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Real Exchange Rate Volatility and Exports: A Study for Four Selected Commodity Exporting Countries

Summary: Commodity exports depend on global demand and prices, but the increasing volatility of real exchange rates (RER) introduces an additional factor. Thus, this paper studies the RER volatility dynamics, estimated through GARCH and IGARCH models for Brazil, Chile, New Zealand, and Uruguay from 1990 to 2013. We study the impact of RER volatility on total exports using Johansen's methodology, including proxies for global demand and international prices. The results suggest that exports depend positively on global demand and international prices for all countries; however, conditional RER volatility resulted significant and negative only for Uruguay, in the short- and long-run.

Key words: Exports, Real exchange rate, GARCH, Co-integration.

JEL: C55, F31, F41.

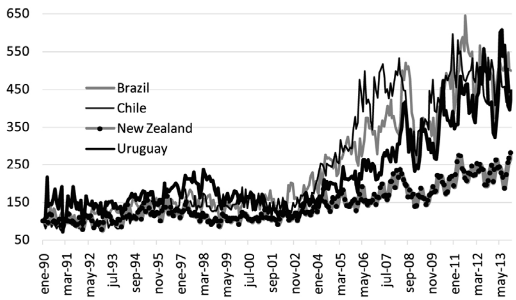
Raw materials exports depend on global demand and prices, but the increasing volatility of real exchange rates (RER) introduces an influence whose impact varies according to the situation and the country. The main argument is as follows: greater RER volatility leads to higher costs for risk-averse traders which implies less foreign trade. This is because the exchange rate is agreed at the time of the commercial contract, but the payment is not made until delivery actually takes place. If changes in exchange rates become unpredictable, this creates uncertainty about the benefits and, therefore, reduces international benefits from trade. Even if hedging in the forward markets were possible, there are limits and costs (Michael D. McKenzie 1999). On the other hand, if exporters are risk takers enough, increased exchange rate volatility raises expected marginal utility of export earnings inducing them to increase exports (Paul De Grauwe 1988).

This paper seeks to estimate the impact of RER volatility on exports for Brazil, Chile, New Zealand, and Uruguay, selected as commodity exporting countries. Due to the heteroscedasticity of RER volatility (Robert F. Engle 1982), we use a generalized autoregressive conditional heteroskedasticity (GARCH) model, according to Tim Bollerslev (1986), or integrated GARCH (IGARCH). Then we study for each country the possible impact of estimated RER volatility on exports, using Søren Johansen (1988, 1992) methodology. The reporting period goes from January 1990 to December 2013.

The main contributions of this paper are first, robust evidence about the connection between RER uncertainty and exports both in the short- and long-run, for four different commodities exporting economies. Second, this analysis is performed for a lengthy data set, which includes several economic events for each country. Third, this work contributes to the empirical literature discussion of the impact of RER volatility over exports. Finally, for the case of Uruguay, this work provides original export demand estimation, including RER volatility modeled as a GARCH process. The article includes in Section 1 an exports characterization, in Section 2 a survey of the main background on the issue, in Section 3 we discuss the methodology, in Section 4 we define data sources, in Section 5 we report the main results and final remarks are in Section 6.

1. Exports Characterization

Figure 1 shows Brazil, Chile, New Zealand, and Uruguay total exports throughout the period 1990-2013. Until early 2000, exports are stagnated, and in the case of Uruguay and New Zealand, they fell since the late 1990s. Since 2003, exports grew for all countries, although all recorded a fall in 2008-2009 due to the international crisis. Particular events in the export series of the different countries are shown in Appendix C.



Source: Bureau of Labor Statistics (BLS 2015)[1], Centro de Economía Internacional (CEI 2015)[2], and International Monetary Fund (IMF 2015a)[3].

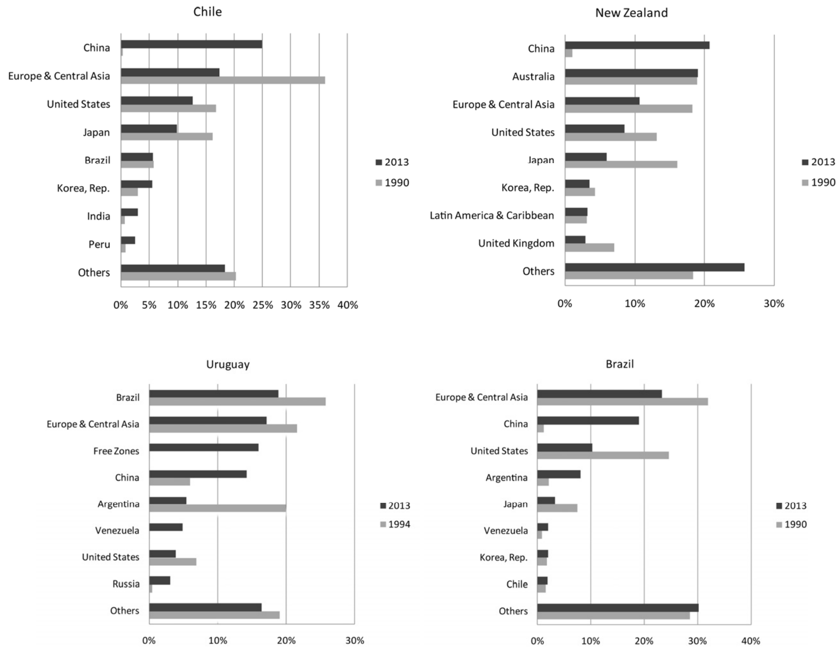
Figure 1 Exports in Constant Dollars 1990-2013 (Jan-90 = 100)

In Figure 2, we can see the main export destinations selected by countries. We compare the year 1990 to 2013.

One of the most relevant features of all selected countries is the large share of exports to China at the end of the period, despite during the first years, sales to this

[1] **Bureau of Labor Statistics (BLS).** 2015. Archived Consumer Price Index Supplemental Files. http://www.bls.gov/cpi/tables/supplemental-files/home.htm (accessed August 19, 2015).
[2] **Centro de Economía Internacional (CEI).** 2015. Comercio Exterior - Total. http://www.cei.gob.ar/es/comercio-exterior-total (accessed August 19, 2015).
[3] **International Monetary Fund (IMF).** 2015a. Direction of Trade Statistics (DOTS). http://data.imf.org/?sk=9D6028D4-F14A-464C-A2F2-59B2CD424B85 (accessed August 19, 2015).

country were virtually non-existent. On the other hand, we can see a decline in the share of exports to Europe and the United States (US) during the period. In the case of Uruguay, in addition to the high participation of China in its exports, the share loss of sales to its big neighbors (Argentina and Brazil) is highlighted. It is also interesting to notice the importance acquired by the Free Trade Zones (FTZ), which in 2013 were placed in the third position as a destination of Uruguay's exports but had no participation in 1990 (also see Table A1 in Appendix A).



Notes: World Bank data for Uruguay is from 1994.

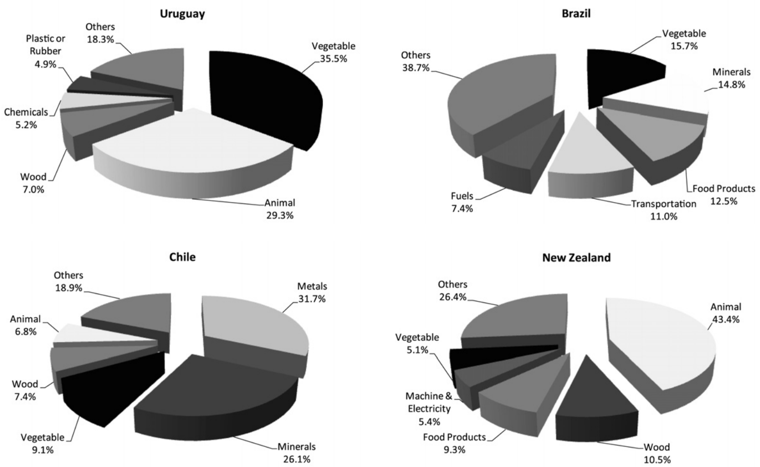
Source: Own calculations based on World Bank (2015)[4].

