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Exchange Rate Volatility and U.S. Auto-Industry Exports: A Panel Cointegration Approach

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ABSTRACT: Intermediate goods are often neglected in the empirical studies of the impact of exchange rate volatility on bilateral trade flows. Using import unit values of 58 motor vehicle products and 193 auto-parts, which are classified by the 10-digit level of Harmonized Tariff Schedule (HTS), this study examines the impact of exchange rate volatility on the U.S. automotive industry exports and imports (both motor-vehicle products and auto-parts) from 37 major trading partners for the period of 1996.01 to 2008.4 by using panel data cointegration techniques. We obtain substantial heterogeneity in terms of the impact of exchange rate volatility for final and intermediate goods. We also find support for the positive hypothesis that exchange rate volatility may lead to greater levels of trade.

Keywords: Exchange rate volatility; the U.S. auto-parts industry; Fragmentation; Panel econometrics
JEL Classifications: C33; F14; F31; F40

1. Introduction

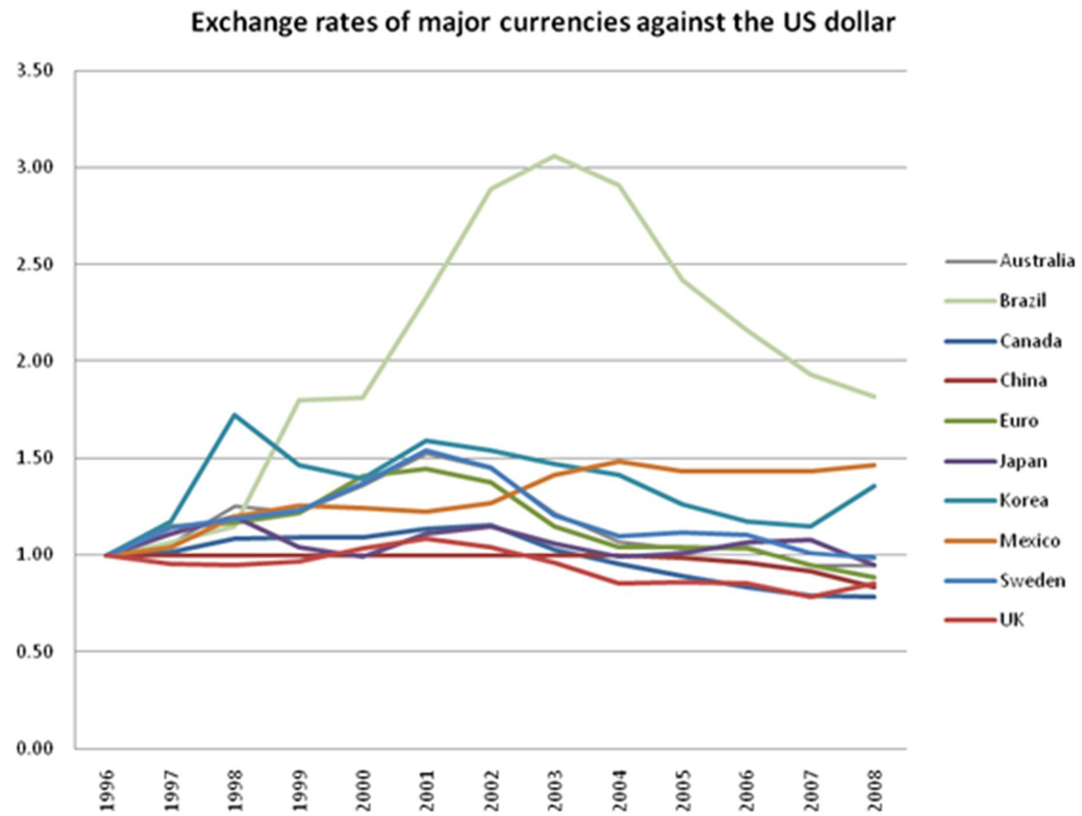
In recent years, the value of the US dollar declined against the currencies of the major trading partners of the US, such as Europe, Canada, Japan, and China. A look at the value of the US dollar shows that the U.S. dollar has gained value before the beginning of 2000, and then it has steadily lost over 25 % of its value for the last decade (Figure1). Meanwhile, the global auto industry has been undergoing significant structural transformation.[1] Among the most important and often cited trends is the increasing use of outsourcing or international fragmentation of production, leading to an increase in trade of intermediate goods as goods are designed, produced and assembled in different locations.[2]

With the increase in globalization, many automakers in the US and Europe, such as General Motors (GM), Ford, Toyota, Honda, and Volkswagen have outsourced an increasing proportion of automotive production to developing countries and emerging economies in order to reduce production costs and improve profitability.[3] Outsourcing helps companies to reduce marginal costs, but it also generates extra costs of service links between the production blocks: links in the form of communication, transportation, coordination, accounting and others (Jones and Kierzkowski, 2001). The decision to fragment production depends on a tradeoff between its extra service costs and the cost saving that can be achieved by outsourcing some of the production stages into countries where factor prices are cheaper.

[1] For a more complete analysis of trends in auto-industry, see Sadler (1999), Diehl (2001), Corswant and Fredriksson (2002), Humphrey (2003), Lall et al. (2004), and Cooney and Yacobucci (2005).
[2] International product fragmentation can be defined as division of production process into different locations across different countries.
[3] Several studies recorded the growing importance of trade in intermediate goods in the US auto-industry. See Türkcan and Ates (2010, 2011).

However, with revenues denominated in the US dollar and costs denominated in other currencies, significant and/ or sudden changes in exchange rates may reduce the locational benefits of fragmentation (Thorbecke, 2008). For instance, US companies with offshore sourcing and offshore affiliates in North America, Europe or China will loose from a declining US dollar because of soaring input, labor and shipping costs in dollar terms, which poses a remarkable risk to profits and competitive positions. In this scenario, exchange rate risk/volatility adversely affects the fragmentation based trade between the US and host country because exchange rate shifts creates risk exposures across the supply chains and thereby increase costs of service links between the production blocks.

Figure 1: The exchange rates of major currencies against the US dollar, 1996-2008. The series are indexed change in the exchange rates based on the value of 1996 (1996=1).



Empirical studies on the impact of exchange rate volatility abound the literature.[4] To the best of our knowledge, with the exception of Thorbecke (2008a, 200b), and Hayakawa and Kimura (2009), intermediate goods trade is often neglected in both theoretical and empirical studies examining the impact of exchange rate volatility or uncertainty on trade flows.[5] However, trade of intermediate goods as well as final goods increased substantially in recent years due to increase in production sharing activities. Despite the increase in intermediate goods trade, empirical evidence on the impact of the exchange rate uncertainty on the intermediate goods trade remains sparse.[6] In this study, we try to fill this gap by investigating the sensitivity of export volumes to exchange rate changes for intermediate as well as final goods in the US auto-industry, where the importance of trade based on production sharing (i.e. fragmentation) is growing, thus both final and intermediate goods trade is increasing.

