IS A MONETARY UNION FEASIBLE FOR LATIN AMERICA? EVIDENCE FROM REAL EFFECTIVE EXCHANGE RATES AND INTEREST RATE PASS-THROUGH LEVELS*

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Resumen: Se evalúa la viabilidad de formar una unión monetaria en América Latina. En primer lugar, analizamos los tipos de cambio reales efectivos mediante cointegración y causalidad de Granger, encontramos evidencia que respalda una unión monetaria compuesta por Argentina, Bolivia, Brasil, Chile, Colombia, México y Paraguay. En segundo lugar, se investiga el grado de heterogeneidad en la transmisión de la política monetaria dentro de la unión monetaria hipotética. Se encuentran considerables asimetrías en los niveles de *pass-through* de las tasas de interés lo que indica la necesidad de reformas sustanciales antes de considerar una unión monetaria en América Latina.

Abstract: This paper assesses the feasibility of forming a common currency in Latin America. First, we examine the cointegration and Granger causality of real effective exchange rates and find evidence supporting a monetary union comprised of Argentina, Bolivia, Brazil, Chile, Colombia, Mexico, and Paraguay. Second, we examine the degree of heterogeneity in the transmission of monetary policy within the hypothetical monetary union. Considerable asymmetries in the pass-through levels of interest rates are found to exist indicating the need for substantial reforms before a Latin American monetary union could take place.

Clasificación JEL/JEL Classification: E43, E52, F15, F33, F36

Palabras clave/keywords: unión monetaria, cointegración, causalidad de granger, pass-through de tasa de interés, política monetaria, monetary union, cointegration, granger causality, interest rate pass-through, monetary policy.

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1. Introduction

Since the launch of the European single currency, the feasibility of forming other currency unions has been widely investigated. While debates continue about the formation of monetary unions in East Africa (*e.g.* Bholla, Aziakpono and Snowball, 2011), East Asia (*e.g.* Sun and Simons, 2011), and between nations of the Gulf Cooperation Council (*e.g.* Louis, Balli and Osman, 2012), the existing literature has typically argued against a common currency in Latin America due to a lack of evidence that the region satisfies the basic criteria of an optimum currency area.[1] The purpose of this paper is to reexamine the feasibility of creating a monetary union in Latin America with the aim of rekindling discussion among academics and policymakers on this important issue.

The largest cost associated with entering into a currency union relates to the loss of monetary independence and nominal exchange-rate flexibility. Consequently, the literature on optimum currency areas has stressed the importance of *shock symmetry* among participants, as a crucial factor for the success of any monetary union (see, *e.g.*, the seminal papers of Mundell, 1961 and McKinnon, 1963). As shocks between countries become more symmetric, a common monetary policy response becomes more appropriate. One important strand of the existing literature has attempted to evaluate the degree of shock symmetry among Latin American countries. Using output and price data, and employing the Blanchard and Quah (1989) decomposition method to uncover the degree of demand and supply shocks between countries, Bayoumi and Eichengreen (1994), Eichengreen (1998), Licandro (2000), Hallwood, Marsh and Scheibe (2006), and Foresti (2007) all found low or negative shock correlations, suggesting that the formation of a monetary union is not feasible in Latin America.[2]

[1] For a broad discussion of the general issues and potential problems of creating a monetary union in Latin America, see Berg, Borensztein and Mauro (2002) and Hochreiter, Schmidt-Hebbel and Winckler (2002). Edwards (2006) provides an excellent review of the existing literature.

[2] The sample of Bayoumi and Eichengreen (1994) consisted of 11 Latin American countries: Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Paraguay, Peru, Uruguay, and Venezuela. Eichengreen (1998) considered countries from the Mercosur region: Argentina, Brazil, Paraguay, and Uruguay, whereas Licandro (2000) added Bolivia and Chile to this group. The sample of Hallwood, Marsh and Scheibe (2006) was comprised of Argentina, Brazil, Chile, Uruguay, and Venezuela. Forseti (2007) considered the same sample of countries

In this paper we take a different approach to test the degree of symmetry to real macroeconomic shocks. Using monthly real effective exchange rate data for the period 1996(1)-2012(11) for 15 Latin American countries, we follow Enders and Hurn (1994) and Sun and Simons (2011) in employing the techniques of cointegration and Granger causality.[3] As discussed by Sun and Simons (2011), the macroeconomic fundamentals of countries participating in an *optimal* currency union share common stochastic trends. Thus, real exchange rates, which are driven by these fundamentals, must also share common stochastic trends. Consequently, if two real effective exchange rates are found to be cointegrated, then this implies shock symmetry, and a monetary union can be considered as a viable option. Causality tests, on the other hand, yield useful information on short-run relationships.

The cointegration and Granger causality analyses suggest the potential of forming a monetary union comprised of the region's two largest economies, Brazil and Mexico, along with Bolivia, Chile, and Colombia. In addition, Argentina and Paraguay are found to be tied to this main group and thus could also be included. The hypothetical monetary union would account for 69.1% (80% if Argentina and Paraguay are included) of the region's total GDP and 66.2% (74.4%) of its total population. Our results indicate that potential membership of the common currency should depend on country size rather than a formation based on existing regional trade agreements and/or geographic proximity. These findings complement recent research concerned with the formation of currency unions in Latin America. For the Mercosur countries, Neves, Stocco and da Silva (2008), using quarterly real exchange rate data for the period 1973-2006, found evidence of cointegration when all member countries were analyzed together, although bivariate relationships highlighted large cross-country dissimilarities, calling into question the viability of a monetary union. We also find little evidence to support the formation of a currency union for Mercosur. Using survey evidence on the determinants of life satisfaction, Hofstetter (2011) argued that the benefits from forming a monetary union comprised of Brazil, Chile, Colombia, Mexico, and Peru could outweigh the costs of increased volatility from relinquishing monetary autonomy. With the exception of Peru, we also find

as Bayoumi and Eichengreen (1994) except for Mexico.

[3] As discussed by Kennedy (2008) and Enders (2010), cointegration avoids the removal of important information concerning long-run relationships that can occur under the Blanchard-Quah structural autoregression approach.

evidence to support the entry of these countries into a Latin American currency union.

Given the possibility of a feasible monetary union in Latin America, the paper then proceeds to consider the important issue of whether the new common currency could be anchored against either the U.S. dollar or the Chinese yuan. Recent cases of dollarization in Ecuador (in 1999) and El Salvador (in 2000) combined with strong trade relations emphasize the importance of the United States for the region. On the other hand, the importance of China has increased dramatically, as highlighted by an average annual export (import) growth rate of 24.8% (24.5%) during the period 2005-2009. Indeed, projections suggest that China will displace the European Union as the region's second largest trading partner within the next 10 years (Rosales and Kuwayama, 2012). However, we find no robust evidence that either the U.S. dollar or the Chinese yuan would be a suitable anchor for the new common currency of the hypothetical monetary union.

