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A MONETARY UNION FOR LATIN AMERICA? AN EMPIRICAL INVESTIGATION.

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Abstract

In this paper we assess the possibility of creating a currency union for the Latin American region based on satisfying some of the OCA criteria. Following the analysis of Sun and Simons (2011), we applied cointegration and Granger causality techniques to analyze the long-run and short-run interactions of the real effective exchange rates. We found strong long-run ties between Colombia, Mexico, Chile, Brazil, and Bolivia; as well as weaker ties between these five countries and Argentina, Paraguay and Peru. Such a monetary union would comprise 85% of Latin America's GDP and 78% of its total population. The hypothetical monetary union is not composed of current regional trading blocs, but rather membership is based on "size matters". We also examined the degree of financial or monetary convergence within the hypothetical monetary union following Bholla, et al. (2011). Considerable dissimilarities in the interest rate pass-through levels were found, both in the short-run and the long-run. A monetary union may be one possible option to protect against macroeconomic shocks and enhance economic integration within Latin America. Nevertheless, the two regional giants, Mexico and Brazil, would be needed to drive the process of further economic integration.

Key Words: monetary union, optimum currency areas, cointegration, Granger causality, interest rate pass-through, error-correction models.

Contents

SECTION 1. INTRODUCTION

Is the formation of a monetary union feasible in Latin America? The aim of this paper is to assess the possibility of creating a currency union for the region of Latin America based on economic criteria. In light of the launch of the European monetary union, and the ongoing debates of forming currency unions in East Africa, East Asia and the nations of the Gulf Cooperation Council,[1] the motivation behind this paper is to re-stimulate debate among academics and policymakers about the possibility of forming a Latin American monetary union.

In Latin America, the debt crisis of the 1980s delayed trade intensification until the 1990s with the appearance of regional trade agreements like the Southern Common Market (Mercosur, 1991) and the North American Free Trade Agreement (NAFTA, 1994).[2] In 2002, the Inter-American Development Bank (IADB) published the report *Beyond Frontiers: The New Latin American Regionalism*. In this twelve chapter report, a monetary union was suggested as a possible stabilization option for countries that had already entered into regional trade agreements. In 2006, Central American and the Caribbean activism to promote trade secured an important victory when the U.S. ratified a long-awaited free trade agreement (the DR-CAFTA) with Costa Rica, El Salvador, Guatemala, Honduras, Nicaragua and Dominican Republic. Since then, the strongest rhetorical statements about forming a monetary union have come from the Central American and the Caribbean region.

The purpose of this paper is two-fold. First, the paper investigates which Latin American countries could feasibly enter into a currency union. The criteria

[1] See, for example, Bholla et al (2011), Louis et al (2012) and Sun and Simons (2011).
[2] There were also renewed integration energies in the Caribbean Community and Common Market (CARICOM, 1973), the Centre American Common Market (CACM, 1960), and the Andean Community (CAN, 1969).

are based on both the degree of symmetry of real macroeconomic shocks, and the degree of similarity in economic structure. To do this we use real effective exchange rate data and employ the empirical techniques of cointegration and Granger causality. While these empirical techniques are standard for time series analysis, their usefulness in identifying solid long-term relationships between time series variables yield important insights concerning the feasibility of monetary unions. The second purpose of the paper is to investigate the level of financial integration among Latin American countries by testing the degree of interest rate pass-through. Using lending, deposit and discount rate data, we assess the current level of financial integration amongst the candidate countries identified by the cointegration and Granger causality analysis. The pass-through analysis is, essentially, an error-correction model technique, which helps us to estimate both short-term and long-term effects of a (theoretically independent) time series on another (theoretically dependent) one. The usefulness of this technique lies in its ability to shed some insight into the differences of the monetary policy transmission mechanism currently experienced among the countries, and therefore, signal the variety of rigidities in the financial systems that a common central bank would face.

To date the number of empirical studies that have investigated the feasibility of forming a monetary union in Latin America is small, especially when compared to the literature that has focused on the political and cultural feasibility, and theoretical costs and benefits of monetary unions. The existing literature has typically focused on evaluating the symmetry of fundamental shocks, using GDP and price data, and applying the Blanchard-Quah Structural Vector Autoregression (SVAR) technique (e.g. Foresti, 2007). This strategy, while standard and can uncover the correlation analysis of shocks, is likely to remove any information about the long-run equilibrium relationship by not considering the possibility that the variables share long-run ties that leads them to similar growth paths.

There are two main contributions of this paper. First, by applying the cointegration and Granger causality techniques we make a first attempt to uncover the long-run relationships among Latin American countries to help determine the presence of symmetry of macroeconomic shocks across the region. Additionally, the use of the real effective exchange rate allows us to consider a large number of countries with a good sample of observations focusing on a period when patterns of regional trade have been enhanced, i.e. since the Latin American countries began following open strategies (promote its export qualities, resort less to policies that reduce imports, and essentially see incorporating to the world market as a strategy that promotes sustained economic growth). Second, we extend the empirical study to evaluate whether the countries identified in the hypothetical Latin American monetary union have a similar degree of financial integration. This exercise not only acts as a robustness test, but at the same time it also identifies (or rejects) the existence of another important precondition for forming a feasible monetary union.

An important issue relates to what currency might be adopted by the potential currency union in Latin America. Recent cases of dollarization in Ecuador in 2000, and El Salvador in 2001 highlight the importance of the U.S. in the region. Thus one conceivable possibility is that the dollar could be employed as an anchor for the new common currency.[3] On the other hand, the rapid economic growth and global integration of China has triggered debates on its future implication for Latin America. China's average annual growth rate of exports and imports to and from the Latin American region during 2005-2009, was 24.8% and 24.5%, respectively. Additionally, projections suggest that China

[3] We do not consider the adoption of a foreign currency *per se*, but rather the creation of a new currency heavily fixed to a foreign currency or to a weighted basket of important existing ones.

could displace the European Union as the region's second largest trading partner in the middle of the coming decade (ECLAC, 2012). Therefore, we also assess the suitability of the Chinese yuan to act as a solid anchor for any common currency in Latin America.

The main results are as follows. First, the cointegration and Granger causality analysis suggest the potential of forming a monetary union comprised of Mexico, Chile, Colombia, Bolivia and Brazil. Argentina and Paraguay are also loosely tied to this main group. Second, the hypothetical monetary union is not composed of current regional trading blocs, but rather membership is based on "size matters". This contrasts with the existing literature which favors the formation of a monetary union based on trade agreements and/or geographic proximity. This analysis suggests that a monetary union would need to include, at least, the region's two largest economies (Brazil and Mexico) and would account for 69.7% (84.57%) of the Latin American region's total GDP and 64.94% (78.13%) of its total population, depending on the countries involved. Third, we did not find evidence that dollarization or adopting the Chinese yuan would be suitable. Fourth, there is little evidence of financial convergence amongst the members of the hypothetical monetary union: the analysis suggests that the level of interest rate pass-through between countries is very low and dissimilar, both in the short-run and the long-run. This suggests that substantial banking reforms would be needed in order to overcome these financial rigidities and remove one important obstacle to creating a successful monetary union in the region.

The empirical results also suggest the following important policy implications. A monetary union for Latin America should not be discarded as one feasible option to protect against global macroeconomic shocks and enhance regional economic integration. Nevertheless, the two regional giants, Mexico and Brazil, would be required to drive this process forward. Greater political cooperation and economic integration between these two countries would enhance the feasibility of any future successful monetary union. Further, countries engaged in free trade agreements should reevaluate the effects of these agreements in generating trade enhancement and economic similarities. Finally, we hope that this paper encourages future research and informed public debate on the issue of a potential monetary union in Latin America.

The rest of the paper is organized as follows. Section 2 reviews the relevant literature, with particular focus on the theory of Optimum Currency Areas, and the empirical work on Latin America. Section 3 performs the cointegration and Granger causality analysis using real effective exchange rate data, whereas Section 4 performs the interest rate pass-through analysis using error-correction model techniques. Finally, Section 5 briefly concludes.

SECTION 2. LITERATURE REVIEW

2.1 The Theory of Optimum Currency Areas

The formation of a monetary union demands its member countries to irrevocably peg their exchange rates rigidly to each other. When would a country benefit from giving up both control over monetary policy setting and its ability to use the exchange rate as a shock absorber? The seminal papers of Mundell (1961) and McKinnon (1963) develop the basic theoretical foundations of what constitutes an Optimum Currency Areas (OCAs). Three key requirements are highlighted as preconditions to the success of a monetary union. First, a high degree of factor mobility (capital and labor) between member countries. Second, symmetry of macroeconomic shocks, so that policy choices of one country is appropriate for the others. Third, the economies being closely linked by trade.

Over the 1960s and 1970s, the prerequisites of an OCA grew significantly, following its original Keynesian foundations.[4] After that, the number of articles on OCAs declined dramatically during the 1980s (Bayoumi and Eichengreen, 1998). The debate in Europe to increase regional integration during the 1990s renewed interest on the subject of monetary unification. However, given the new heavily formalized approach in macroeconomics, the verbal arguments of previous contributions were somewhat unsatisfactory. In that sense, the contribution of Bayoumi (1994) using a general equilibrium approach and stronger microeconomic foundations, was very important. From this paper, we can emphasize two conclusions of Bayoumi's model. First, the correlations of the underlying disturbances are clearly important for the success of a monetary union. Second, while a monetary union might raise the welfare of the regions within the union, it unambiguously lowers welfare for regions outside the union. These conclusions justified the necessity of empirical literature focusing on the nature and correlation of macroeconomic shocks within a potential monetary union. Additionally, the results signaled possible implicit ex-post benefits of monetary unions because of future enhanced trade and avoiding the losses of staying outside the monetary union.

Deregulation policies of the financial sector in developed countries and free capital mobility during the 1990s, combined with fixed or mixed exchange rate regimes, resulted in systemic crises and heavy losses in many developing countries due to "sudden stops" and "speculative attacks". Many articles emphasized the lack of central bank independence and government's strong incentives to inflate, as causes of self-fulfilling panics (see e.g. Sachs, et al.,

[4] See the summarized Decalogue in Edwards (2006).

1996). These raised questions on how costly the loss of monetary independence really was for developing countries. Suggested losses of adopting a common currency with other countries focalized on losing both an element of national identity and the seignorage gains of the government, while potential gains where true credibility and stabilizing inflation (Edwards, 2006). Nevertheless, if potential members face different allocation costs for taxes due to a less developed tax system, and hence different optimal inflation rates, a common currency that forces convergence in inflation rates could be extremely costly for that country with less chances of raising public funds through taxes (Goldberg, et.al., 1993). Academics viewed these new reflections as evidence of the importance to achieve previous financial convergence as a prerequisite for joining a monetary union: hence the need for empirical financial research on the OCA criteria.