[4] See McKenzie (1999) and Bahmani-Oskooee and Hegerty (2007) for an excellent review on this issue.

[5] Both Thorbecke (2008b) and Hayakawa and Kimura (2009) have investigated the relationship between exchange rate volatility and intermediate goods trade, focusing on East Asia, and found evidence that exchange rate volatility significantly reduced the flows of intermediate goods within East Asian production networks.

[6] There is a small group of papers that empirically investigate the impact of exchange rate volatility on trade flows using sectoral disaggregated data, For a survey, see Byrne (2008).

We chose automobile industry for several reasons. First, engineering industries discrete production processes that can be separated economically, so production-sharing is possible across countries. In addition, empirical evidence has shown that auto-parts and components trade has drastically increased in both exports and imports in the US (See Türkcan and Ates, 2011). The nominal value of imported parts tripled from \$31.5 billion to \$93 billion and of exported parts from \$17 billion to \$53 in the last two decades. These increases suggest that fragmentation became more prevalent in this sector. Second, the US is ranked as the world's second-largest exporter of auto-parts, following only Germany.[7] Finally, auto-parts is the one of the most important manufacturing and export sectors for the US.[8] Therefore, given its crucial importance in the global auto-industry and in the US economy, the US auto-industry has become an appropriate case to study the impacts of exchange rate volatility on the fragmentation based trade.

The aim of this study is to examine the impact of the exchange rate volatility on exports by disentangling auto-industry exports into final and intermediate goods for the US trade with 37 selected trading partners over the period 1996.02-2008.12. Hence, disaggregation of total trade volumes into final (motor vehicle products) and intermediate goods (auto-parts) component could help us to understand the implication of the volatility of the US dollar on the fragmentation based trade. This is the basic question we try to answer in this paper. In addition, the impacts of exchange rate volatility on both motor vehicle products and auto-parts exports are investigated by using non-stationary panel estimation techniques and tests for cointegration. The present study shows that the effect of exchange rate volatility is negative and significant for both motor vehicle products and auto parts, although it is strongest for exports of motor vehicle products.

The remainder of the paper is organized as follows. Next section presents the empirical model and data used in the analysis and addresses the key issues in the estimation. The regression results of the empirical model are given in Section 4. The final section draws some concluding remarks.

2. Empirical Model, Data, and Estimation

2.1. The Empirical Model

A vast majority of papers that have examined the effects of exchange rate volatility on exports have generally employed a simple model in which the quantity of exports is a function of the importing country's income, relative prices, and exchange rate volatility (See Bini-Smaghi, 1991; Chowdhury, 1993; Choudhry 2005; Baak et al., 2007; Bahmani-Oskooee and Hegerty, 2008; Arize et al., 2008; Thorbecke, 2008b).Following the convention, we specify the following equation specified to quantifythe relationship between the US auto-industry exports and exchange rate volatility:

$$x_{ijt} = \alpha_{ij} + \mu_t + \beta_1 y_{jt} + \beta_2 p_{ijt} + \beta_3 v_{ijt} + \varepsilon_{ijt} \tag{1}$$

where x_{ijt} denotes real auto-industry exports (both motor-vehicle products and auto-parts) from a country i (the US) to its trading partner j; y_{jt} is a measure of real economic activity of its trading partner j; p_{ijt} is the relative prices between the US and its trading partner and measured by the bilateral real exchange rate between the US and its trading partner j; and v_{ijt} denotes the volatility of the bilateral exchange rates between the US and its trading partner j. In addition, α_{ij} is the country effect; μ_t is the time effect; and finally ε_{ijt} is the disturbance term. All variables except the volatility of the bilateral exchange rates are in natural logarithm and the subscript t symbolizes the time.

Based on the traditional trade theory, the real economic activity of the importer is positively related to the volume of bilateral trade. (Choudhury, 2005; Arize, 2008). Therefore, we expect a positive sign for β_1. On the other hand, as Choudhury (1993) shows, an increase in the price level in

[7] US Automotive Parts Industry Annual Assessment, Department of Commerce, April 2009.

[8] In 2006, the US auto-parts industry shipments account for 4.2 per cent of total manufacturing shipments, while auto-parts employment represents 5.2 per cent of total manufacturing industries employment. Out of total manufacturing exports, the US auto-parts exports is 5.8 per cent in 2006 (US Automotive Parts Industry Annual Assessment).

the exporting country relative to the importing one will cause domestic goods to become less competitive than foreign goods, so exports are expected to be negatively related to the relative price variable. Thus, we expect the value of β_2 to be negative.

As discussed before, theoretical studies show that exchange rate volatility might have a negative effect, a positive effect or no effect at all on trade volume. A survey of theoretical models suggests that the relationship between exchange rate volatility and trade volume depend on assumptions about the nature of the risk aversion parameter, functional forms, type of trader, presence of adjustment costs, market structure, and availability of hedging opportunities (see Chit et al., 2010). Likewise, the empirical studies shows mixed, but mostly negative results on the effect of exchange rate volatility on trade volumes, partly due to the absence of a uniform definition of volatility and estimation methods (see Bahmani-Oskooee and Hegerty, 2007). Hence, the sign of β_3 could be either be positive or negative.

2.2 Data and Definition of Variables

The dependent variables in the models, measured at the 10-digit Harmonized Tariff Schedule (HTS) of the US, were derived from the web site of the United States International Trade Commission (USITC): http://www.usitc.gov. All other data are collected from the IMF International Financial Statistics.

