What Explains Nominal Exchange Rate Volatility? Evidence from the Latin American Countries[†]

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Abstract

This paper investigates the short-run and long-run impact of the determinants of nominal exchange rate volatility in three Latin American countries during the period 1979-2009. We estimate a multivariate GARCH model and include the covariances of those determinants, which have been ignored in the prior relevant literature. In combination with the role of financial openness and alternative exchange rate regimes, we find that nominal variability, namely variability in the money supply and inflation, explains exchange rate volatility. Output variations are found to be important as well, but only in floating countries. Financial openness seems to affect significantly the volatility of nominal exchange rate in all countries under examination. Finally, flexible exchange rate regimes tend to increase exchange rate volatility only in fixed and floating countries.

Key Words: exchange rate volatility, multivariate GARCH, BEKK, Granger-causality

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1. Introduction

An important challenge to exchange rate theory is the solution to the puzzle that exchange rates are less volatile in more financially open economies. However, there is no uniform solution to this puzzle. On one hand, more open economies tend to have mean-reverting exchange rates, implying less volatility (Hau, 2002; Bleaney, 2008), while, on the other hand, more open countries nay exhibit greater exchange rate volatility (Amor, 2008). Moreover, there is a great disagreement in the finance literature about the behavior of nominal exchange rate volatility under alternative exchange rate arrangements. Flood and Rose (1995) highlight empirically a positive link between exchange rate volatility and flexible exchange rate regimes while Valachy and Kocenda (2003) provide either positive or negative link according to the countries under investigation. Friedman (1953) argues that exchange rate volatility cannot be reduced by switching from floating to fixed exchange rates. Lastly, there is a strand of literature that supports that the stabilization of exchange rates may be achieved through macroeconomic stability (Morana, 2009; Devereux and Lane 2003) while other studies suggest the opposite (Friedman, 1953; Flood and Rose, 1995).

The majority of the existing literature investigates the effects of exchange rate volatility on a number of macroeconomic variables, e.g. trade (Baum and Caglayan, 2009). However, there is a lack of studies examining the determinants of exchange rate volatility. Methodologically, the volatility of macroeconomic aggregates has been computed using the standard deviation or variance of the aggregate (Flood and Rose, 1995; Hau, 2002) while the conditional volatility has been captured by ARCH and GARCH models (Stancik, 2006; Duarte et al., 2008). However, all of these studies neither focus on the specific set of determinants of nominal exchange rate volatility as a whole nor incorporate the covariances of the potential determinants of exchange rate volatility in their investigation. It is possible that the combination of changes in two macroeconomic variables could impact on the volatility of exchange rate. This combination or co-movement of the variables is captured by the conditional covariance terms, which we also take into consideration in our model. Furthermore, some of the aforementioned studies explore the volatility of exchange rate in a panel of countries, and not each country individually.

Therefore, the contribution of the present paper is along several lines: First, we examine the impact of output and monetary volatility determinants on nominal exchange rate volatility.

Second, we incorporate in the group of exchange rate volatility determinants the covariances of shocks to variables that we consider important exchange rate volatility determinants.[1] Third, we estimate multivariate GARCH models in order to capture the volatility. More specifically, in this paper we purport to provide empirical evidence of the determinants of nominal exchange rate variability for three Latin American countries, namely Argentina, Bolivia, and Chile.

The selection of these countries is based on the exchange rate regime that these countries exhibited most of the time. Argentina is characterized as a fixed exchange rate regime country, Bolivia as intermediate and Chile as floating.[2] Among the determinants, we include important domestic macroeconomic variables and international influences, captured by financial openness, proxied as deviation from the Uncovered Interest Parity (UIP), and the exchange rate regime.[3] Lastly, the inclusion of the covariances is considered of great importance because it can capture the joint effect of shocks to specific macroeconomic variables on exchange rate volatility.

To measure volatility, we estimate an Autoregressive Asymmetric Diagonal BEKK (AR-AS-DiagBEKK), a type of a multivariate GARCH model. We estimate the Diagonal BEKK in order to avoid computational difficulties of jointly estimating all model parameters in the first stage. Further, we estimate a reduced-form model of exchange rate volatility and in order to address short-run causality issues we conduct Granger-causality tests. Lastly, the computation of the long-run effects of the determinants of exchange rate volatility is addressed following a variant of the approach of Grier and Smallwood (2007).

The remainder of the paper proceeds as follows: Section 2 presents the determinants of exchange rate volatility concerning the relationship among the variables of interest. Section 3 describes the econometric model and methodology proposed. Section 4 discusses the data. Section 5 presents the empirical results. Section 6 reports the main conclusions and draws some policy implications.

[1] High inflation in Latin American countries stimulated the specific set of included covariances in the model.

[2] The classification is according to Reinhart and Rogoff (2002).

[3] Financial openness is considered a more direct determinant of exchange rate volatility since the effect of capital transactions does not take long to be absorbed. Trade openness is, in principal, included as determinant, but it proved insignificant and thus it is removed from the model. The insignificance may be due to non-linearities in the relationship unable to be captured by our model or because of the fact that the effects of trade openness are not observed immediately.

2. Determinants of Exchange Rate Volatility

2.1. The effect of output volatility on exchange rate volatility

The effect of output volatility on exchange rate volatility is ambiguous from a theoretical point of view. An increase in domestic output, caused by an increase in government spending, leads to an increase in imports causing the exchange rate to increase (depreciation). According to Friedman's (1953) perspective, exchange rate instability may be a symptom of macroeconomic instability concluding that there is a positive relationship between exchange rate volatility and macroeconomic volatility. However, Friedman claims that it is possible for higher macroeconomic volatility to lead to lower exchange rate volatility. This effect may be insignificant according to Morana (2009) in the sense that countries that do not heavily depend on foreign trade suffer less, if any at all, from fluctuations in the exchange rate.[4]

2.2. The effect of monetary policy volatility on exchange rate volatility

A shock in money supply is positively associated with exchange rate volatility. This can be due to nominal devaluations or/and increases in prices (Carrera and Vuletin, 2002). The positive relationship is also supported by Morana (2009), but he finds also a negative relationship in one country of his sample. Calderon (2004) finds that higher money supply volatility leads to higher exchange rate volatility, while Amor (2008) argues that the aforementioned effect is not significant in a sample of developing countries.