Figure 2 Exports Destinations by Countries

According to the classification used by the World Bank as the main type of products exported in 2013, the case of Uruguay are mainly concentrated in the category of "vegetable" (soybeans and cereals) and "animal" (beef and dairy), representing about 65% of total exports. Meanwhile, Brazil has a more diversified distribution as vegetable exports, which represented 15.7% (soybeans, corn, and sugar), minerals (iron) contributed 14.8%, food products 12.5%, transport (cars and boats) 11.0% and fuels (oil) 7.4%, representing approximately 62% of total exports. Chile's main exports are metals (31.7%) and minerals (26.1%), particularly copper and its by-products, reflecting the importance of the mining sector exports, which represent more than 50%.

[4] **World Bank.** 2015. World Integrated Trade Solution (WITS). http://wits.worldbank.org/countrystats.aspx?lang=es (accessed July 09, 2015).

Finally, for New Zealand, the most significant share in exports is animals (43.4%), mainly frozen meat and dairy products. Below, we present in Figure 3 the exports composition by country.



Source: Own calculations based on World Bank (2015).

Figure 3 Exports by Country

2. Background

The adoption of floating exchange rate regimes since 1973 has increased the importance of the studies associated with exchange rate volatility influence on international trade, both nominal and real exchange rate. The theoretical and empirical literature is inconclusive regarding the effects of such an impact. The evidence shows positive, negative, and neutral results, a combination of the previous three, and no significant results (Ilhan Ozturk 2006; Bruno Ćorić and Geoff Pugh 2010). This can be a consequence of methodological differences in terms of the number of countries considered, the specification of the exchange rate volatility used, or the sample periods (Ozturk 2006).

Peter Hooper and Steven W. Kohlhagen (1978) developed one of the first works that explores the relationship between nominal exchange rate volatility (measured by standard deviation (S. D.)) and trade. The study is for developed countries and covers the period from the mid-1960s to the mid-1970s. The results shed no significant evidence of the sign of the impact. From this work, David O. Cushman (1983) advances in a similar line, but analyzing instead real exchange rate volatility on trade. He finds that an unexpected movement in the RER has a significant and negative effect on trade. M. A. Akhtar and R. Spence Hilton (1984) also found a negative correlation, but unlike previous studies, they use the standard deviation of effective exchange rate as a measure of volatility. The study considered bilateral trade between the United States and Germany in the period 1974-1981. Similar results were obtained by Abdur R.

Chowdhury (1993) who found a negative impact of exchange rate volatility on the exports volume to the G-7 countries for the period 1973-1990. They built a volatility temporary variable through a moving average of standard deviation of real exchange rate growth rate.

For Asian countries, the evidence of exchange rate volatility impact on exports points to the predominance of adverse effects (Saang Joon Baak, M. A. Al-Mahmood, and Souksavanh Vixathep 2007; Tajul Arrifin Masron and Mohd Naseem Niaz Ahmad 2009; Myint Moe Chit, Marian Rizov, and Dirk Willenbockel 2010; Norimah Ramli and Jan M. Podivinsky 2011; Yin-Wong Cheung and Rajeswari Sengupta 2013). Cheung and Sengupta (2013) studied the effect of RER and RER volatility in export shares of non-financial Indian companies for the period 2000-2010. The empirical analysis shows that there has been a significant negative impact of exchange rate volatility in exports of Indian companies. Moreover, Baak, Al-Mahmood, and Vixathep (2007) found negative impact of exchange rate volatility (measured by the standard deviation of RER) on exports to four East Asian countries (Hong Kong, South Korea, Singapore, and Thailand) and its bilateral trade with Japan and the United States for the period 1990-2001. Similarly, Ramli and Podivinsky (2011) conducted a study, but unlike the previous ones, they first consider five countries of the Association of South East Asian Nations (ASEAN) (Malaysia, Singapore, Philippines, Indonesia, and Thailand) and its bilateral trade with the United States for the period 1990-2010; and then they consider the RER volatility estimated through a GARCH(1,1) process. As a result, they find that the volatility of the bilateral real exchange rate has mainly a significant negative impact on exports, except for the case of Indonesia, which is positive.

Chit, Rizov, and Willenbockel (2010), unlike previous studies, consider a greater number of countries, taking into account bilateral trade in some East Asian countries (China, Indonesia, Malaysia, Philippines, and Thailand), as well as 13 industrial countries. They use a panel with information for the period 1982-2006, specifying three measures of volatility for the real exchange rate (standard deviation, the moving average of the standard deviation and conditional volatility – GARCH). They get that regardless of the proxy used as exchange rate volatility, the impact generates a negative effect on exports from emerging countries of East Asia. Meanwhile, Masron and Niaz Ahmad (2009) noted that the exchange rate volatility (GARCH(1,1)) has a negative effect on exports demand from Malaysia and Turkey for the period 1970-2004. However, when studies incorporate a variable reflecting regional economic integration (Malaysia with the ASEAN and Turkey with the European Union (EU)), the negative impact becomes insignificant for the case of Turkey. Moreover, Zukarnain Zakaria (2013) points out that the effect is ambiguous for Malaysia, and this is explained on the basis that the exchange rate volatility has a negative effect on bilateral trade with the US, while with Japan, it is positive. Malaysia's exports to the United Kingdom (UK) and Singapore show no evidence of any connection with exchange rate volatility. The period considered was 2002-2012, and exchange rate volatility is modeled through a GARCH(1,1). Jianbin Situ (2015) examined the effects of real exchange rate volatility on exports from developed countries and export-oriented less-developed countries (LDCs) to the US for the period 1994-2014. This empirical investigation found a significant and negative impact of real exchange rate volatility on exports from developed countries over all the period. However, the effects are less clear for LDCs in the period.

Empirical evidence for Latin America can be found in Antonio Aguirre, Afonso Ferreira, and Hilton Notini (2007), Daniel Berrettoni and Sebastián Castresana (2007), Aldo Otero Adamo and Mario Silva Arteta (2008), and Mohsen Bahmani-Oskooee, Hanafiah Harvey, and Scott W. Hegerty (2013). They study the exchange rate fluctuations impact on manufacturing exports. The first explores the case of Peru over the period 1994-2004; the second studies Brazil from 1986 to 2002; the third examine Brazil from 1971-2010; and the fourth analyzes Argentina from 1992 to 2006. To model exchange rate volatility, they use the standard deviation; additionally, Aguirre, Ferreira, and Notini (2007) and Adamo and Arteta (2008) specify a model with the conditional variance (GARCH). Additionally, Augustine C. Arize, Thomas Osang, and Daniel J. Slottje (2008) study exchange rate volatility impact on global exports of eight countries in Latin America: Bolivia, Colombia, Costa Rica, Dominican Republic, Ecuador, Honduras, Peru, and Venezuela, for the period 1973-2004. They use an ARCH(1) specification to model the exchange rate volatility, with the exception of Honduras for which they use a linear moment (LM) model. With the exception of the empirical analysis of Aguirre, Ferreira, and Notini (2007) and Bahmani-Oskooee, Harvey, and Hegerty (2013) that find conflicting results, they all get similar results: a significant and negative effect of exchange rate volatility on exports. Marilyne Huchet-Bourdon and Jane Korinek (2012) analyzed the impact of exchange rates and their volatility on bilateral trade between Chile and New Zealand (small and open economies) with China, Euro Area and the US (large partners) for two sectors: agriculture and nonmanufacturing-mining. They found that exchange rate volatility (moving standard deviation and GARCH model) has a greater impact on trade flows of the small and open economies than those of larger economies. However, findings do not clearly indicate the direction of the impact (increases or decreases) in all countries and sectors considered.

There are also a number of works that find positive effects of exchange rate volatility on exports. Don Bredin, Stilianos Fountas, and Eithne Murphy (2003) studied the impact of exchange rate volatility, both short- and long-term, global and sectoral, on Irish exports to the European Union for the period 1979 to 1992 (Irish national and multinational companies sectors). RER volatility is modeled *via* mobile standard deviation of the growth rate of RER. In the short-term, they found that volatility has a negative effect on multinational firms' exports and has no effect on national firms, generating a negative impact on overall exports. In the long-term, the exchange rate volatility has no effect on multinational firms' exports, but they found a positive effect on exports of domestic firms, and therefore, the effect on global exports is positive. Similarly, Ahmed A. A. Asseery and David A. Peel (1991) analyze five developed countries and also found a positive effect of exchange rate volatility on exports over the period 1972-1987.