The USITC database provides detailed monthly bilateral trade data for product exports in values (US$ at current prices) at the 10-digit level of the HTS. As briefly mentioned in the Introductory section, Equation 1 will be estimated for three product groupings: motor vehicle products (final goods), auto-parts (intermediate goods), and auto-industry (total). For the measurement of export volumes for three product groupings, we employed the list provided by the Office of Aerospace and Automotive Industries' Automotive Team. This can be found at http//www.ita.doc.gov/td/auto.html. In this study, 59 items are selected as motor vehicle products and 193 items are considered as auto-parts from the ten-digit product level of HTS. Furthermore, following the same procedure as Chit et al. (2010), data on exports of auto-industry exports denominated in current US$ are deflated using the US GDP deflator (base 2005=100) to define them in real terms.[9]

In the earlier empirical studies, the real GDP of importing country is generally used as a measure of the level of economic activity of the importing country (see Sauer and Bohara, 2001; Baak et al., 2007; and Chit, 2010). Since such a measure is not easily available on monthly base, the level of economic activity of the importing country is proxied by the industrial production index (base 2005=100) (See, for instance, Koray and Lastrapes, 1989; Choudhry, 2005; and Zhang et al., 2006).[10]

The relative prices of exports can be measured by either bilateral real exchange rate or the terms of trade. Following Baak et al. (2007) and Chit et al. (2010), in equation (1), we proxy the relative prices of exports (p_{ijt}) by the bilateral real exchange rate between the US and its trading partner. The bilateral real exchange rate in the export demand function is calculated using following equation:

$$p_{ijt} = \ln\left(e_{ijt} \times \frac{cpi_{it}}{cpi_{jt}} \right) \tag{2}$$

where p_{ijt} denotes the log of the bilateral real quarterly exchange rate between the US and its trading partner, e_{ijt} is the nominal monthly exchange rate and defined as country j currency per U.S. dollars, and cpi_{it} and cpi_{jt} and represents the monthly consumer price index (CPI) of the US and its trading partner j at time t, respectively. In terms of this definition, the increase in p_{ijt} can be interpreted as the real appreciation of the US dollar. All price indexes use 2005 as the base year.

[9] Due to lack of monthly data on the US GDP deflator, quarterly series is converted to a monthly basis by using quadratic interpolation method.

[10] Whenever the monthly series of industrial production indices are not available, quarterly series of industrial production indices are used as proxies by employing quadratic interpolation method.

Various measures of exchange rate volatility have been used in the empirical literature.[11] Initially, the standard deviation or a moving standard deviation of nominal or real exchange rates as the measure of exchange rate volatility seems to be the most commonly used method in the empirical literature. However, in recent years, a number of studies employ multiple proxies in order to ensure robustness (see, for instance, Clark et al., 2004; Chit et al., 2010). Likewise, we construct two measures of exchange rate volatility (v_{ijt}): the moving average standard deviation (MASD) of the monthly real exchange rate and the conditional variance of the first difference of the log of the real exchange rate estimated using a well-known Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model. In addition, one of the main discussions on the calculation of exchange rate volatility is whether the nominal or the real exchange rate should be used. Although the nominal exchange rate was often used at first in the literature, we employ only bilateral real exchange rates in the calculation of the volatility for both measures. [12] Thurbsy and Thursby (1987) state that that the impact of real exchange rate volatility on trade flows is not much different from that of nominal exchange rate volatility.

Following Kenen and Rodrik (1986), Koray and Lastrapes (1989), and Chowdhury (1993), the first measure for volatility used in this paper is a moving standard deviation of the monthly real exchange rate, and expressed as:

$$v_{ijt} = \left[\frac{1}{m} \sum_{k=1}^{m} \left(r_{ijt-k-1} - r_{ijt-k-2} \right)^2 \right]^{1/2} \tag{3}$$

where r_{ijt} is the bilateral real exchange rate and m is the order of moving average. The order of moving average is set equal to 12. Figure 2 presents the volatility (moving standard deviation of the monthly real exchange rate) for US-its 5 trading partner's exchange rate series. We can see that large volatility was apparent in Japan until the end of 1999 while volatility in Europe began to rise in 1999. However, by 2008 Mexico has the highest exchange rate volatility among the selected countries. On the other hand, as shown in Figure 2, China's currency has been relatively less volatile though it becomes notably more volatile in recent years.

Volatility is also proxied by the conditional variance of the first difference of the log of the real exchange rate using a GARCH (1,1) model, proposed by Bollerslev (1986).[13] As in Siregar and Rajan (2004), Chowdhury (2005) and Chit et al. (2010), we assume that the log difference of monthly real exchange rates follow a random walk with a drift:[14]

$$\Delta r_{ijt} = \alpha_o + \alpha_1 r_{ijt-1} + \mu_{ijt} \tag{4}$$

where Δr_{ijt} is the first difference of the bilateral real exchange rate between the US and its trading partner j at time t, μ_{ijt} is the error term. Conditioned on an information set at time t, denoted Ω_{ijt-1}, the distribution of the error term, μ_{ijt}, is assumed to be:

$$\mu_{ijt} \left| \Omega_{ijt-1} \sim N\left(0, h_{ijt}\right)\right. \tag{5}$$

where h_{ijt} denotes the conditional variance of the error term, μ_{ijt}, and represented by the following expression:

$$h_{ijt} = \lambda_0 + \lambda_1 \mu_{ijt-1}^2 + \lambda_3 h_{ijt-1} \tag{6}$$

where μ_{ijt-1}^2 is the one-period lag of the squared residuals generated from equation (4) and measures information about volatility from the previous period and called as the ARCH term, and h_{ijt-1} is the

[11] See McKenzie (1998) and Bahmani-Oskooee and Hegerty(2007) for the survey of the different measures of exchange rate volatility employed in the empirical literature.

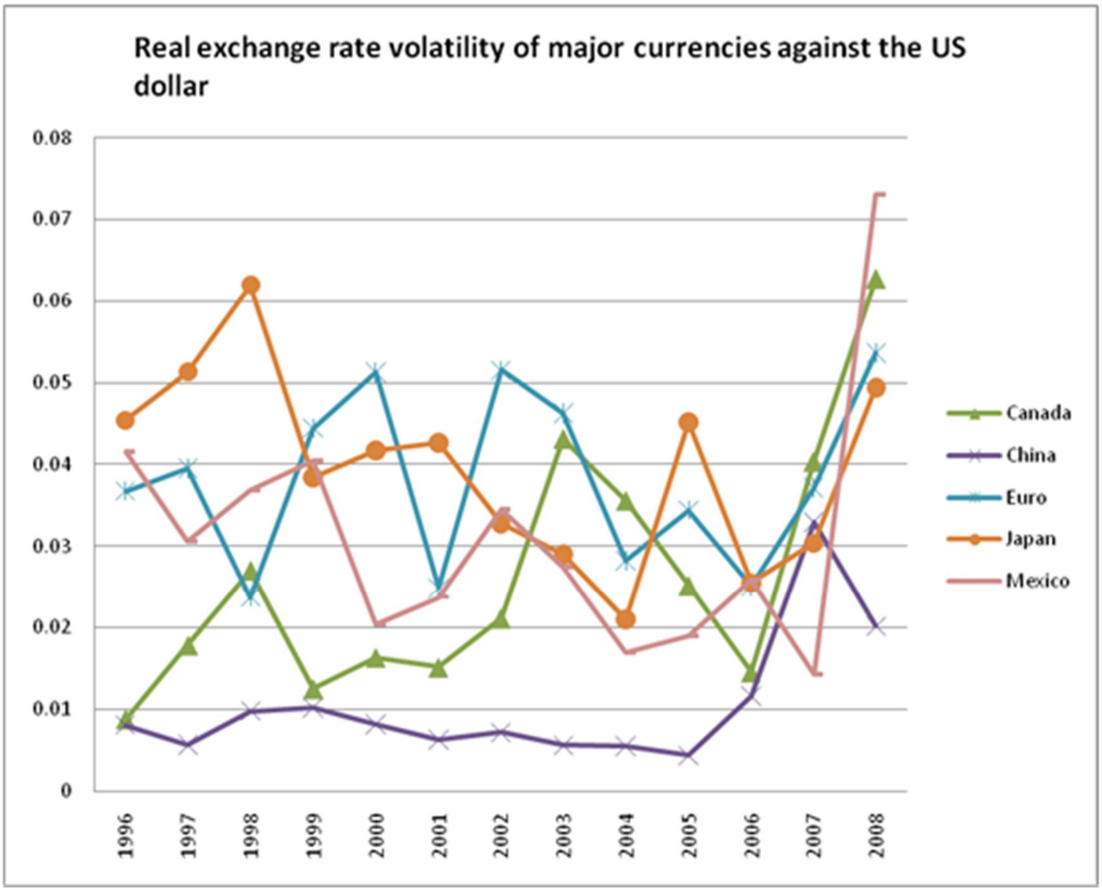
[12] Similar practice is carried out in Koray and Lastrapes (1989), Siregar and Rajan (2004), and Zhang et al. (2006).