With respect to inflation volatility, the literature provides mixed results. Under high (low) inflation, central banks raise (lower) interest rates causing an appreciation (depreciation) of the currency. Therefore the impact of inflation volatility on exchange rate volatility can be either positive or negative depending on the central bank's goal. Besides, there is evidence (Edwards, 1993) that countries that suffer from hyperinflation may be characterized by increased exchange rate volatility in their attempt to stabilize inflation.

[4] The insignificance of this impact is also confirmed by Flood and Rose (1995), who claim that macroeconomic volatility is not an important source of exchange rate volatility.

2.3. The effect of financial openness on exchange rate volatility

The relationship between financial openness and exchange rate volatility is expected to be ambiguous. Dornbusch (1976) argues that freely operating foreign exchange rate markets (perfect capital mobility) may lead to overshooting in nominal exchange rates in the presence of nominal shocks and hence causing an increase in exchange rate volatility. However, evidence on emerging countries (Kose et al., 2006) reveals that during financial globalization the degree of risk sharing has increased decreasing the relative consumption volatility for the countries having more significant financial flows. Therefore, more financial integration leads to a decrease in exchange rate volatility (Hau, 2002; Devereux and Lane, 2003).

2.4. The effect of exchange rate regimes on exchange rate volatility

Regarding the impact of various exchange rate regimes we consider in the present study on the exchange rate volatility, the literature provides divergent evidence. According to Friedman (1953), systemic volatility of exchange rate cannot be reduced by changing exchange rate regimes. However, Mussa (1986), Eichengreen (1988), Flood and Rose (1995) argue that flexible exchange rates are more volatile than fixed. In the same line, Carrera and Vuletin (2002) point to a positive relationship of intermediate (and fixed) exchange rate regime and exchange rate volatility, while Valaschy and Kocenda (2003) claim that a floating regime contributes to increased or decreased exchange rate volatility according to the sample of countries.[5]

The above analysis provides a guideline of the expected effects of the aforementioned factors on nominal exchange rate volatility based on economic theory. However, we perform a short-run and long-run analysis of the determinants of exchange rate volatility which highly depends on the sample size, the time period and the lag length of the explanatory factors.

[5] To define the type of exchange rate regime that holds in a country is a difficult task since the announced (de jure) regime is different from the actual (de facto) one. Taking into account the de facto classification of exchange rate regimes presented in Levy-Yeyati and Sturzenegger (2005) and Reinhart and Rogoff (2002) one can derive an overall classification of exchange rate regimes. Corbo (2002) defines three categories of exchange rate systems: fixed or hard pegs (dollarization, currency unions and currency boards), intermediate regimes (fixed-but-adjustable pegs, crawling pegs and crawling bands) and floating regimes (managed floats and free floats). Definitions on exchange rate regimes can be found analytically in International Monetary Fund (IMF) Annual Report Exchange Rate Arrangements and Anchors of Monetary Policy (2001).

Therefore, we cannot provide a priori expectations of the sign of the effects on exchange rate volatility and the effects may appear to be ambiguous in the short- and long run.

3. Model Specification

3.1. Modeling Volatility

To model conditional variances and covariances we estimate a multivariate GARCH model. We have chosen to estimate a BEKK model (Engle and Kroner, 1995) because it ensures that the conditional variance-covariance matrix is always positive definite. However, there are several specifications for multivariate GARCH models.[6] We estimate the Diagonal BEKK in order to avoid computational difficulties of jointly estimating all model parameters in the first stage.[7] Further, we include Autoregressive (AR) terms in order to account for autocorrelation. The introduction of asymmetric terms is of great importance in order to take into account "bad" or "good" news that affect the conditional variance of the variables under investigation.

We present the Autoregressive-Asymmetric-Diagonal BEKK (AR-AS-DiagBEKK) model as it is used for the computation of the conditional variance (or/and covariance) of the variables under investigation. To illustrate this we present four variables of interest, that is, (i) nominal exchange rate (XR), (ii) output (IPI), (iii) money supply (M1) and (iv) inflation (INF) (i.e., system of four conditional mean equations) and the simple GARCH(1,1) model for the conditional variance because according to the related literature it explains volatility in a satisfactory way.

The selection of the aforementioned variables is based on the monetarist model of Mussa (1976) and Frenkel (1976) and justified in the theoretical work of Obstfeld and Rogoff (1996) who consider that the exchange rate is affected by the monetary policy and output of the domestic and foreign economy, under different assumptions. In addition, Driskill and

[6] A common specification of multivariate GARCH models is the VECH model introduced by Bollerslev et al. (1988). As the number of variables employed in the model increases, the estimation of this model seems infeasible since a large number of parameters are needed to be estimated. For this reason, Bollerslev et al. (1988) restricted the conditional variance-covariance matrix by assuming that matrices of ARCH and GARCH coefficients are diagonal. Other specifications of multivariate GARCH models include the Constant Conditional Correlation (CCC) specification (Bollerslev, 1990) and the Dynamic Conditional Correlation (DCC) model (Engle, 2002; Tse and Tsui, 2002). An analytical survey of multivariate GARCH models can be found in Bauwen et al. (2006).

[7] EVIEWS 6 has been used for the estimation of BEKK models and there is the restriction of estimating the Diagonal form of multivariate GARCH models.

McCafferty (1980) and Flood and Rose (1999) support the validity of output and monetary policy variables as determinants of exchange rate while Grydaki and Fountas (2009), extending the previous works, introduce the role of selected covariances, apart from monetary and real shocks, as determinants of exchange rate volatility.