Christopher F. Baum, Mustafa Caglayan, and Neslihan Ozkan (2004) found that on average, the effects of exchange rate volatility on exports is positive for a sample of 13 developed countries in the period 1980-1998. The originality of his analysis is to model the exchange rate volatility from daily frequency by an AR(2) process. Additionally, Baum and Caglayan (2007) analyzed the effect of exchange rate volatility (GARCH specification) on bilateral trade and fluctuations in trade flows for a group of developed countries in the period 1980-1998. They found that the volatility of the exchange rate has both positive and negative impacts on bilateral trade; however, the effect is mainly positive with respect to fluctuations of trade.

The literature also registers the existence of no significant impact of exchange rate volatility on global exports. Dimitrios Serenis and Paul Serenis (2008, 2010) pointed out that overall, exchange rate volatility has no effects on aggregate levels of exports by analyzing the case of four European countries: Norway, Poland, Hungary, and Switzerland for the period 1973-2004 (Serenis and Serenis 2010) and for the period 1973-2006 (Serenis and Serenis 2008). Jamal Bouoiyour and Refk Selmi (2014a) used several econometric methods: ordinary least squares (OLS), instrumental variables, autoregressive distributed lag (ARDL), spectral analysis, the evolution and decomposition of wavelet to estimate the relation between the volatility of the exchange rate and exports for Tunisia. Regarding exchange rate volatility measures, they used the moving average of standard deviation and GARCH model. The overall result is a significant negative effect in the short-term, but no significance in the long-term. Dimitrios Asteriou, Kaan Masatci, and Keith Pilbeam (2016) also studied the impact of nominal and real exchange rate volatility on exports and imports in the long-run. They analyzed the cases of Mexico, Indonesia, Nigeria, and Turkey for the period 1995-2012 and use the GARCH specification to model the exchange rate volatility. The empirical analysis shows similar results as in Bouoiyour and Selmi (2014a) in the long-run, except for Turkey, in which case volatility has a small significant impact on international trade.

A set of meta-analysis present an extensive study regarding the effects of exchange rate volatility on international trade, including Ozturk (2006), Ćorić and Pugh (2010), and Bouoiyour and Selmi (2014b). Most of them show a negative effect of exchange rate volatility on international trade. Ozturk (2006) reviews 42 documents that go from 1978 to 2005. Ćorić and Pugh (2010) consider a set of 58 studies published from 1978 to 2003, and notes that the dummy variable representing the exchange rate regime was regularly found significant. Finally, Bouoiyour and Selmi (2014a) analyzed 59 publications from 1984 to 2014. The evidence regarding the impact is 29 (negative), 6 (positive), 6 (not significant), and 18 (ambiguous).

We found no works linking RER volatility and exports in Uruguay; however, Fernando Lorenzo, Nelson Noya, and Christian Daude (2000) study the Uruguay bilateral RER volatility (with Argentina and Brazil) to explain the evolution of real exchange rate. They found that bilateral RER presents conditional heteroscedasticity and can be modeled through a GARCH.

3. Methodology

3.1 Theoretical Model Specification

This section presents the main methodological aspects considered in the models, following similar steps as Kosta Josifidis, Jean-Pierre Allegret, and Emilija Beker-Pucar

(2009). First, we made a univariate analysis of the series analyzing the presence of unit roots through the Augmented Dickey-Fuller (ADF) test, in order to determine the order of integration of each series.

Subsequently, taking into account the integration order of the series, applying Johansen methodology, we made a multivariate analysis to try to capture the effects of short- and long-run relationships between exports and the considered determinants. For this, we inquired about the existence of co-integration relationships in the case of countries with all its series integrated of first order, I(1).

Following Walter Enders (1995), co-integration analysis is based on autoregressive vector with a vector error correction (VEC) specified in a model of endogenous variables. The VEC modeling can be represented as:

$$\Delta X_{it} = A_1 \Delta X_{it-1} + \cdots + A_k \Delta X_{it-k+1} + \prod X_{it-k} + \mu + \Gamma D_t + \varepsilon_t, t = 1, \ldots, T, \quad (1)$$

where Δ denotes difference variables, X is a vector of endogenous variables, μ is a vector of constants, D_t contains a set of instrumental variables (dummies) seasonal and interventions, and ε_t is the error term and is distributed $N(0, \sigma^2)$. The dummy variables were included to correct outliers in the different series and resulted additive outliers (AO) or transitory changes (TC) to better fit the model (Appendix C).

Information on long-term relationships is included in the $\prod = \alpha\beta'$ matrix, where β is the vector of coefficients for the existing equilibrium relations and α is the coefficients' vector of the adjustment mechanism in the short-term. Identifying the range of the $\prod$ matrix determines the total co-integrating relationships between the variables.

Then, we performed the exclusion tests for β coefficients, and as suggested in Erdal Demirhan and Banu Demirhan (2015) or Asseery and Peel (1991), between others, the variables with "t" not significant were omitted in the final models, for each country.

Having examined the long-term relationship, we proceed to the analysis of the short-term, showing the adjustment mechanism of the variables in the short-run to the long-term equilibrium relationship. In this case, we also performed exclusion tests for α coefficients, excluding them from the short-run equation, when the tests show no significant values.

Afterwards, we calculated impulse-response functions (IRF) in the vector error correction model (VECM), which traces the effect of a one-time shock to one of the innovations on current and future values of the endogenous variables, to determine and compare the magnitude, significance, and sign (always variables are considered in the first difference). IRF are applied to analyze the dynamic influence effects of different macroeconomics variables on real export within a period of 24 months.

Finally, we perform Granger causality tests (Wald test) between groups of variables included in the different models, as correlations do not necessarily imply causation in any meaningful sense of that word. In this case, the null hypothesis is that one variable does not cause, in the Granger sense, the other variable. Also, we included long-run causality test (t-test) through the statistical significance of the coefficients of the previous period of the error correction term (ECT) (in levels). It represents the deviation of the dependent variables from the long-run equilibrium.

3.2 Real Exchange Rate Volatility Measurement

In order to analyze the RER impact and its volatility on exports, we estimated co-integration models for each country exports measured in constant dollars (deflated by the US consumer price inflation (CPI)), the overall RER with the major trading partners in each country, its volatility estimated through a GARCH model and world demand, estimated by global imports measured also in constant dollars.

Regarding the literature on how to measure the RER volatility, there is no consensus; therefore, there are multiple approaches in the literature. Among the most common specifications can be found the standard deviation, moving average of the standard deviation, and the conditional variance specified by the squared residuals of ARIMA (Autoregressive Conditional Heteroscedasticity (ARCH) processes, Engle 1982; Generalized ARCH (GARCH) Bollerslev 1986 or some variant GARCH).

In this work, we consider the RER conditional volatility as a measure of uncertainty, and we estimate it through a GARCH process variant, called IGARCH introduced by Engle and Bollerslev (1986). The GARCH model has been widely used to model volatility in the literature on time series models. Overall, the GARCH model for the exchange rate can be represented as follows:

$$y_t = \delta_0 + \sum_{i=1}^{k} \delta_i . y_{t-i} + \varepsilon_t;\ \varepsilon_t \sim N(0, \sigma_t^2), \tag{2}$$

$$GARCH(p,q): \sigma_t^2 = \alpha_0 + \sum_{i=1}^{q} \alpha_i . \varepsilon_{t-i}^2 + \sum_{i=1}^{p} \beta_i . \sigma_{t-i}^2. \tag{3}$$

Equation (2) represents a process autoregressive (AR) of order k, AR(k), where y_t is the real exchange rate, expressed in logarithm; the parameter δ_0 is the constant, k is the number of delays; and ε_t is the heteroscedastic term of error of the conditional variance (σ_t^2).

Equation (3) specifies a conditional variance $GARCH(p,q)$, where q is the number of ARCH terms, p is the number of GARCH terms. The conditional variance is represented by three terms: (i) the average, α_0; (ii) the ARCH term, which measures the previous period volatility by squared residuals delays in the first equation, ε_{t-i}^2; and (iii) the GARCH term, which captures the previous error variance prediction, σ_{t-i}^2. As the conditional variance is positive, it is required that the parameters α_0, α_i and β_i to be ≥ 0; and further that $\sum \alpha_i + \sum \beta_i < 1$ to ensure the process to be stationary in covariance.