2.2 Empirical Research

Novel empirical methodologies in macroeconomics during the 1980s; particularly, the development of Vector Autoregression (VAR) models (Sims, 1980) and Structural Vector Autoregression (SVAR) models (Blanchard and Quah, 1989), were adopted as the standard approach to find evidence of symmetry of shocks. Bayoumi and Eichengreen (1994), using data for the period 1969-89, employed the Blanchard and Quah decomposition method to uncover similarities of supply and demand disturbances between regions that were theoretical candidates for a monetary union. In particular, they were interested in three regions: Europe, Asia, and the Americas.

Bayoumi and Eichengreen's (1994) results found no evidence to support Latin America as an OCA, even if the U.S. or Canada were included in the hypothetical monetary union.[5] One interesting discovery was the lack of correlation evidence between the NAFTA and Southern Common Market (*Mercosur*) members.[6] In subsequent work, Eichengreen (1998) focused on the *Mercosur* region, where he evaluated four of the OCA criteria using the bilateral exchange rate: similarity of disturbances, commodity composition of trade, bilateral trade, and country size. Eichengreen observed higher than predicted exchange-rate volatility, but even the predicted levels were higher than those for most European countries. Despite rejecting, once again, *Mercosur* as an OCA, he was confident that the choice of exchange rate regime (especially the Convertibility Plan in Argentina) could change countries' characteristics in the medium-run, hence turning feasible a monetary union. However, the Brazilian crises of 1998/99 and the 2001/02 crisis in Argentina showed that such exchange rate regimes were far from stabilizing.

Most literature on the feasibility of forming a monetary union in Latin America has relied either on analyzing political and cultural obstacles, or on changes in the degree of trade, within the region. Still, interesting empirical work has been done, in the spirit of Bayoumi and Eichengreen's methodology. Alesina, Barro, and Tenreyro (2002) used yearly data on outputs, bilateral trade, and prices from 1960 to 1997 to form a panel of countries. They asked what "strong" currency could work as the best "monetary anchor"; the options were: the U.S. dollar, the euro, and the yen. They also adopt an instrumental variables approach to deal with endogeneity of some OCA criteria that might bias some results when using the SVAR methodology. Their results could not define a yen area, and suggested heterogeneous policy implications for the Latin American

[5] The sample consisted of 11 Latin American countries: Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Paraguay, Peru, Uruguay, and Venezuela.

[6] The North American Free Trade Agreement (NAFTA) includes Canada, the U.S., and Mexico. In 1994, *Mercosur* members were Argentina, Brazil, Paraguay and Uruguay.

area (e.g. Mexico and Ecuador should favor the euro, while Argentina the dollar). Hallwood, Marsh, and Scheibe (2006) used quarterly data on GDP and CPI from 1995Q1-2001Q4, combined with VAR techniques to analyze if five Latin American countries (Argentina, Brazil, Chile, Uruguay and Venezuela) would benefit from a monetary union either through official dollarization or by adopting a common currency arrangement. In this study the results were also negative, as they concluded that macroeconomic shocks are highly asymmetric between the Latin American countries and the U.S. While more positive evidence of shock correlation is found within the set of Latin American countries alone, this was not strong enough to consider recommending the formation of a monetary union. Finally, Foresti (2007) gathers annual data on the GDP and GDP deflator for ten Latin American countries from 1962-2003 and also estimates supply and demand disturbances using the Blanchard and Quah decomposition method. To our knowledge, this is one of the most detailed empirical works focusing solely on the Latin American region. The results suggest that a high percentage of output variability is due to permanent shocks. Moreover, the variability on short-run inflation was significantly different across countries. Foresti recommended a higher degree of policy coordination and economic integration prior to any attempt of creating a monetary union.

Recently, the empirical literature has stressed the importance of financial convergence as a prerequisite for joining a monetary union. Specifically, we focus on the transmission of interest rate shocks both domestic and internationally. De Bondt, Mojon and Valla (2005) can be considered the pioneering empirical paper on interest rate pass-through. They used data for European countries on 41 monthly retail interest rate series from April 1994 to December 2002. The techniques employed where Granger causality and error-correction models. They were interested in finding possible sluggishness in the monetary policy transmission mechanism due to the difference in maturity between retail bank products and money market interest rates. It was shown that retail rates depend on long-term market interest rates, and that there exist incomplete pass-through because short-term market interest rates movements are not fully transmitted to market interest rates with longer maturity. On the other hand, Edwards (2009) used weekly data on interest rate differentials for seven countries from Latin America and Asia, for the period January 2000 – September 2008. The results suggest a negative pass-through from Fed policy rates to interest rate differentials with a coefficient of -0.5. Also important, is that in Latin America there is an immediate overreaction to changes in Fed policy rates, ending with a very fast and cyclical adjustment.

To end this section, we summarized the two key papers that we will follow in the next two sections. Sun and Simons (2011) investigates the feasibility of forming a monetary union in East Asia. They used monthly observations on the real effective exchange rate for eleven countries, from 1981-2007. Using real effective exchange rates they employ the techniques of cointegration and Granger causality, to uncover possible long-run common trends, and short-run competitive adjustments and interdependence. Their results suggest that there might be five countries with potential for success for further monetary integration; namely, South Korea, the Philippines, Thailand, Indonesia, and Malaysia. The motivation in following their approach is twofold. First, its use of an informative variable not yet fully explored for OCA analysis in Latin America, which additionally has observations for a large number of countries in the region. Second, by using a cointegration approach we consider the possible removal of important information about the long-run relationships that might occur when using the Blanchard-Quah SVAR methodology.

Bholla, Aziakpono and Snowball (2011) investigate whether the five member countries of the East African Community (EAC); namely, Burundi, Kenya, Rwanda, Tanzania, and Uganda, have achieved a similar pass-through of monetary policy. To this end, they used cointegration, error-correction, and asymmetric error-correction frameworks. According to economic theory, differences in the structure of the banking system and in the level of information technology seen in the different countries are the basic factors that explain dissimilarities in the pass-through adjustment processes.[7] Their results indicate a low degree of financial convergence amongst the countries and considerable sluggishness in the area to fully realize the monetary policy transmission mechanism. The motivation in following their approach in this paper is to obtain some information on the current level of financial convergence in Latin America, and to uncover some of the monetary and financial reforms that are required to make a Latin American monetary union a feasible possibility.

SECTION 3. COINTEGRATION AND GRANGER CAUSALITY ANALYSIS

3.1 Data description and the use of the real effective exchange rate

A fundamental prerequisite for a successful monetary union requires the symmetry of common shocks across member countries. Using real effective exchange rate (REER) data, this section attempts to evaluate the degree of similarity in the economic structure of Latin American countries with respect to their adjustment to real macroeconomic shocks. There are three main advantages of using the REER in the OCA analysis. The first advantage relates to the issue

[7] See Sander & Kleimeier (2004).

of data availability: a total of fifteen Latin American economies can be analyzed, including the nine largest countries and four Central American and Caribbean countries which are typically ignored from a multi-regional analysis. The second advantage is the theory underpinning the REER. A measure of the real competitiveness of each country is needed to help us uncover implicit economic adjustments to common exogenous shocks. In essence a real effective exchange rate can be considered a good proxy of that real competitiveness.[8] As discussed by Sun and Simons (2011), the third advantage of REERs is that they often contain richer information for countries that follow open strategies and share similar patterns of trade and trade composition. Many Latin-American countries: (i) import capital equipment and technology from abroad, (ii) export raw materials with which many countries are well-endowed, (iii) trade parts and intermediate goods between each other that they use to export final goods, (iv) take advantage of relatively low-wages. Accordingly, REERs may represent possible common stochastic trends, driven by macroeconomic fundamentals, in a better way than simple bilateral real exchange rates.

The cointegration technique allows us to characterize the historical adjustments of two REERs through a significant linear approximation: if two REERs are found to be cointegrated then this suggests evidence of real shock symmetry. A similar economic structure might necessitate similar countercyclical policies in two different countries. On the other hand, Granger causality tests permit us to conclude whether adjustments in one country (reflected in its REER) trigger significant adjustments in another country, hence revealing interdependence between the two economies.

[8] A caveat is worth mentioning for the Latin-American region. The REER can only be considered a good proxy of international competitiveness since the mid-1990s, when an open globalized strategy became the norm (Berg, et al., 2002).

For both the cointegration and Granger causality analyses we use monthly data for the real effective exchange rate (REER). The series covers the period January 1996- November 2012 and includes fifteen Latin-American countries[9]: Mexico, Belize, Nicaragua, Costa Rica, Dominican Republic, Colombia, Venezuela, Ecuador, Bolivia, Brazil, Chile, Paraguay, Uruguay, Peru, and Argentina. With the exception of Argentina and Peru, all the series were obtained from the *International Financial Statistics* (IFS) database of the International Monetary Fund (IMF).[10] The maximum number of missing values for a country is 23 for Paraguay, so we do not envisage problems because of missing data.[11] In what follows an increase (decrease) in the REER represents an appreciation (depreciation).

3.2 Econometric Methodology

Following Sun and Simons (2011), the econometric methodologies employed in this section are
cointegration and Granger causality. We briefly discuss each in turn.

Cointegration (Johansen methodology)

All variables are pretested to assess their order of integration. Straightforward regression analysis cannot be implemented if the variables are found to be non-

[9] By starting the series in 1996 we sidestep the worst effects of the *Tequila Crisis* and focus on the post-crisis adjustments in the region.
[10] The REERs for Argentina and Peru where obtained from their respective central bank's websites, with no important concerns about missing values.
[11] We used standard practices to reduce the number of missing values. If the missing value was unique, it was filled with the average of the two closer values. Also we did not encounter long straight series of missing values, so important biases due to recognized trends (appreciation or depreciation) were solved by interpolating.

stationary (i.e., have unit roots). After a visual analysis of the time series, each of the monthly REER series was checked for the presence of a unit root using the Dickey-Fuller with Generalized Least Squares detrending (DFGLS) test. The Augmented Dickey-Fuller and Phillips-Perron tests were also performed for robustness. Only the Dominican Republic's REER series was not found to be an I(1) process (stationary after one difference), so this country was omitted from the rest of the analysis.[12]

The notion of cointegration (Engle & Granger, 1987) allows regression analysis using I(1) variables to be potentially significant. For illustration, suppose $\{y_t\}_{t=0}^{\infty}$ and $\{x_t\}_{t=0}^{\infty}$ are two I(1) processes. Then (generally), $y_t - \gamma x_t$ is an I(1) process for any number γ. Nevertheless, it is possible that for some $\gamma \neq 0$, $y_t - \gamma x_t$ is an I(0) process. If that is the case, we say that *y* and *x* are cointegrated and γ is the parameter of cointegration.