The conditional mean equation is specified as:

$$Y_t = \mu + \sum_{i=1}^{p} \Gamma_i Y_{t-1} + \varepsilon_t , \quad \varepsilon_t \mid \psi_{t-1} \sim \left(0, H_t\right) \tag{1}$$

where, Y, μ and ε are 4x1 vectors of dependent variables, intercepts and the innovation vector, respectively; i denotes the autoregressive term of each dependent variable. It is mentioned that each dependent variable is regressed only on its lagged values and not on the lagged values of all dependent variables.

The conditional variance equation is given by:

$$H_t = C_0^{*\prime} C_0^* + A_{11}^{*\prime} \varepsilon_{t-1} \varepsilon_{t-1}^{\prime} A_{11}^* + \mathrm{B}_{11}^{*\prime} H_{t-1} \mathrm{B}_{11}^* + D_{11}^{*\prime} \xi_{t-1} \xi_{t-1}^{\prime} D_{11}^* \tag{2}$$

where C_0^*, A_{11}^*, B_{11}^*, D_{11}^* are 4x4 matrices of parameters, H_t is a 4x4 conditional variance-covariance matrix and ξ_t is a 4x1 vector which accounts for asymmetries in the conditional variance-covariance matrix. More specifically, ξ_t represents the effect of bad news on the conditional variance of the variables. We assume that bad news in terms of nominal exchange rate, money supply and inflation are considered higher than expected magnitudes and, hence, correspond to a positive residual. On the other hand, we assume that bad news in terms of output correspond to lower than expected levels and, hence, lead to a negative residual. In our model $\xi_{XR,t}$, $\xi_{M1,t}$ and $\xi_{INF,t}$ is $\max\left(\varepsilon_{XR,t}, 0\right)$, $\max\left(\varepsilon_{M1,t}, 0\right)$ and $\max\left(\varepsilon_{INF,t}, 0\right)$, respectively, reflecting the positive innovations (bad news) on the specific variables; $\xi_{IPI,t}$ is $\min\left(\varepsilon_{IPI,t}, 0\right)$ indicating the negative innovations on the particular variable. The positive definiteness of the covariance matrix is ensured because of the quadratic nature of the terms on the right hand side of equation (2). Matrices A_{11}^*, B_{11}^*, D_{11}^* are diagonal by definition, while C_0^* is an upper triangular matrix.[8]

[8] Under the assumption that the error terms follow the Multivariate Student's t conditional distribution, the parameters of the multivariate GARCH model can be estimated by maximizing the log likelihood function:

$$L_{Stud} = T\, log\left\{ \left(\Gamma\left(v+m\right)/2 \right) v^{m/2} \Big/ \left(v\pi\right)^{m/2} \Gamma\left(v/2\right)\left(v-2\right)^{m/2} \right\} - 1/2 \sum_{t=1}^{T} \left\{ log\left(\left|H_t\right|\right) + \left(v+m\right) log\left[1 + \varepsilon_t^{\prime} H_t^{-1} \varepsilon_t / v - 2\right] \right\}$$

where m is the number of conditional mean equations, and v is the estimated degrees of freedom with $2 < v \leq \infty$. Laurent and Peters (2002) provide details on the log likelihood functions of multivariate GARCH models.

3.2. Reduced-Form Model

Having derived the conditional variance vectors from Diagonal BEKK models and computed the moving covariance of the selected variables, our final step consists of estimating the impact of the volatilities as well as of (selected) covariances of domestic macroeconomic variables on nominal exchange rate volatility for every country in our sample.[9] Financial openness and exchange rate regimes are also included in the model. To assess the impact of every regressor on exchange rate volatility, we apply Granger-causality tests.

Equation (3) below regresses the volatility of nominal exchange rate on the lagged values of the volatilities (σ^2) of nominal exchange rate, output, money supply and inflation, the lags of two covariances (cov): money supply and inflation, and output and inflation, the lagged values of financial openness *(FO)* and intermediate and floating exchange rate regime dummies, *(REG_INTERM, REG_FLOATING)* letting the fixed exchange rate regime be the reference group. We select the specific covariances because Latin American countries exhibited hyperinflation for part of our sample and it is of great interest to examine the impact of the co-movement of inflation with the selected macroeconomic variables on exchange rate volatility. In addition, we do not include a trade openness variable in specification (3), because it is statistically insignificant in all cases. Its exclusion enhances the statistical significance of the rest of the regressors.

$$\begin{aligned}\sigma^2_{XR,t} = c + \sum_{p=1}^{P} \{ & \phi_{1p}\sigma^2_{XR,t-p} + \phi_{2p}\sigma^2_{IPI,t-p} + \phi_{3p}\sigma^2_{M1,t-p} + \phi_{4p}\sigma^2_{INF,t-p} + \phi_{5p}FO_{t-p} \\ & + \phi_{6p}cov(M1, INF)_{t-p} + \phi_{7p}cov(IPI, INF)_{t-p} + \phi_8 REG\_INTERM_t \\ & + \phi_9 REG\_FLOATING_t \} + \varepsilon_t \end{aligned} \tag{3}$$

where ϕ_{ip} is the coefficient of the ith variable for p lags and ε_t is the error term. A regressor in equation (3) does not Granger cause exchange rate volatility when $\phi_{i1} = ... = \phi_{iP} = 0$ where $i = 2,...,7$. To test these hypotheses a Wald test is conducted and we use the heteroskedasticity and correlation consistent standard errors suggested by Newey and West (1987). It is noted that this model is not subject to the generated regressors critique advanced by Pagan (1984), as in equation (4) we include the lagged variance and covariance terms, and not the contemporaneous

[9] For the theoretical justification of the selection of the covariances included in equation (4), see Grydaki and Fountas (2009).

ones. In addition, we include lagged values of financial openness in order to avoid endogeneity as the specific independent variable is correlated with the dependent one by construction.[10]

Prior to our estimation, we undertake the following steps: (i) check for stationarity of the data and discussion of the descriptive statistics, (ii) test for heteroskedasticity, and (iii) specification tests on the estimated BEKK models. These steps, together with the description of our variables and their sources, are presented in the next section.