Moreover, in the literature regarding the empirical relationship between exports and exchange rate volatility, GARCH(1,1) is widely used for its significant results. However, in some applications, it has been found that the estimates of $\hat{\alpha}_1$ and $\hat{\beta}_1$ tend to approach $\hat{\alpha}_1 + \hat{\beta}_1 = 1$, indicating that the GARCH(1,1) process is no longer stationary. That is why it is more appropriate to specify a regressive process IGARCH(1,1) to model the conditional variance, whose expression has the following form:

$$IGARCH(1,1): \sigma_t^2 = \alpha_0 + \sigma_{t-1}^2 + \alpha_1(\varepsilon_{t-1}^2 - \sigma_{t-1}^2). \tag{4}$$

The peculiarity of the regressive process $IGARCH(1,1)$ is the presence of a unit root, that is, the process is I(1), indicating the persistence of the conditional variance over time (Bollerslev, Engle, and Daniel B. Nelson 1994).

4. Definitions and Data Sources

In this paper, we consider four countries: Brazil, Chile, New Zealand, and Uruguay. The criteria for selecting these countries responds, in first place, to the fact that they are all commodity exporting countries; in second place, Uruguay, Chile, and New Zealand are small and open economies – small domestic markets that depend on international prices – in contrast to Brazil, a large and relatively closed economy, included to contrast both kinds of economies; and in the third place, to show the results for some Latin American countries against a similar country elsewhere in the world, we include New Zealand.

The starting point considered for this analysis is 1990 when Latin American economies return to growth, after the 1980s, "the lost decade". At the beginning of the 1980s, many Latin American countries faced a profound financial crisis that was only solved at the end of the decade with the implementation of external opening measures and the return of international capitals, so the 1990s was a period of growth mainly led by exports (Carlos Quenan and Sébastien Velut 2014).

The series used are total good exports, world imports, international prices of countries' most important raw materials exported and real exchange rate (used also to build the real exchange rate volatility). In all cases, it is considered the monthly frequency of the series for the period 1990-2013 (288 observations) and log transformations.

First, as a world demand proxy, we use world imports, measured in constant dollars, deflated by the US CPI. World imports data are from the IMF (2015a), and the US CPI data are from the US Bureau of Economic Analysis (BLS 2015).

Secondly, export data correspond to total good exports in constant dollars, deflated by the US CPI, exports data are provided by the CEI (2015) and the IMF (2015a).

The Brazilian real exchange rate is from the Instituto de Pesquisa Econômica Aplicada (IPEA 2015)[5], and volatility is our estimation through GARCH(1,1) methodology. Chile's real exchange rate is from the United Nations Economic Commission for Latin America and the Caribbean (UNECLAC 2015)[6], and we estimated its volatility through an IGARCH(1,1). For New Zealand, the real exchange rate is from the Reserve Bank of New Zealand (2014)[7], and volatility is calculated using an IGARCH(1,1). Finally, for Uruguay, the real exchange rate is our own calculations, using retail prices (CPI) and official exchange rates of nine major trading partners. The volatility of the real exchange rate was estimated from an IGARCH(1,1) process.

Finally, a proxy of the most important export prices for each country was included. In the case of Brazil, New Zealand, and Uruguay, we used the food price index

[5] **Instituto de Pesquisa Econômica Aplicada (IPEA).** 2015. http://www.ipeadata.gov.br/Default.aspx (accessed March 13, 2015).

[6] **United Nations Economic Commission for Latin America and the Caribbean (UNECLAC).** 2015. CEPALSTAT Databases and Statistical Publications. http://estadisticas.cepal.org/cepalstat/WEB_CEPALSTAT/estadisticasIndicadores.asp?idioma=i (accessed March 11, 2015).

[7] **Reserve Bank of New Zealand.** 2014. Exchange Rates and TWI - B1. http://www.rbnz.govt.nz/statistics/b1 (accessed December 28, 2014).

compiled by the IMF (2015b)[8] and, in the case of Chile, the metal's price index compiled by the Central Bank of Chile (2015)[9]. This last one was included, because in Chile, more than 50% of exports are metals.

4.1 Series Analysis

First, we analyzed the series stationarity through the Augmented Dickey-Fuller test. We can see the results in Table 1 and Table 2 below.

Table 1 Unit Root Test

Unit root test - ADF
H_0 = There is a unit root

	Statistical value of the series in levels	Reject H_0 at 95%	Statistical value of the series in first differences	Reject H_0 at 95%
World imports (LM)	2.5865	No	-5.0464	Yes
International food prices index (LPR)	0.4090	No	-11.3406	Yes
International metals price index ($LPRM$)	0.5116	No	-11.9865	Yes

Notes: The numbers of lags was determined according to the Akaike criterion.

Source: Authors' estimations.

From the information provided in Table 1, we conclude that all the variables are first-order integrated, I(1).

Table 2 Unit Root Test

Unit root test - ADF
H_0 = There is a unit root

	Statistical value of the series in levels	Reject H_0 at 95%	Statistical value of the series in first differences	Reject H_0 at 95%
Uruguay				
Exports (X_URU)	1.8648	No	-4.8120	Yes
RER (RER_URU)	-1.8584	No	-8.5237	Yes
RER volatility ($RERV\_URU$)	-0.6302	No	-15.4245	Yes
Brazil				
Exports (X_BRA)	1.8675	No	-3.9901	Yes
RER (RER_BRA)	0.5596	No	-11.8911	Yes
RER volatility ($RERV\_BRA$)	-5.4063	Yes	-9.5735	Yes
Chile				
Exports (X_CHI)	1.2434	No	-3.4074	Yes
RER (RER_CHI)	-2.6344	No	-10.2483	Yes
RER volatility ($RERV\_CHI$)	-0.4942	No	-5.8877	Yes
New Zealand				
Exports (X_NZEL)	1.2617	No	-4.1282	Yes
RER (RER_NZEL)	0.4558	No	-7.4189	Yes
RER volatility ($RERV\_NZEL$)	0.4459	No	-7.4459	Yes

Notes: The numbers of lags were determined according to the Akaike criterion.

Source: Authors' estimations.

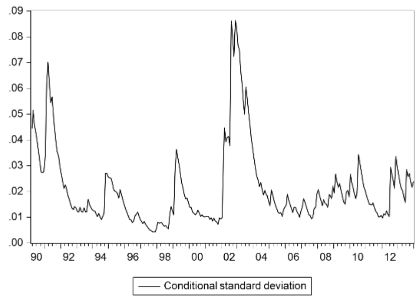
[8] **International Monetary Fund (IMF).** 2015b. IMF Primary Commodity Prices. http://www.imf.org/en/Research/commodity-prices (accessed April 23, 2015).
[9] **Central Bank of Chile.** 2015. Prices Statistics. http://si3.bcentral.cl/Siete/secure/cuadros/arboles.aspx (accessed April 28, 2015).

In the case of Uruguay, New Zealand, and Chile, all the series studied were not stationary, integrated of first order, I(1). As a result, the study was done *via* the Johansen (1988) methodology, trying to find a long-term relationship through a co-integrating vector, estimating an VEC model (Engle and Clive W. J. Granger 1987; Johansen 1992).

In the case of Brazil, the exports series and the real exchange rate were non-stationary and integrated of first order, although the exchange rate volatility was stationary, so it was included as exogenous in the model as Kyriacos Aristotelous (2001). Nevertheless, RER volatility was not significant, so it was excluded from the model.

4.2 Conditional Heteroscedasticity Models

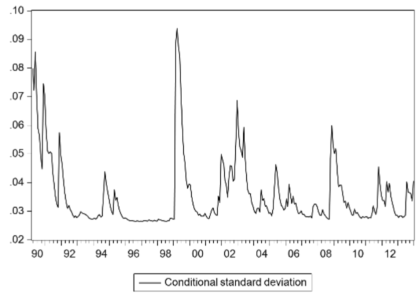
We found that the best $GARCH(p,q)$ model for Uruguay, Chile, and New Zealand was an $IGARCH(1{,}1)$ and a $GARCH(1{,}1)$ for Brazil. These models were estimated through maximum likelihood method, assuming Gaussian innovations. The results for these models, the plots and equation of the conditional volatility estimations over the sample period are in Figures 4.1 to 4.4.



Source: Authors' estimations.

