International Journal of Business and Finance Research
Vol. 9, No. 4, 2015, pp. 37-49
ISSN: 1931-0269 (print)
ISSN: 2157-0698 (online)

IBFR
www.theIBFR.com

THE IMPACT OF EXCHANGE RATE VOLATILITY ON COMMODITY TRADE BETWEEN THE UNITED STATES AND SPAIN

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ABSTRACT

In this paper we analyze the effects of the real exchange rate volatility on disaggregated sectoral data on the trade flows between the United States and Spain. This study uses monthly trade flows on United States exports to and imports from Spain over the period from January 1993 to December 2012 and the method of bounds testing or the Autoregressive Distributed Lag (ARDL) approach to cointegration analysis. Our results reveal that exports depend positively on the levels of foreign economic activity but negatively on relative prices. However, the exchange rate volatility tends to provide mixed effects. In addition, imports depend positively on the levels of domestic economic activity but negatively on relative prices. As in the case of exports, the exchange rate volatility tends to provide mixed effects. Furthermore, in the case of both exports and imports, the effects of exchange volatility are found to yield mixed effects in the short-run and in the long-run.

JEL: F14, F31

KEYWORDS: Spain, Imports, Exports, Exchange Rate Volatility

INTRODUCTION

There are a large number of studies in the area of exchange volatility and trade. Despite the sizeable number of studies conducted, no real consensus about the impact of exchange rate volatility on trade has emerged. While a large number of studies find that exchange rate volatility tends to reduce the level of trade, others find either weak or insignificant or positive relationships. For example, Onafowara and Owoye (2008), Byrne, Darby, and MacDonald (2008), Choudhry (2005), Bahmanee-Oskooee (2002), Arize, *et al.* (2000), Arize (1995), Chowdhury (1993), Pozo (1992), and Bahmani-Oskooee and Ltaifa (1992), find evidence for negative effects. According to these scholars, exchange rate volatility may affect exports directly through uncertainty and adjustment costs for risk-averse exporting investors. Further, it may have an indirect effect through its impact on the structure of output, investment and government policy. On the other hand, Doyle (2001), Chou (2000), McKenzie and Brooks (1997), Qian and Varangis (1994), Kroner and Lastrapes (1993), and Asseery and Peel (1991) find evidence for a positive effect for volatility on export volumes of some developed countries because exchange rate volatility makes exporting more attractive to risk-tolerant exporting firms. However, other scholars such as Aristotelous (2001), Bahmani-Oskooee and Payestch (1993), Bahmani-Oskooee (1991), and Hooper and Kohlhagen (1978) have reported no significant relationship between exchange rate volatility and exports.

Reasons for contradictory results by different studies may be due to a variety of factors, among them: different methods used to measure exchange rate volatility; the use of different price deflators; the differential use of sample data, for example, the use of aggregate export data versus sectoral export data; different time-frames; ignoring import dependency on intermediate and capital goods of the receiving country, as is the case with many developing countries; and the absence of complex econometric methods for studying these variations. As a result, scholars stopped investigating the exchange rate volatility-export nexus by the late 1990's. However, with better access to sectoral data and the development of more sophisticated econometric models, recent studies have begun evaluating the exchange rate volatility-export connection from a sectoral perspective. The rationale behind this is that different trade sectors would be impacted differentially by exchange rate volatility, and therefore may be more revealing than aggregate studies. This study focuses on disaggregated trade flows between the United States and Spain to uncover the nature and sensitivity of the relationship between exchange rate volatility and trade flows. We use the method of bounds testing or the Autoregressive Distributed Lag (ARDL) approach to cointegration analysis for this purpose. Using this approach we investigate the effects of exchange rate volatility on United States sectoral exports to and sectoral imports from Spain over a period of 20 years using monthly data from January 1993 to December 2012.

We provide a brief review of the literature in the next section. Thereafter, we lay the empirical framework of our study by specifying our model. In the section following that we discuss variable definitions and outline our data sources. Empirical results from the bounds testing approach to cointegration, and error-correction model estimates are presented in the penultimate section. The final section presents a summary and conclusion of the results obtained in this study.

LITERATURE REVIEW

In this section we present a brief overview of studies that examine the exchange rate volatility-trade nexus. We begin by discussing the most recent and sophisticated studies, employing cointegration techniques and error-correction models, to older, less complex studies. Bahmani-Oskooee and Harvey (2011) investigate the effects of exchange rate fluctuations on trade flows between the U.S. and Malaysia using disaggregated, industry-level annual export and import data for 17 export industries and 101 importing industries from 1971 to 2006. They conclude that while exchange rate volatility exerts short-run effects in trade flows of almost two-thirds of the industries, these effects last into the long-run in 38 U.S. exporting industries and in 10 U.S. importing industries.