[13] This measure is similar to those employed by Siregar and Rajan (2004), Choudhury (2005), and Chit et al. (2010).

[14] As pointed out by Bollerslev et al. (1992), the GARCH (1,1) process is generally sufficient to capture volatility dynamics for most financial and economic series.

GARCH term and represents the last period's forecast error variance. The estimated value of the conditional variance in equation (6) is then can be applied in the estimation of equation (1) as a measure of the real exchange rate volatility.[15] Descriptive statistics of both GARCH (1, 1) and MASD measurements of the real exchange rate volatility along with the real exports, relative prices, and bilateral real exchange rate are reported in Table A1.

Figure 2. MASD of the monthly real exchange rate of five major currencies against the US dollar, 1996-2008.



2.3. Econometric Procedure

a. Panel Unit Root Tests

We employ two different methods for panel cointegration estimation –fully modified ordinary least squares (FMOLS) techniques and dynamic ordinary least squares (DOLS) techniques- to the panel data from 28 major trading partners covering a period from 1996.02 to 2008.12, as in the studies of Faruqee (2004), Thorbecke (2008b), and Arize et al. (2008).

The first step in applying panel cointegration analysis is to investigate panel nonstationarity of the variables used in the equation (1). Four types of panel unit root tests, Levin, Lin, and Chu (1993) (LLC), Im, Pesaran, and Shin (2003) (IPS), Breitung (2000) and Hadri (2000) unit root tests, are employed in this paper.

With the exception of IPS test, all of the aforementioned tests assume there is a common unit root process across the relevant cross sections.[16] LLC and Breitung tests assume that the variable y is determined by following stochastic process:

$$y_{i,t} = \rho_i y_{i,t-1} + \gamma_i z'_{i,t} + \varsigma_{i,t} \qquad i = 1,.., N \quad and \quad t = 1,...T \tag{7}$$

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[15] In order to save space, the results of GARCH (1,1) model are not reported or discussed here, but interested readers can obtain the results from the authors upon request. However, it is worth noting that the GARCH(1,1) volatility measures of the real exchange rate volatility are not significantly different from the MASD volatility measures although the GARCH exchange rate volatility are generally smaller than the MASD volatility.

[16] In the literature, this is referred to as pooling the residuals along the within-dimension.

where ρ_i is an autoregressive (AR) coefficient, $z_{i,t}$ is the deterministic component and $\varsigma_{i,t}$ is the error term. The deterministic component, $z_{i,t}$, could be zero, one, units and/or time effects. It is assumed that each AR coefficient is the same for all units, $\rho_i = \rho$, that the error term $\varsigma_{i,t}$ is stationary process. By contrast, the less restrictive IPS test allows for individual unit root process; i.e. ρ_i may vary across cross-sections.[17] LLC, Breitung and IPS tests have null hypothesis of unit root (i.e. H_0: nonstationary) against alternative hypothesis that all individual series in the panel data are stationary.

Hadri (2000) proposes a Lagrange multiplier test (LM) based on the residuals. Similar to the Kwiatkowski et al. (1992) unit root test, the Hadri LM test assumes that each time series is stationary around a deterministic level or around a unit specific deterministic trend against the alternative hypothesis of a unit root in panel data (i.e. H_1: nonstationary).

Table 1. Panel Unit Root Tests

	LLC	IPS	Breitung	Hadri
Levels				
X_{ijt} (Total)	0.809 (0.790)	-2.640*** (0.004)	-1.774** (0.038)	46.521*** (0.000)
X_{ijt} (Motor Vehicles)	0.488 (0.687)	-4.237*** (0.000)	-4.371*** (0.000)	39.355*** (0.000)
X_{ijt} (Auto-Parts)	-0.618 (0.268)	-4.728*** (0.000)	-2.342*** (0.009)	32.630*** (0.000)
Y_{jt}	-0.458 (0.323)	1.138 (0.872)	0.333 (0.630)	74.399*** (0.000)
$REER_{ijt}$	-0.041 (0.516)	-1.101 (0.135)	-2.108*** (0.017)	34.801*** (0.000)
VOL_{ijt} (Real exchange rate volatility)	-2.790*** (0.002)	-13.494*** (0.000)	-6.534*** (0.000)	7.206*** (0.000)
VOL_{ijt} (GARCH volatility)	4.965 (1.000)	-4.796*** (0.000)	-4.758*** (0.000)	11.129*** (0.000)
First Differences				
X_{ijt} (Total)	-33.132*** (0.000)	-58.204*** (0.000)	-5.248*** (0.000)	-1.590 (0.944)
X_{ijt} (Motor Vehicles)	-32.186*** (0.000)	-57.559*** (0.000)	-5.258*** (0.000)	-2.881 (0.998)
X_{ijt} (Auto-Parts)	-31.070*** (0.000)	-54.680*** (0.000)	-4.432*** (0.000)	-2.207 (0.986)
Y_{jt}	-40.183*** (0.000)	-55.005*** (0.000)	-9.107*** (0.000)	0.551 (0.709)
$REER_{ijt}$	-53.157*** (0.000)	-50.403*** (0.000)	-6.279*** (0.000)	5.225*** (0.000)
VOL_{ijt} (Real exchange rate volatility)	-16.318*** (0.000)	-27.879*** (0.000)	-8.752*** (0.000)	-3.934 (1.000)
VOL_{ijt} (GARCH volatility)	-56.223*** (0.000)	-54.693*** (0.000)	-8.100*** (0.000)	-3.472 (0.999)

Notes: LLC, IPS and Breitung use the null hypothesis of unit root while the Hadri test uses null of no unit root. The test statistics are from a model that includes a constant. The choice of lag length is based on the modified AIC. The p-values are given parentheses. *,**,*** indicate statistical significance at 10, 5, and 1 % levels respectively.