Denoting the bivariate vector of two REERs, countries *i* and *j*, as $R = \left(r_i, r_j\right)'$, we estimate the following vector autoregressive model:

$$R_t = \beta_0 + \beta_1 R_{t-1} + \beta_2 R_{t-2} + \cdots + \beta_p R_{t-p} + \varepsilon_t \qquad (1)$$

where β_0 is *2x1* vector of constants; β_i, *i= 1, 2,…,p* is the *2x2* coefficient matrix of R_t lagged by *i* periods, and ε_t is the *2x1* vector of residuals that satisfies standard Gaussian properties. The optimum lag length, *p*, is determined using the likelihood ratio test.[13]

[12] Detailed results of the unit root tests are available upon request.
[13] This estimator generally favors a larger number of lags than the Akaike and Standard Bayesian Information Criteria (AIC and SBC, respectively) multivariate generalizations**.** We favored a large number of lags to control for possible autocorrelation issues in the residuals of the VAR models employed here.

One problem with our period of analysis is the possibility of structural breaks that may lead towards cointegration biases. The structural breaks that are most likely to cause noise are the 1998/99 crisis in Brazil, the 2001/02 crisis in Argentina, and the 2007/08 global financial crisis. Ultimately, we only considered impulse dummies for the Brazilian crisis and the recent global financial crisis in our estimation of the optimum number of lags in the VAR model. An impulse dummy for the Argentinian crisis was only included for Argentina, as recent evidence suggests that this crisis was locally concentrated and did not generate systemic problems to the region as a whole (see, e.g. Allegret & Sand-Zantman, 2009).

It can be shown (see, e.g. Enders, 2010), that model (1) can be rewritten in vector error-correction (VEC) form as:

$$\Delta R_t = \beta_0 + \pi R_{t-1} + \sum_{i=1}^{p-1} \pi_i \Delta R_{t-i} + \varepsilon_t \qquad (2)$$

$$\pi = -\left(I - \sum_{i=1}^{p} \beta_i\right) \quad , \quad \pi_i = -\sum_{j=i+1}^{p} \beta_j \quad , \quad I\text{: identity matrix.}$$

The key feature to note in (2) is the rank of matrix π, which is *2x2* in our case. The rank of π is equal to the number of independent cointegrating vectors. Equivalently, the rank of π is equal to the number of its non-zero characteristic roots (or eigenvalues). We use the following test for the number of cointegrating relationships in our bivariate model:

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{2} \ln\left(1 - \hat{\lambda}_i\right), \qquad (3)$$

where $\hat{\lambda}_i$ is the estimated values of the characteristic roots (eigenvalues) obtained from the estimated matrix π, and T is the number of usable observations. This statistic tests the null hypothesis that the number of distinct cointegrating vectors is less than or equal to r against a general alternative. Note that the λ_{max} statistic is not used because it requires a large sample size (about 300 observations) to be reliable (Kennedy, 2008). If we reject the null, that is $Rank(\pi) \geq 1$, the two REERs are cointegrated and may share long-run common trends. The final considerations we must take into account are stability and white noise residuals of the VEC model. After accepting cointegration, the stability of the implicit model is tested using the eigenvalue stability condition.[14] Finally, we perform the Lagrange-multiplier (LM) test for the null hypothesis of no autocorrelation in the residuals of the VEC model.

Granger Causality

In a VAR, each random disturbance influences all the endogenous variables. The point of estimating a VAR or VEC (when accounting for the possibility of cointegration) is to characterize the joint distribution of the elements of the vector of variables. Nonetheless, random disturbances may exhibit their influence on some of the endogenous variables earlier. Granger causality (Granger, 1969; Granger, 1980) helps to identify evidence of this temporal ordering. This is tested by examining whether lagged values of one variable improve the forecast of another variable, after lagged values of that other variable are already taken into account.

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[14] It is *implicit* because we do not perform any numerical analysis, such as forecasting or impulse-response functions, using our VEC model.

Specifically, we test if the lagged values of the REER of one country are useful in forecasting the other REER in our optimum bivariate systems. Suppose we find evidence of Granger causality from the REER of country B on the REER of country A. Three possible explanations can explain this finding. First, country A could have an intervention policy on the world market and considers country B as a close competitor in trade, hence, it uses the exchange rate as its policy tool to maintain competitiveness whenever it feels threatened. Second, currency A is pegged to currency B.[15] Third, there exists strong complementarities in the products that the two countries export. For example, country B might export an important raw material that country A uses as an input to produce an important export good of its own.

First consider the case when the REERs are I(1) and thus cointegrated. In this case there is an error-correction term that is significant. To overcome this problem, we follow Dolado and Lutkepohl (1996), whose methodology leads to valid Wald tests with asymptotic χ^2 distributions. First recall the basic VAR system given in (1), which is estimated after finding the appropriate number of lags *p*. Now we refit the data with a VAR(*p* + 1), such that:

$$r_{i,t} = \delta_{0,i} + \sum_{k=1}^{p+1} \delta_{41,k} r_{i,t-k} + \sum_{k=1}^{p+1} \delta_{42,k} r_{j,t-k} + \varepsilon_{i,t} \qquad (4)$$

$$r_{j,t} = \delta_{0,j} + \sum_{k=1}^{p+1} \delta_{51,k} r_{i,t-k} + \sum_{k=1}^{p+1} \delta_{52,k} r_{j,t-k} + \varepsilon_{j,t} \qquad (5)$$

[15] We should not think only in cases where the pegging of the currency is strictly enforced. While many Latin American countries currently follow a floating exchange rate regime, central banks commonly intervene in foreign exchange markets. Also note that because we are working with a real weighted exchange rate, there might be strong dissimilarities between the REER and the bilateral nominal exchange rate commonly reported.

where $\delta_{ls,k}$ is the coefficient of the *kth* lag of variable *s* on equation *l*; $\delta_{0,s}$ is the intercept when *s* is the dependent variable; and $\varepsilon_{s,t}$ is the white noise residual when *s* is the dependent variable. Even though we now have least squares estimators of the coefficients that are consistent and asymptotically efficient, the Granger causality test only uses the first *p* optimum number of lags. Consider equation (4). To see if the REER of j Granger causes the REER of i, test: $H_0\colon \delta_{42,k} = 0, \quad \forall\, k = 1, 2, \dots, p$; that is, we test the null hypothesis that the REER of country *j* does not Granger cause the REER of country *i*. We use an estimated F-test statistic: if it is greater than the critical value, then the null hypothesis is rejected, and we conclude that country's *j* REER Granger causes the REER of country *i*.

Now consider the case when the REERs are I(1) but non-cointegrated. We estimate the following VAR model:

$$\Delta r_{i,t} = \delta_{0,i} + \sum_{k=1}^{p} \delta_{61,k} \Delta r_{i,t-k} + \sum_{k=1}^{p} \delta_{62,k} \Delta r_{j,t-k} + \varepsilon_{i,t} \qquad (6)$$

$$\Delta r_{j,t} = \delta_{0,j} + \sum_{k=1}^{p} \delta_{71,k} \Delta r_{i,t-k} + \sum_{k=1}^{p} \delta_{72,k} \Delta r_{j,t-k} + \varepsilon_{j,t}. \qquad (7)$$

Notice that the coefficients are analogous to the case in (4) and (5) but are differenced. Let's illustrate the Granger-causality test with equation (6). The null hypothesis of no Granger causality is: $H_0\colon \delta_{62,k} = 0, \quad \forall\, k = 1, 2, \dots, p$.

3.3 Cointegration Results

Table 1 reports the λ-trace test results for the 91 bivariate models estimated following the Johansen methodology. Despite using impulse dummies when estimating the correct VAR model, as mentioned in the methodology subsection, we did not consider evidence of cointegration when rejecting the null hypothesis at the 10% significance level (λ-trace critical value of 13.33) to further reduce the danger of cointegration biases. The hypothesis of no cointegration is rejected in 9 out of the 91 cases.

Table 1. Cointegration Analysis for Latin American Countries

	MEX	BEL	NIC	CR	COL	VEN	ECU	BOL	BRA	CHI	PAR	URU	PER
BEL	9.61												
NIC	14.12	5.85											
CR	12.12	8.18	5.04										
COL	**19.82***	5.17	10.19	8.63									
VEN	12.43	9.75	7.74	8.37	7.66								
ECU	17.68#	7.74	18.43#	10.16	12.19	9.47							
BOL	9.89	4.5	9	6.72	**16.53***	8.23	10.8						
BRA	15.35	4.59	9.25	8.41	**24.19****	7.23	8.38	**16.03***					
CHI	**21.04****	3.95	13.63	8.05	**20.09****	9	11.09	14.54	8.81				
PAR	11.03	3.49	8.97	9.44	11.97	12.87	10.83	**16.25***	10.78	7.34			
URU	13.3	3.47	9.67	14.89	14.56	8.86	9.21	9.2	12.73	9.02	14.92		
PER	13.69	7.19	9.83	6.96	11.01	6.8	10.83	9.04	9.44	12.7	**16.38***	9.06	
ARG	**19.93***	7.44	14.1	2.65	9.77	5.26	12.7	3.12	8.58	11.9	6.47	2.95	5.08

Note: Each number is the λ-trace statistic estimated for testing cointegration between row country's REER and column country's REER. Critical values (Drift case considered) are: 15.41 at 5% significance level; 20.04 at 1% significance level. Hence; *, **, denote significance at 5% and 1%, respectively. # denotes cases where the λ-trace statistic signaled evidence of cointegration but the implicit VEC model was not stable or had not white noise residuals (or both). Abbreviations: MEX=Mexico, BEL=Belize, NIC=Nicaragua, CR=Costa Rica, COL=Colombia, VEN=Venezuela, ECU=Ecuador, BOL=Bolivia, BRA=Brazil, CHI=Chile, PAR=Paraguay, URU=Uruguay, PER=Peru, ARG=Argentina

To summarize the cointegration relationships, Figure 1 illustrates an "Extended Diamond" of the results. The main diamond shows the only group that exhibits strong long-run co-movements. Colombia works as the *heart* of the diamond, as evidence of cointegration was found between Colombia and Mexico, Chile, Brazil, and Bolivia. At the same time, these four countries exhibit evidence of cointegration with one extra member of the group.

Cointegration is found between Mexico-Chile, and Brazil-Bolivia. The "extension" links in the main group come from Mexico and Bolivia. The former shows evidence of long-run ties with Argentina, while the latter contains a link with two countries: Paraguay and Peru. Paraguay exhibits a weak link with the main group because of cointegration evidence with Bolivia. In turn, Peru exhibits an even weaker link because its only relationship is with Paraguay.