Bahmani-Oskooee and Hegerty (2009) investigate the effects of exchange rate fluctuations on trade flows between the U.S. and Mexico using disaggregated, industry-level annual export and import data for 102 industries from 1962 to 2004. They analyze both the short- and long-term effects of volatility in the peso/dollar real exchange rate on Mexican-United States trade. They conclude that in the short-term increased volatility negatively affects trade flows in most industries. Long-term effects however, are significant for only one-third of the industries studied, and of this, only two-thirds are negative. They speculate that increased Mexican integration and liberalization of economic policies allow for greater adjustments in the long-term so that volatility is less of a problem in the long-term than in the short-term.

Byrne, Darby, and MacDonald (2008) analyze the impact of exchange rate volatility on the volume of bilateral U.S. trade flows using homogenized and differentiated sectoral annual data over the period 1989-2001 for a cross-section of 6 EU countries and 22 industries. Their study finds that clustering all industries together provides evidence of a negative effect on trade from exchange rate volatility, which confirms findings of other studies using aggregate data. However, when investigating sectoral trade differences, the effects of exchange rate volatility on trade is negative and significant for differentiated goods and insignificant for homogeneous goods, confirming recent studies that sectoral differences are in fact crucial to explaining the differential impact of volatility on trade. They suggest that a greater degree of disaggregation at the industry level may provide more worthwhile results, which is what we do in this study. Bahmani-Oskooee and Kovyryalova (2008) investigate the effect of exchange rate fluctuations on trade flows between the U.S. and the United Kingdom using disaggregrated annual export and import data for 177 commodities industries from 1971 to 2003. They analyze both the short- and long-term effects of real exchange rate volatility on trade between the U.S. and the UK. Their results reveal that the volatility of the real dollar–pound rate has a short-term significant effect on imports of 109 industries and on exports of 99 industries. In most cases, such effects are unfavorable. In the long run, however, the number of significant cases is somewhat reduced: only 62 import and 86 export industries are significantly and adversely affected by exchange rate volatility. The industries affected involve both durable and non-durable goods, and include small as well as large industries, supporting findings by aggregate studies.

In another study, Bahmani-Oskooee and Mitra (2008), investigate the effects of exchange rate volatility on trade flows between the U.S. and India, an emerging economy. Using annual data from 40 industries from 1962–2004, their results demonstrate that exchange rate volatility has more short-run than long-run effects. In the short-run, 17 industries were affected on the import side and 15 on the export side. The industries affected show India's increasing ability to produce import substitutable goods. However, in the long run, only a few industries are affected because the increasing dependence on trade between India and the US cause industries to respond inelastically to exchange rate volatility.

Using both the nominal and the real exchange rate between the United States dollar and the currencies of Canada and Japan, Choudhury (2005) investigates the influence of exchange rate volatility on U.S. real exports to Canada and Japan using aggregate monthly data ranging from January 1974 to December 1998. The study uses conditional variance from the GARCH (1, 1) model as a measure of exchange rate volatility, and finds significant and mostly negative effects of exchange rate volatility on real exports.

As in the above studies, Sukar and Hassan (2001) investigate the relationship between U.S. trade volume and exchange rate volatility using cointegration and error-correction models. Their study uses quarterly aggregate data covering the period 1975Q1 – 1993Q2 and a GARCH model to measure the exchange rate volatility. Paralleling other studies, the authors find evidence for a significantly negative relationship between U.S. export volume and exchange rate volatility. However, unlike other findings, they reveal that the short-run dynamics of the exchange rate volatility -trade relationship is insignificant. They argue that this result may be due to the existence of avenues for hedging against exchange risks so as to neutralize the negative impact of exchange rate volatility. Other scholars argue that this short-run insignificant relationship may be because of the investigators' use of aggregate data, which ignores sectoral differences. For example, while one sector may exhibit a negative relationship, another may exhibit an equal but opposite effect so that they offset each other.

Arize (1995), using monthly series from February 1978 to June 1986 analyzes the effects of real exchange rate volatility on the proportions of bilateral exports of nine categories of goods from the U.S. to seven major industrial countries. The volatility measure employed is the standard deviation of the monthly percentage change in the bilateral exchange rate between the U.S. and the importing country from t to $t-12$. The study reveals differential effects of exchange rate volatility across different categories of exports. The study also concludes that exchange rate uncertainty has a negative effect on U.S. real exports, and that it may have a major impact on the allocation of resources to different industries depending on trade elasticities. Lastrapes and Koray (1990) analyze the interrelationships among exchange rate volatility, international trade, and macroeconomic variables using the vector autoregression (VAR) model. The model estimates U.S. multilateral trade from 1973 to 1990 and includes a moving standard deviation measure of real exchange rate volatility. While the results reveal some evidence of a statistically significant relationship between volatility and trade, the moving average representation of the model implies a rather small quantitative effect. The study concludes that exchange rate volatility is influenced by the state of the economy, a factor ignored in a variety of other studies.

Finally, Klein (1990) is one of the first few scholars to analyze the effects of exchange rate volatility on the proportion of disaggregated bilateral exports of nine categories of goods from the U.S. to seven major industrial countries using fixed effects framework. Using monthly series data from February 1978 to June

1986, the study reveals that in six categories of exports exchange rate volatility significantly affects the volume of exports and in five of these categories the effect is positive, suggesting that real exchange rate volatility may in fact increase exports by risk-taking firms.