Table 1 presents the results for the panel unit root tests at levels and first difference. The test for the levels and first difference is applied with a constant and the lengths of the lags included in the tests are determined by the modified Akaike information criterion. In general, the panel unit root tests produced mixed results for both dependent and explanatory variables at levels. In particular, the results

[17] In the literature, this is referred to as pooling the residuals along the between-dimension.

of the IPS test and Breitung test at the levels indicate that the null of non-stationarity is rejected except for the level of economic activity of the importing country, y_{jt}, and the relative price variable, p_{ijt}. By contrast, the LLC test and the Hadri LM test reported in Table 1 indicate that all variables are non-stationary at the level form except the MASD volatility indices. Despite the mixed results at the level form, the test results of the all four panel unit root tests using first differences indicate that all series are stationary and integrated of order one. Therefore, it is safe to conclude that series in question are not stationary and OLS estimation will result in biased and inconsistent estimates.

b. Panel Cointegration Tests

Having confirmed that variables are nonstationary and exhibit unit roots, next step is to determine whether there exists some long-run equilibrium relationship (cointegration) among auto-industry exports, foreign economic activity, relative price, and exchange rate volatility. In order to test cointegration relationship in the variables in equation (1), the methodology proposed by Pedroni (1999) is employed.

Pedroni (1999) has developed seven tests based on the residuals from the cointegrating panel regression under the null hypothesis of nonstationarity.

The method utilizes the residuals from the panel cointegration regression given by:

$$y_{i,t} = \alpha_i + \delta_{it} + \gamma_t + X'_{it}\beta_i + e_{it} \qquad i = 1,..., N \quad t = 1,..., T \tag{8}$$

$$\hat{e}_{it} = \rho_i \hat{e}_{it-1} + \xi_{it} \tag{9}$$

First four tests are based on pooling along the within-dimension and the remaining three are based on pooling along the between-dimensions. Within-dimension based statistics are known as panel cointegration statistics, which are a variance ratio test (v-statistic), a panel version of Phillips-Perron (1988) (PP) ρ-statistic and t-statistic, and ADF t-statistic. The null hypothesis is $\rho_i = 1$ against $\rho_i = \rho < 1$. Between-dimension based statistics are known as group-mean panel cointegration statistics. Group panel statistics are Phillips-Perron (1988) (PP) ρ-statistic and t-statistic and ADF t-statistic. The null hypothesis is $\rho_i = 1$ against $\rho_i < 1$.

The results of Pedroni panel cointegration tests with an intercept and a time trend using both MASD volatility measure and GARCH volatility measure in equation (1) are reported in Table 2, respectively. As shown in Table 2, all test statistics reject the null of no cointegration in every case. Consequently, it can be concluded there exists a long-run relationship among the variables used in equation (1).

Table 2. Pedroni's Cointegration Tests

	Using real exchange rate volatility			Using GARCH volatility		
Statistics	Total Auto-Industry	Motor Vehicles	Auto-Parts	Total Auto-Industry	Motor Vehicles	Auto-Parts
Panel v-stat	9.105***	8.471***	8.188***	8.120***	8.012***	-7.799***
Panel ρ-stat	-36.241***	-34.702***	-39.289***	-33.261***	-32.874***	-37.753***
Panel PP-stat	-26.996***	-26.558***	-28.738***	-25.434***	-25.637***	-27.905***
Panel ADF-stat	-9.735***	-7.540***	-9.368***	-9.591***	-6.895***	-8.258***
Group ρ-stat	-37.068***	-33.535***	-40.998***	-36.372***	-32.875***	-39.408***
Group PP-stat	-29.452***	-27.251***	-32.064***	-29.031***	-26.751***	-32.001***
Group ADF-stat	-8.723***	-9.057***	-11.144***	-8.439***	-8.680***	-8.247***

Notes: The null hypothesis is no cointegration. An intercept and a time trend were included in the cointegration regression. *, **, *** indicate statistical significance at 10, 5, and 1 % levels respectively.

c. DOLS and FMOLS

In order to estimate and test the magnitude of this long run relationship between auto-industry exports and explanatory variables, we employ two types of estimators: between-dimension (group-mean) panel fully modified ordinary least squares (FMOLS) estimator and between-dimension (group-mean) panel dynamic OLS (DOLS) estimator from Pedroni (2000, 2001).

According to Kao and Chiang (2000), conventional OLS estimator under panel cointegration cannot be used because it has a non-neglible bias in finite samples due to serial correlation and endogeneity. The problem is amplified in a panel setting by the potential dynamic heterogeneity over the cross sectional dimension. To circumvent these problems, several alternative estimation procedures such as FMOLS estimation and DOLS estimation have been proposed.

Pedroni (1996, 2000) introduce both the within dimension and between-dimension (group-mean) FMOLS estimator to overcome the endogeneity in the regressors and also to account for the dynamic heterogeneity amongst panel members. On the other hand, Kao and Chiang (1997) recommend the use of panel within-dimension DOLS estimator. The within-dimension panel DOLS estimation involves regressing dependent variable on a constant, regressors, and leads and lags of their first differences. The use of lag and lead values of the first differenced explanatory variables as additional regressors is to correct the endogeneity and serial correlation problems. In a series of Monte Carlo experiments, Kao and Chiang (2000) study the asymptotic distributions for the FMOLS, and the DOLS estimators in cointegrated regression models of panel data. Their Monte Carlo simulation results show that the within-dimension DOLS estimator exhibits beter small sample properties than the within-dimension FMOLS estimator.

More recently, Pedroni (2001) introduce between-dimension (group-mean) panel DOLS estimator and evaluates the asymptotic properties of within-dimension estimators and between-dimension estimators of the FMOLS and the DOLS. Pedroni (2001) concluded that between-dimension estimators have relatively lower small sample distortions and more flexibility in terms of hypothesis testing.