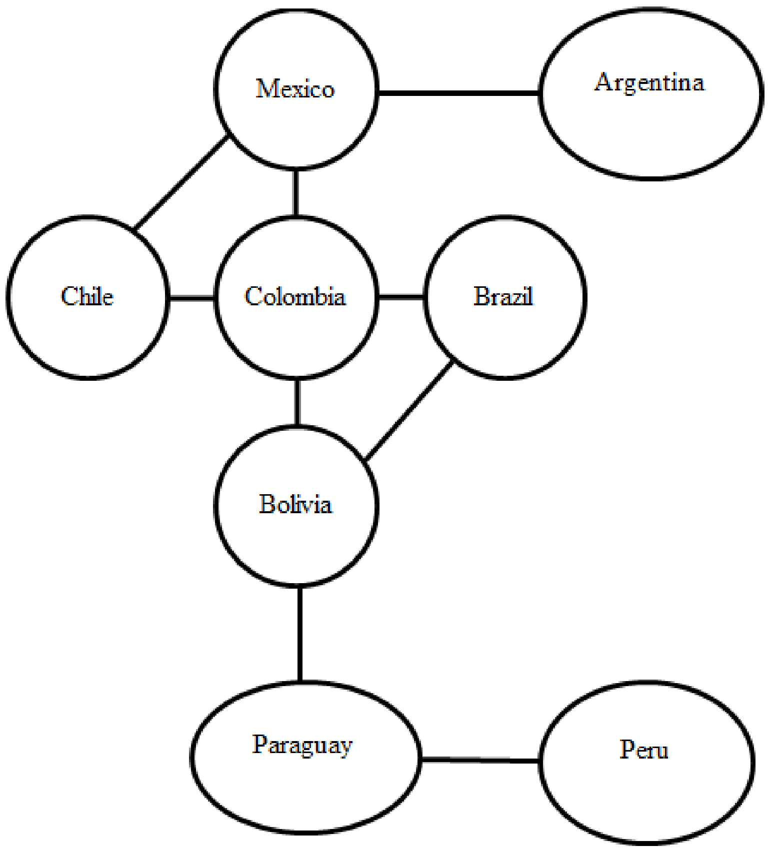

Figure 1: The "Extended Diamond".
Note: Line sections links two countries that are cointegrated.

Perhaps unsurprisingly, an immediate implication of our results is that a uniform monetary union for all Latin America seems implausible. The cointegration analysis sheds some light on the heterogeneous ways in which

Latin American countries have immersed in more open-oriented economic strategies since the mid-1990s.[16] The low number of cointegration relationships found (around 10%) highlights the still rather low level of economic integration the region has achieved. One particularly interesting finding is the lack of integration between the three Central American countries in the sample: Belize, Nicaragua and Costa Rica. Furthermore, there is no evidence of Central American integration with Mexico and/or Colombia. This signals possible complications that the "Mesoamerican" region will encounter in accomplishing the objectives of trying to promote complementarities and cooperation among the member countries.[17] Also, despite the long standing agreements of CARICOM and CACM, the DR-CAFTA, and the work of the Central American Monetary Council,[18] more convergence is necessary before a monetary union in Central America and the Caribbean becomes a feasible possibility.

Another surprising finding is the weak links among the Mercosur countries. Evidence of integration, if any, is indirect, and Uruguay and Venezuela do not even feature. One possible explanation is that these countries are of the "open-shy" type, meaning that global competitive attitudes and synchronization are erratic and short-lived. Therefore, evidence of possible integration would be more difficult to find in a long-run analysis like cointegration.

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[16] Appearances and some formal evidence around the world support this fact (see, e.g. IADB, 2002). Cases go from open-economies examples like Mexico or Chile; to "open-shy" countries like Brazil; ending with re-close tendencies countries like Venezuela or Argentina.

[17] The Mesoamerican Project (or Puebla-Panama plan) intends to remedy the lack of investment and trade in the southern states of Mexico, all the countries of Central America, Dominican Republic, and Colombia. See http://www.proyectomesoamerica.org/ for further details.

[18] The Central American Monetary Council is integrated by the Presidents of the central banks of the same countries of the DR-CAFTA (except the U.S.). The Council defined in 2007 as its strategic core objectives to encourage coordination and convergence in monetary, exchange and financial policies.

Overall, we did not find evidence of strong long-run ties based on regional trade blocs or underlying geographical bias. This suggests that the existence of trade agreements and improved political and cultural relations are not sufficient conditions to recommend forming a monetary union. Indeed, if one of the objectives of regional trade agreements is to enhance greater regional integration, at least in terms of increasing the degree of symmetry to macroeconomic shocks, it is still far from being accomplished. For the small economies in the sample the message is clear: serious and consistent steps to achieve further regional integration should be taken to enhance their economic role.

While the percentage of cointegration cases found is low, the relative size of the hypothetical monetary union is not. Table 2 reports GDP (PPP-adjusted) and population data for each country. Panel A only includes the potential members of the "main diamond", while Panel B includes all the countries of the "extended diamond". If the five countries of the main diamond where to form a monetary union, this bloc would account for around 65% of the total Latin American population, and would be responsible for nearly 70% of its total GDP. For the case where Argentina (the third biggest economy in the region), Paraguay, and Peru, are also members of the hypothetical monetary union, the extended bloc would account for nearly 85% of total Latin American GDP and 78% of its population. In short, size appears to matter a lot in Latin America in its ability to achieve symmetry in macroeconomic shocks. For example, the main diamond includes the two giants, Brazil and Mexico, the fourth largest economy in the region, Colombia, and Chile, the richest of the five based on GDP per capita figures. [19]

[19] The importance of including the biggest regional economies in a Latin American monetary union is similar to the conclusions of Sato, et al. (2009) who found that the ASEAN economies were not a feasible group to form a monetary union without the inclusion of Japan.

TABLE 2. RELATIVE SIZE OF THE POTENTIAL MONETARY UNION.

Country	GDP (Billions)	Percentage of Population	GDP per capita
PANEL A			
Bolivia	45.57	1.76	4,603.53
Brazil	2,178.53	32.81	11,272.96
Chile	258.54	2.82	15,039.90
Colombia	413.87	7.9	9,592.91
Mexico	1,564.87	19.65	14,405.93
%Latin American GDP	69.67%		
% Latin American population		64.94%	
PANEL B			
Bolivia	45.57	1.76	4,603.53
Brazil	2,178.53	32.81	11,272.96
Chile	258.54	2.82	15,039.90
Colombia	413.87	7.9	9,592.91
Mexico	1,564.87	19.65	14,405.93
Argentina	644.30	6.93	15,901.24
Paraguay	33.34	1.15	5,207.70
Peru	276.54	5.11	9,357.95
%Latin American GDP	84.57%		
% Latin American population		78.13%	

Sources: World Economic Outlook of the International Monetary Fund for the GDP and GDP per capita data. The U.S. Census Bureau's International Database for population data.
Notes: GDP is for 2010 data, except for Bolivia and Colombia which is 2009 data. GDP per capita is for 2010 data; Argentina, Bolivia, Chile, Colombia, Paraguay and Peru are IMF staff estimates.

3.4 Granger causality results

Table 3 reports the Granger causality results for all 15 countries in the sample. The number of unidirectional Granger causality cases is 50 out of 182 (around 28%). This signals a higher number of relationships in the short-run than in the long-run. However, the relationships are not clearly concentrated within a particular group of countries. For the Central American countries we find evidence of stronger ties compared to the cointegration results. Three out of six unidirectional causality relationships are found. This may signal that at least in the short-run, similar macroeconomic shocks have induced similar stabilization policies. However, still no links are uncovered between these countries and Mexico or Colombia. Thus, the general Mesoamerican region appears weakly integrated.

The number of causality relationships present for the majority of the small economies in the sample is puzzling when compared with the weak evidence of similarities exposed using the cointegration analysis. This suggests that the former may not be evidence of regional integration but rather of economic dependence in a few varieties of goods and services. The price change in a good or service in which a country heavily depends can trigger strong but short-lived monetary policy actions in the region to maintain competitiveness. A deeper analysis is needed to understand if the number of Granger causality results is evidence of monetary policy synchronization opportunities, constant beggar-thy-neighbor policies, or weak and unreliable data. [20, 21]

[20] As mentioned earlier, Latin American countries share similar patterns of trade and trade composition, but the specific raw materials in which they are well-endowed differ and in some case are substitutes. A high dependence in these goods can be an incentive for beggar-thy-neighbor policies.

[21] The apparent evidence of short-run interactions suggests that future tests should be carried to see if they become long-run relationships (cointegration, vector error-correction models, etc., using different variables that may reflect symmetry in real shocks from the recent integration process of the region onwards). Unfortunately, a lot of the data which can be used (like GDP or inflation) is published quarterly for a large number of the sample countries, and this makes it difficult to uncover evidence of the recent regional integration process. Clearly, uncovering synchronous common business cycles going from the short-run to the long-run is beyond the scope of the current paper.

Table 3. Granger Causality analysis for Latin American countries.

	MEX	BEL	NIC	CR	COL	VEN	ECU	BOL	BRA	CHI	PAR	URU	PER	ARG
MEX		0.547	0.635	0.061	0.574	**0.025***	0.296	0.164	0.146	0.365	**.016***	**.020***	0.301	0.875
BEL	0.783		**0.015***	**0.042***	0.585	**0.000****	0.068	0.135	0.814	0.961	**.003****	**.067***	0.175	0.718
NIC	0.918	0.656		0.773	0.226	0.168	**0.000****	0.006#	**.010****	0.243	0.174	0.572	0.172	0.228
CR	0.132	**0.001****	0.462		0.395	0.228	**0.049***	**0.025***	0.695	0.148	0.491	0.312	**.046***	0.527
COL	0.387	0.358	0.071	0.495		0.586#	**0.001****	0.07	0.169	**.004****	0.797	0.147	0.137	0.999
VEN	0.092	0.083	0.755	**0.026***	.066#		1	0.981	0.819	0.895	**.000****	0.854	0.479	1
ECU	0.528	0.979	0.319	0.677	0.491	0.945		0.660#	0.315	0.074	0.878	0.329	0.471	0.834
BOL	0.679	**0.011***	0.001#	**0.030***	**.002****	**0.030***	0.605#		0.782	0.177	**0.038***	**.007****	0.133	0.096
BRA	0.153	0.152	0.517	0.578	**.000****	0.555	**0.005****	0.374		**.012***	0.17	**.000****	0.561	**.000****
CHI	**0.048***	0.348	0.152	0.311	0.13	**0.025***	0.949	0.767	0.276		**.000****	0.51	0.525	0.153
PAR	0.679	0.373	**0.016***	**0.010****	**0.001****	0.233	**0.001****	**.003****	0.123	0.797		0.256	0.921	0.153
URU	0.687	0.675	0.734	**0.008****	**0.000****	**.000****	0.108	**.030***	0.553	**.017***	0.333		0.099	0.27
PER	0.428	0.279	**0.016***	**.000****	0.088	0.714	0.077	0.088	0.282	0.457	0.053	0.591		0.842
ARG	0.982	0.96	0.892	0.713	**0.000****	**.000****	0.291	**.002****	**.000****	0.383	**.012***	**.000****	0.77	

Notes: Each number is the p-value for testing the null hypothesis of no Granger causality. The direction of causality runs from the row REER to the column REER. For example, looking at Mexico (MEX) in the row and at Costa Rica (CR) in the column, we conclude that "Mexico's REER Granger causes Costa Rica's REER because the p-value is .042, hence we reject the null hypothesis at the 5% significance level. *, **, #; denotes Granger causality at the 5%, Granger causality at the 10%, and either instability or serial autocorrelation, respectively. Abbreviations: MEX=Mexico, BEL=Belize, NIC=Nicaragua, CR=Costa Rica, COL=Colombia, VEN=Venezuela, ECU=Ecuador, BOL=Bolivia, BRA=Brazil, CHI=Chile, PAR=Paraguay, URU=Uruguay, PER=Peru, ARG=Argentina.