One major problem with most of the studies above is that the sample period includes the period prior to the end of the fixed exchange regime, so results may include the lag effects of fixed exchange rates on trade before 1973 lingering on during the transition period after the implementation of the floating exchange rate regime. The current study corrects for this potential bias by using United States monthly industry trade data covering a 20-year period from January 1993 to December 2012. The methodology used in this study incorporates the recent developments in the literature, namely, the ARDL approach to cointegration analysis, which may uncover the nature and sensitivity of the exchange rate volatility-trade nexus.

METHODOLOGY

Model Specification

The objective of this study is to assess the effects of exchange rate volatility on the trade flows disaggregated at the 2-digit Harmonized System (HS) industry level. The study uses trade data on U.S. exports to and imports from Spain. Drawing on the existing empirical literature in this area, we specify that a standard long-run export demand function for commodity i to take the following form (see, for example, Ozturk and Kalyonku, 2009; Choudhry, 2005; Arize, 1998, 1996, 1995; and Asseery and Peel, 1991):

$$\ln X_{it} = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln P_{it} + \beta_3 \ln VOL_t + \varepsilon_t \qquad (1)$$

Where X_{it} is the real export volume of commodity i in period t, Y_t is the real income of Spain in period t, P_{it} is the relative price of exports of commodity i in period t, VOL_t is a measure of exchange rate volatility, and ε_t is a white-noise disturbance term.

Economic theory posits that the real income level of the domestic country's trading partners would have a positive effect on the demand for its exports. Therefore, *a priori*, we would expect that $\beta_1 > 0$. On the other hand, if the relative price of exports rise (fall), domestic goods become less (more) competitive than foreign goods, causing the demand for exports to fall (rise). Therefore, *a priori*, one would expect that β_2, which measures the competitiveness of U.S. exports relative to Spanish domestic production, is negative. The last explanatory variable is a measure of exchange rate volatility. Various measures of real VOL have been proposed in the literature. Some of these measures include (1) the averages of absolute changes, (2) the standard deviations of the series, (3) the deviations from the trend, (4) the squared residuals from the ARIMA or ARCH or GARCH processes, and (5) the moving sample standard deviation of the growth rate of the real exchange rate. Since the effects of VOL on exports have been found to be empirically and theoretically ambiguous (Bredin, *et al.* 2003), β_3 could be either positive or negative.

Equation (1) shows the long-run relationships among the dependent and independent variables in our model. Given the recent advances in time-series analysis, in estimating the long-run model outlined by equation (1), it is now a common practice to distinguish the short-run effects from the long-run effects. For this purpose, equation (1) should be specified in an error-correction modeling (ECM) format. This method had been used in many recent studies including Bahmani-Oskooee and Hegerty (2009), Bahmani-Oskooee and Wang (2008, 2009), Bahmani-Oskooee and Mitra (2008), Bahmani-Oskooee and Kovyryalova (2008), and Bahmani-Oskooee and Ardalani (2006). According to Bahmani-Oskooee and

Wang (2008), such an approach is warranted given that the measure of exchange rate volatility is a stationary variable (see, for example, De Vita and Abbot, 2004; Bahmani-Oskooee & Payesteh, 1993; and Doyle, 2001), whereas the other variables in equation (1) are non-stationary. Therefore, following Pesaran, Shin, and Smith (2001) and their method of bounds testing or the Autoregressive Distributed Lag (ARDL) approach to cointegration analysis, we rewrite equation (1) as an error-correction model in equation (2) below.

$$\Delta \ln X_t = \alpha_0 + \sum_{i=1}^{n} \beta_i \Delta \ln X_{t-i} + \sum_{i=0}^{n} \gamma_i \Delta \ln Y_{t-i} + \sum_{i=0}^{n} \delta_i \Delta \ln P_{t-i} + \sum_{i=0}^{n} \varphi_i \Delta \ln VOL_{t-i} + \lambda_0 \ln X_{t-1} + \lambda_1 \ln Y_{t-1} + \lambda_2 \ln P_{t-1} + \lambda_3 \ln VOL_{t-1} + \omega_t \qquad (2)$$

Where Δ is the difference operator and the other variables are as defined earlier. Pesaran, Shin, and Smith's (2001) bounds testing approach to cointegration is based on two procedural steps. The first step involves using an F-test or Wald test to test for joint significance of the no cointegration hypothesis $\mathrm{H}_0 : \lambda_0 = \lambda_1 = \lambda_2 = \lambda_3 = 0$ against an alternative hypothesis of cointegration, $\mathrm{H}_1 : \lambda_0 \neq 0$, $\lambda_1 \neq 0$, $\lambda_2 \neq 0$, $\lambda_3 \neq 0$. This test is performed using equation (2). The advantage of this approach is that there is no need to test for unit roots, as is commonly done in cointegration analysis. Pesaran, Shin, and Smith (2001) provide two sets of critical values for a given significance level with and without time trend. One assumes that the variables are stationary at the levels or I(0), and the other assumes that the variables are stationary at the first difference or I(1). If the computed F-values exceed the upper critical bounds value, then H_0 is rejected signaling cointegration among the independent variables. If the computed F-value is below the critical bounds values, we fail to reject H_0. Finally, if the computed F-statistic falls within the boundary, the result is inconclusive. After establishing cointegration, the second step involves estimation of the long-term elasticities and the error-correction model. Similar to the export model, the import model is also specified drawing on the existing empirical literature in this area. We specify that a standard long-run import demand function for commodity i to take the following form (see, for example, Ozturk and Kalyonku, 2009; Choudhry, 2005; Arize, 1998, 1996, 1995; and Asseery and Peel, 1991):