4. Data and Estimation of Volatilities

4.1. Data

To model exchange rate volatility we compute the volatilities of the following seasonally adjusted variables: (i) Nominal Exchange Rate (XR): We use the official rate, the market rate and the principal rate defined as the number of US dollars per unit of domestic currency for Argentina, Bolivia and Chile, respectively; (ii) Output (Industrial Production Index, IPI): We collect data on crude petroleum production index, crude oil production index and manufacturing production index for Argentina, Bolivia and Chile, respectively, as a proxy of output; (iii) Money supply (M1): Money base has been used as a proxy for money supply for all countries; (iv) Inflation (INF): We use the percentage change of consumer price index (CPI); (v) Trade Openness (TR): We compute the ratio of imports (c.i.f.) and exports (f.o.b.) over IPI as a proxy of trade openness; (vi) Financial Openness (FO): We construct the deviation from Uncovered Interest Parity (UIP), which accounts for country risk premium due to capital controls, international differences in the tax treatment of interest income and risk-averse investors. [11]The mathematical expression for the proxy of financial openness is: $FO = i - i^{*} - \left(\left(XR_t - XR_{t-1} \right) / XR_{t-1} \right)$, where i and i^{*} represent short-term interest rates in the domestic (Latin American) countries and the U.S., respectively; XR_t and XR_{t-1} reflect the

[10] We also account for structural breaks in each economy. These breaks refer to outliers in nominal exchange rate volatility and financial openness. Some of the outliers may be affected by the liberalization of capital accounts. The date of liberalization account for Argentina, Bolivia and Chile is November 1989, January 1990 and January 1992, respectively. De jure liberalization dates of capital accounts for emerging economies can be found in Alper et al. (2007). The liberalization of capital accounts is considered by the measure of financial openness that we introduce (deviation from UIP) and thus an inclusion of a dummy reflecting that liberalization is not necessary.

[11] The same proxy for financial variable is also used in Goh et al. (2006).

nominal exchange rate in current period and previous one, respectively.[12] We use data of money market rate for Argentina and deposit rate for Bolivia and Chile as a proxy for domestic interest rate, while federal funds rate is used for the U.S.; (viii) Regimes (REG): We construct two dummy variables for the three exchange rate regimes. The dummy for the fixed regime takes the value one for the period that countries were operating under fixed exchange rate regime and zero otherwise. In the same way, the other exchange rate regime dummy is constructed.[13]

The corresponding data for all aforementioned variables are retrieved from IMF (2009) except for the data for exchange rate regime classification that are obtained from Levy-Yeyati and Sturzenegger (2005), Bubula and Otker-Robe (2002) and IMF Annual Report Exchange Rate Arrangements and Anchors of Monetary Policy. The three countries in our sample are representative of fixed exchange rate regimes (Argentina), intermediate (Bolivia) and flexible (Chile). The classification of the countries according to their exchange rate regime is found in Corbo (2002). The empirical analysis employs seasonally adjusted monthly data for the period 1979:01-2009:05 and the starting point of the sample is selected according to the availability of the data for the countries under investigation.

4.2. Preliminary Tests

The next step of our analysis involves the examination of the stationarity property in our data. The findings reveal that XR, IPI, M1 and TR are I(1) and INF and FO are I(0) in all countries.[14] Exceptions are INF and TR which turn out to be I(1) and I(0) in Chile and Bolivia, respectively. According to the four unit root tests we use the first difference of the logs (except for FO) of all variables apart from INF for which we use its level. Moreover, all the stationary

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[12] The negative effect corresponds to a positive sign of the coefficient of our measure of financial integration. An increased (decreased) level of capital mobility is associated with smaller (larger) deviations from UIP. The more mobile capital is, the more substitutable financial assets are, and the more difficult it is for a country to set its interest rate independently of world interest rates.

[13] The exact periods of implementation of the three exchange rate regimes in the three Latin American countries can be provided upon request.

[14] We apply the following stationarity tests to the logs of our variables (apart from INF): (i) Generalized Least Squares (GLS)-detrended Dickey-Fuller (DF-GLS) (Elliot et al., 1996), (ii) Ng-Perron (NP) (Ng and Perron, 2001), (iii) Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979) and (iv) Phillips and Perron (PP) (Phillips and Perron, 1988). For tests (i) and (iii) the selection of the number of lags in the test equations is according to the Schwartz Information Criterion (SIC). For tests (ii) and (iv) the lag length has been selected by the Spectral GLS-detrended AR estimator based on SIC and the kernel-based estimator of the frequency zero spectrum, which is based on a weighted sum of the covariances, respectively.The stationarity is tested at 5% significance level and the time trend has been taken into account in the test equation. The unit root test results are not reported for space considerations but can be provided upon request.

series are non-normally distributed according to the descriptive statistics, in all countries. Thus, the estimation of BEKK models assuming multivariate t-distributed errors is suggested. [15]

Having removed the unit root from the variables, what follows is the examination of clustering volatility (the presence of ARCH effects) or any form of heteroskedasticity. To do so, we apply the ARCH Lagrange Multiplier (ARCH-LM) test and the White (1980) test in the cases that ARCH-LM is not sufficient to detect heteroskedasticity.[16] The estimated mean equation for each variable includes a number of lags of the corresponding variable in order for autocorrelation to be corrected. The results are shown in Table 1.

[Table 1 HERE]

In Argentina, IPI and INF appear to support the presence of clustering volatility at 1% and M1 at 5% significance level. With respect to XR, it exhibits heteroskedasticity according to the White test at 5%. In Bolivia, all variables are characterized by ARCH effects at 1% or 5% significance levels. Lastly, all variables in Chile are volatility clustered at 1% significance level, except of the exchange rate which is at 5%.

4.3. Specification Tests and Residual Diagnostics of Diagonal BEKK Models

Our next step proceeds with the estimation of the Diagonal BEKK model as it is described in equations (1) and (2). To account for possible autocorrelation, we use AR terms in all variables in the Diagonal BEKK conditional mean equation considering 1 to 12 lags. The selection of the best model is based on the residual properties, i.e. no remaining autocorrelation and remaining GARCH effects. Taking into account the residual properties of all 36 estimated models, we end up estimating an AR(*p*)-AS-DiagBEKK model, where $p = 9$ for Argentina and $p = 11$ for Bolivia and Chile. Table 2 reports the specification tests for the three AR(*p*)-AS-DiagBEKK models in order to ensure that models fit the data well.