Finally, the paper investigates the current level of financial integration amongst the candidate countries identified by the cointegration and Granger causality analyses. Many authors (see, *e.g.*, Edwards, 2006) have stressed the importance of financial convergence as a prerequisite for entering a monetary union. Using monthly lending, deposit, and discount rate data, we test the degree of interest rate pass-through for the period 1996(1)-2012(11). The pass-through analysis is used to uncover how changes in monetary policy over time are passed on to bank lending and deposit rates. As discussed by Sander and Kleimeier (2004), this helps reveal any asymmetries that would exist across countries under a common monetary policy. Following Cottarelli and Kourelis (1994), De Bondt, Mojon and Valla (2005), and Bholla, Aziakpono and Snowball (2011), we estimate an error-correction model and find that the level of pass-through among countries of the hypothetical monetary union is low and dissimilar, with little evidence of convergence over time. Our results indicate that substantial banking reforms would be needed in order to reduce financial rigidities and remove one important obstacle to creating a successful currency union within the region.

Overall, while this paper finds no evidence to recommend a monetary union comprising all countries of Latin America, a common currency for a subset of countries appears possible. For this to become a reality, greater political cooperation and economic integration between the candidate countries would be required.[4] Furthermore, as highlighted by this paper, significant reform of the banking system

[4] See Dorrucci *et al.* (2005) for a discussion on the European lessons for regional

should be one important prerequisite to membership. Future research in this area, and many others that are beyond the scope of this paper, would be highly desirable to help stimulate political and public debate on the issue of forming a Latin American monetary union.

The rest of the paper is organized as follows. Section 2 presents the key findings from the cointegration and Granger causality analyses, while section 3 discusses the main results from the interest rate pass-through analysis. Section 4 briefly concludes.

2. Cointegration and Granger causality analyses

2.1. *Data description*

As emphasized by the optimum currency area literature, a fundamental prerequisite for a successful monetary union is the symmetry of shocks across member countries. Using real effective exchange rate (REER) data, this section attempts to evaluate the degree of shock similarity in Latin America employing the techniques of cointegration and Granger causality. Cointegration allows us to characterize the historical adjustments of two REERs through a significant linear approximation: cointegration of two REERs suggests real shock symmetry. On the other hand, Granger causality tests determine whether adjustments in the REER of one country trigger significant REER adjustments in another country, thereby revealing short-run interdependence either due to exchange rate policy coordination or the existence of strong complementarities in the exports of the two countries.

As discussed by Sun and Simons (2011), REERs can contain richer information than simple bilateral real exchange rates especially for countries that follow similar development strategies (*e.g.* export-orientated growth) and share similar patterns of trade and trade composition. The series used for the REER covers the period January 1996-November 2012 and includes fifteen Latin American countries: Argentina, Belize, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, Mexico, Nicaragua, Paraguay, Peru, Uruguay, and Venezuela.[5] With the exception of Argentina and Peru, all the series were obtained from the International Financial Statistics

integration in Latin America.

[5] By starting the series in 1996 we avoid the worst effects of the *Tequila Crisis*.

(IFS) database of the International Monetary Fund (IMF).[6] The maximum number of missing values for a country was 23 (for Paraguay) although we do not envisage problems because of missing data.[7] In what follows, an increase (decrease) in the REER represents an appreciation (depreciation).

2.2. *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.

2.2.1. Cointegration - Johansen (1988) and Johansen and Juselius (1990) methodology

All variables are first pretested to assess their order of integration. 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 (DFGLS) detrending test. The Augmented Dickey-Fuller and Phillips-Perron tests were also performed for robustness. The REER series of the Dominican Republic was the only series that was not found to be an I(1) process (stationary after one difference) and was consequently omitted from the rest of the analysis.[8]

The notion of cointegration (Engle and 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 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 for countries i and j as $R = (r_i, r_j)'$, we estimate the following vector autoregressive (VAR) model:

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[6] The REERs for Argentina and Peru where obtained from their respective central bank websites.

[7] We used standard practices to reduce the number of missing values. If the missing value was unique it was approximated using the average of the two closest values. Biases due to recognized trends (appreciation or depreciation) were solved by interpolating.

[8] Detailed results of the unit root tests are available upon request.

$$R_t = \beta_0 + \beta_1 R_{t-1} + \beta_2 R_{t-2} + \ldots + \beta_s R_{t-s} + \varepsilon_t \tag{1}$$

where β_0 is a 2×1 vector of constants; β_k, $k = 1, 2, ..., s$ is a 2×2 coefficient matrix of R_t lagged by k periods; and ε_t is a 2×1 vector of residuals that satisfies standard Gaussian properties. The optimum lag length, s, is determined using the likelihood ratio test.[9]

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 1998/1999 crisis in Brazil, the 2001/2002 crisis in Argentina, and the 2007/2008 global financial crisis. Impulse dummies for the Brazilian crisis and the recent financial crisis were included in our estimation of the optimal 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 and Sand-Zantman, 2009).

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

$$\Delta R_t = \beta_0 + \pi R_{t-1} + \sum_{k=1}^{s-1} \pi_k \Delta R_{t-k} + \varepsilon_t; \tag{2}$$

$$\pi = -\left(I - \sum_{k=1}^{s} \beta_k\right); \pi_k = -\sum_{p=k+1}^{s} \beta_p$$

where I denotes an identity matrix. The key feature to note in (2) is that the rank of the 2×2 matrix π is equal to the number of independent cointegrating vectors, or equivalently, the number of its non-zero characteristic roots. We use the following test for the number of cointegrating relationships in our bivariate model:

9 This estimator generally favors a larger number of lags than both the Akaike and Standard Bayesian Information Criteria multivariate generalizations. We favored a large number of lags to control for possible autocorrelation issues in the residuals of the VAR models employed.

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

where $\hat{\lambda}_i$ is the estimated values of the characteristic roots 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 z 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 $Rank(\pi) \geq 1$, then the two REERs are cointegrated. The final considerations we take into account relate to the stability and white noise residuals of the VEC model. If cointegration is not rejected, then the stability of the implicit model is tested using the eigenvalue stability condition.[10] Finally, we perform a Lagrange-multiplier (LM) test for the null hypothesis of no autocorrelation in the residuals of the VEC model.

2.2.2. 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 have influenced some of the endogenous variables earlier. Granger causality (Granger, 1969; 1980) helps to identify evidence of this temporal ordering. Specifically, we test whether the lagged values of the REER of one country are useful in forecasting the other REERs in our optimum bivariate systems.