$$\ln M_{it} = \theta_0 + \theta_1 \ln Y_t + \theta_2 \ln P_{it} + \theta_3 \ln VOL_t + \varepsilon_t \qquad (3)$$

Where M_{it} is the real import volume of commodity i in period t, Y_t is the real income of the United States in period t, P_{it} is the relative price of exports of commodity i in period t, VOL_t is a measure of exchange rate volatility, and ε_t is a white-noise disturbance term.

Economic theory posits that the real income level of the domestic (importing) country would have a positive effect on the demand for its imports. Therefore, *a priori*, we would expect that $\theta_1 > 0$. On the other hand, if the relative price of imports rise (fall), foreign goods become less (more) competitive than domestic goods, causing the demand for imports to fall (rise). Therefore, *a priori*, one would expect that θ_2, which measures the competitiveness of Spanish exports relative to the U.S. domestic production, is negative. The last explanatory variable is a measure of exchange rate volatility. Since the effects of VOL on imports have been found to be empirically and theoretically ambiguous (Bredin, *et al.* 2003), θ_3 could be either positive or negative. Equation (3) shows the long-run relationships among the dependent and independent variables in our model. As it was done in the case of the export model, in estimating the long-run model outlined by equation (3), in order to distinguish the short-run effects from the long-run effects, equation (3) should also be specified in an error-correction modeling (ECM) format. Therefore, following Pesaran, Shin, and Smith (2001) and their method of bounds testing or the Autoregressive

Distributed Lag (ARDL) approach to cointegration analysis, we rewrite equation (3) as an error-correction model in equation (4) below.

$$\begin{aligned}\Delta \ln M_t = \psi_0 + \sum_{i=1}^{n}\theta_i \Delta \ln M_{t-i} + \sum_{i=0}^{n}\rho_i \Delta \ln Y_{t-i} + \sum_{i=0}^{n}\sigma_i \Delta \ln P_{t-i} \\ + \sum_{i=0}^{n}\omega_i \Delta \ln VOL_{t-i} + \mu_0 \ln M_{t-1} + \mu_1 \ln Y_{t-1} + \mu_2 \ln P_{t-1} + \mu_3 \ln VOL_{t-1} + \omega_t\end{aligned} \quad (4)$$

Where Δ is the difference operator and the other variables are as defined earlier. Pesaran, Shin, and Smith's (2001) bounds testing approach to cointegration is based on two procedural steps. The first step involves using an F-test or Wald test to test for joint significance of the no cointegration hypothesis $H_0 : \mu_0 = \mu_1 = \mu_2 = \mu_3 = 0$ against an alternative hypothesis of cointegration, $H_1 : \mu_0 \neq 0$, $\mu_1 \neq 0$, $\mu_2 \neq 0$, $\mu_3 \neq 0$. This test is performed using equation (2). The advantage of this approach is that there is no need to test for unit roots, as is commonly done in cointegration analysis. Pesaran, Shin, and Smith (2001) provide two sets of critical values for a given significance level with and without time trend. One assumes that the variables are stationary at the levels or I(0), and the other assumes that the variables are stationary at the first difference or I(1). If the computed F-values exceed the upper critical bounds value, then H_0 is rejected signaling cointegration among the independent variables. If the computed F-value is below the critical bounds values, we fail to reject H_0. Finally, if the computed F-statistic falls within the boundary, the result is inconclusive. After establishing cointegration, the second step involves estimation of the long-term elasticities and the error-correction model.

DATA SOURCES AND VARIABLES

Our export and import data series span a 20-year period from January 1993 through December 2012, leading to 240 monthly observations. Monthly data on real export volume and prices are taken from the Global Trade Information Services, *World Trade Atlas Database*. Monthly data on real export and import volumes and prices have been converted into export volume indices and export price indices with 2005 serving as the base (=100). The study focuses on the top twenty export commodities and top twenty import commodities defined at the 2-digit Harmonized System (HS) codes level, and selected based on their average export and import values between 1993 and 2012. The top 20 export products from the U.S. to Spain are: Pharmaceutical Products (HS30); Miscellaneous Grain, Seed, Fruit (HS12); Mineral Fuel and Oil (HS27); Machinery (HS84); Organic Chemicals (HS29); Edible Fruit and Nuts (HS08); Aircraft and Spacecraft (HS88); Electrical Machinery (HS85); Optical and Medical Instruments (HS90); Passenger Vehicles (HS87); Iron and Steel Products (HS73); Paper and Paperboard (HS48); Plastic (HS39); Miscellaneous Chemical Products (HS38); Woodpulp Etc. (HS47); Food Waste and Animal Feed (HS23); Perfumery, Cosmetic, Etc. (HS33); Iron and Steel (HS72); Copper and Articles Thereof (HS74); and Rubber (HS40).