[Table 2 HERE]

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[15] The considered descriptive statistics include the mean, the coefficient of variation, skewness, kurtosis and Jarque-Bera statistics. Their values are not reported for space considerations and can be provided upon request.

[16] The complexity of many economic time series generates non-linear dynamics that make classical econometric techniques to fail (Kyrtsou, 2005, 2008).

The results of Table 2 indicate that in all models the matrices of ARCH and GARCH and asymmetric terms are jointly significant at 1% significance level. The joint significance of the diagonal elements of A_{11}^{*}, B_{11}^{*} and D_{11}^{*} implies that shocks to monetary variables and to output, as well as, to the exchange rate combined with their volatilities, tend to influence with a lag nominal exchange rate volatility. In addition, the joint statistical significance of the asymmetric terms, indicated by the joint significance of the diagonal elements of D_{11}^{*}, implies that the conditional variance of nominal exchange rate and its determinants is affected by "bad" or "good" news. This is the case for the three countries of our sample.

In addition, we test for joint significance of the covariances of the determinants of nominal exchange rate volatility in the three models. We find that in no case the covariances are statistically significant. Therefore, we compute the twelve-period moving covariance, as we are working with monthly data.

The next step is to provide residual diagnostics for the three Diagonal BEKK models in order to test for remaining autocorrelation and ARCH effects. The tests are conducted at 1% significance level and Table 3 reports the results.

[Table 3 HERE]

In order to test for remaining autocorrelation, we calculate the Ljung-Box Q-statistic of the standardized residuals for two lag lengths, 6 and 12. Conducting the test at 1% significance level, we find that in two out of three Diagonal BEKK models (cases of Argentina and Bolivia) none of the variables included exhibits remaining autocorrelation, since the Q-statistic is not statistically significant. In the third one (Chile), inflation appears to exhibit remaining autocorrelation along 12 lags. Moreover, in order to detect any remaining ARCH effects we compute the squared standardized residuals. The Q-statistic is not statistically significant failing to reject the null hypothesis of no remaining ARCH effects along the two lag lengths in Argentina and Bolivia.[17]

[17] Other AR-AS-DiagBEKK specifications up to 18 lags were tried in order to remove remaining autocorrelation and ARCH effects, but without success.

5. Results

5.1 Short-Run Effects

We can now provide the short-run effects on nominal exchange rate volatility for Argentina, Bolivia and Chile. Table 4 reports the short-run effects along 6, 12 and 18 lags (up to one-and-a-half years). We take into account the aforementioned lag lengths because the effects on exchange rate volatility take very short time to materialize, as compared for example with the effects on output volatility.

[Table 4 HERE]

Our results show that there is a negative, but insignificant, effect of output volatility on exchange rate volatility in Argentina and Bolivia (along all lag lengths). The insignificance of this effect is in line with Morana (2009) and Friedman's view (1953). However, output variability is found to have a positive and significant impact on exchange rate volatility in Chile (along all lag lengths). This effect is confirmed by Friedman (1953) where output volatility amplifies exchange rate volatility.

Regarding the effect of money volatility, findings reveal that there is a negative conjunction in Argentina along 12 and 18 lags and in Chile across 18 lags. However, the effect is positive in Argentina along 6 lags and insignificant in Bolivia. The insignificant effect is in line with Amor (2008) while the positive one is confirmed by Calderon (2004). The negative effect is associated with increased interest rates which lead to a decrease in money supply and therefore a decrease in exchange rate volatility (Carrera and Vuletin, 2002).

Continuing with the effect of inflation volatility we detect that there is a positive and significant impact on nominal exchange rate volatility in almost all cases apart from Argentina (18 lags) and Chile (6 and 12 lags). These results are consistent with the aforementioned studies of Morana (2009) and Friedman (1953). According to our findings an increase of inflation volatility inflation leads to a very small increase in exchange rate volatility in Argentina, Bolivia and Chile.

Commenting on the effect of financial openness, we report a negative association (the sign of the corresponding coefficient is positive) with exchange rate variability across all countries and lag lengths. The only exception is the case of Bolivia in which the effect is positive (negative coefficient) along 6 lags at 15% significance level and insignificant along 18 lags. The sign and significance of the effect agrees with Kose et al. (2006) and Devereux and Lane (2003). However, the magnitude of the effect is extremely small in the cases of Argentina and Chile. A possible explanation is that what matters is the degree of capital account liberalization. Since financial openness is associated with the degree of capital account liberalization, a low degree of capital mobility may reflect a small impact of financial openness on exchange rate volatility while a high degree may result in a large impact on nominal exchange rate volatility.

Regarding the covariances of money supply and inflation, and output and inflation, they are found to be overall statistically significant across countries. More specifically, the covariance of money supply and inflation is found to be statistically significant in Chile along 18 lags while in Argentina and Bolivia it is significant along all lag lengths.[18] The effect of the covariance of money supply and inflation is expected to be either positive or negative. The reasoning is that higher inflation causes an increase in interest rates leading to increased capital inflows and appreciated exchange rate (exchange rate decreases). On the other hand, higher money supply pushes interest rates down and the domestic currency depreciates (exchange rate increases). Given the positive relationship between exchange rate and exchange rate volatility, the sign of the effect of the covariance on exchange rate volatility depends on whether the change in inflation or in money supply dominates (Grydaki and Fountas 2009). According to the findings, a negative effect holds in Argentina and Bolivia at 18 lags, while the positive impact is valid in the remaining lags of Argentina and Bolivia while it holds for all lags in Chile.