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 the methodology of Dolado and Lütkepohl (1996) which 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 s. Now we refit the data with a VAR$(s+1)$, such that:

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

$$r_{i,t} = \delta_{0,i} + \sum_{v=1}^{s+1} \delta_{41,v} r_{i,t-v} + \sum_{v=1}^{s+1} \delta_{42,v} r_{j,t-v} + \varepsilon_{i,t} \tag{4}$$

$$r_{j,t} = \delta_{0,j} + \sum_{v=1}^{s+1} \delta_{51,v} r_{i,t-v} + \sum_{v=1}^{s+1} \delta_{52,v} r_{j,t-v} + \varepsilon_{j,t} \tag{5}$$

where $\delta_{la,v}$ is the coefficient of the vth lag of variable a on equation l; and $\delta_{0,a}$ and $\varepsilon_{a,t}$ denote the intercept and white noise residuals, respectively, when a is the dependent variable. The least square estimators of the coefficients are now consistent and asymptotically efficient. The Granger causality test uses the first s optimum number of lags. Consider equation (4). We test the null hypothesis that the REER of country j does not Granger cause the REER of country $i: \ H_0 : \delta_{42,v} = 0, \forall v = 1, 2, \ldots, s$. We use an estimated F-test statistic: if it is greater than the critical value then the null hypothesis is rejected.

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

$$\Delta r_{i,t} = \delta_{0,i} + \sum_{v=1}^{s} \delta_{61,v} \Delta r_{i,t-v} + \sum_{v=1}^{s} \delta_{62,v} \Delta r_{j,t-v} + \varepsilon_{i,t} \tag{6}$$

$$\Delta r_{j,t} = \delta_{0,j} + \sum_{v=1}^{s} \delta_{71,v} \Delta r_{i,t-v} + \sum_{v=1}^{s} \delta_{72,v} \Delta r_{j,t-v} + \varepsilon_{j,t} \tag{7}$$

Notice that the coefficients are analogous to the case in (4) and (5) above except that they are differenced. Let us illustrate the Granger-causality test using equation (6). In this case, the null hypothesis of no Granger causality is: $H_0 : \delta_{62,v} = 0, \forall v = 1, 2, \ldots, s$.

2.3. *Cointegration results*

Table 1 reports the λ-trace test results for the 91 bivariate models estimated following the methodology of Johansen (1988) and Johansen and Juselius (1990). In addition to including impulse dummies when estimating the correct VAR model, 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 non-cointegration is rejected in 9 out of the 91 cases.

To help summarize the cointegration relationships, Figure 1 illustrates our key findings. The main *diamond* shows the only group that exhibits strong long-run co-movements. Colombia is at the core of the diamond, as evidence of cointegration was found between Colombia and four other countries: Bolivia, Brazil, Chile, and Mexico. At the same time, these four countries exhibit evidence of cointegration with one additional member of the group: Chile-Mexico and Bolivia-Brazil.

An extension to the diamond arises due to links between Bolivia and Mexico with other countries. Mexico shows evidence of long-run ties with Argentina, while Bolivia contains a link with both Paraguay and Peru. Paraguay exhibits a weak link with the main group, only showing evidence of cointegration with Bolivia. In turn, Peru exhibits an even weaker link as its only relationship is with Paraguay.

An immediate implication of our results is that a monetary union comprised of all Latin American countries seems implausible. The limited number of cointegration relationships found (around 10%) highlights the low level of monetary integration within the region. The analysis finds a clear lack of integration between the three Central American countries in the sample: Belize, Costa Rica, and Nicaragua. Furthermore, we find no evidence of significant integration between these Central American countries and either Colombia or Mexico.

The launch of the Mesoamerican Integration and Development Project (MIDP) in 2008 to strengthen economic integration in the Mesoamerican region is an important step towards rectifying this situation.[11] Overall, despite the long standing agreements of the Caribbean Community and Common Market (CARICOM) and the Central American Common Market (CACM), the ratification of the Dominican Republic - Central America - United States Free Trade Agreement (CAFTA-DR),[12] and the work of the Central American Council, a common currency still seems a highly unfeasible option for the Central American and Caribbean region.

[11] The MIDP is comprised of Belize, Colombia, the Dominican Republic, Guatemala, Honduras, Mexico, Nicaragua, Panama, and El Salvador.

[12] The five Central American countries involved are Costa Rica, Guatemala, Honduras, Nicaragua, and El Salvador.

Similarly, we find little evidence of integration among Mercosur countries. Indeed, the only direct link found was between Bolivia and Brazil. This complements the conclusions of the existing literature (see, *e.g.*, Eichengreen, 1998; Licandro, 2000; Hallwood, Marsh and Scheibe, 2006; Neves, Stocco and da Silva, 2008) that Mercosur does not satisfy the criteria for forming a successful monetary union.

Overall, we do not find evidence of strong long-run ties based on regional trade blocs or geographical proximity. While regional trade agreements have resulted in greater economic cooperation amongst members, there is little evidence to suggest that this has enhanced greater monetary integration (at least in terms of reducing the degree of asymmetry to macroeconomic shocks). Although the percentage of cointegration cases found is low among our sample of 14 countries, 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 countries of the main diamond of figure 1, while Panel B includes all the countries of the extended diamond (without Peru). If the five countries of the main diamond were to form a monetary union, this bloc would account for over 66% of the total Latin American population in 2012 and would be responsible for nearly 70% of its total GDP. For the case where Argentina (the third biggest economy in the region) and Paraguay are also members of the hypothetical monetary union, the extended bloc would account for nearly 80% of total Latin American GDP and over 74% of its population.

In summary, country size appears to be an important factor in the formation of a monetary union in Latin America. Large countries within the region appear to have greater shock symmetry than small countries. For instance, the main diamond includes the two regional 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.[13]

[13] The importance of including the biggest regional economies in a Latin American monetary union is similar to the conclusions of Sato, Zhang and Allen (2009) who found that a monetary union among ASEAN economies was only feasible with the inclusion of Japan.

Table 1

Cointegration analysis for 14 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	4.50	9.00	6.72	**16.53***	8.23	10.80						
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****	9.00	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	9.20	12.73	9.02	14.92		
PER	13.69	7.19	9.83	6.96	11.01	6.80	10.83	9.04	9.44	12.7	**16.38***	9.06	
ARG	**19.93***	7.44	14.10	2.65	9.77	5.26	12.70	3.12	8.58	11.9	6.47	2.95	5.08

Notes: 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 unstable or showed evidence of autocorrelated 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.