Table 4 summarizes the Granger causality p-values for the countries in the “extended diamond”. Mexico only presents two cases of Granger causality, but it’s causality with Chile reinforces the relationship between these two relatively open-oriented countries. On the other hand, Colombia strengthens its position as the core for the union of cointegrated countries, where it is Granger caused in four occasions. This signals Colombia as a country that adjusts its competitiveness (purposely or not) after observed changes in other countries’ REERs.

Table 4. Granger causality results for cointegrated countries.

	MX	ARG	CHI	COL	BRA	BOL	PAR	PER
MX		0.875	0.365	0.574	0.146	0.164	**0.016***	0.301
ARG	0.982		0.383	**0.000****	**0.000****	**0.002****	**0.012***	0.77
CHI	**0.048***	0.153		0.13	0.276	0.767	**0.000****	0.525
COL	0.387	0.999	**0.004****		0.169	0.07	0.797	0.137
BRA	0.153	**0.000****	**0.012***	**0.000****		0.374	0.17	0.561
BOL	0.679	0.096	0.177	**0.002****	0.782		**0.038***	0.133
PAR	0.679	0.153	0.797	**0.001****	0.123	**.003****		0.921
PER	0.428	0.842	0.457	0.088	0.282	0.088	0.053	

Notes: Each number is the p-value for testing the null hypothesis of no Granger causality for the countries in Figure 1. *, **, #; denotes Granger causality at the 5%, Granger causality at the 10%, and either instability or serial autocorrelation, respectively. Abbreviations: MEX=Mexico, BEL=Belize, NIC=Nicaragua, CR=Costa Rica, COL=Colombia, VEN=Venezuela, ECU=Ecuador, BOL=Bolivia, BRA=Brazil, CHI=Chile, PAR=Paraguay, URU=Uruguay, PER=Peru, ARG=Argentina.

One interesting feature is the now important impact of Argentina on other South American countries, particularly Brazil, which also has a number of causality relationships. The Argentina-Brazil case is one of the few cases of bidirectional causality in the whole sample. This suggests that the two biggest economies of the South America region generate important short-term impacts in the synchronization reactions of other South American countries.

Finally, we consider the situation of the two other countries outside of the "main diamond". On one side Paraguay shows a total of six Granger causality relationships, while Peru shows none. This reveals that Paraguay may truly be considered a member of a hypothetical OCA, while Peru does not appear to present sufficient integration to be included.

3.5 The dollar or the yuan?

In this subsection we analyze the possible anchoring of the new currency of our hypothetical monetary union to the U.S. dollar or the Chinese yuan. In essence, the adoption of the new currency involves irrevocably fixing the exchange rate between all the members of the currency union. One possibility, is anchoring all member country currencies to a foreign currency, at least initially. There could be a wide-range of reasons for this; from sharing a lot of trade with the foreign country, to wanting to import low and stable inflation. Whatever the reason, sharing strong long-run and short-run trends signal a potential gain in anchoring the new currency (either completely or giving it a stronger weight in a basket of other important currencies) to a solid foreign currency.

The importance of the U.S. for the Latin American region has induced researchers to analyze possible costs and benefits of strict dollarization when considering monetary regime options for Latin American countries (e.g. Berg, et al. 2002). During the 1980s and 1990s, the vast majority of Latin American countries fixed their currencies against the dollar. Nowadays, even if most countries use more flexible exchange rate regimes, they keep most of their foreign exchange reserves in dollars and there is informal evidence of foreign exchange intervention when the terms of trade become too volatile (Mohanty and Turner, 2006). On the other hand, Latin America's trade relations with

China have steadily increased over the last decade: in 2000 exports to China accounted for only 1.1% of total exports, whereas by 2010 they now accounted for 8.3%. The United Nations Economic Commission for Latin America and the Caribbean (ECLAC) is particularly optimist on the trade opportunities between the region and the Asian giant, forecasted total exports and imports in 2020 to be 19.3% and 16.2%, respectively (ECLAC, 2012). The size and reach of the China and the U.S. make them potential candidates to serve as an anchor for the new currency of the hypothetical Latin American monetary union.

Table 5 reports the results from the cointegration and Granger causality tests of the bivariate models for each Latin American country against the U.S. and China. Regarding the cointegration analysis, the U.S. shows evidence of sharing long-run co-movements with three out of fourteen countries: Mexico, Colombia and Argentina. On the other hand, China shows cointegration evidence with just Costa Rica.

In the Granger causality analysis, the U.S. causes competitive adjustments in three countries: Nicaragua, Venezuela and Paraguay, and China also exhibit's three cases of Granger causality for Belize, Nicaragua and Bolivia. Overall, a strong and consistent influence of both the U.S. and China over the Latin American region is not favored using REER data. Nevertheless, the U.S. still appears to be a considerable stronger force despite the increasing influence of China. However, China already appears to have an important role in Central America, where one relationship was found for each of the three Central American countries in the sample.

Table 5. Impact of the U.S. and China in Latin America.

Latin American country	Foreign country	λ-trace	Cointegration ?	Causality test's p-value	Granger causality ?
MEXICO	USA	**16.75***	**YES**	0.852	NO
	CHINA	9.94	NO	0.746	NO
BELIZE	USA	9.58	NO	0.107	NO
	CHINA	6.7	NO	**0.002****	**YES**
NICARAGUA	USA	14.84	NO	**0.03***	**YES**
	CHINA	4.22	NO	**0.011***	**YES**
COSTA RICA	USA	9.43	NO	0.719	NO
	CHINA	**16.14***	**YES**	.359#	NO
COLOMBIA	USA	**16.56***	**YES**	0.657	NO
	CHINA	13.83	NO	0.73	NO
VENEZUELA	USA	11.57	NO	**0.017***	**YES**
	CHINA	11.93	NO	0.154	NO
ECUADOR	USA	10.74	NO	0.088	NO
	CHINA	9.93	NO	0.15	NO
BOLIVIA	USA	9.01	NO	0.748	NO
	CHINA	5.43	NO	**.008***	**YES**
BRAZIL	USA	10.49	NO	0.688	NO
	CHINA	13.75	NO	0.147	NO
CHILE	USA	13.68	NO	0.858	NO
	CHINA	11.63	NO	0.809	NO
PARAGUAY	USA	8.99	NO	**0.002****	**YES**
	CHINA	13.83	NO	0.446	NO
URUGUAY	USA	14.78	NO	0.166	NO
	CHINA	8.2	NO	0.183	NO
PERU	USA	11.22	NO	0.494	NO
	CHINA	10.31	NO	0.058	NO
ARGENTINA	USA	**19.14***	**YES**	0.355	NO
	CHINA	2.83	NO	0.327	NO

Notes: Each Latin American country (first column) is analyzed for evidence of cointegration and Granger causality with both USA and China (second column). The third column reports the λ-trace statistic estimated for testing cointegration and the fourth column concludes. Critical values (Drift case considered) are: 15.41 at 5% significance level; 20.04 at 1% significance level. The fifth column reports the p-value for testing the null hypothesis of no Granger causality from the foreign country to the Latin American country and the sixth column concludes. *, **, #; denotes significance at the 5% level, significance at the 1% level, and instability and/or autocorrelation, respectively.

Finally, Table 6 reports the results of cointegration and Granger causality relationships just for the hypothetical monetary union supported by the cointegration analysis. There is no evidence that the dollar or yuan are influential enough to either form a monetary bloc in Latin America, or act as an anchor currency.

Table 6. Impact of the USA and China on the hypothetical Latin American Monetary Union's members.

	MEX	ARG	CHI	COL	BRA	BOL	PAR	PER
USA (cointegration)	**YES**	**YES**	NO	**YES**	NO	NO	NO	NO
CHINA (cointegration)	NO	NO	NO	NO	NO	NO	NO	NO
USA (causality)	NO	NO	NO	NO	NO	NO	**YES**	NO
CHINA (causality)	NO	NO	NO	NO	NO	**YES**	NO	NO

Notes: The table reports the conclusion of cointegration (rows 3 and 4) and Granger causality (rows 5 and 6) of the USA and China with each of the Latin American countries for which we got evidence of cointegration. Acronyms: MEX=Mexico, ARG=Argentina, CHI=Chile, COL=Colombia, BRA=Brazil, BOL=Bolivia, PAR=Paraguay, PER=Peru.

To end this section, we summarize our key findings from the cointegration and Granger causality analysis. A monetary union comprising all the Latin American countries is not recommended: our analysis suggests a high degree of asymmetry in responses to real macroeconomic shocks. Nevertheless, based on the cointegration analysis, we find strong long-run ties between Colombia, Mexico, Chile, Brazil, and Bolivia; as well as weaker ties between these five countries and Argentina, Paraguay and Peru. Granger causality analysis uncovered important short-run interactions, perhaps because of enhanced economic dependence not accompanied with regional integration. Further analysis is needed to determine if these short-run interactions are causing closer regional economic integration. Overall, both exercises find strong symmetries between the eight countries in the "extended diamond": the possible exception being Peru who does not yet exhibit significant short-run interactions with another country in the hypothetical monetary union. Even if the quantity of countries is small, the relative size of the hypothetical monetary union would be large, comprising 85% of Latin America's GDP and 78% of its population. Size appears to matter in Latin America in finding evidence of long-run symmetries between countries. Finally, we employed the cointegration and Granger causality tests to see if there was evidence of a U.S. or China currency bloc forming in the region. While the U.S. is still the bigger force in the region, it is not influential enough to either consider dollarization of the Latin American monetary union, or the anchoring of any new currency against the dollar.