The same holds for the determination of the sign of the covariance of output and inflation. Regarding the effect of the covariance of output and inflation, we find ambiguous results for Argentina and Chile while we provide a negative impact of the covariance on nominal exchange rate volatility for Bolivia. However, along the significant coefficients in each case there is consistency in sign. The impact is positive in Argentina and Chile along 6 and 18 (at 15% significance level) lags, respectively, while it is negative in Bolivia (6, 12 lags). An increase in inflation is associated with a decrease in output, which in turn increases exports leading to an improvement of trade balance making exchange rate to decrease (exchange rate appreciates). An exogenous increase in output increases imports leading to a deterioration of trade balance, and hence, to increase of exchange rate (exchange rate depreciates). Given the positive relationship

[18] The covariance is significant in Argentina at 15% significance level along 12 lags.

between exchange rate and exchange rate volatility, the sign of the effect of the covariance on exchange rate volatility depends on whether the change in inflation or in output prevails (Grydaki and Fountas 2009).

Finally, regarding the impact of intermediate and floating exchange rate regimes, we detect that under intermediate and floating exchange rate regimes nominal exchange rate volatility tends to increase more than under fixed regime in Argentina and Chile, while in Bolivia such effect is insignificant. The increased exchange rate volatility under intermediate and floating exchange rate regimes is consistent with Flood and Rose (1995) while the insignificant effect of the aforementioned regimes compared to fixed is confirmed by Friedman (1953) who suggests that exchange rate volatility cannot be reduced by switching exchange rate regimes.

5.2. Long-Run Effects

As we find ambiguous short-run effects for at least one determinant of exchange rate volatility (e.g. money supply volatility) for at least one country of the sample (Argentina), we next examine the long-run causal relationships. We follow a variant of the approach of Grier and Smallwood (2007) in order to compute the long-run effects of the volatility of the determinants of exchange rate volatility.[19] These long-run effects are calculated as:

$$Long\ Run\ Effect = \frac{\sum_{p=1}^{P} \phi_{i,p}}{\left(1 - \sum_{p=1}^{P} \phi_{1,p}\right)} \sigma_{u,i}^2 \quad (4)$$

where $\sigma_{u,i}^2$ denotes the unconditional variance of the ith regressor, with $i = 2,...,7$. When we compute the long-run effect of the covariances, we substitute $\sigma_{u,i}^2$ with the unconditional covariance, $cov_{u,ij}$, in equation (4).

In principle, the long-run effect could be mixed. That means that at least one significant lag length is positive and one negative. According to Grier and Smallwood (2007) it turns out to be positive or negative according to the sign of the significant coefficient that corresponds to the longest lag. Table 5 reports the long-run effects of the determinants of nominal exchange rate volatility in Argentina, Bolivia and Chile.

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[19] Grier and Smallwood (2007) use the unconditional standard deviation in formula (4).

[Table 5 HERE]

With the inclusion of covariances, the volatility of output impacts positively on the volatility of nominal exchange rate in the long run for the case of Chile, while this effect is insignificant in Argentina and Bolivia. Regarding monetary variables, money supply volatility impacts negatively on nominal exchange rate volatility in the cases of Argentina and Chile while inflation volatility impinges a positive effect in Bolivia and Chile. The latter effect is negative in Argentina. Lastly, financial openness appears to have a negative effect in Argentina and Bolivia whereas a positive one in Chile.[20] With respect to the long-run effect of the selected covariances we provide evidence in favor of a negative long-run impact of the covariance of money supply and inflation on nominal exchange rate volatility in Argentina and Bolivia. The aforementioned covariance has a positive effect in Chile. Moreover, the covariance of output and inflation affects positively the volatility of nominal exchange rate in all countries.

In comparison with the short-run effects on nominal exchange rate volatility, the insignificant impact of output volatility in Argentina and Bolivia and the positive one in Chile is confirmed both in the short- and long run. The short-run impact of money supply volatility is the same in the long run as well for all countries, while inflation volatility affects positively nominal exchange rate volatility in the three countries in the short run. The aforementioned effect is positive only in Bolivia and Chile, in the long run. Lastly, the short-run impact of financial openness on nominal exchange rate volatility is negative in Argentina and Chile, while ambiguous in Bolivia. The aforementioned effect is the same for Argentina and Bolivia in the long run, while it appears to be positive in Chile. With respect to the covariances, the short-run effect of the covariance of money supply and inflation is ambiguous in Argentina and Bolivia and positive in Chile, while the long-run effect is negative in Argentina and Bolivia and positive in Chile. Moreover, the covariance of output and inflation is found to have a short-run ambiguous affect on nominal exchange rate volatility in Argentina and Chile, whereas a negative one in Bolivia. In the long run, the effect is positive in all cases.

Having detected the effects of the determinants of nominal exchange rate volatility, it is useful to conduct a comparison with the results of related empirical studies. Our findings support a positive and significant effect of output volatility on nominal exchange rate volatility in Chile

[20] The positive effect in Bolivia is justified from the sign of the coefficient with the largest lagged value (Grier and Smallwood, 2007).

while the aforementioned effect is insignificant in Argentina and Bolivia. The positive and significant result agrees with Morana (2009) for a different time period and sample of countries. Moreover, our empirical evidence shows a significant and negative effect of money supply volatility (Flood and Rose, 1995) in Argentina and Chile, while an insignificant effect holds in Bolivia. The latter is in line with Bleaney (2008) for a different time period and sample of countries. Regarding inflation volatility, we find that the latter affects significantly and positively nominal exchange rate volatility across the three Latin American countries of our sample, being in line with Morana (2009). Furthermore, our findings reveal that the greater the financial openness in the countries under investigation is, the lower the nominal exchange rate volatility appears to be (Duarte et al., 2008). With respect to exchange rate regimes, we find that under intermediate and floating exchange rate regimes, nominal exchange rate volatility exacerbates in fixed (Argentina) and floating (Chile) countries (Flood and Rose, 1995; Deviatov and Dodonov, 2005; Stancik, 2006).