Figure 1
The extended diamond

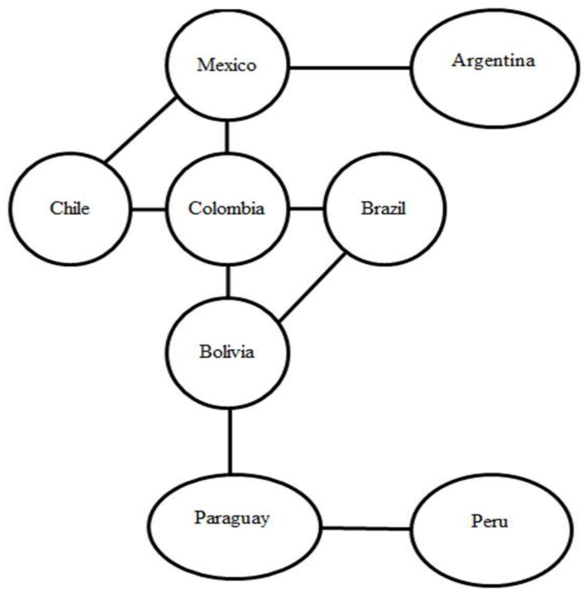

Table 2
Relative size of the potential monetary union

Country	*GDP (PPP-Billions)*	*Percentage of population*	*GDP per capita (PPP)*
Panel A			
Bolivia	55.23	1.85	5 099.27
Brazil	2 355.59	33.86	11 875.26
Chile	320.54	2.97	18 419.04
Colombia	502.87	7.95	10 791.73
Mexico	1 758.90	19.61	15 311.77
% Latin American GDP	69.12		
% Latin American population		66.24	

Table 2
(*continued*)

Country	*GDP (PPP-Billions)*	*Percentage of population*	*GDP per capita (PPP)*
Panel B			
Bolivia	55.23	1.85	5 099.27
Brazil	2 355.59	33.86	11 875.26
Chile	320.54	2.97	18 419.04
Colombia	502.87	7.95	10 791.73
Mexico	1 758.90	19.61	15 311.77
Argentina	743.12	7.00	18 112.33
Paraguay	40.87	1.14	6 136.46
% Latin American GDP	79.97		
% Latin American population		74.38	

Source: *World Economic Outlook* of the International Monetary Fund.
Notes: GDP, population and GDP per capita are for 2012 data. All figures are IMF staff estimates.

2.4. *Granger causality results*

Table 3 reports the results from the Granger causality analysis for all 14 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 analysis as three out of six unidirectional causality relationships were found. This suggests that at least in the short-run similar macroeconomic shocks have induced similar stabilization policies. However, there is still no evidence of links between these countries and either Colombia or Mexico. Thus, the general Mesoamerican region appears weakly integrated.

Compared to the relatively low number of cointegration relationships, the increase in the number of causality relationships found for the majority of the small economies in the sample suggests that the latter may not be evidence of regional monetary integration, but rather of economic dependence on a low number of goods and services. Price changes in a good or service on which a country is heavily dependent on 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 found is evidence of monetary policy synchronization opportunities, constant beggar-thy-neighbor policies, or weak and unreliable data.[14]

Table 4 summarizes the Granger causality p-values for the countries in the extended diamond illustrated in figure 1. For the cointegrated pairs Mexico-Chile, Chile-Colombia, Brazil-Colombia, Colombia-Bolivia, and Bolivia-Paraguay, we also find evidence of causality. While no Granger causality interaction was found for the cointegrated pair Mexico-Argentina, one interesting feature is the causality interactions found for Argentina for a number of non-cointegrated pairs: Argentina-Brazil, Colombia-Argentina, Bolivia-Argentina, and Paraguay-Argentina. Indeed, the Argentina-Brazil case is one of the few cases of bidirectional causality in the sample. This suggests that the two biggest economies of South America generate important short-term policy impacts in the synchronization reactions of other South American countries. For Paraguay and Peru, Paraguay shows a total of five Granger causality relationships (including one case of bidirectional causality), whereas Peru shows none (not even for the cointegrated pair Paraguay-Peru). This suggests that Paraguay may be considered a member of the hypothetical monetary union, whereas there is not sufficient evidence of monetary integration for Peru to be included.

2.5. *The dollar or the yuan?*

In this subsection we analyze the possibility of anchoring the common currency of the hypothetical monetary union to either the U.S. dollar or the Chinese yuan. In essence, the adoption of the common currency involves irrevocably fixing the exchange rate between all participants. Therefore, one possibility is to anchor the currencies of all participant countries to a foreign currency.

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[14] The apparent evidence of short-run interactions suggests that future tests should be carried out to see if they become long-run relationships. 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 14 Latin American countries

	MEX	*BEL*	*NIC*	*CR*	*COL*	*VEN*	*ECU*	*BOL*	*BRA*	*CHI*	*PAR*	*URU*	*PER*	*ARG*
MEX		.547	.635	.061	.574	**.025***	.296	.164	.146	.365	**.016***	**.020***	.301	.875
BEL	.783		**.015***	**.042***	.585	**.000****	.068	.135	.814	.961	**.003****	**.067***	.175	.718
NIC	.918	.656		.773	.226	.168	**.000****	.006#	**.010****	.243	.174	.572	.172	.228
CR	.132	**.001****	.462		.395	.228	**.049***	**.025***	.695	.148	.491	.312	**.046***	.527
COL	.387	.358	.071	.495		.586#	**.001****	.070	.169	**.004****	.797	.147	.137	.999
VEN	.092	.083	.755	**.026***	.066#		1	.981	.819	.895	**.000****	.854	.479	1
ECU	.528	.979	.319	.677	.491	.945		.660#	.315	.074	.878	.329	.471	.834
BOL	.679	**.011***	.001#	**.030***	**.002****	**.030***	.605#		.782	.177	**.038***	**.007****	.133	.096
BRA	.153	.152	.517	.578	**.000****	.555	**.005****	.374		**.012***	.170	**.000****	.561	**.000****
CHI	**.048***	.348	.152	.311	.130	**.025***	.949	.767	.276		**.000****	.510	.525	.153
PAR	.679	.373	**.016***	**.010****	**.001****	.233	**.001****	**.003****	.123	.797		.256	.921	.153
URU	.687	.675	.734	**.008****	**.000****	**.000****	.108	**.030***	.553	**.017***	.333		.099	.270
PER	.428	.279	**.016***	**.000****	.088	.714	.077	.088	.282	.457	.053	.591		.842
ARG	.982	.960	.892	.713	.000**	.000**	.291	**.002****	**.000****	.383	**.012***	**.000****	.770	

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 Chile (CHI) in the row and at Mexico (MEX) in the column, we conclude that Chile's REER Granger causes Mexico's REER because the p-value is .048; hence we reject the null hypothesis at the 5% significance level. *, **, #; denotes Granger causality at the 5%, Granger causality at the 1%, and either instability or serial autocorrelation, respectively. Abbreviations: same as table 1.