SECTION 4. INTEREST RATE PASS-THROUGH ANALYSIS

4.1 Data description and monetary policy pass-through

The main purpose of this section is to look at the issue of financial or monetary convergence. The previous section looked at issues relating to whether real shocks had a symmetrical impact across Latin American countries, and found evidence that a monetary union comprised of Colombia, Chile, Mexico, Brazil, Bolivia, Argentina and Paraguay was feasible. In this section, we check the robustness of this hypothetical monetary union by investigating the degree of financial convergence between these seven countries. In particular, we focus on the degree of power the monetary authority has in affecting the private financial sector. To this end, our analysis is conducted using interest rate pass-through theory.

By definition, there is complete pass-through if the central bank can transfer all the cost associated with an increase in its policy rate onto the retail rates. If the central bank can only transfer part of the cost, then there is incomplete pass-through; while if the central bank transfers more than the cost then it is considered to be over pass-through. Most empirical studies (see, e.g. Wang & Lee, 2009; Haughton & Iglesias, 2012) find evidence of consistent rigidities and heterogeneity in the pass-through levels in various countries. A high level of pass-through reflects a strong effectiveness of monetary policy, because the private retail rates truly influence the market demand and supply of loans and deposits; and consequently, inflation, investment and production.

In addition to the level of pass-through, the strength of a central bank's monetary policy in affecting financial markets is determined by evidence of a long-run relationship between the central bank's policy rate and the retail interest rates. Therefore, we are interested in two things: the level (size) of the transmission, and the possible speed of the adjustment process that signals important links between the central bank and the private sector. In short, we can characterize the pass-through by the impact level of the central bank (discount rate) on the private sector (retail rate), and by uncovering the presence of long-run ties between the interest rates controlled by the different agents.

We use monthly interest rate data from January 1999 – November 2012 based on the data availability of Mexico, Chile, Colombia, Brazil, Bolivia, Argentina, Paraguay and Peru. All series were obtained from the *International Financial Statistics* (IFS) database of the *International Monetary Fund* (IMF). Two of the countries present serious problems of data availability. The discount rates of Argentina and Paraguay could not be accurately obtained for the whole sample period and where consequently omitted. Our initial incline was to eliminate Peru from the analysis as well given the lack of robustness of evidence of integration with the other countries.[22] But after the omission of Paraguay we decided to include it, given that Peru exhibits some degree of integration with Paraguay. In summary, the interest rate pass-through analysis is conducted for the following six countries: Mexico, Chile, Colombia, Brazil, Bolivia and Peru.

[22] As previously discussed its only cointegration relationship was with Paraguay, and no other Granger causality relationship was found with another member of the theoretical monetary union.

The discount rate represents the central bank policy rate; the series for Bolivia, Chile and Colombia are complete; while the series of Brazil and Peru miss only three and one observation(s), respectively. For Mexico, the series begins only from 2008, when Mexico's central bank stated the official objective discount rate. Consequently, we used the Treasury-Bill as a proxy from the beginning of the period to 2007.[23] For the retail rates the lending and deposit interest rates are used. The series for both the lending rate and the deposit rate are complete for all countries except Peru which misses a single observation.

4.2 Econometric methodology

The econometric techniques employed follows Enders (2010) and Bholla, et al. (2011). The standard methods needed for the empirical pass-through analysis are Autoregressive Distributed Lag (ADL) and Error-Correction models. In these cases we explicitly specify a dependent variable and a joint set of independent variables, using economic theory. In particular, we exploit the links of weak exogeneity, cointegration, error-correction models, and ADL models, to disentangle valid numerical information from the data.

All the interest rate variables where tested for a unit root using the DFGLS and Phillips-Perron tests. The results suggest that all the interest rates were I(1) processes.[24] Then we tested for cointegration between the discount rate (DR) and each of the private retail rates (PR1 for the lending rate and PR2 for the deposit rate). The important thing here is that we exploit economic theory and do not treat the variables symmetrically (as in the Johansen methodology).

Suppose that we have a cointegrated system. If a variable does not respond to the discrepancy from the long-run equilibrium relationship, then we

[23] Note that for Argentina and Paraguay, the Treasury-Bill is also not reported.
[24] As with the REER tests, the unit root and cointegration tests are available upon request.

can say that the variable is weakly exogenous. The argument for our particular case, is that the discount rate is weakly exogenous because it can be controlled (theoretically) by the central bank, and therefore does not experience any type of feedback from the private retail rates when disturbances need to be adjusted. In essence, our theory dictates a causal relationship of interest, but we want to know if this relationship involves a long-run combination (cointegration), or if the relationship is simpler and only involves dynamic short-term and long-term effects of one series on the other.

First, the following error-correction model is estimated:

$$\Delta PRi_t = \alpha_0 + \alpha_1 PRi_{t-1} + \alpha_2 DR_{t-1} + \alpha_3 \Delta DR_t + \sum_{m=1}^{p} \alpha_{PRi,m} \Delta PRi_{t-m} + \sum_{k=1}^{q} \alpha_{DR,k} \Delta PRi_{t-k} + v_t \qquad (8)$$

where i = 1, 2 (lending, deposit); Δ denotes a difference; *PR* denotes Private Rate; *DR* denotes Discount Rate; and v_t is the error term. Since we are not treating all variables symmetrically, the lag length (p) for the lag values of the private rate (dependent variable) is not needed to be the same lag length (q) as the lag values of the discount rate (independent and weak exogenous variable). We estimate the most appropriate model using the Akaike Information Criterion (AIC) which generally favors a greater number of lags (Stock & Watson, 2002), but is preferred over a greater probability of having correlation issues with the error term.[25]

[25] Only when the number of lags is “sufficient”, can we expect to avoid endogeniety problems, as we have included enough past information in the error-correction or ADL model.

Equation (8) is the error-correction model in ADL form when using the concept of weak exogeneity, which allows our model to be identified. This equation enables us to obtain estimates of the short-run pass-through and adjustment velocity, if and only if, these two variables are cointegrated. It can be shown (see, e.g. Enders, 2010), that a proper test for cointegration is to use the *t*-statistic for the null hypothesis $\alpha_1 = 0$ in (8). If we accept the null, then there is no error-correction term. Hence, the proper alternative hypothesis is $\alpha_1 < 0$, so that we can accept convergence, with the estimated value of α_1 being the velocity of adjustment . The appropriate critical values depend on the number of I(1) regressors in the model (denoted by k), the adjusted sample size N^{adj}, and if there is an intercept (*d*). Then, $N^{adj} = N - (2k - 1) - 1$; where N is the number of observations in the sample.[26]

If evidence of cointegration is found, then α_1 is, as already mentioned, the velocity of adjustment parameter, α_3 is the short-run pass-through (impact multiplier), and also it is valid to get the long-run pass-through from a simple OLS regression of the private retail rate on the discount rate. That is; when we found evidence of cointegration, the following model is estimated:

$$PRi_t = \theta_0 + \theta_1 DR_t + u_t \qquad (9)$$

where i = 1, 2 (lending, deposit); u_t is the error term which satisfies standard Gaussian properties; and θ_1 is the long-run pass-through (long-term multiplier or propensity).

In the case where there is no cointegration, a standard rational ADL model is employed:

[26] We use Statistical Table F in Enders (2010) for the hypothesis tests.

$$\Delta PRi_t = \alpha_0 + \alpha_3 \Delta DR_t + \sum_{m=1}^{p} \alpha_{PRi,m} \Delta PRi_{t-m} + \sum_{k=1}^{q} \alpha_{DR,k} \Delta PRi_{t-k} + v_t. \quad (10)$$

This simply removes the (separated) no-significant error-correction term in model (8). Since our differenced variables are stationary, we should not worry about additional information relating to long-run relationships, since we have rejected cointegration. The ADL specification avoids problems of spurious regression but a simple OLS model like (9) is not correct to identify the long-run pass-through of interest rates. From equation (10), α_3 is our short-run pass-through estimate and from the ADL model we obtain the long-run pass-through. First rewrite the ADL(p, q) model (10), using the lag operator *L* as:

$$\alpha_{PRi}(L) PRi_t = \alpha_0 + \alpha_{DR}(L) DR_t + v_t \quad (11)$$

$$\alpha_{PRi}(L) = 1 - \alpha_{PRi,1} L - \alpha_{PRi,2} L^2 - \cdots - \alpha_{PRi,p} L^p$$
$$\alpha_{DR}(L) = \alpha_3 + \alpha_{DR,1} L + \alpha_{DR,2} L^2 + \cdots + \alpha_{DR,q} L^q.$$

Since we have stationarity, the long-run solution can be found as:

$$PRi_t = \alpha^{-1}{}_{PRi}(L)\alpha_0 + \alpha^{-1}{}_{PRi}(L)\alpha_{DR}(L) DR_t + \alpha^{-1}{}_{PRi}(L) v_t.$$

The expected value is:

$$E[PRi_t] = \alpha^{-1}{}_{PRi}(L)\alpha_0 + \alpha^{-1}{}_{PRi}(L)\alpha_{DR}(L) E[DR_t].$$

When all the variables take their long-run values then:

$$E[PRi_t] = E[PRi_{t-1}] = LE[PRi_t] = E[PRi_{t-2}] = L^2E[PRi_t] = \cdots$$
$$= E\left[PRi_{t-p}\right] = L^pE[PRi_t]$$
$$E[DR_t] = E[DR_{t-1}] = LE[DR_t] = E[DR_{t-2}] = L^2E[DR_t] = \cdots = E\left[DR_{t-q}\right]$$
$$= L^qE[DR_t].$$

Therefore, the long-run pass-through, when interest rates are not cointegrated, is:

$$LRPT_{No\ Cointegration} = \alpha^{-1}{}_{PRi}(1)\alpha_{DR}(1)$$
$$= \frac{\alpha_3 + \alpha_{DR,1} + \alpha_{DR,2} + \cdots + \alpha_{DR,q}}{1 - \alpha_{PRi,1} - \alpha_{PRi,2} - \cdots - \alpha_{PRi,p}}. \qquad (12)$$

Rolling-window technique analysis

We carried out the analysis described above for the whole sample period (167 observations). Then we divided the entire sample into seven periods and redone the analysis for each of the subsamples. The first six periods have 84 observations each; we call them windows 1999-2005, 2000-2006, 2001-2007, 2002-2008, 2003-2010, and 2004-2011. The last period has 95 observations; called window 2005-2012. The purpose is to capture the dynamic development of the interest rate pass-through over time for both the lending and deposit rates. This analysis should shed some light on how effective is the control of each country central bank's monetary policy on the private financial sector, and if there is some evidence of a more homogenous situation between the candidate countries. In essence, we graphically examine two characteristics that make the formation of a monetary union more feasible: convergence in the magnitude of the pass-through among the candidate countries and a strong monetary transmission (high pass-through).