Due to these advantages, panel FMOLS and panel DOLS between-dimension (group-mean) estimators developed by Pedroni (2001) are employed in this study.[18] All the DOLS regressions were carried out using the same basic structure; i.e. adding one lag and one lead of the first difference of the right-hand variables to the equations.[19] In addition, time dummies are included in the estimation of both DOLS and FMOLS estimations. The estimation results are summarized and discussed in the next section. [20]

3. Empirical Results

The regression results of the panel group-mean DOLS and FMOLS for each of the three product groupings (total auto-industry exports, motor vehicle products exports, and auto-part exports) are documented in Tables 3 and 4 using two alternative measures of exchange rate volatility, i.e. a moving average standard deviation (MASD) measure and GARCH measure, respectively. [21] For comparison, the results of the conventional OLS estimation are also provided in Table 3 and 4. Following the arguments made in the previous section, in the remainder of the analysis discussion of the results for motor vehicle products and auto-parts will focus on those obtained using the both DOLS and FMOLS methods.[22] In general, the three methods provide similar results, suggesting that the bias introduced by neglecting non-stationary of involved variables is rather small. In addition, most of the parameters are statistically significant and have the expected signs. Furthermore, as shown in Tables 3 and 4, the estimated coefficients are qualitatively same for both volatility measures, suggesting that the results are robust across different measures of the exchange rate volatility.

Focusing first on motor vehicle products, the results reported in Tables 3 and 4 indicate that the coefficients of the importing country's income (y_{jt}) are all positive and significant in all regressions which is consistent with our predictions and also in line with the findings in previous studies. Given that the importing country's income is expressed in logarithms, the magnitude of the coefficients using the DOLS and FMOLS method reported in Table 3suggests that a one percent increase in economic activity at abroad should raise motor vehicle products exports by about 0.02-0.03 percent. It should be noted that luxury items such as motor vehicle products have a relatively high

[18] Details of these methods are available in Pedroni (1996, 2000, and 2001).

[19] The results, available on request, are robust to using higher order of leads and lags and inclusion of more than one lead and lag makes no qualitative change in the findings.

[20] Recently several studies, such as Faruqee (2004), Thorbecke (2008), and Arize et al. (2008) have used DOLS technique to investigate the impact of exchange-rate volatility on the trade flows.

[21] Country-specific DOLS and FMOLS results are available upon request from the authors.

[22] Recall that one of the question we have explored in this paper whether exchange rate volatility has any different impact on motor vehicle products exports and auto-parts exports or not.

income elasticity of demand. As such, they are very responsive to changes in incomes. However, the magnitudes of the importing country's income coefficient found in our study are quite small compared to the ones in other studies. For instance, Choudhry (2005), which examines the influence of exchange rate volatility on the real exports of the US to Canada and Japan during the period from 1974 and 1998, report income elasticities ranging from 0.30 percent to 6.2 percent.[23] Similarly, using industry level data, Bahmani-Oskooee and Hegerty (2008) evaluates the short-run and long-run impact of exchange rate volatility on the Japanese imports from the US over the period 1973-2006 and find that the long-run income elasticity is close to unity in the road motor vehicles industry.[24]

Table 3. Panel Estimation of the US Auto-Industry Exports using GARCH Volatility, 1996:02-2008:12

	Total Auto-Industry			Motor Vehicles			Auto-Parts		
Variables	**OLS**	**DOLS**	**FMOLS**	**OLS**	**DOLS**	**FMOLS**	**OLS**	**DOLS**	**FMOLS**
Y	1.042*** (6.56)	0.020*** (8.423)	0.01*** (7.01)	2.326*** (13.68)	0.027*** (7.160)	0.02*** (6.26)	0.577*** (3.30)	0.011*** (7.642)	0.01*** (5.77)
REER	-0.046*** (-3.17)	-0.383*** (-7.864)	-0.26*** (-6.87)	-0.058*** (-3.89)	-0.351*** (-8.081)	-0.23*** (-6.78)	-0.017 (-1.12)	-0.085*** (-5.743)	-0.10*** (-5.84)
VOL	-23.060*** (-8.59)	-2.852*** (-4.057)	-4.82*** (-3.08)	-33.007*** (-9.29)	-22.689*** (-5.320)	-16.74*** (-4.00)	-21.742*** (-8.15)	-5.618*** (-3.577)	-8.00*** (-2.96)
Observations per country	155	155	155	155	155	155	155	155	155
Total no. of observations	4340	4340	4340	4340	4340	4340	4340	4340	4340

Notes: In parenthesis, t-values are given. *, **, *** indicate statistical significance at 10%, 5 %, and 1 % levels respectively. FMOLS reports Pedroni (1996) between-dimension (group-mean) panel FMOLS while DOLS reports between-dimension (group-mean) panel DOLS introduced in Pedroni (2001b). One lag and one lead were selected for DOLS estimations. The coefficients for lag and lead are not reported. Time dummies are included in the estimation of both DOLS and FMOLS (not reported).

Table 4. Panel Estimation of the US Auto-Industry Exports using Real Exchange Rate Volatility, 1996:02-2008:12

	Total Auto-Industry			Motor Vehicles			Auto-Parts		
Variables	**OLS**	**DOLS**	**FMOLS**	**OLS**	**DOLS**	**FMOLS**	**OLS**	**DOLS**	**FMOLS**
Y	0.013*** (7.62)	0.022*** (9.076)	0.02*** (7.06)	0.028*** (14.84)	0.030*** (7.853)	0.02*** (6.29)	0.007*** (4.04)	0.011*** (8.240)	0.01*** (6.13)
REER	-0.047*** (-3.35)	-0.211*** (-9.825)	-0.18*** (-7.45)	-0.063*** (-4.27)	-0.722*** (-12.257)	-0.44*** (-8.40)	-0.017 (-1.13)	-0.296*** (-5.580)	-0.01*** (-5.29)
VOL	-5.754*** (-6.12)	0.903 (1.158)	0.17* (1.59)	-7.786*** (-7.88)	-4.283*** (-2.470)	-1.92 (-0.30)	-6.479*** (-6.23)	1.091** (2.029)	0.05 (0.76)
Observations per country	155	155	155	155	155	155	155	155	155
Total no. of observations	4340	4340	4340	4340	4340	4340	4340	4340	4340

Notes: In parenthesis, t-values are given. *, **, *** indicate statistical significance at 10%, 5 %, and 1 % levels respectively. FMOLS reports Pedroni (1996) between-dimension (group-mean) panel FMOLS while DOLS reports between-dimension (group-mean) panel DOLS introduced in Pedroni (2001b). One lag and one lead were selected for DOLS estimations. The coefficients for lag and lead are not reported. Time dummies are included in the estimation of both DOLS and FMOLS (not reported).