Table 4

Granger causality results for cointegrated countries

	MEX	*ARG*	*CHI*	*COL*	*BRA*	*BOL*	*PAR*	*PER*
MEX		.875	.365	.574	.146	.164	**.016***	.301
ARG	.982		.383	**.000****	**.000****	**.002****	**.012***	.770
CHI	**.048***	.153		.130	.276	.767	**.000****	.525
COL	.387	.999	**.004****		.169	.070	.797	.137
BRA	.153	**.000****	**.012***	**.000****		.374	.170	.561
BOL	.679	.096	.177	**.002****	.782		**.038***	.133
PAR	.679	.153	.797	**.001****	.123	**.003****		.921
PER	.428	.842	.457	.088	.282	.088	.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% level, and Granger causality at the 1% level, respectively. Abbreviations: MEX = Mexico, ARG = Argentina, CHI = Chile, COL = Colombia, BRA = Brazil, BOL = Bolivia, PAR = Paraguay, PER = Peru.

On the one hand, the importance of the United States for the Latin American region has induced researchers to analyze the possible costs and benefits of strict dollarization when considering monetary regime options for Latin American countries (see, *e.g.*, Alesina, Barro and Tenreyro, 2002; Berg, Borensztein and Mauro, 2002; Hochreiter, Schmidt-Hebbel and Winckler, 2002; Panizza, Stein and Talvi, 2003; Hofstetter, 2011). On the other hand, Latin America's trade with China has increased dramatically. For example, during the period 2005 - 2009 the average annual growth rate of exports (imports) from China to Latin America was 24.8% (24.5%). Projections suggest that China will displace the European Union as the region's second largest trading partner within the next 10 years (Rosales and Kuwayama, 2012). Consequently, we investigate whether either the U.S. dollar or the Chinese yuan could serve as an anchor for the hypothetical new common currency.[15]

Table 5 summarizes the results from the cointegration and Granger causality tests of the bivariate models for each country against the United States and China. By inspection, there is evidence that the United States causes short-run competitive adjustments in Paraguay, Nicaragua, and Venezuela, whereas China exhibits Granger causality relationships with Belize, Bolivia, and Nicaragua. While the cointegration analysis shows evidence of long-run co-movements for the United States with Argentina, Colombia, and Mexico, and China shows evidence with Costa Rica, the cointegration results are not robust as no evidence of Granger causality was found for any of these relationships. Table 6 summarizes the results of cointegration and Granger causality relationships for the hypothetical monetary union. By inspection, there is no significant evidence that the dollar or the yuan are influential enough to act as an anchor currency.

To end this section, we summarize our key findings so far. Based on the cointegration and Granger causality analyses, we find evidence supporting a monetary union comprised of Bolivia, Brazil, Chile, Colombia, and Mexico; as well as the potential for including Argentina and Paraguay. The analyses suggest that this bloc has five cointegrated-causality pairs (Mexico-Chile, Chile-Colombia, Brazil-Colombia, Colombia-Bolivia, and Bolivia-Paraguay) and eight non-cointegrated-causality pairs (including two cases of bidirectional causality). While the quantity of countries involved in the hypothetical monetary union is small, its relative size would be large, com-

[15] In addition, we also tested the degree of cointegration between Spain and our sample of Latin American countries but did not find any significant relationship.

prising approximately 80% of the region's 2012 GDP and 74% of its population. Finally, we employed the techniques of cointegration and Granger causality to test if there was evidence to support the formation of a United States or China currency bloc. No evidence exists to support either full dollarization or yuanization within the hypothetical monetary union or the anchoring of the new common currency against either of these currencies.

Table 5
Impact of the United States and China in Latin America

Latin American country	*Foreign country*	*λ-trace?*	*Co-integration?*	*Causality test's p-value*	*Granger causality?*
MEXICO	USA	**16.75***	**YES**	.852	NO
	CHINA	9.94	NO	.746	NO
BELIZE	USA	9.58	NO	.107	NO
	CHINA	6.70	NO	**.002****	**YES**
NICARAGUA	USA	14.84	NO	**.030***	**YES**
	CHINA	4.22	NO	**.011***	**YES**
COSTA RICA	USA	9.43	NO	.719	NO
	CHINA	**16.14***	**YES**	.359#	NO
COLOMBIA	USA	**16.56***	**YES**	.657	NO
	CHINA	13.83	NO	.730	NO
VENEZUELA	USA	11.57	NO	**.017***	**YES**
	CHINA	11.93	NO	.154	NO
ECUADOR	USA	10.74	NO	.088	NO
	CHINA	9.93	NO	.150	NO
BOLIVIA	USA	9.01	NO	.748	NO
	CHINA	5.43	NO	**.008***	**YES**
BRAZIL	USA	10.49	NO	.688	NO
	CHINA	13.75	NO	.147	NO
CHILE	USA	13.68	NO	.858	NO
	CHINA	11.63	NO	.809	NO
PARAGUAY	USA	8.99	NO	**.002****	**YES**
	CHINA	13.83	NO	.446	NO
URUGUAY	USA	14.78	NO	.166	NO
	CHINA	8.20	NO	.183	NO

Table 5
(*continued*)

Latin American country	*Foreign country*	*λ-trace?*	*Co-integration?*	*Causality test's p-value*	*Granger causality?*
PERU	USA	11.22	NO	.494	NO
	CHINA	10.31	NO	.058	NO
ARGENTINA	USA	**19.14***	**YES**	.355	NO
	CHINA	2.83	NO	.327	NO

Notes: Each Latin American country is analyzed for evidence of cointegration and Granger causality with both the USA and China. 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.

Table 6
Impact of the United States and China on the members of the hypothetical monetary union

	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: Abbreviations are the same as in table 4.

3. Interest rate pass-through analysis

3.1. *Data description and monetary policy pass-through*

In this section, we investigate the degree of financial convergence between the seven countries identified as possible members of a Latin

American monetary union. In particular, the analysis focuses on the degree of interest rate pass-through in each country. The greater the degree of interest rate pass-through, the greater the impact of a change in central bank interest rates on retail rates.[16] In addition, our analysis also reveals how fast this transmission occurs. These two issues are important because if the retail rates of the candidate countries respond quickly and in a similar pattern to changes in their policy rates, then nominal convergence likely exists, which is a necessary condition for a monetary union (Bholla, Aziakpono and Snowball, 2011).

We use monthly data for the period January 1999 - November 2012 for Bolivia, Brazil, Chile, Colombia, and Mexico. The discount rate represents the central bank policy rate, and lending and deposit interest rates are used for the retail rates.[17] All series were obtained from the IFS database of the IMF. The discount rates of Argentina and Paraguay could not be accurately obtained for the whole sample period and these countries were consequently omitted.