6. Conclusions

We examine the determinants of nominal exchange rate volatility focusing on the volatility of several macroeconomic variables over the period 1979-2009 for three Latin American countries, Argentina, Bolivia and Chile. Allowing for the potential impact of the volatilities of the determinants of nominal exchange rate volatility, as well as, the impact of the covariances of these determinants, we estimate (i) an Autoregressive Asymmetric Diagonal BEKK model, in order to capture the volatility of the variables under investigation, and (ii) a reduced-form model, in order to perform Granger causality tests, including the role of financial integration and the implementation of alternative exchange rate regimes.

Our results are summarized as follows: In the short run, volatility in nominal variables reflected by changes in money supply and inflation explains exchange rate volatility in most countries. Output volatility is found to be important, but to a lesser extent. This is in line with the findings of Blanchard and Simon (2001), who argue that monetary policy stabilizes the economy better because of transparent policy goals. However, this inference partially agrees with the perspective of Flood and Rose (1995) that macroeconomic volatility is not considered as an important stabilizer of exchange rates. Financial integration seems to affect quite significantly the volatility of nominal exchange rate in all countries under examination. Finally, the covariances of the selected macroeconomic variables seem to be significant contributors of exchange rate volatility in most cases. In the long run, the picture remains similar, as the nominal shocks play a dominant role.

Important policy implications arise from the above results. The exchange rate stabilization strategy should include the reduction of inflation volatility, increased financial openness, and the correct choice of exchange rate regime. In other words, it seems that fixed exchange rate countries maintain a fixed regime in order to reduce nominal exchange rate volatility, while floating countries is possible to consider the adoption of a fixed regime to reduce exchange rate volatility. For intermediate countries, a switch to an alternative exchange rate regime seems not conducive to curtailing nominal exchange rate volatility. The above implications should be taken with caution since the empirical evidence is based on a small number of countries.

Appendix

Table 1: Heteroskedasticity tests (*LM*-statistic)

	XR	IPI	M1	INF
Argentina				
ARCH LM	7.0248 (12) [0.8560]	64.5620 (12) [0.0000]	23.1028 (12) [0.0269]	102.0651 (12) [0.0000]
WHITE	74.7136 [0.0325]	----	----	----
Bolivia				
ARCH LM	48.1154 (12) [0.0000]	12.4976 (4) [0.0140]	48.4390 (12) [0.0000]	158.7001 (12) [0.000]
Chile				
ARCH LM	17.7969 (8) [0.0228]	32.5300 (12) [0.0011]	27.5643 (12) [0.0064]	119.6169 (12) [0.0000]

Notes: Figures represent the value of LM-statistic (Obs*R^2). The numbers in parentheses and brackets are the lags for ARCH test and probability values, respectively.
---- denote that ARCH LM test is sufficient to justify the existence of heteroskedasticity.

Table 2: Specification tests (Chi-square)

Argentina			
	No GARCH	$H_0: \alpha^*_{11} = \alpha^*_{22} = \alpha^*_{33} = \alpha^*_{44} = \beta^*_{11} = \beta^*_{22} = \beta^*_{33} = \beta^*_{44} = \delta^*_{11} = \delta^*_{22} = \delta^*_{33} = \delta^*_{44} = 0$	57504.19 [0.0000]
AR(9)-Asymmetric	No Asymmetry	$H_0: \delta^*_{11} = \delta^*_{22} = \delta^*_{33} = \delta^*_{44} = 0$	61.9893 [0.0000]
	No Covariance significance	$H_0: c^*_{12} = c^*_{13} = c^*_{14} = c^*_{23} = c^*_{24} = c^*_{34} = 0$	3.2815 [0.7728]
Bolivia			
	No GARCH	$H_0: \alpha^*_{11} = \alpha^*_{22} = \alpha^*_{33} = \alpha^*_{44} = \beta^*_{11} = \beta^*_{22} = \beta^*_{33} = \beta^*_{44} = \delta^*_{11} = \delta^*_{22} = \delta^*_{33} = \delta^*_{44} = 0$	14349.91 [0.0000]
AR(11)-Asymmetric	No Asymmetry	$H_0: \delta^*_{11} = \delta^*_{22} = \delta^*_{33} = \delta^*_{44} = 0$	42.3598 [0.0000]
	No Covariance significance	$H_0: c^*_{12} = c^*_{13} = c^*_{14} = c^*_{23} = c^*_{24} = c^*_{34} = 0$	2.5549 [0.8623]
Chile			
	No GARCH	$H_0: \alpha^*_{11} = \alpha^*_{22} = \alpha^*_{33} = \alpha^*_{44} = \beta^*_{11} = \beta^*_{22} = \beta^*_{33} = \beta^*_{44} = \delta^*_{11} = \delta^*_{22} = \delta^*_{33} = \delta^*_{44} = 0$	10890.26 [0.0000]
AR(11)-Asymmetric	No Asymmetry	$H_0: \delta^*_{11} = \delta^*_{22} = \delta^*_{33} = \delta^*_{44} = 0$	38.8291 [0.0000]
	No Covariance significance	$H_0: c^*_{12} = c^*_{13} = c^*_{14} = c^*_{23} = c^*_{24} = c^*_{34} = 0$	5.9741 [0.4261]

Note: Figures and numbers in brackets reflect the value of Chi-square statistic and the corresponding probability value, respectively.