These 20 export products accounted for 83.1% of total U.S. exports to Spain in 2012. The top 20 import products of the U.S. from Spain are: Mineral Fuel and Oil (HS27); Machinery (HS84); Pharmaceutical Products (HS30); Organic Chemicals (HS29); Electrical Machinery (HS85); Beverages (HS22); Iron and Steel Products (HS73); Rubber (HS40); Preserved Food (HS20); Plastic (HS39); Passenger Vehicles (HS87); Optical and Medical Instruments (HS90); Edible Fruit and Nuts (HS08); Iron and Steel (HS72); Footwear (HS64); Aircraft and Spacecraft (HS88); Miscellaneous Art of Base Metal (HS83); Ceramic Products (HS69); Precious Stones and Metals (HS71); and Leather Art, Saddlery, and Bags (HS42). These 20 import products accounted for 72.8% of total U.S. imports from Spain in 2012. The real income variable for Spain is proxied by the industrial production index (2005=100) of Spain while the real income variable for the U.S. is proxied by the industrial production index (2005=100) of the U.S. The underlying series are obtained from the International Monetary Fund's *International Financial Statistics database* and from the Organization for Economic Cooperation and Development's online database.

The relative price ratio for U.S. exports is calculated as the ratio of the export price index of each commodity to the price level, proxied by the consumer price index (2005=100) of Spain. The export price index for each of the export products is computed using the unit prices taken from the Global Trade Information Services, *World Trade Atlas Database*, while the consumer price index is also obtained from the International Monetary Fund's *International Financial Statistics database*. Following Bahmani-Oskooee and Wang (2008, 2009), and Sekkat and Varoudakis (2000), the real exchange rate, RER_t, is constructed as:

$$RER_t = \left(\frac{ER_t^{US-SP} \times P_t^{SP}}{P_t^{US}} \right) \quad (3)$$

where RER_t is the real exchange rate, ER_t^{US-SA} is the bilateral nominal exchange rate between the United States and Spain at time t, P_t^{SP} is the consumer price index (2005=100) of Spain at time t, and P_t^{US} is the consumer price index (2005=100) of the U.S. at time t. The monthly data on nominal exchange rates are taken from the IMF, *International Financial Statistics database*. Finally, our measure of volatility is constructed following Bredin, Fountas, and Murphy (2003), Weliwita, Ekanayake, and Tsujii (1999), Chowdhury (1993), Lastrapes and Koray (1990), and Koray and Lastrapes (1989). Following these authors the real exchange rate volatility measure is constructed as:

$$VOL_t = \left[\frac{1}{m} \sum_{i=1}^{m} (\ln RER_{t+i-1} - \ln RER_{t+i-2})^2 \right]^{1/2} \quad (6)$$

where VOL_t is the volatility of real exchange rate, RER_t is the real exchange rate and m = 4 is the order of the moving average. According to Koray and Lastrapes (1989), this measure can capture general movements in real exchange rate volatility and exchange rate risk over time.

EMPIRICAL RESULTS

Applying the ARDL approach to cointegration to monthly data from January 1993 to December 2012, we assess the top twenty U.S. export products to and the top twenty U.S. import products from Spain. First, we estimate equations (2) and (4). Following Bahmani-Oskooee and Mitra (2008) we impose a maximum of four lags on each first differenced variable and employ Akaike's Information Criterion (AIC) to select the optimum lag length. Choosing a combination of lags that minimizes the AIC, we then test whether the variables for each industry are cointegrated. The results of the exports model are shown in Table 1 while the results of the imports model are presented in Table 2. Table 1 reveals that fifteen of the twenty industries encompass an F-statistic above the upper bound of 4.35, implying that these industries' four variables are cointegrated. The other five industries reveal an F-statistic below the lower bound of 3.23, indicating no cointegration among variables. Therefore, only those fifteen industries that exhibit cointegrating relationships among variables are used to analyze the effects of volatility on exports.

Table 1: Cointegration Test Results of Top Twenty Export Commodities from the U.S. to Spain