To end this subsection, note that we have included dummies for the 1998/99 crisis of Brazil, the 2001/02 crisis of Argentina, and the 2007/08 global financial crisis. As in the REER analysis of the previous section, we want to reduce the possibility of cointegration biases due to structural breaks. In this case we focus on internal financial characteristics for each of the countries, and these three dummies plausibly capture the main breaks for each of them.[27] We test joint significance of the dummies at the 10% significance level, and see which of them might be important to consider when estimating the best models.

4.3 Results

Table 7 summarizes the main results of this section for each country. Additionally, to ease the analysis of convergence and similarities for the whole period, we also present graphical plots of the short-run pass-through and the long-run pass-through for both the lending (Figure 2) and deposit (Figure 3) rates. An important caveat is worth mentioning for the case of the lending rate. We included neither Bolivia nor Chile because of the lack of significance of both the short-run and long-run pass-through for the last subsample periods, which makes it impossible to draw inference of possible convergence or similarities. In the rest of the cases, we did not find strong restrictions and only warn whenever necessary.

[27] In the previous section we argued that the Argentinian crisis of 2001/02 had concentrated real effects instead a systematic effect on the whole region. Therefore, it was only considered when testing cointegration between Argentina and another country. That argument is more difficult to sustain on nominal grounds, but testing significance for the dummy variables in each country's model allows us to consider important biases that might affect only that country.

Table 7. Interest rates Pass-Through main results.

Country	Interest Rate	Period	t-statistic for Cointegration Test	Cointegration?	Short-run Pass Through	Long-run Pass Through	Velocity of Adjustment
Mexico	Lending	1999-2005	-3.32*	YES	0.933	1.093	0.3163
		2000-2006	-3.30*	YES	0.952	1.092	0.2688
		2001-2007	-2.76	NO	0.887	0.994	
		2002-2008	-2.28	NO	1.024	1.070	
		2003-2009	-1.89	NO	1.027	1.124	
		2004-2010	-1.32	NO	1.032	1.251	
		2005-2012	-1.42	NO	0.918	1.295	
		1999-2012	-3.12	NO	0.933	1.040	
	Deposit	1999-2005	-1.79	NO	0.293	0.462	
		2000-2006	-2.29	NO	0.292	0.472	
		2001-2007	-2.81	NO	0.288	0.215	
		2002-2008	-1.08	NO	0.205	0.341	
		2003-2009	-1.01	NO	0.177	0.337	
		2004-2010	-1.10	NO	0.184	0.309	
		2005-2012	-1.67	NO	0.149	0.326	
		1999-2012	-2.79	NO	0.289	0.464	
Bolivia	Lending	1999-2005	-1.62	NO	0.198#	2.452#	
		2000-2006	-1.91	NO	0.076#	2.228#	
		2001-2007	-3.61*	YES	0.402#	1.261	0.5309
		2002-2008	-2.62	NO	0.406#	1.810#	
		2003-2009	-1.70	NO	0.242#	1.276	
		2004-2010	-1.64	NO	-0.106#	0.381#	
		2005-2012	-2.59	NO	0.247#	0.698#	

Table 7 Continued.

		1999-2012	-2.73	NO	-0.068#	0.717#	
	Deposit	1999-2005	-2.84	NO	0.027#	0.453#	
		2000-2006	-2.76	NO	0.014#	0.424	
		2001-2007	-3.16	NO	-0.098#	0.335	
		2002-2008	-3.12	NO	-0.071#	0.455	
		2003-2009	-1.23	NO	0.348	0.403	
		2004-2010	-3.88**	YES	0.310	0.514	0.1674
		2005-2012	-1.90	NO	0.328	0.359	
		1999-2012	-1.81	NO	0.288	0.462	
Colombia	Lending	1999-2005	-2.58	NO	0.408	0.490	
		2000-2006	-2.73	NO	0.366	0.466	
		2001-2007	-1.67	NO	0.109#	0.702	
		2002-2008	-0.43	NO	0.093#	0.457	
		2003-2009	-3.22	NO	0.248	0.804	
		2004-2010	-2.50	NO	0.330	0.975	
		2005-2012	-2.97	NO	0.384	0.992	
		1999-2012	-4.06**	YES	0.352	0.794	0.1836
	Deposit	1999-2005	-2.02	NO	0.218	0.205	
		2000-2006	-2.37	NO	0.095	0.216	
		2001-2007	-0.74	NO	0.064#	0.715#	
		2002-2008	-0.28	NO	0.011#	-0.000#	
		2003-2009	-3.53*	YES	0.142	0.533	0.1216
		2004-2010	-3.10	NO	0.143	0.701	
		2005-2012	-3.27*	YES	0.180	0.630	0.1538
		1999-2012	-4.70**	YES	0.206	0.699	0.1700
Peru	Lending	1999-2005	-1.84	NO	0.096#	0.750	
		2000-	-1.96	NO	0.098#	0.678	

2006

Table 7 Continued

		2001-2007	-1.65	NO	0.255#	0.543#	
		2002-2008	-2.88	NO	0.805	0.314	
		2003-2009	-2.04	NO	0.613	0.532	
		2004-2010	-0.50	NO	0.599	0.433	
		2005-2012	-1.44	NO	0.544	0.523	
		1999-2012	-1.43	NO	0.178	0.559	
	Deposit	1999-2005	-1.45	NO	-0.043#	0.336	
		2000-2006	-1.84	NO	-0.034#	0.351#	
		2001-2007	-1.56	NO	0.314	0.563	
		2002-2008	-3.71*	YES	0.175	0.390	0.1699
		2003-2009	-3.10	NO	0.159	0.511	
		2004-2010	-0.92	NO	0.170	0.526	
		2005-2012	-0.03	NO	0.159	0.509	
		1999-2012	-2.54	NO	0.002#	0.380	
Chile	Lending	1999-2005	-3.94**	YES	0.541	1.012	0.6152
		2000-2006	-3.15	NO	0.515	0.808	
		2001-2007	-3.29*	YES	0.380	1.037	0.3427
		2002-2008	-3.23	NO	-0.054#	1.713#	
		2003-2009	-1.88	NO	0.204#	1.660	
		2004-2010	-1.71	NO	0.098#	1.618#	
		2005-2012	-1.28	NO	0.171#	1.599#	
		1999-2012	-4.42**	YES	0.481	1.028	0.2602
	Deposit	1999-2005	0.01	NO	0.533	1.504#	
		2000-2006	-0.85	NO	0.532	0.637	
		2001-2007	-1.19	NO	0.459	0.798	
		2002-2008	-4.88**	YES	0.784	1.044	0.6529

Table 7 Continued

		2003-2009	-4.08**	YES	0.836	1.026	0.4831
		2004-2010	-4.22*	YES	0.877	1.020	0.6245
		2005-2012	-3.65*	YES	0.903	1.024	0.4347
		1999-2012	-3.00	NO	0.620	0.799	
Brazil	Lending	1999-2005	-2.13	NO	0.797	1.358	
		2000-2006	-2.55	NO	0.752	1.340	
		2001-2007	-2.70	NO	0.643	1.032	
		2002-2008	-2.60	NO	0.645	1.039	
		2003-2009	-2.39	NO	0.904	1.250	
		2004-2010	-2.89	NO	0.495	1.004	
		2005-2012	-2.56	NO	0.691	1.809	
		1999-2012	-3.40*	YES	0.837	1.824	0.1477
	Deposit	1999-2005	-2.71	NO	0.484	0.825	
		2000-2006	-2.01	NO	0.514	0.843	
		2001-2007	-2.04	NO	0.522	0.837	
		2002-2008	-3.26*	YES	0.650	0.970	0.4911
		2003-2009	-2.97	NO	0.448	0.808	
		2004-2010	-3.55*	YES	0.544	0.946	0.4151
		2005-2012	-4.01**	YES	0.724	0.923	0.4458
		1999-2012	-3.63*	YES	0.612	0.993	0.2644

Notes: Cointegration tests are based on the critical values in Table F of Enders (2010). The critical values for an adjusted sample size of 100 (Case of the rolling-windows) are -3.247, and -3.874, at the 5%, and 1% significance level, respectively. The critical values for an adjusted sample size of 200 (complete 1999-2012 period) are -3.231, and -3.834, at the 5%, and 1% significance level, respectively. Hence; *, **, indicates significance at the 5%, and 1% for the t-statistic reported in column 4. # Indicates that the pass-through (short-run or long-run) reported in columns 6 or 7 are not significant at least at the 10% level.

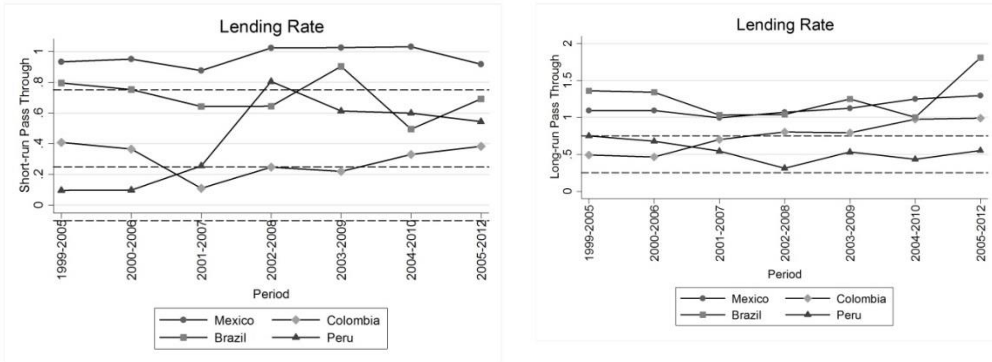

Figure 2. Pass-through of the Lending Rate. Left graph is the short-run pass-through and the right graph is the long-run pass-through.

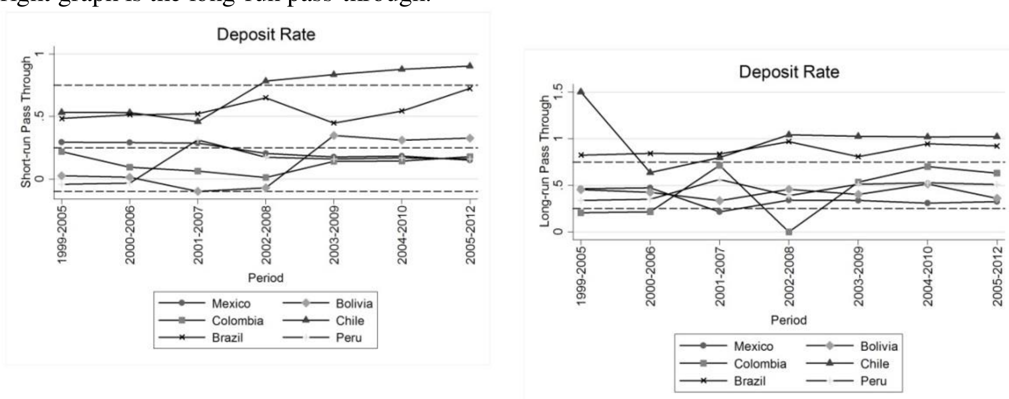

Figure 3. Pass-through of the Deposit Rate. Left graph is the short-run pass-through and the right graph is the long-run pass-through.