Figure 4.1 Estimated Conditional Volatility - Uruguayan Model

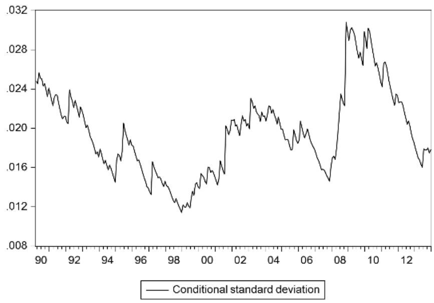
$$\underset{}{GARCH_t} = \underset{(0.0220)}{0.2273} * RESID_{t-1}^2 + \underset{(0.0220)}{(1 - 0.2273)} * GARCH_{t-1}. \tag{5}$$



Source: Authors' estimations.

Figure 4.2 Estimated Conditional Volatility - Brazilian Model

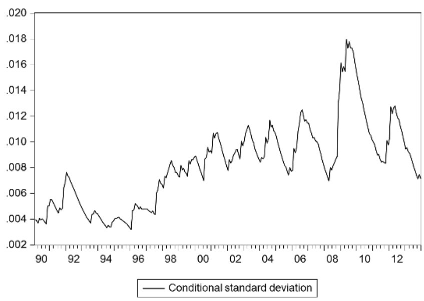
$$GARCH_t = \underset{(0.0000)}{0.0003} + \underset{(0.0575)}{0.1692} * RESID^2_{t-1} + \underset{(0.0658)}{0.6317} * GARCH_{t-1}. \quad (6)$$



Source: Authors' estimations.

Figure 4.3 Estimated Conditional Volatility - Chilean Model

$$GARCH_t = \underset{(0.0127)}{0.0615} * RESID^2_{t-1} + (1 - \underset{(0.0127)}{0.0615}) * GARCH_{t-1}. \quad (7)$$



Source: Authors' estimations.

Figure 4.4 Estimated Conditional Volatility - New Zealand Model

$$GARCH_t = \underset{(0.158)}{0.0885} * RESID^2_{t-1} + (1 - \underset{(0.158)}{0.0885}) * GARCH_{t-1}. \quad (8)$$

The Uruguayan chart shows that the exchange rate volatility was moderately high and persistent through the 1990-2013 period. There was a significant increase in the exchange rate volatility since 2002 as a consequence of the Uruguayan crisis and its devaluation impact.

Brazilian exchange rate volatility was also relatively high from 1990 to 1992, but decreased between 1994 and 1998, with the "Plan Real". Since 1999, the Russian moratorium had a negative effect on real exchange rate volatility, which returned to higher levels with the different international crisis.

As illustrated by the Chilean and New Zealand charts, exchange rate volatility in both countries was quite low and persistent throughout the 1990-2013 period.

Although in both cases, exchange rate volatility had an important growth since 2008 as result of the international financial crisis. Comparing the two country pairs, in the case of Uruguay and Brazil, volatility levels were considerably higher than Chile and New Zealand.

5. Main Results

Table 3 summarizes the main results of the Johansen co-integration test for Brazil, Chile, New Zealand, and Uruguay.

Table 3 Johansen Co-Integration Test Results

Country H_0	$r = 0$	$r \leq 1$	$r \leq 2$
Trace test			
Brazil	34.9103**	5.4568	0.1300
Chile	53.2480**	20.2940	7.6023
New Zealand	57.9399**	27.7703	9.7670
Uruguay	63.8075***	26.9797	6.6223
Maximum eigenvalue test			
Brazil	29.4535***	5.4568	0.1300
Chile	32.9540***	12.6917	6.9466
New Zealand	30.1696**	18.0033	6.8549
Uruguay	36.8279***	20.3574	5.5397

Notes: r represents the number of co-integrating vectors. The significance level for rejecting H_0 at * 10%, ** 5%, and *** 1%.

Source: Authors' estimations.

Since co-integration tests indicate a long-term relationship for each country exports equation, we estimated a VECM for each country, the coefficients t-statistics are shown in parentheses below Equations (9)-(12). The residual diagnostic tests for each model are summarized in Appendix B.

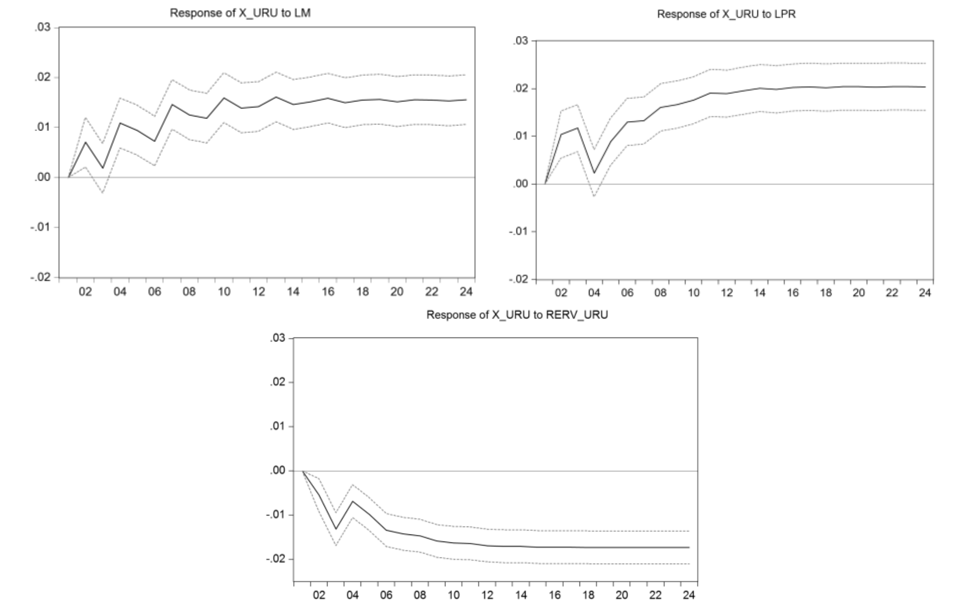
In the case of Uruguay, the final adjustment was for the period January 1993 to December 2013, because the Uruguayan economy had strong adjustments in the early 1990s, due to high inflation and the beginning of the stabilization plan with exchange rate anchor. After adjusting the residuals, including seasonal dummies, and dummies to correct outliers in the series (see Appendix C). After exclusion tests performed for β coefficients, the RER was not significant in the equation, and it was not significant as an exogenous variable, so the co-integrating vector for Uruguay was:

$$\begin{aligned} X_{URU_t} &= \underset{(7.4364)}{0.5744LM_t} + \underset{(6.3847)}{0.7103LPR_t} - \underset{(-6.5929)}{0.2025RERV_{URU_t}} - 4.9993. \end{aligned} \tag{9}$$

From the tests of weak exogeneity, LM and LPR coefficients were zero, so only RER volatility acts in the short-term on exports, with a coefficient of -0.29.

The co-integrating equation represents the long-term relationship between the variables, where the impact of global demand, represented by world imports (LM), is positive and equal to 0.57, while international food prices (LPR) impact with a 0.71 coefficient and RER volatility ($RERV_{URU}$) has a negative impact, with a 0.20 coefficient.

According to this result, the existence of a co-integrating vector among the variables is not rejected, and the sign of the coefficients are as expected. In addition, we made the corresponding exclusion tests for β and α for weak exogeneity. We found a significant and negative short-term adjustment to the co-integration equation between RER volatility and exports, with a 0.29 coefficient (see Table B1 in Appendix B). In the same line, Demirhan and Demirhan (2015) found a negative and significant impact from exchange rate volatility on exports in the long-run, close to 0.63. This negative effect is only 0.017 in the short-run adjustment. Also, Glauco De Vita and Andrew Abbott (2004) found a negative influence on exchange rate volatility on UK aggregate exports to EU member countries in the long-run. However, they did not find a short-term connection between the variables.



Source: Authors' estimations.

Figure 5 Uruguay Impulse-Response Functions - Response to Cholesky One S. D. Innovations

Figure 5 shows the impulse-response functions analysis from international food prices, global demand, and exchange rate volatility to exports. A positive and permanent effect of the first two variables is confirmed, reaching 1.9% in the first 12 months for international food prices and 1.4% for global demand. Moreover, the impact is negative and permanent when there is a shock on RER volatility, reaching 1.7% after 12 months. Its result suggests some adverse response of exports from real exchange rate uncertainty.