3.2. *Econometric methodology*

The econometric techniques employed follow Enders (2010) and Bholla, Aziakpono and Snowball (2011). All variables are tested for a unit root using the DFGLS and Phillips-Perron tests. The results suggest that all the interest rates were I(1) processes.[18] 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 to note here is that we exploit economic theory and do not treat the variables symmetrically (as in the Johansen, 1988, methodology of section 2.2.). Let us suppose that we have a cointegrated system. If a variable does not respond to the discrepancy from the long-run equilibrium relationship, then we can say that the variable is weakly exogenous. In our particular case, the discount

[16] Recent empirical studies (see, *e.g.*, Wang and Lee, 2009; Haughton and Iglesias, 2012) find evidence of consistent rigidities and heterogeneity in the pass-through levels in various countries.

[17] For Mexico, the discount rate series only begins from 2008 when the central bank of Mexico published the official objective discount rate. Consequently, we used the Mexican treasury bill interest rate as a proxy from the beginning of the sample period until the end of 2007.

[18] The unit root and cointegration tests are available upon request.

rate is weakly exogenous because it can be controlled by the central bank, and therefore it does not experience any type of endogenous feedback from the private retail rates when disturbances need to be adjusted for. In essence, the weak exogeneity assumption dictates a causal relationship from the monetary policy (discount) rate to the retail (private) rate 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 the discount rate on either the lending or deposit rate.

The following error-correction model is first estimated:

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

where $i = 1$, 2 (lending, deposit); Δ denotes a first-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 (the dependent variable) is not needed to be the same lag length, q, as the lag values of the discount rate (the independent and weakly exogenous variable). We estimate the most appropriate model using the Akaike Information Criterion (AIC), which generally favors a greater number of lags (Stock and Watson, 2002) but reduces the probability of correlation issues with the error term.[19]

Equation (8) is the error-correction model in Autoregressive Distributed Lag (ADL) form when using the concept of weak exogeneity which allows our model to be identified. Equation (8) 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 do not reject the null then there is no error-correction term. Hence, the proper alternative hypothesis is $\alpha_1 < 0$ so that we do not reject convergence, with the estimated value of α_1 being the velocity of adjustment. The appropriate critical values depend on the number of

[19] Only when the number of lags is "sufficient" can we expect to avoid endogeneity problems, since we will have included enough past information in the error-correction model.

I(1) regressors in the model (denoted by k), the adjusted sample size N^{adj}, and the inclusion of an intercept. Then, $N^{adj} = N-(2k-1)-1$, where N is the number of observations in the sample.[20]

If evidence of cointegration is found, then, as already mentioned, α_1 is the velocity of adjustment parameter and α_3 is the short-run pass-through (impact multiplier). It is then valid to obtain the long-run pass-through from a simple OLS regression of the private retail rate on the discount rate. That is, if there is evidence of cointegration, then the following model can be estimated:

$$PRi_t = \theta_0 + \theta_1 DR_t + u_t \tag{9}$$

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:

$$\begin{aligned}\Delta PRi_t = \alpha_0 + \alpha_3 \Delta DR_t + \sum_{m=1}^{p} \alpha_{PRi,m} \Delta PRi_{t-m} \\ + \sum_{k=1}^{q} \alpha_{DR,k} \Delta DRi_{t-k} + v_t\end{aligned} \tag{10}$$

This simply removes the (separated) non-significant error-correction term in model (8). Since our differenced variables are stationary, we need not worry about additional information relating to long-run relationships, as we have rejected cointegration. The ADL specification avoids problems of spurious regression but a simple OLS model like (9) is no longer 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:

$$\begin{aligned}&\alpha_{PRi}(L)\, PRi_t = \alpha_0 + \alpha_{DR}(L)\, DR_t + v_t; \\ &\alpha_{PRi}(L) = 1 - \alpha_{PRi,1} L - \alpha_{PRi,2} L^2 - \ldots - \alpha_{PRi,p} L^p; \\ &\alpha_{DR}(L) = \alpha_3 + \alpha_{DR,1} L + \alpha_{DR,2} L^2 + \ldots + \alpha_{DR,q} L^q.\end{aligned} \tag{11}$$

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

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

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

with expected value:

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

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

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

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

$$\begin{gathered} LRPT_{No\ cointegration} = \alpha^{-1}_{PRi}\ (1)\ \alpha_{DR}\ (1) = \\ \frac{\alpha_3 + \alpha_{DR,1} + \alpha_{DR,2} + \ldots + \alpha_{DR,q}}{1 - \alpha_{PRi,1} - \alpha_{PRi,2} - \ldots - \alpha_{PRi,p}} \end{gathered} \tag{12}$$

3.2.1. Rolling-window technique analysis

We carried out the analysis described above for the whole sample period (167 observations). We then divided the entire sample into seven periods and redid 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. The analysis should shed some light on two characteristics concerning the current feasibility of our hypothetical monetary union: convergence in the magnitude of the pass-through among the member countries and a strong transmission of monetary policy (i.e. a high pass-through level).

To end this subsection, note that we have included dummies for the 1998/1999 crisis of Brazil, the 2001/2002 crisis of Argentina, and the 2007/2008 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 the internal financial characteristics of each of the countries, and these three dummies plausibly capture the main breaks for each of them.[21] We test the joint significance of the dummies at the 10% significance level and see which dummies might be important to consider when estimating the best models.

3.3. *Results*

Table 7 summarizes the main results of this section. In addition, 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.[22]

First, consider the case of short-run pass-through; i.e., the immediate adjustment of retail rates to changes in the discount rate. For the lending rate (the left-hand diagram given in figure 2), strong asymmetries are found across the period.

---

[21] In the previous section, we argued that the Argentinian crisis of 2001/2002 had concentrated national effects instead of 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 the significance of the dummy variables in each country's model allows us to consider important biases that might affect only that country.

[22] For the case of the lending rate, Bolivia and Chile are omitted from figure 2 because of the lack of significance of most of the rolling windows.

Table 7

Interest-rate pass-through 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**	.933	1.093	.3163
		2000-2006	**-3.30***	**YES**	.952	1.092	.2688
		2001-2007	-2.76	NO	.887	.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	.918	1.295	
		1999-2012	-3.12	NO	.933	1.040	
	Deposit	1999-2005	-1.79	NO	.293	.462	
		2000-2006	-2.29	NO	.292	.472	
		2001-2007	-2.81	NO	.288	.215	
		2002-2008	-1.08	NO	.205	.341	
		2003-2009	-1.01	NO	.177	.337	
		2004-2010	-1.10	NO	.184	.309	
		2005-2012	-1.67	NO	.149	.326	
		1999-2012	-2.79	NO	.289	.464	
Bolivia	Lending	1999-2005	-1.62	NO	.198#	2.452#	
		2000-2006	-1.91	NO	.076#	2.228#	

Table 7
(*continued*)