HS	Industry	F-Statistic	ECM	Cointegrated?
30	Pharmaceutical Products	1.07	-0.024 (0.55)	No
12	Miscellaneous Grain, Seed, and Fruit	8.28	-0.708 (9.44)	Yes
27	Mineral Fuel and Oil, etc.	20.47	-0.749 (9.09)	Yes
84	Machinery	17.27	-0.600 (8.15)	Yes
29	Organic Chemicals	4.80	-0.288 (3.91)	Yes
08	Edible Fruit and Nuts	5.54	-0.199 (4.58)	Yes
88	Aircraft and Spacecraft	10.09	-0.911 (9.09)	Yes
85	Electrical Machinery	5.62	-0.311 (4.64)	Yes
90	Optical and Medical Instruments	10.07	-0.442 (6.35)	Yes
87	Passenger Vehicles	2.73	-0.227 (3.11)	No
73	Iron and Steel Products	19.56	-0.693 (8.88)	Yes
48	Paper and Paperboard	5.62	-0.309 (4.89)	Yes
39	Plastic	2.92	-0.182 (3.25)	No
38	Miscellaneous Chemical Products	13.19	-0.452 (7.15)	Yes
47	Woodpulp, etc.	8.86	-0.579 (5.78)	Yes
23	Food Waste and Animal Feed	5.06	-0.372 (4.36)	Yes
33	Perfumery, Cosmetic, etc.	2.45	-0.253 (3.12)	No
72	Iron and Steel	5.31	-0.381 (4.18)	Yes
74	Copper and Articles Thereof	2.04	-0.155 (2.72)	No
40	Rubber	5.67	-0.328 (4.50)	Yes

Note: This table summarizes the results of the bounds testing approach to cointegration. The figures in parentheses are absolute value of the t-statistic. ECM represents the error-correction term. The upper bound critical value for the F-statistic with unrestricted intercept and no trend at the 5% level of significance is 4.35. The lower bound critical value is 3.23. These values are taken from Pesaran, Shin, and Smith (2001, Table CI(iii) Case III, p. 300).

Similarly, Table 2 reveals that fourteen of the twenty industries encompass an F-statistic above the upper bound of 4.35, implying that these industries' four variables are cointegrated. The other six industries reveal an F-statistic below the lower bound of 3.23, indicating no cointegration among variables. Therefore, only those fourteen industries that exhibit cointegrating relationships among variables are used to analyze the effects of volatility on imports. The estimated coefficients for the fifteen cointegrated export industries are presented in Table 3. Following the studies by Bahmani-Oskooee and Harvey (2011), Bahmani-Oskooee and Hegerty (2009), Bahmani-Oskooee and Wang (2008, 2009), Bahmani-Oskooee and Mitra (2008), Bahmani-Oskooee and Kovyryalova (2008), and Bahmani-Oskooee and Ardalani (2006), we report only the short-run volatility coefficients and all the long-run coefficients.

Short-Run Effects of Exchange Rate Volatility on Exports: The short-run estimated coefficients on exchange rate volatility presented on the left panel in Table 3 reveal a mixture of negative and positive signs. There is also a significance variation of the exchange rate volatility on exports among industries in the short-run. Some of the coefficients are positive and statistically significant. These industries include miscellaneous grain, seed and fruit, mineral fuel and oil, etc., machinery, electrical machinery, iron and steel products, and paper and paperboard. The products that have negative coefficients show that these coefficients are statistically insignificant in the short-run.

Long-Run Effects of Exchange Rate Volatility on Exports: The long-run coefficient estimates are shown in the right panel of Table 3. As economic theory postulates, the real income variable renders a positive sign in all cases. This coefficient is statistically significant at the 1% level in ten of the industries and significant at the 5% level in one industry. The relative price variable displays the expected negative sign in all industries and is statistically significant at the 1% level in eight of the fifteen industries, and at the 5% level in three industries. This result is similar to those of Bahmani-Oskooee and Harvey (2011), Bahmani-Oskooee and Mitra (2008), Bahmani-Oskooee and Kovyryalova (2008), and Bahmani-Oskooee and Ardalani (2006). Finally, the estimated coefficients on VOL show a mixture of negative and positive signs and only seven of the fifteen are statistically significant. Our findings are somewhat similar to those of Bahmani-Oskooee and Hegerty (2009) and Bahmani-Oskooee and Wang (2008, 2009). In general, in the long-run, exchange rate volatility appears to have mixed effect on the U.S. exports to Spain.

Table 2: Cointegration Test Results of Top Twenty Import Commodities from Spain to the U.S.

HS	Industry	F-Statistic	ECM	Cointegrated?
27	Mineral Fuel and Oil, etc.	12.73	-0.695 (7.11)	Yes
84	Machinery	4.70	-0.314 (3.44)	Yes
30	Pharmaceutical Products	1.75	-0.072 (1.70)	No
29	Organic Chemicals	4.56	-0.218 (3.42)	Yes
85	Electrical Machinery	4.93	-0.219 (3.64)	Yes
22	Beverages	21.51	-0.637 (9.33)	Yes
73	Iron and Steel Products	17.26	-0.891 (9.83)	Yes
40	Rubber	8.75	-0.354 (5.79)	Yes
20	Preserved Food	19.40	-0.480 (8.81)	Yes
39	Plastic	4.88	-0.272 (3.56)	Yes
87	Vehicles, Not Railway	2.88	-0.159 (3.41)	No
90	Optical and Medical Instruments	4.89	-0.166 (3.93)	Yes
08	Edible Fruit and Nuts	19.30	-0.702 (8.80)	Yes
72	Iron and Steel	2.60	-0.187 (2.79)	No
64	Footwear	6.18	-0.341 (4.95)	Yes
88	Aircraft and Spacecraft	2.82	-0.259 (3.29)	No
83	Miscellaneous Art of Base Metal	9.84	-0.395 (6.16)	Yes
69	Ceramic Products	2.09	-0.099 (2.87)	No
42	Leather Art, Saddlery and Bags	2.61	-0.157 (3.12)	No
71	Precious Stones and Metals	5.99	-0.405 (4.81)	Yes

Note: This table summarizes the results of the bounds testing approach to cointegration. The figures in parentheses are absolute value of the t-statistic. ECM represents the error-correction term. The upper bound critical value for the F-statistic with unrestricted intercept and no trend at the 5% level of significance is 4.35. The lower bound critical value is 3.23. These values are taken from Pesaran, Shin, and Smith (2001, Table CI(iii) Case III, p. 300).