Table 3: Residual diagnostics

	$\varepsilon_{XR,t}$	$\varepsilon_{IPI,t}$	$\varepsilon_{M1,t}$	$\varepsilon_{INF,t}$
Argentina				
Q(6)	2.297	1.373	1.056	5.743
	[0.890]	[0.967]	[0.983]	[0.453]
$Q^2(6)$	0.216	0.746	0.037	9.231
	[1.000]	[0.993]	[1.000]	[0.161]
Q(12)	6.650	3.980	3.181	23.224
	[0.880]	[0.984]	[0.994]	[0.026]
$Q^2(12)$	0.345	1.104	0.069	22.943
	[1.000]	[1.000]	[1.000]	[0.028]
Bolivia				
Q(6)	1.730	2.741	1.395	14.765
	[0.943]	[0.841]	[0.966]	[0.022]
$Q^2(6)$	0.039	0.592	0.182	1.840
	[1.000]	[0.997]	[1.000]	[0.934]
Q(12)	4.454	6.151	3.425	16.527
	[0.974]	[0.908]	[0.992]	[0.168]
$Q^2(12)$	0.082	1.592	1.314	1.881
	[1.000]	[1.000]	[1.000]	[1.000]
Chile				
Q(6)	10.157	11.680	9.094	7.192
	[0.118]	[0.069]	[0.168]	[0.303]
$Q^2(6)$	4.328	10.534	8.839	17.297
	[0.632]	[0.104]	[0.183]	[0.008]
Q(12)	21.539	20.114	14.709	68.307
	[0.043]	[0.065]	[0.258]	[0.000]
$Q^2(12)$	8.014	17.795	14.803	25.689
	[0.784]	[0.122]	[0.252]	[0.012]

Notes: Probability values are in brackets. *Q(p)* and $Q^2(p)$ are the Ljung-Box test statistic for *p*th order serial correlation for standardized residuals and squared standardized residuals, respectively.

Table 4: Short-run effects on exchange rate volatility (Granger causality tests)

	σ^2_{IPI}	σ^2_{M1}	σ^2_{INF}	FO	$cov(M1, INF)$	$cov(IPI, INF)$	REG_INTERM	$REG\_FLOATING$
Argentina								
	-0.2574	0.0036	0.0000	0.0000	0.0006	0.0037	0.0119	0.0032
6 lags	(0.9040)	(40.5199)	(46.6589)	(67.3849)	(25.2832)	(138.7730)	(2.1503)*	(1.7881)*
	[0.9890]	[0.0000]	[0.0000]	[0.0000]	[0.0003]	[0.0000]	[0.0323]	[0.0748]
	-0.5243	-0.4505	0.0000	0.0000	0.0006	0.0043	0.0100	0.0016
12 lags	(1.7753)	(87.5680)	(37.9456)	(54.6986)	(17.5828)	(6.3617)	(1.7854)*	(0.8504)*
	[0.9997]	[0.0000]	[0.0002]	[0.0000]	[0.1290]	[0.8968]	[0.0754]	[0.3959]
	-0.9790	-0.4170	0.0000	0.0000	-0.0001	-0.0017	0.0138	0.0016
18 lags	(7.6205)	(77.4422)	(18.7720)	(28.9212)	(29.2187)	(12.5405)	(1.4116)*	(0.5881)*
	[0.9838]	[0.0000]	[0.4060]	[0.0494]	[0.0458]	[0.8181]	[0.1596]	[0.5571]
Bolivia								
	-2.6806	0.6311	0.0000	-0.0013	0.0003	-0.0069	0.0145	0.0172
6 lags	(1.8901)	(9.0835)	(18.3438)	(10.0735)	(17.3854)	(11.7140)	(0.7197)*	(0.9842)*
	[0.9295]	[0.1689]	[0.0054]	[0.1216]	[0.0080]	[0.0687]	[0.4723]	[0.3258]
	-9.0467	5.9038	0.0000	0.0228	0.0090	-0.1430	-0.0393	0.0301
12 lags	(2.7680)	(8.2465)	(101.7132)	(19.6334)	(179.2684)	(27.1542)	(-1.1458)*	(0.5751)*
	[0.9970]	[0.7656]	[0.0000]	[0.0743]	[0.0000]	[0.0073]	[0.2530]	[0.5657]
	-3.7797	4.4482	0.0000	0.1528	-0.0116	-0.3684	-0.0516	0.5262
18 lags	(2.4587)	(4.6697)	(43.7015)	(11.2477)	(35.7328)	(7.8659)	(-1.0678)*	(1.1921)*
	[1.0000]	[0.9993]	[0.0006]	[0.8836]	[0.0076]	[0.9806]	[0.2869]	[0.2346]

Table 4 (continued)

	σ^2_{IPI}	σ^2_{M1}	σ^2_{INF}	FO	$cov(M1, INF)$	$cov(IPI, INF)$	REG_INTERM	$REG\_FLOATING$
Chile								
	0.0640	-0.0068	0.0000	0.0000	0.0017	-0.0004	0.0001	0.0001
6 lags	(19.6229)	(4.6829)	(8.6477)	(11.3016)	(5.7933)	(3.2860)	(2.0465)*	(1.9388)*
	[0.0032]	[0.5851]	[0.1944]	[0.0795]	[0.4467]	[0.7722]	[0.0416]	[0.0535]
	0.0385	-0.0207	0.0000	0.0000	0.0029	-0.0008	0.0002	0.0002
12 lags	(32.9017)	(13.1182)	(10.8135)	(20.7988)	(10.9249)	(7.0726)	(2.2998)*	(1.9918)*
	[0.0010]	[0.3605]	[0.5450]	[0.0534]	[0.5354]	[0.8528]	[0.0223]	[0.0475]
	0.0138	-0.0129	0.0000	0.0000	0.0043	0.0033	0.0002	0.0002
18 lags	(30.4419)	(26.5713)	(28.0903)	(26.0852)	(24.3156)	(24.2918)	(1.5725)*	(1.3464)*
	[0.0334]	[0.0874]	[0.0607]	[0.0978]	[0.1450]	[0.1457]	[0.1174]	[0.1797]

Notes: Figures are the sum of the lagged coefficients of the causing variable. Figures in parentheses and brackets represent the value of Chi-square statistic and the corresponding probability value, respectively. * denotes that figures in parentheses and brackets reflect the value of t-statistic and the corresponding probability value, respectively.

Table 5: Long-run effects on exchange rate volatility

	Argentina	**Bolivia**	**Chile**
σ^2_{IPI}	insignificant	insignificant	positive
σ^2_{M1}	negative*	insignificant	negative
σ^2_{INF}	negative	positive	positive
FO	negative	negative*	positive
$cov(M1, INF)$	negative*	negative*	positive
$cov(IPI, INF)$	positive	positive	positive

Note: * indicates that the result is in principle mixed.

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