	21.036* 2.633 1.237	22.273* 0.570 1.077	2.280 1.126 1.366
DOM. REP.	XXXXX	2.212 1.373 1.396	1.427 0.772 1.370	0.025 0.983 1.278	XXXXX
HONDURAS	XXXXX	XXXXX	**22.189*** **5.600*** **1.082**	1.508 3.130 1.391	XXXXX
EL SALV.	XXXXX	XXXXX	XXXXX	3.800 0.640 1.119	0.640 2.453 1.221
GUATEMALA	XXXXX	XXXXX	XXXXX	XXXXX	0.099 0.854 1.164

N.A. means not applicable. The first two values refer to the test statistics for Hx and Hy respectively using the Hausman test of Marinucci and Robinson (2001). The third value is the estimated value of d^*. $\chi_1^2(5\%) = 3.84$.
In bold and with an asterisk, those cases where we reject the null hypothesis of no cointegration at the 5% level.