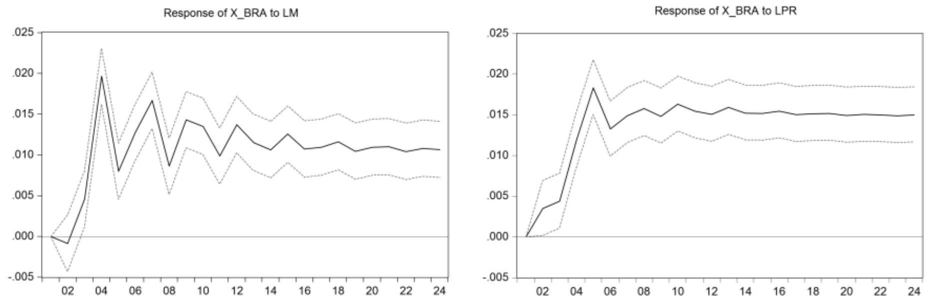
In the case of Brazil, as the real exchange rate volatility was stationary, I(0), the model was estimated for the other four variables (X_BRA, RER_BRA, LM, and LPR), with RER volatility as an exogenous variable. However, we could not detect statistically significant relationship between RER or RER volatility and exports in the Brazilian VEC model, and we did not include these variables. Similar results found

Aguirre, Ferreira, and Notini (2007) in the case of GARCH model for the RER volatility, so the resulting equation for Brazil is:

$$\underset{}{X_{BRA_t}} = \underset{(15.1712)}{1.0171LM_t} + \underset{(4.8952)}{0.5017LPR_t} - 7.1327. \tag{10}$$

According to the coefficients estimated value, the greatest impact in the case of Brazil comes from global demand, with an elasticity near 1, while food international prices have a significant impact, although smaller, with a coefficient of 0.50. The error correction equation shows, in the short-run, a significant adjustment between international prices and exports.

In Figure 6, the impulse-response functions are represented.



Source: Authors' estimations.

Figure 6 Brazil Impulse-Response Functions - Response to Cholesky One S. D. Innovations

The impact of one standard deviation shock of both variables on exports is positive after two periods with an impact below 2% for global demand, and with a final effect near 1.5% for prices. The response turns to be positive over time, suggesting that increased global demand has a positive long-term impact on exports.

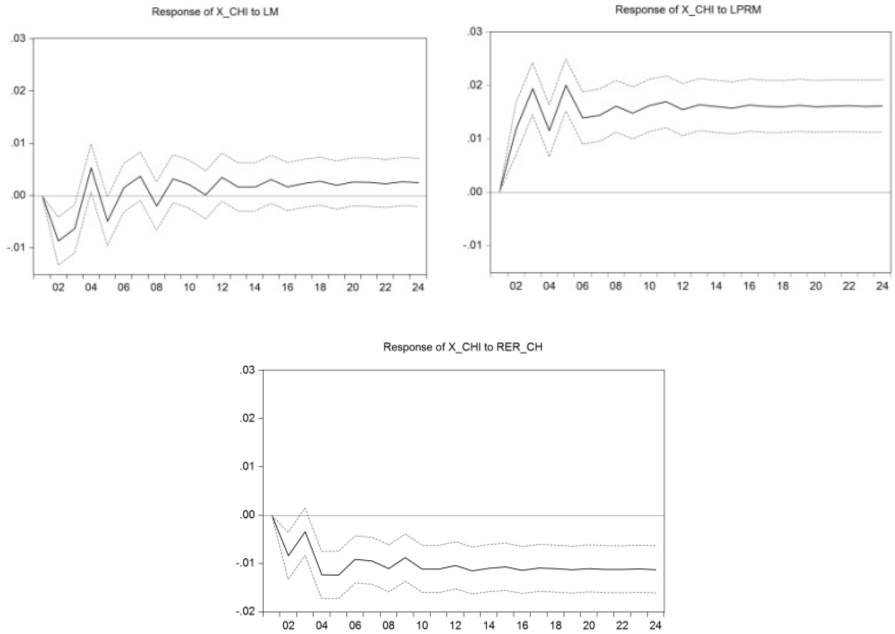
In the case of Chile, after analyzing the series stationarity included in the model (Tables 1 and 2), all variables were not stationary, so we investigated the existence of co-integration between variables and estimated a co-integration vector, where volatility was not significant and the real exchange rate was significant, but with negative sign.

The resulting equation is:

$$X_{CHI_t} = \underset{(12.8374)}{1.2227LM_t} + \underset{(2.8220)}{0.2155LPRM_t} - \underset{(-3.5788)}{0.5727RER_{CHI_t}} - 6.9014. \tag{11}$$

The co-integrating vector estimated for Chilean exports shows a positive coefficient 1.2 for world imports, near 0.2 for international metals prices, and negative near 0.6 for RER. Similar results found Huchet-Bourdon and Korinek (2012) for the RER impact on exports in Chile. They explained this result by Chilean copper exports to China, relatively price inelastic. As Chile is the world's largest copper exporter, this would explain why we found a negative relationship between RER and exports. According to the error correction equation, in the short-run, international metal prices and real exchange rate are significant to explain exports adjustment.

Figure 7 illustrates the impulse-response functions.



Source: Authors' estimations.

Figure 7 Chile Impulse-Response Functions - Response to Cholesky One S. D. Innovations

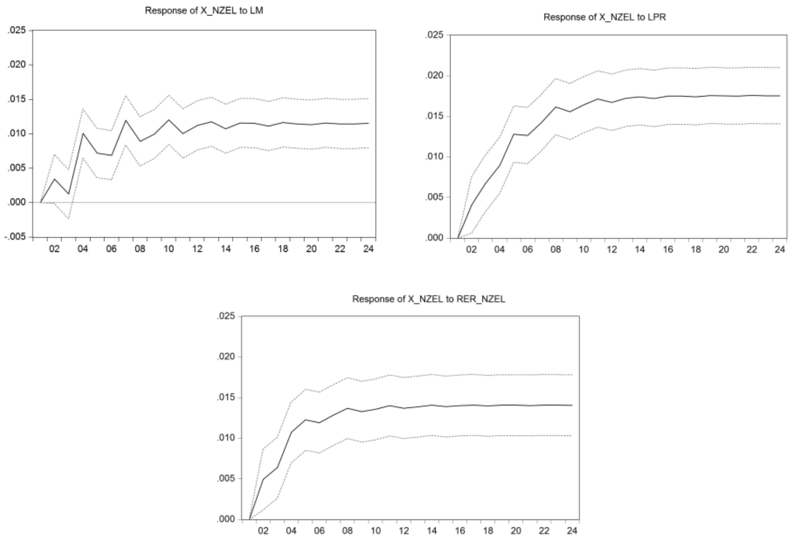
The impulse-response functions, which show the shock of different variables on exports, indicate that despite an initial fluctuating impact on global demand, afterwards, it is not significantly different from zero. In the case of metal prices, the impact remains positive from the beginning, somewhat below 2%. In the case of the real exchange rate, the impact is negative, and from the fourth month onward, the impact is near 1%.

For New Zealand exports co-integrating vector, estimated RER volatility is not significant:

$$X_{NZEL_t} = \underset{(11.3963)}{0.4384LM_t} + \underset{(4.3750)}{0.3702LPR_t} + \underset{(2.1741)}{0.2166RER_{NZEL_t}} - 1.6359. \quad (12)$$

The global demand coefficient is barely higher than 0.4, the international food prices coefficient is close to 0.37, while the exports response elasticity to changes in the real exchange rate was estimated slightly above 0.2. Our results are similar to those found by Huchet-Bourdon and Korinek (2012): positive effects of RER on exports (as predicted by economic theory) and insignificant effects of RER volatility on exports. This is due to the fact that in this type of economies, commodities exports depend mainly on international prices. In addition, in the short-run adjustment, the only significant variables to explain export movements are international prices.

Figure 8 presents the impulse-response functions.



Source: Authors' estimations.

Figure 8 New Zealand Impulse-Response Functions - Response to Cholesky One S. D. Innovations

Impulse-response functions show a moderate impact of the three variables, but positive and sustained over time, slightly higher in the case of food prices than international demand and real exchange rate.

Empirical results in this paper for Chile, Brazil, and New Zealand are similar to the conclusions obtained in Aristotelous (2001) by analyzing the real exchange rate volatility impact of UK exports to US.

Summarizing, real exchange rate volatility has no significant effect on exports in the short- and long-run, except for the Uruguayan case.

Interpreting our findings, in the cases of Brazil, Chile, and New Zealand, capital and futures markets allowed hedge to the cost of risk; however, in the case of Uruguay, the absence of these markets prevented to mitigate the effects of RER uncertainty on exports. A theoretical support of this argument is presented in Jean-Marie Viaene and Casper G. de Vries (1992).