First consider the case of the short-run pass through; i.e., the immediate adjustment of the private rates to changes in the discount (objective) rate. For the lending rate (left graph of figure 2), strong asymmetries are found across the period. Mexico exhibits the strongest and most significant short-run pass-through. [28] Brazil shows considerable variance in its pass-through level from a minimum of almost 0.5 to a maximum of around 0.9. Peru initially has no significant levels of pass-through but this conclusion changes from the 2002-2008 rolling-window, exhibiting strong financial rigidities. Finally, Colombia shows systematically low levels of short-run pass-through onto the lending rate

[28] This does not diminish even when the sample has a greater weight for the official discount rate instead of the proxy Treasury-Bill rate (2008 onwards).

(2001-2007 and 2002-2008 are not significant), suggesting the presence of strong rigidities.

Overall, the evidence for the short-run pass-through on the lending rate indicates strong rigidities for all sample countries, except Mexico and, to a lesser degree, Brazil. Some degree of convergence could be argued only when considering the non-significant cases of Bolivia and Chile, and even in this case convergence is concentrated at low levels of short-run pass-through. Therefore, with the exception of Mexico, Latin American countries experience strong, but different levels, of rigidities. This suggests the presence of strong market imperfections and makes difficult the task of central banks to counteract shocks in a rapid and efficient lapse of time.

The insight of strong rigidities is supported by the short-run pass-through on the deposit rates results. The only country that has closely fulfilled complete pass-through for the deposit rate is Chile with a value of 0.9 for the rolling-window of 2005-2012. Brazil presents similar pass-through levels as its lending rate case. Bolivia only has significant results for the last three rolling-windows, at rather low levels, which may signal that until recently, the government had no real impact on the private sector, maybe because of a lack of trust from investors. The three countries that exhibit some degree of convergence for the final subsample periods are Mexico, Colombia and Peru, but at very low levels of below 0.2 for short-run pass-through. One possible explanation for the different degrees of short-run pass-through is how strongly financial institutions can use their market power. For example, the Mexican central bank may have a stronger interest in controlling the lending rate and therefore uses greater regulation strategies to influence that interest rate. Consequently, banks and other financial institutions have incentives to maximize their spread by offering a very low and unresponsive deposit rates to savers. An opposite situation might be occurring in Chile.

Now we turn to analyze the situation of the long-run pass-through for the members of the hypothetical monetary union. For the lending rate case, we see higher values relative to the short-run case**,** except Peru. These higher long-run pass-through values imply that Latin American countries now have a stronger impact on the financial sector, but the strong rigidities suggest that their impact occurs with considerable lag. Mexico and Brazil present puzzling high values of long-run pass through (even if we consider the period 2005-2012 of Brazil as an outlier). One possibility for the last subsample periods is that domestic banks increased lending rates because of perceived information asymmetries triggered by the financial bubbles during the recent crisis. It is also possible that because of the very low interest rates in the U.S. and Europe, banks may have received significant foreign resources which they are rationing the supply of credit because of credit risk factors. Further research is clearly necessary to get a more plausible answer.

The results for the long-run pass-through of the deposit rate, is given in the right panel of Figure 3. Once again, no evidence of convergence is found among the countries. Chile and Brazil exhibit near perfect long-run pass-through, while the other countries present strong, but different, rigidities. Again, Latin American countries seem to have a better control over the financial sector in the long-run but there are consistently different pass-through levels across the subsamples and no final convergence can be assured, unless, arguably, at very low levels (i.e. the situation of Mexico and Bolivia at around 0.35).

To complement the analysis of the long-run pass-through situation between the countries, we analyzed the results concerning the evidence of cointegration and the speed of adjustment. Table 8 reports the results for the cases where cointegration was found and adds the calculated months it takes for the long-run pass-through to be fully realized.[29]

Table 8. Further Analysis of Long-run Pass-Through Results.

Country	Interest Rate	Period	Long-run Pass Through	Velocity of Adjustment	Months to full Pass Through realization.
Mexico	Lending	1999-2005	1.093	0.3163	12
		2000-2006	1.092	0.2688	14
Bolivia	Lending	2001-2007	1.261	0.5309	7
	Deposit	2004-2010	0.514	0.1674	16
Colombia	Lending	1999-2012	0.794	0.1836	17
	Deposit	2003-2009	0.533	0.1216	20
		2005-2012	0.630	0.1538	18
		1999-2012	0.699	0.1700	18
Peru	Deposit	2002-2008	0.390	0.1699	14
Chile	Lending	1999-2005	1.012	0.6152	6
		2001-2007	1.037	0.3427	11
		1999-2012	1.028	0.2602	14
	Deposit	2002-2008	1.044	0.6529	5
		2003-2009	1.026	0.4831	7
		2004-2010	1.020	0.6245	5
		2005-2012	1.024	0.4347	8
Brazil	Lending	1999-2012	1.824	0.1477	25
	Deposit	2002-2008	0.970	0.4911	7
		2004-2010	0.946	0.4151	9
		2005-2012	0.923	0.4458	8
		1999-2012	0.993	0.2644	13

Notes: The sixth column reports the calculated months it takes to fully realize the long-run pass-through of an interest rate in case of cointegration. Full realization is considered when no more than 0.00 of the pass through is yet to be realized.

The total number of cointegration relationships found is 21 out of the 96 total periods analyzed (12 complete samples and 84 rolling-window subsamples). This means that for 22% of cases, a strong long-run equilibrium

[29] The error-correction term in our model with evidence of cointegration is: $-\alpha_1(PRi_{t-1} - \theta_1 DR_t)$. Hence, every unit impact of the discount rate causes a long-run pass through of θ_1. This disruption of the long-run equilibrium relationship between the interest rates is spread over future time periods at a rate of the velocity of adjustment, α_1, per time period.

relationship that involves an adjustment mechanism process is present. The relatively recent gained credibility in most Latin American central banks, the market power in the financial sector with dominance of foreign control, and the strong dependence on external flows could still undermine domestic financial control in Latin America at least in a consistent and long-horizon period of time.

Of the 21 cointegration relationships, 8 correspond to the lending rate and 13 to the deposit rate. The average number of months for the long-run pass-through to be completely realized is 13.25 for the lending rate and 11.38 for the deposit rate. The evidence of strong rigidities is increased when Chile is omitted, in which case the months to full realization increase to 15 and 13.67 for the lending and deposit rates, respectively. The adjustment lags indicate that even if the Latin American countries present a higher degree of impact in the long-run than in the short-run, as stated previously, they still face strong resistance from the private sector to react and may need to resort to an extra number of monetary tools to fulfill their monetary goals.

Overall, there exist considerable dissimilarities in the pass-through levels, both in the short-run and the long-run, a lack of convergence in high values of pass-through, a relatively low number of cointegration relationships, and adjustment lags, between our candidate countries to form a monetary union. We know that these countries show some evidence of symmetric reactions to external shocks and share evidence of affecting each other in the international market. However, strong steps are required to achieve financial convergence. In the case of forming a monetary union, a common central bank would likely affect the member countries in very different ways. In addition, it would have to fight a variety of rigidities in different countries and financial markets. In essence, any attempt to create a common central bank would first require far-reaching financial reforms to combat current rigidities and achieve nominal convergence. Since, any monetary union should include the regions two biggest economies, Brazil and Mexico we should expect these two countries to be at the forefront of financial reforms.

SECTION 5. CONCLUSIONS

This paper has investigated the possibility of creating a currency union for the Latin American region. We first examined the degree of symmetry of macroeconomic shocks using real effective exchange rate data for of fifteen Latin American countries. Following the analysis of Sun and Simons (2011), we applied cointegration and Granger causality techniques to analyze the long-run and short-run interactions. We found strong long-run ties between Colombia, Mexico, Chile, Brazil, and Bolivia; as well as weaker ties between these five countries and Argentina, Paraguay and Peru. Such a monetary union would comprise 85% of Latin America's GDP and 78% of its total population. Additionally, it was shown the neither the dollar nor the yuan are suitable anchor currencies for the hypothetical monetary union.

We also examined the degree of financial or monetary convergence within the hypothetical monetary union. We used monthly interest rate data of the discount, lending and deposit rates, and followed Bholla, et al. (2011), in applying error-correction models based on the theory of interest rate pass-through. Considerable dissimilarities in the pass-through levels where found, both in the short-run and the long-run. The lack of convergence in high values of pass-through, the relatively low number of cointegration relationships, and adjustment lags, between our candidate countries suggest that strong steps are needed to achieve financial convergence, before a feasible Latin American monetary union can take place.

Our analysis offers some interesting policy implications. A monetary union may be one possible option to protect against macroeconomic shocks and enhance economic integration within Latin America. Nevertheless, the two regional giants, Mexico and Brazil, would be needed to drive the process of further economic integration, both on real and nominal grounds. In this case, greater political and economic relationships between these two countries would be welcomed.

Two final caveats are worth mentioning. First, a deeper analysis is needed to understand if short-run interactions between Latin American countries are actual evidence of closer economic integration, rather than economic dependence. Second, for a feasible monetary union to take place will require significant political and public support. Future research should be directed towards quantifying the economic costs and benefits for each country of participating in a Latin American monetary union.

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APPENDIX

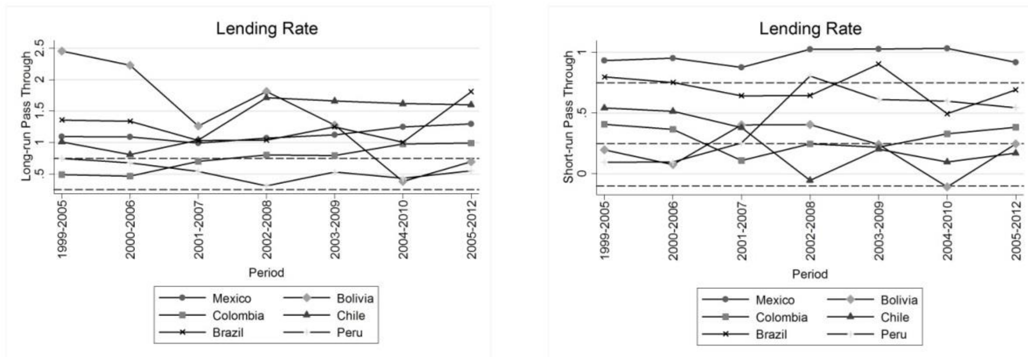

Figure A. Pass-through of the Lending Rate. Left graph is the short-run pass-through and the right graph is the long-run pass-through. Results for Chile and Bolivia are not-significant for most of the rolling-windows. See Table 7.