When it comes to the relative price variable (p_{ijt}), proxied by the bilateral real exchange rate between the US and its trading partner, we obtain a negative and significant impact on bilateral motor vehicle exports between the US and its trading partners for both MASD volatility measure and GARCH volatility measure and are clearly parallel to the theoretical expectations of the model. In other words, the volume of the US motor vehicle products exports tends to be lower when the foreign substitute products are relatively cheaper (meaning an increase in p_{ijt}). To gauge economic

[23] Arize et al. (2008) provide several explanations for the relatively high income elasticities found in the empirical analysis of export demand functions.

[24] Using export demand equations, Bahmani-Oskooee and Kovyryalova (2008) also report high income elasticities in the US road motor vehicle exports to the UK, respectively.

significance, consider a one percent increase in the bilateral real exchange rate (real appreciation of the US dollar). Such an increase would lead to a reduction in the exports of the US motor vehicle products ranging from 0.05 percent to 0.7 percent, depending on the estimation method and measure of volatility used. The coefficient on the relative price ratio is again quite small in our study compared to the ones obtained in Choudhry (2005) and Bahmani-Oskooee and Hegerty (2008). Bahmani-Oskooee and Hegerty (2008) show that the depreciation of the Japanese Yen would lead to a 2.89 percent reduction in the Japanese road motor vehicles imports originating from the US.

More importantly, as documented in Tables 3 and 4, the coefficient estimates for both the real exchange rate volatility indices (v_{ijt}) are statistically significant and negative in all regressions except when the panel FMOLS method is estimated using the MASD volatility measure, thus generally confirming our expectations regarding the effects of exchange rate volatility. The estimated coefficients of the volatility measure for panel DOLS using the MASD volatility measure (-4.28) indicate that a one standard deviation increase in exchange rate volatility (0.026, see Table A1) around its mean leads to 11 percent decrease in the motor vehicle exports.[25] For GARCH volatility measure, one standard deviation reduction in volatility will raise motor vehicle exports by 24 percent... It should be noted that GARCH volatility measure does tend to produce more large negative impact on the exports than MASD measure, suggesting that GARCH model systematically overstate the magnitude of the impact of the exchange rate volatility on exports as often encountered in the empirical analysis.[26] Furthermore, the evidence indicates that motor vehicle products exports react faster to changes in exchange rate volatility than changes in foreign income and relative price changes. This result is also quite consistent with the results of Arize et al (2008), Chit el al. (2010). The negative coefficients of the volatility found in this study thus support for the negative hypothesis that exchange rate volatility may depress the international trade if hedging is not possible or is costly. As mentioned earlier, this is because higher risk leads to higher cost for risk-averse traders and hence reduces the benefits of international trade (See McKenzie, 1999). In other words, an increase in exchange rate volatility tends to reduce the profits of risk-averse traders and thereby induce them to shift from risky export markets to less risky domestic markets.

Moreover, the importing country's income again consistently imposes a significant positive impact on the auto-parts exports of the US in all cases (see Tables 3 and 4), which is in line with our expectation and many recent studies. The range of the estimates is between 0.007 and 0.577. It is interesting to note that, the estimated coefficient of Y_{jt} in motor vehicle products is relatively larger than in auto-parts, implying that the impact of the income changes are more pronounced in the motor vehicle exports than auto-parts exports, which lead us to believe that intermediate goods trade is less sensitive to changes in income. This partially support commonly held knew that the income elasticity of the final goods exports is significantly larger than that of the intermediate goods exports due to its cyclical sensitivity. For instance, by decomposing trade data into components and final goods, Athukorala and Menon (2010) examine the impact of global production sharing on the price and income elasticities of world exports over the years1992 to 2006 and find that, for the machinery and transport equipment industry (SITC 7), the world income elasticity of demand is much smaller in the parts and component equation (0.19) compared to final exports (0.51). They argue that demand for components is mainly governed by the production process of user industries, whereas final demand is more closely linked with domestic income. Furthermore, reflecting the new reality of global production linkages common among today's international carmakers, Athukorala and Menon (2010) report that the income elasticity of demand for total machinery and transport equipment is 0.52 %

[25] Note that volatility measures in this paper are not defined in logarithms; hence we need a method to transform our semi-elasticity estimates of the exchange rate volatility into percent figures. Following Chit et al. (2010), the elasticity of the motor vehicle exports with respect to the exchange rate volatility measure is computed by multiplying the estimated coefficient of MASD volatility measure with one standard deviation of the MASD volatility measure over the sample period (0.026, see Table A1) and then multiply the calculated value by 100 to convert into percent.

[26] For example, the size of estimated coefficients of the GARCH volatility measure in both Siregar and Rajan (2004) and Chit et al. (2010) are considerably larger than the coefficients of the other measures of exchange rate volatility. Chit et al (2010) suggested that GARCH volatility measure from high frequency data, such as hourly or daily data, are more suitable than the low-frequency data such as monthly, quarterly or annually data.

compared with 0.82 % for total manufacturing exports and therefore confirms the hypothesis that a rapid expansion of global production sharing in the automobile industry tends to dampen the degree of sensitivity of trade flows to changes in income. The same conclusions was also reached by Thorbecke (2008a), who find that the size of the importer income coefficient for capital goods (final) exports from Japan to other Asian countries is considerably larger than for intermediate goods exports.

As was the case with motor vehicle products, an increase in the relative price variable, i.e. appreciation of the US dollar, generally leads to a decline in the US auto–parts exports. The corresponding estimates of the relative price variable ranges between -0.01 and -0.72, all being statistically significant at 5 % level, except when the OLS method is estimated using the MASD volatility measure. Moreover, the results indicate that the size of the estimated coefficient of the relative price variable for auto-parts exports is quite lower than for motor vehicle products exports, which implies that auto-parts trade is also less sensitive to the bilateral real exchange rates. This finding similar to the findings of Athukorala and Menon (2010), who report that, for machinery and transport equipment, the price elasticity of demand for components in the components and parts equation is 0.08 % compared with 0.18 % for final goods.[27] This evidence hence supports the hypothesis that global production sharing may also reduce the sensitivity of trade flows to relative price changes. [28] Athukorala and Menon (2010) provide several explanations for the relatively low sensitivity of parts and components trade to changes in relative prices.[29] First, and foremost, setting up global production networks requires high fixed costs. Once incurred, relative price and cost changes become less important in business decision making because the cost of these investments is irreversible even if the firm decides to stop the exporting activities. Second, firms engaging in production sharing exhibit a lower elasticity of substitution between home and foreign inputs relative to other firms because within global production networks, production stages located in different countries perform specific tasks which are not easily substituted elsewhere. When the auto-parts are highly substitutable, a price increase is likely to induce the motor vehicle producers in the trading partners to switch to an alternate. However, the increase in the multinational activity of the US motor vehicle producers in recent years across the globe and their preference to obtain intermediate goods from their home country may explain relatively small sensitivity of auto-parts exports to exchange rate changes. As indicated by Türkcan and Ates (2009), the low sensitivity of the US auto-parts trade to the exchange rate changes may be due to an increased prevalence of the MNCs in the US auto-industry. MNCs might have more leverage than independent firms in responding exchange rate uncertainty due to their world-wide networks.