Country	*Interest rate*	*Period*	*t-statistic for cointegration test*	*Cointegration?*	*Short-run pass through*	*Long-run pass through*	*Velocity of adjustment*
		2001-2007	**-3.61***	**YES**	.402#	1.261	.5309
		2002-2008	-2.62	NO	.406#	1.810#	
		2003-2009	-1.70	NO	.242#	1.276	
		2004-2010	-1.64	NO	-.106#	.381#	
		2005-2012	-2.59	NO	.247#	.698#	
		1999-2012	-2.73	NO	-.068#	.717#	
	Deposit	1999-2005	-2.84	NO	.027#	.453#	
		2000-2006	-2.76	NO	.014#	.424	
		2001-2007	-3.16	NO	-.098#	.335	
		2002-2008	-3.12	NO	-.071#	.455	
		2003-2009	-1.23	NO	.348	.403	
		2004-2010	**-3.88****	**YES**	.310	.514	.1674
		2005-2012	-1.90	NO	.328	.359	
		1999-2012	-1.81	NO	.288	.462	
Colombia	Lending	1999-2005	-2.58	NO	.408	.490	
		2000-2006	-2.73	NO	.366	.466	
		2001-2007	-1.67	NO	.109#	.702	
		2002-2008	-.43	NO	.093#	.457	

Table 7
(*continued*)

Country	*Interest rate*	*Period*	*t-statistic for cointegration test*	*Cointegration?*	*Short-run pass through*	*Long-run pass through*	*Velocity of adjustment*
		2003-2009	-3.22	NO	.248	.804	
		2004-2010	-2.50	NO	.330	.975	
		2005-2012	-2.97	NO	.384	.992	
		1999-2012	**-4.06****	**YES**	.352	.794	.1836
	Deposit	1999-2005	-2.02	NO	.218	.205	
		2000-2006	-2.37	NO	.095	.216	
		2001-2007	-.74	NO	.064#	.715#	
		2002-2008	-.28	NO	.011#	-.000#	
		2003-2009	**-3.53***	**YES**	.142	.533	.1216
		2004-2010	-3.10	NO	.143	.701	
		2005-2012	**-3.27***	**YES**	.180	.630	.1538
		1999-2012	**-4.70****	**YES**	.206	.699	.1700
Chile	Lending	1999-2005	**-3.94****	**YES**	.541	1.012	.6152
		2000-2006	-3.15	NO	.515	.808	
		2001-2007	**-3.29***	**YES**	.380	1.037	.3427
		2002-2008	-3.23	NO	-.054#	1.713#	
		2003-2009	-1.88	NO	.204#	1.660	
		2004-2010	-1.71	NO	.098#	1.618#	

Table 7
(*continued*)

Country	*Interest rate*	*Period*	*t-statistic for cointegration test*	*Cointegration?*	*Short-run pass through*	*Long-run pass through*	*Velocity of adjustment*
		2005-2012	-1.28	NO	.171#	1.599#	
		1999-2012	**-4.42****	**YES**	.481	1.028	.2602
	Deposit	1999-2005	.01	NO	.533	1.504#	
		2000-2006	-.85	NO	.532	.637	
		2001-2007	-1.19	NO	.459	.798	
		2002-2008	**-4.88****	**YES**	.784	1.044	.6529
		2003-2009	**-4.08****	**YES**	.836	1.026	.4831
		2004-2010	**-4.22***	**YES**	.877	1.020	.6245
		2005-2012	**-3.65***	**YES**	.903	1.024	.4347
		1999-2012	-3.00	NO	.620	.799	
Brazil	Lending	1999-2005	-2.13	NO	.797	1.358	
		2000-2006	-2.55	NO	.752	1.340	
		2001-2007	-2.70	NO	.643	1.032	
		2002-2008	-2.60	NO	.645	1.039	
		2003-2009	-2.39	NO	.904	1.250	
		2004-2010	-2.89	NO	.495	1.004	
		2005-2012	-2.56	NO	.691	1.809	
		1999-2012	**-3.40***	**YES**	.837	1.824	0.1477

Table 7
(*continued*)

Country	*Interest rate*	*Period*	*t-statistic for cointegration test*	*Cointegration?*	*Short-run pass through*	*Long-run pass through*	*Velocity of adjustment*
	Deposit	1999-2005	-2.71	NO	.484	.825	
		2000-2006	-2.01	NO	.514	.843	
		2001-2007	-2.04	NO	.522	.837	
		2002-2008	**-3.26***	**YES**	.650	.970	.4911
		2003-2009	-2.97	NO	.448	.808	
		2004-2010	**-3.55***	**YES**	.544	.946	.4151
		2005-2012	**-4.01****	**YES**	.724	.923	.4458
		1999-2012	**-3.63***	**YES**	.612	.993	.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%, respectively, for the levels 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

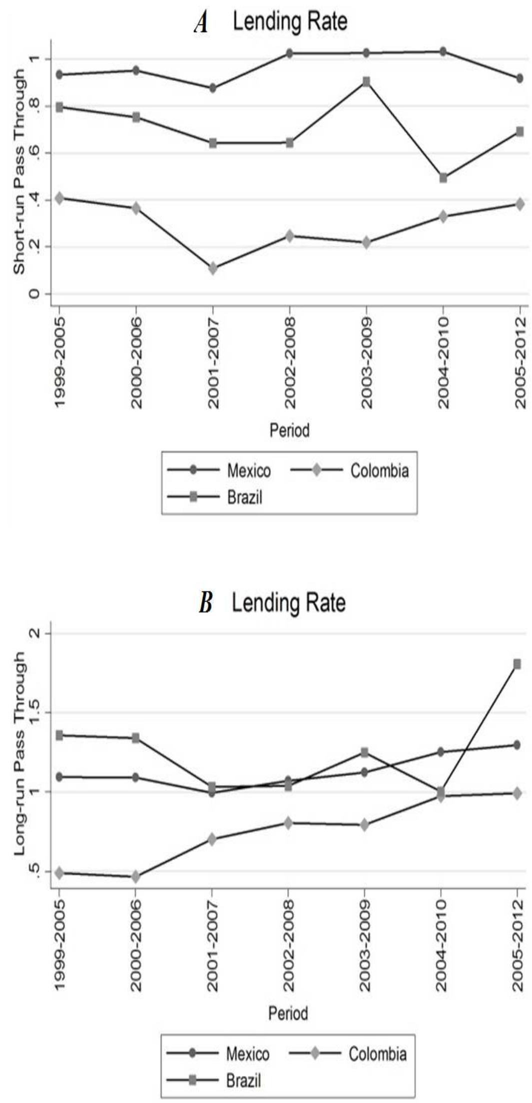

Diagram *A* illustrates the short-run pass-through, while diagram *B* illustrates the long-run pass-through.