The residual diagnostic tests used to determine the robustness of the error-correction model of the empirical estimation for each model are reported in Appendix. The adjusted R for all four models estimated is over 0.50. Jarque-Bera test for residuals normality shows that error terms are multivariate normal distributed at the significance level of 5%, except for Uruguay (see Table B2 in Appendix B). In addition, Breusch-Godfrey LM test of residual serial correlation (F-version) does not present autocorrelation problems at the significance level of 5%. Moreover, the null homoscedasticity hypothesis cannot be rejecting in the ARCH-LM test for heteroscedasticity at the 5% level (see Table B3 in Appendix B).

Finally, it is necessary to check whether the direction of causality is from the variables (world imports [LM], international food prices index [LPR] or international metals price index [$LPRM$], real exchange rate [RER], and RER volatility [$RERV$]) to exports or *vice versa*.

The Granger causality F-statistics tests for the variables in levels are presented in Table 4.

Table 4 Granger Causality Test (Short-Run)

Variable/Exports	Brazil	Chile	New Zealand	Uruguay
LM	21.6881*	14.3699*	4.7940	4.5566
$LPR/LPRM$	11.0942*	14.3300*	0.0609	5.6313
RER	-----	3.2780	3.2698	-----
$RERV$	-----	-----	-----	5.3183
Exports/Variable	$\boldsymbol{LM}$	$\boldsymbol{LPR/LPRM}$	$\boldsymbol{RER}$	$\boldsymbol{RERV}$
Brazil	4.9118	13.0705*	-----	-----
Chile	4.7717	8.1718	1.8390	-----
New Zealand	2.6510	0.4001	1.3242	-----
Uruguay	2.1184	5.3534	-----	9.2050*

Notes: H_0: X_{it} does not Granger cause X_{jt} $(i \neq j)$, where X_{it} represents the variables in the row and X_{jt} represents the variables in the column. 12 lags. 276 observations in each series. Sample: 1990M01-2013M12. * denotes rejection of the hypothesis at the 0.05 level. Variables are in first difference.

Source: Authors' estimations.

According to the Granger causality test, in the case of the Brazilian economy, we can reject the hypothesis that global demand and international food prices does not Granger cause Brazilian exports. According to impulse-response function results, exports increase 1.2% in approximately 1 year after a shock on global demand, and 1.5% in about 6 months after a shock on international prices. Non-causality of Brazilian exports to international prices is also rejected. However, causality of exports to global demand is rejected. Therefore, it appears that Granger causality runs two-way from international prices to exports and *vice versa*. This result may be due to Brazil being a large economy, and therefore, its exports affect international food prices.

In the case of Chile, Granger test runs one way, so we found causality of global demand and international metals prices for exports, but causality in the other direction is rejected. Considering impulse-response functions, exports are not modified after a shock on global demand, and 1.5% in about six months after a shock on international prices.

For New Zealand, all options are not rejected (in the sense of Granger), so we cannot reject neither the hypothesis that international food prices, global demand, and real exchange rate do not cause the country's exports, nor the opposite. According to impulse-response functions, exports increase 1.0% in about 8 months after a shock on global demand, and 1.8% in about 10 months after a shock on international prices.

Finally, for Uruguay, according to the Granger test, it is not rejected that real exchange rate volatility, global demand, and international food prices do not cause in the sense of Granger Uruguayan exports, although the fact that exports do not Granger cause real exchange rate volatility is rejected. Considering impulse-response functions, exports increase 1.5% in about 10 months after a shock on global demand, near 2.0% in also 10 months after a shock on international prices, and exports diminish near 1.5% after a shock on RER volatility.

Table B4 (Appendix B) shows the results of the long-run causality test. We found a statistically significant and negative coefficient in each model at the 5% level by the *t*-test. These results indicate the endogeneity of the dependent variables in the long-run.

6. Final Remarks

The importance of the relationship between exchange rates and exports comes from the theoretical Mundell-Fleming models. Nevertheless, the empirical work analyzing this relationship is not conclusive. While in the economic literature there is consensus regarding the sign of global demand and international prices effect on exports, the evidence is less conclusive when including RER or even its volatility in the analysis.

According to this paper results, we can conclude that global demand and international prices influence good exports for all the selected countries. But only in the case of Uruguay, the impact of RER volatility was significant both in the short- and the long-run, and with a negative sign, implying that a higher volatility tends to diminish exports. Chile and New Zealand have a considerably lower RER volatility, so it is possible that the exchange rate uncertainty does not affect the export decisions in these countries, as it does in Uruguay. However, in the case of New Zealand and Chile, we found a significant impact of real exchange rate on exports, although the sign of the coefficient was negative in the case of Chile. This result may be due to the fact that Chile is the world's largest copper exporter, and Chilean exports only depend on the external demand.

For the Brazilian case, probably, we must need to adjust a different RER volatility model or consider a longer data series to capture the RER volatility on Brazilian exports. To Uruguay, which is a small open economy, the RER volatility has a negative impact on exports and may be due to the absence of capital markets and futures markets where hedging against sudden changes in the exchange rate, so economic policy must be taken it into account. For the other three countries, economies with more options for exporters to hedge against sudden changes, this fact could help to explain these results.

There could be also other reasons for these results related to the impact of imports in the production of the exported goods, as imported machinery, packing material, or other imported inputs, that depending on their weight in the cost functions of exporters could counteract the positive impact of exchange rate changes on exports.

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Appendix A - Exports Characterization

Table A1 Exports Share by Country or Destination Area (%)

Origin: Uruguay Destination	1994	2013
Brazil	25.8	18.9
Europe & Central Asia	21.6	17.2
Free Zones	0.1	16.0
China	6.0	14.2
Argentina	20.0	5.4
Venezuela	0.1	4.9
United States	6.9	3.9
Russia	0.5	3.1
Others	19.1	16.4

Origin: Chile Destination	1990	2013
China	0.4	24.8
Europe & Central Asia	36.0	17.4
United States	16.8	12.7
Japan	16.2	9.9
Brazil	5.8	5.7
Korea, Rep.	3.0	5.5
India	0.7	3.0
Peru	0.9	2.5
Others	20.3	18.4

Origin: Brazil Destination	1990	2013
Europe & Central Asia	31.9	23.3
China	1.2	19.0
United States	24.6	10.3
Argentina	2.1	8.1
Japan	7.5	3.3
Venezuela	0.9	2.0
Korea, Rep.	1.7	2.0
Chile	1.5	1.9
Others	28.6	30.2

Origin: New Zealand Destination	1990	2013
China	1.0	20.7
Australia	19.0	19.0
Europe & Central Asia	18.2	10.6
United States	13.1	8.5
Japan	16.1	5.9
Korea, Rep.	4.2	3.4
Latin America & Caribbean	3.0	3.2
United Kingdom	7.0	2.9
Others	18.4	25.7

Notes: The term "Others" refers to the rest of the World.

Source: Own elaboration based on data from World Bank (2015).

Appendix B - VEC Residual Test Results

Table B1 VEC Error Correction Equation

Brazil	$D(X\_BRA)$	$D(LM)$	$D(LPR)$	
Coefficient	-0.1692	0.0000	0.0588	
Std. error	(0.0527)	(0.0000)	(0.0165)	
t-statistics	[-3.2107]	[NA]	[3.5603]	
Chile	$D(X\_CHI)$	$D(LM)$	$D(LPRM)$	$D(RER\_CH)$
Coefficient	-0.1589	0.0000	0.1003	-0.0356
Std. error	(0.0675)	(0.0000)	(0.0292)	(0.0142)
t-statistics	[-2.3550]	[NA]	[3.4406]	[-2.5031]
New Zealand	$D(X\_NZEL)$	$D(LM)$	$D(LPR)$	$D(RER\_NZEL)$
Coefficient	-0.3171	0.0000	0.0663	0.0000
Std. error	(0.0686)	(0.0000)	(0.0273)	(0.0000)
t-statistics	[-4.6194]	[NA]	[2.4277]	[NA]
Uruguay	$D(X\_URU)$	$D(LM)$	$D(LPR)$	$D(RERV\_URU)$
Coefficient	-0.2425	0.0000	0.0000	-0.2861
Std. error	(0.0463)	(0.0000)	(0.0000)	(0.0769)
t-statistics	[-5.2389]	[NA]	[NA]	[-3.71978]

Source: Authors' estimations.