MNCs can shield themselves against unfavorable exchange rate shocks by employing different strategies such as using intra-corporate exchange rates, manipulating the prices charged on the intra-firm transfers, rescheduling the timing of the payments on those transfers and invoicing contracts in selected currencies in order to minimize the effects of exchange rate uncertainty on the trade flows. As a result of these well established relationships between the US auto firms and their affiliates in the world, the US motor vehicle producers in the host country would not need to switch to

[27] Using import demand functions, Athukorala and Menon (2010) also investigates the sensitivity of the US machinery and transport equipment imports to changes in relative prices during 1996-2007 and find that parts and components are considerably less sensitive to changes in relative prices than final goods. Similarly, Arndt and Huemer (2007) examine the sensitivity of the US-Mexico bilateral manufacturing trade to relative price and home and foreign income changes for the period of 1989-2002 and conclude that auto-parts exports do not respond to relative price changes and exclusively determined by domestic and foreign income changes. In contrast, Swenson (2000) analyses the sensitivity of firms located in the U.S. foreign trade subzones to a dollar depreciation and found a decline in the usage of imported inputs in the production by the U.S. firms as a response to depreciation.[27]

[28] An alternative view, Byrne (2008) argues that in the case of differentiated goods it is not easy for firms to switch foreign suppliers or find new buyers in response to changes in the exchange rate because of the search costs. In contrast, for the intermediate goods search costs are minimal due to fact that intermediate goods do not vary between suppliers and can be substituted quickly and therefore the response to changes in the exchange rate will be large.

[29] For a more detailed discussion of the impact of the global production sharing on the trade flows, see Athukorala and Menon (2010).

foreign substitute products.[30] Finally, global production sharing tends to reduce the link between the domestic cost of production and export competitiveness due to the fact that any adverse changes in the exchange rates is, for instance, quickly offset by the reduction in costs that occurs as the price of imported inputs fall. [31] This may be particularly true in the automobile industry where the level of fragmentation is substantially increased in recent years.

4. Conclusion

In conclusion, the empirical evidence presented in this paper yields a mixed picture on the trading effects of exchange rate volatility on the US auto-parts exports. In particular, the estimates of FMOLS and DOLS yield quite different results across different measures of uncertainty. However, as we emphasized earlier, the GARCH measure of volatility is not suitable for low frequency data such as monthly data, and consequently tends to reflect the underlying relationship between exchange rate volatility and trade volume incorrectly. Since we used monthly exchange rate to calculate GARCH volatility in our paper, the results and thereby discussions based on the MASD measure seems to be more appropriate. As documented, the estimation results indicate that exchange rate volatility have a positive and significant effect on the US auto parts using the DOLS method, but an insignificant effect using the FMOLS method. The estimation results of for panel DOLS using the MASD measures suggest that a one standard deviation increase (0.026, see Table A1) in volatility would increase the auto-parts exports between the US and its trading partners by around 2.8 percent. These results seem to be inconsistent with recent findings of Thorbecke (2008b) who provided empirical evidence that exchange rate volatility reduces the flow of electronic components and thereby the level of fragmentation within East Asia.[32] This finding is contrary to the view that exchange rate volatility reduces the locational benefits of fragmentation and hence firm's incentives for shifting tasks across borders (Thorbecke, 2008b). However, the results presented here for the US auto-parts industry generally support Athukorola and Menon (2010) assertions, as stated above, that global production sharing tends to weaken the link between the exchange rate changes and trade volumes.[33]

There are several implications in this study. First, the findings of this present study imply that the ongoing process of product fragmentation that has shaped and continues to affect the US auto industry has also had a major influence on the relationship between volatility and trade in terms of both direction and magnitude. The results from the present study suggest that empirical studies should take this phenomenon into consideration. Last but not least, our findings imply that the impact of exchange rate volatility should be tested in the context of disaggregated trade data. However, previous empirical studies have tended to focus on the total volume of trade and ignored the characteristics of the industry the despite the fact that the impact of exchange rate volatility is not uniform between countries and industries/commodities both in terms of direction and size (McKenzie, 1999). This practice has produced ambiguous results in the previous studies. However, the results of the present study indicate that differences do exist across industries. The empirical analysis in this paper, therefore, demonstrates that the use of disaggregated bilateral trade at the product/industry level increases the likelihood of capturing the true relationship between the exchange rate volatility and trade volumes. The findings in this paper also support the theoretical claim that volatility can increase the potential gains from trade, i.e. income effects dominate the substitution effects, and thereby leading to an increase in the volume of trade accordingly.

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[30] For example, the results in Blonigen (2001) show a strong positive relationship between Japanese automobile production in the US and imported Japanese automobile imports.

[31] Greenaway et al. (2010) also find evidence of both a negative effect from appreciation and an offsetting effect through imported intermediate goods.

[32] In contrast, using the GARCH measure of volatility, results in Table 4 indicate that the effect of exchange rate volatility on the US auto exports has a negative sign and is statistically significant across different estimation techniques, which is consistent with our expectation and previous studies.

[33] Similarly, using US sector-level data over the period 1989-2001, Byrne et al (2008) shows that the effect of exchange rate uncertainty is negative and significant for the exports of the differentiated goods but insignificant for homogenous/intermediate goods.

This heterogeneity in the effect of exchange rate volatility can also be due to the transfer pricing between the US auto-makers and their affiliates in its trading partners. While it was not our focus in this work, it is an attractive area for future research.

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Appendix

Table A1. Descriptive Statistics of the Main Variables

Variable	Mean	St. Dev.	Min.	Max.	Obs.
X_{ijt} (Total)	12.601	1.886	8.230	17.695	4340
X_{ijt} (Motor Vehicles)	11.330	2.053	4.645	17.021	4340
X_{ijt} (Auto-Parts)	12.009	2.001	7.860	17.240	4340
Y_{jt}	4.539	0.154	3.846	5.080	4340
$REER_{ijt}$	1.310	1.797	-0.702	7.419	4340
VOL_{ijt} (Real exchange rate volatility)	0.038	0.026	0.001	0.274	4340
VOL_{ijt} (GARCH volatility)	0.021	0.011	0.002	0.210	4340

Notes: All variables except the real exchange rate volatility and GARCH volatility are in logs.