Figure 3
Pass-through of the deposit rate

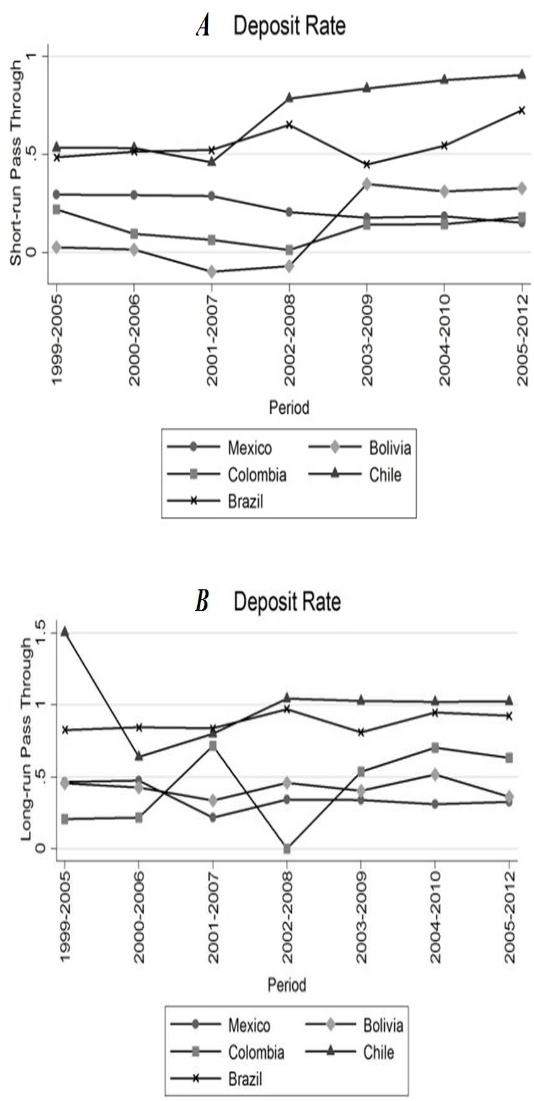

Diagram *A* illustrates the short-run pass-through, while diagram *B* illustrates the long-run pass-through.

Mexico exhibits the strongest and most significant short-run pass -through.[23] Brazil shows considerable variance in its pass-through level from a minimum of almost 0.5 to a maximum of around 0.9, whereas Colombia shows systematically low levels of short-run pass-through onto the lending rate (the rolling-windows 2001-2007 and 2002-2008 are not significant). The results for Bolivia find that none of the rolling-windows are significant, and for Chile the last four rolling-windows are not significant. Overall, the evidence on short-run pass-through for the lending rate indicates strong rigidities for all sample countries, with the exception of Mexico.

The previous finding of strong rigidities is also supported by the short-run results for deposit rates. The only country that has a strong pass-through for the deposit rate is Chile with a value of 0.9 for the rolling-window of 2005-2012. Brazil presents pass-through levels that are similar to lending rate pass-through levels, and results for Bolivia are significant only for the last three rolling-windows, albeit at low levels. The two countries that exhibit some degree of convergence for the final subsample periods are Mexico and Colombia but at very low levels of below 0.2 for short-run pass-through.

We now analyze the situation of long-run pass-through for the members of the hypothetical monetary union. For the lending rate case, there is little evidence of credit rationing with the exception of Colombia. For the deposit rate, Chile and Brazil exhibit near perfect long-run pass-through while the other countries present strong, but different levels, of rigidities. However, there is no evidence that pass-through levels across countries are converging towards high values of pass-through. Considerable differences continue to exist throughout the rolling-window subsamples.

To complement the analysis of long-run pass-through, we also tested for cointegration and estimated the speed of adjustment. Table 8 reports the results for the cases where cointegration was found and adds the calculated number of months it takes for the long-run pass-through to be fully realized.[24] The total number of cointegration relationships found is 21 out of the 96 periods analyzed (12 complete samples and 84 rolling-window subsamples). This means that for 22%

[23] 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).

[24] 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.

of cases, a strong long-run equilibrium relationship that involves an adjustment mechanism process is present. 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, suggesting a slow speed of adjustment for both the lending and deposit rates.

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	.3163	12
		2000-2006	1.092	.2688	14
Bolivia	Lending	2001-2007	1.261	.5309	7
	Deposit	2004-2010	.514	.1674	16
Colombia	Lending	1999-2012	.794	.1836	17
	Deposit	2003-2009	.533	.1216	20
		2005-2012	.630	.1538	18
		1999-2012	.699	.1700	18
Chile	Lending	1999-2005	1.012	.6152	6
		2001-2007	1.037	.3427	11
		1999-2012	1.028	.2602	14
	Deposit	2002-2008	1.044	.6529	5
		2003-2009	1.026	.4831	7
		2004-2010	1.020	.6245	5
		2005-2012	1.024	.4347	8
Brazil	Lending	1999-2012	1.824	.1477	25
	Deposit	2002-2008	.970	.4911	7
		2004-2010	.946	.4151	9
		2005-2012	.923	.4458	8
		1999-2012	.993	.2644	13

Notes: The sixth column reports the calculated number of months it takes to fully realize the long-run pass-through of an interest rate in the case of cointegration. Full realization is considered when no more than .00 of the pass-through is yet to be realized.

Overall, our candidate countries show evidence of considerable dissimilarities in pass-through levels, both in the short-run and in the long-run, a lack of convergence towards high values of pass-through, and a relatively low number of cointegration relationships and adjustment lags. This suggests that a common monetary policy would generate significant heterogeneous effects across member countries. Consequently, for a Latin American monetary union to become feasible, strong steps are required to achieve financial convergence among countries and remove a variety of rigidities in the banking and financial market sectors.

4. Conclusions

This paper has investigated the possibility of forming a currency union within Latin America. We first examined the degree of shock symmetry using real effective exchange rate data for fifteen countries. Following the analysis of Sun and Simons (2011), we applied cointegration and Granger causality techniques to analyze long-run and short-run interactions. We found strong long-run ties between Bolivia, Brazil, Chile, Colombia, and Mexico; as well as weaker ties between these five countries and Argentina and Paraguay. Such a monetary union would comprise 80% of the region's GDP and 74% of its total population. It was shown the neither the U.S. dollar nor the Chinese yuan are suitable anchor currencies for the hypothetical monetary union.

We then examined the degree of monetary policy convergence within the hypothetical currency union. Using monthly interest rate data for the discount, lending, and deposit rates, we followed Bholla, Aziakpono and Snowball (2011) in employing error-correction models to determine the degree of interest rate pass-through for the candidate countries. Considerable dissimilarities in the pass-through levels were found, with little evidence of cross-country convergence. This analysis suggests that significant reforms of the banking and financial sectors are needed before a Latin American monetary union could take place.

In our opinion, future research should be directed towards further investigating the feasibility of a monetary union within the region, with particular emphasis on quantifying the economic costs and benefits for each potential member country.

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