Table B2 VEC Residual Diagnostic Test Results

	Brazil model	Chile model	New Zealand model	Uruguay model
Adj. *R*-squared	0.7235	0.5738	0.7071	0.5085
AIC	-11.7468	-15.2753	-16.3595	-12.1277
Jarque-Bera	5.0960*	10.5723*	4.7420*	66.3175
ARCH-LM test (*F*-version)	344.3568*	565.5398*	493.2296*	666.5134*

Notes: Adj. *R*-squared is the coefficient of determination adjusted. Lag lengths are selected by AIC. Jarque-Bera: orthogonalization Cholesky (Lutkepohl) and null hypothesis residuals are multivariate normal. The *F*-statistic version of the ARCH-LM test statistic for autoregressive conditional heteroscedasticity: White test is with no cross-terms and null hypothesis is no heteroscedasticity. * denotes no rejection of the hypothesis at the 0.95 level. Sample: 1990M01-2013M12.

Source: Authors' estimations.

Table B3 VEC Residual Serial Correlation LM Test Results

Lags	*LM*-stat Brazil	*LM*-stat Chile	*LM*-stat New Zealand	*LM*-stat Uruguay
1	9.0758*	18.4397*	17.7605*	25.1169*
2	7.2909*	17.4125*	18.0516*	24.1004*
3	14.5464*	21.6412*	17.3478*	24.0432*
4	6.4151*	18.2392*	12.5536*	18.8364*
5	4.7677*	23.6375*	14.9433*	21.5659*
6	2.2795*	19.8703*	18.0348*	17.3553*
7	10.1139*	22.0564*	10.9357*	21.2236*
8	10.8366*	16.8030*	10.3992*	9.8823*
9	13.5449*	21.1425*	21.4919*	16.0332*
10	9.7492*	28.4834	22.6213*	17.2687*
11	3.5679*	16.4182*	25.0235*	16.9515*
12	7.8423*	26.0644*	33.5618	16.5304*

Notes: Null hypothesis: no serial correlation at lag order h. Sample: 1990M01-2013M12. * denotes no rejection of the hypothesis at the 0.95 level.

Source: Authors' estimations.

Table B4 Causality Test (Long-Run)

Dependent variable: X	Brazil	Chile	New Zealand	Uruguay
Variable	Coefficient	Coefficient	Coefficient	Coefficient
$ECT(-1)$	-0.3774*	-0.3008*	-0.6229*	-0.2111*
$X(-1)$	-0.5409*	-0.6507*	-0.4297*	-0.3116*
$X(-2)$	0.0959	-0.0359	0.1201	-0.0508
$X(-3)$	0.2167*	0.1007	0.2101*	0.0712
$X(-4)$	0.1194	0.0081	---	---
$LM(-1)$	0.2402*	0.2574*	-0.1531	0.2065
$LM(-2)$	-0.4090*	-0.6636*	-0.1571	0.1544
$LM(-3)$	-0.2455*	-0.3441*	-0.5024*	0.0851
$LM(-4)$	-0.7425*	-0.4296*	---	---
$LPR/LPRM(-1)$	0.1913	0.4314*	0.2985	0.5694*
$LPR/LPRM(-2)$	0.4042	0.3931*	0.2502	0.1127
$LPR/LPRM(-3)$	0.7209*	0.1687	0.1406	-0.0501
$LPR/LPRM(-4)$	0.4374	0.2213	---	---
$RER(-1)$	---	-0.6108	0.0795	---
$RER(-2)$	---	0.5090	0.2748	---

$RER(-3)$	---	-0.4187	0.2670	---
$RER(-4)$	---	-0.0873	---	---
$RERV(-1)$	---	---	---	-0.0145
$RERV(-2)$	---	---	---	-0.0275
$RERV(-3)$	---	---	---	0.0824*

Notes: Variables are in first difference. Values in parentheses are the lags. * is the 5% critical level.

Source: Authors' estimations.

Appendix C - Outliers and Dummy Variables

Table C1 Uruguay

Period	Type	Influence				Brief description
		X	LM	LPR	$RERV$	
1994-10	AO				x	"Tequila effect", due to Mexico 1994 crisis
1996-04	AO	x		x		Impact of mad cow disease in beef exports, and China wool demand
1996-09	AO			x		Oil prices increase, due Gulf War
1998-10	AO				x	1998 Russian financial crisis
1999-02	AO			x	x	Brazilian devaluation
2000-11	AO	x				First outbreak of foot-and-mouth disease (FMD) in Uruguay
2001-04	TC	x	x			Generalization of FMD in Uruguay
2002-02	AO				x	Argentinean crisis and devaluation
2002-08	AO				x	Uruguayan bank crisis and devaluation
2003-06	AO	x				Uruguayan public debt restructuring
2008-08	TC	x		x		International crisis
2008-10	TC		x	x		International crisis and devaluation in Uruguay
2009-01	AO		x	x	x	GDP contraction in developed economies
2009-05	TC	x	x			Recovery of international markets
2010-06	TC			x	x	Appreciation of Uruguayan RER with its trading partners
2012-07	AO	x		x	x	Euro Zone problems and changes in Uruguayan monetary policy
2012-11	AO				x	Euro Zone problems and changes in Uruguayan monetary policy
2013-05	AO	x				Changes in Uruguayan monetary policy
2013-07	TC				x	Monetary policy in Uruguay goes from interest rate to M1

Source: Authors' estimations.

Table C2 Brazil

Period	Type	Influence			Brief description
		X	LM	LPR	
1990-10	AO	x	x		Trade reform (liberalization)
1991-05	AO			x	Widespread decline in international food prices
1991-09	AO	x			Impact of international food prices decline
1992-08	AO		x	x	European Monetary System crisis
1992-10	TC			x	International prices recovery
1993-01	AO		x		Recovery of global demand
1993-05	AO	x	x		Trade liberalization
1995-06	TC			x	Tequila effect
1996-04	AO			x	Increase of international grain prices
1996-09	AO			x	Increase of international grain prices
1998-08	AO	x		x	Russian moratorium
1999-08	AO			x	Sharp fall in the index of international prices of basic food export

2001-04	AO		x		International production slowdown
2002-07	AO	x		x	Financial uncertainty and devaluation of the real
2002-09	AO	x			Devaluation of the real due to presidential elections (Lula)
2003-09	AO	x		x	Exports increase
2005-04	AO			x	Commodity prices increase
2006-07	AO	x			Impact of rising oil prices
2008-03	AO	x			Impact of rising oil prices
2008-08	TC		x	x	International crisis and commodity prices fall
2008-10	AO		x	x	International crisis
2009-01	AO		x	x	International crisis
2009-07	AO			x	Fall in food prices
2010-07	TC			x	Price increase of some relevant products by effects of drought
2011-10	AO		x	x	Falling international oil prices
2012-07	AO			x	Increase in commodity prices, due to weather problems

Source: Authors' estimations.

Table C3 Chile

Period	Type	Influence				Brief description
		X	LM	$LPRM$	RER	
1993-01	AO	x	x			Fall in international prices, mainly copper and cellulose. Imports' license system and application of a countervailing duty
2000-04	AO	x				Entry into force of the Free Trade Agreement with Korea
2001-07	AO				x	Impact of Argentinean crisis
2006-06	AO	x		x		Fall of copper prices
2007-08	AO			x		New increase of copper prices
2008-10	AO	x	x	x	x	International crisis impact
2009-01	AO		x			International crisis impact
2009-11	AO				x	Exchange market volatility
2010-05	AO			x		Increase of copper prices due to China's demand
2011-01	AO			x	x	Influence of US economic recovery instability in the currency

Source: Authors' estimations.

Table C4 New Zealand

Period	Type	Influence				Brief description
		X	LM	$LPRM$	RER	
1990-12	AO	x				Gulf crisis impact on world demand and increase of agricultural commodities supply
1993-01	AO		x			Global recession
1996-04	AO			x		Increase of international grain prices
2000-12	AO				x	Impact of inflation fall on the RER
2007-08	AO				x	US monetary policy impact
2008-10	AO		x	x	x	International crisis impact
2009-01	AO		x	x		International crisis impact
2012-07	AO			x		Increase in commodity prices, due to weather problems

Source: Authors' estimations.

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