Table 3: Short-Run and Long-Run Coefficient Estimates: Exports Model

Industry	Short-Run Coefficient Estimates					Long-Run Coefficient Estimates			
	$\Delta \ln V_t$	$\Delta \ln V_{t-1}$	$\Delta \ln V_{t-2}$	$\Delta \ln V_{t-3}$	$\Delta \ln V_{t-4}$	Constant	$\ln Y_t$	$\ln P_t$	$\ln V_t$
Misc. Grain, Seed & Fruit					0.362* (2.09)	9.916	2.940** (4.14)	-1.753** (5.34)	-0.252* (2.06)
Mineral Fuel and Oil, etc.		0.203* (2.23)				6.514	3.132** (9.55)	-0.508** (5.27)	-0.134* (1.93)
Machinery		0.101* (2.48)			0.094* (2.35)	-1.713	1.622** (5.78)	-0.170 (0.91)	-0.036 (0.73)
Organic Chemicals		-0.127 (1.43)				2.749	0.549 (1.63)	-0.222 (0.97)	0.181 (1.07)
Edible Fruit and Nuts				-0.103 (1.10)		2.673	2.035* (2.21)	-2.821** (3.70)	-0.055 (1.27)
Aircraft and Spacecraft			-0.238 (1.58)			4.406	1.927** (3.60)	-3.020** (3.10)	0.292** (3.11)
Electrical Machinery	0.110* (2.44)					1.734	0.437 (0.98)	-0.406* (2.24)	0.079 (0.97)
Optical and Medical Inst.			-0.027 (1.46)			2.250	0.964** (4.67)	-1.168** (8.09)	0.091* (2.32)
Iron & Steel Products				0.137* (2.15)		-10.524	2.601** (9.17)	-2.975** (8.71)	0.051 (1.31)
Paper and Paperboard	0.117* (2.25)					2.635	1.213 (1.50)	-0.621 (1.48)	0.031 (1.42)
Misc. Chemicals	0.118 (1.62)					-9.695	3.385** (4.11)	-2.051* (2.04)	0.175* (2.16)
Woodpulp, etc.		-0.051 (1.19)				2.370	0.787** (4.16)	-0.258 (1.21)	0.038 (1.10)
Food Waste & Ani. Feed				0.292 (1.43)		22.295	1.245** (4.19)	-4.902** (5.74)	-0.451* (1.99)
Iron & Steel			-0.296 (1.38)			-8.756	1.819 (1.40)	-2.498** (3.44)	-0.503* (1.91)
Rubber			0.072 (1.43)			-6.089	2.503** (6.99)	-0.805* (2.07)	-0.063 (1.84)

*Note: This table summarizes the results obtained using the ARDL model defined in Equation (2). The figures in parentheses are absolute value of t-statistic. ** and * indicate the statistical significance at the 1% and 5% levels, respectively.*

The estimated coefficients for the fourteen cointegrated import industries are presented in Table 4. As in the case of exports, we report only the short-run volatility coefficients and all the long-run coefficients.

Short-Run Effects of Exchange Rate Volatility on Imports: The short-run estimated coefficients on exchange rate volatility presented on the left panel in Table 4 reveal a mixture of negative and positive signs. There is also a significance variation of the exchange rate volatility on imports among industries in the short-run. Some of the coefficients are positive but only two of the coefficients are statistically significant. These industries include optical and medical instruments and footwear. The products that have negative coefficients show that these coefficients are statistically insignificant in the short-run. However, precious metal and stones industry shows a negative and statistically significant effect in the short-run.

Long-Run Effects of Exchange Rate Volatility on Imports: The long-run coefficient estimates are shown in the right panel of Table 4. As economic theory postulates, the real income variable renders a positive sign in all cases. This coefficient is statistically significant at the 1% level in seven of the industries and significant at the 5% level in two industries. The relative price variable displays the expected negative sign in all industries and is statistically significant at the 1% level in eight of the fifteen industries, and at the 5% level in two industries. This result is similar to those of Bahmani-Oskooee and Harvey (2011), Bahmani-Oskooee and Mitra (2008), Bahmani-Oskooee and Kovyryalova (2008), and Bahmani-Oskooee and Ardalani (2006). Finally, the estimated coefficients on exchange rate volatility show a mixture of negative and positive signs and only three of the fourteen are statistically significant. Our findings are somewhat similar to those of Bahmani-Oskooee and Hegerty (2009) and Bahmani-Oskooee and Wang (2008, 2009). In general, in the long-run, exchange rate volatility appears to have mixed effect on the U.S. imports from Spain.

Table 4: Short-Run and Long-Run Coefficient Estimates: Imports Model

Industry	Short-Run Coefficient Estimates					Long-Run Coefficient Estimates			
	$\Delta \ln V_t$	$\Delta \ln V_{t-1}$	$\Delta \ln V_{t-2}$	$\Delta \ln V_{t-3}$	$\Delta \ln V_{t-4}$	Constant	$\ln Y_t$	$\ln P_t$	$\ln V_t$
Mineral Fuel				0.245		1.663	1.586	-1.645**	-0.044
and Oil, etc.				(1.54)			(1.77)	(3.57)	(1.33)
Machinery					0.056	1.835	1.562*	-1.614**	-0.050
					(1.48)		(2.36)	(5.19)	(1.67)
Organic					0.074	2.084	2.404**	-2.594*	0.107
Chemicals					(1.69)		(3.99)	(2.16)	(0.89)
Electrical					-0.066	1.001	2.071*	-1.595**	0.206
Machinery					(1.21)		(2.13)	(3.52)	(1.77)
Beverages					-0.070		1.237	-2.360**	0.018
					(1.56)	18.532	(1.65)	(9.47)	(1.46)
Iron & Steel			-0.096			-9.870	3.864**	-1.809**	0.066
Products			(1.48)				(9.37)	(6.86)	(1.51)
Rubber			-0.018			-1.485	1.314**	-1.053	-0.060
			(1.49)				(3.97)	(1.17)	(1.01)
Preserved			0.016			3.896	1.093	-1.240**	-0.084**
Food			(1.56)				(1.55)	(8.04)	(2.64)
Plastic				0.080		-11.613	2.983**	-1.248	-0.148*
				(1.34)			(8.27)	(1.24)	(1.99)
Optical and		0.092*				7.106	1.593	-1.596	-0.145
Medical Inst.		(2.14)					(1.06)	(1.38)	(1.65)
Edible Fruit	-0.317					-4.993	3.997**	-1.331	-0.101
and Nuts	(1.54)						(3.70)	(1.39)	(1.57)
Footwear		0.073*				-13.336	2.112**	-1.959**	-0.218**
		(1.99)					(4.20)	(8.44)	(3.54)
Misc. Art of					-0.044	-7.710	1.952**	-2.295*	0.083
Base Metal					(1.49)		(4.72)	(2.09)	(1.30)
Precious	-0.092*					5.744	1.601	-3.179**	-0.063
Stones&Met.	(2.20)						(1.36)	(6.82)	(1.09)

*Note: This table summarizes the results obtained using the ARDL model defined in Equation (2). The figures in parentheses are absolute value of t-statistic. ** and * indicate the statistical significance at the 1% and 5% levels, respectively.*

SUMMARY AND CONCLUSIONS

In this paper we have examined the dynamic relationship between exports, imports, and exchange rate volatility in United States' trade with Spain, in the context of a multivariate error-correction model. Estimates of the long-run export and import demand functions were obtained by employing the bounds testing approach to cointegration using monthly data for the period January 1993 - December 2012. The cointegration results clearly show that there exists a long-run equilibrium relationship between real exports, real foreign economic activity, relative prices, and real exchange rate volatility, in fifteen of the twenty commodities selected. Similarly, the cointegration results clearly show that there exists a long-run equilibrium relationship between real imports, real domestic economic activity, relative prices, and real exchange rate volatility, in fourteen of the twenty commodities selected. In the long-run, all the specifications yielded expected signs for the coefficients. Most of our estimated coefficients are statistically significant either at the 1% or 5% levels. There is also a significance variation of the exchange rate volatility on exports among industries in the short-run. Some of the coefficients are positive and statistically significant. These industries include miscellaneous grain, seed and fruit, mineral fuel and oil, etc., machinery, electrical machinery, iron and steel products, and paper and paperboard. The products that have negative coefficients show that these coefficients are statistically insignificant in the short-run. There is also a significance variation of the exchange rate volatility on imports among industries in the short-run. Some of the coefficients are positive but only two of the coefficients are statistically significant. These industries include optical and medical instruments and footwear.

The products that have negative coefficients show that these coefficients are statistically insignificant in the short-run. However, precious metal and stones industry shows a negative and statistically significant effect in the short-run. These results point out to the decreasing competitiveness of U.S. exports in the global economy despite the depreciating value of the dollar over time. It underscores the degree to which a developed country such as Spain has succeeded in finding alternative markets in Europe and especially in Asia in the last decade. One of the limitations of the present study is the limited number of products included in the study. While the current study considered only top 20 export and import products, more meaningful conclusions would have been attained if the number of products are increased. Future research on the topic will cover all export and import products. Future research will also carry out the analysis by breaking the time period into two separate periods: first covering the period from January 1993 to December 1998 and the second covering the period from January 1999 to December 2013. The second period corresponds to the adoption of Euro by Spain in 1999.

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