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Journal of Risk and Financial Management

Provided in Cooperation with:

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Suggested Citation: Lee, Chien-Hui; Li, Shu-Hui; Lee, Jen-Yu (2022) : Examining the asymmetric effects of third country exchange rate volatility on trade between the US and the EU, Journal of Risk and Financial Management, ISSN 1911-8074, MDPI, Basel, Vol. 15, Iss. 8, pp. 1-20,
<https://doi.org/10.3390/jrfm15080321>

This Version is available at:

<https://hdl.handle.net/10419/274843>

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Article

Examining the Asymmetric Effects of Third Country Exchange Rate Volatility on Trade between the US and the EU

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Abstract: This paper aims to examine the symmetric and asymmetric effects of third country exchange rate volatility on the trade flow between the US and EU from January 2003 through March 2021. The monthly disaggregated data of the top twelve export and import industries are the sample frame. We find that separating increased volatility from declines and introducing a nonlinear adjustment to the volatility shows a more significant outcome than symmetric analysis. Different industries carry distinctive behaviors regarding exchange rate risk, and the third country effect plays a vital role in trade. Moreover, increased CNY/USD real exchange rate volatility increases bilateral trade between the US and EU.

Keywords: asymmetric effects; exchange rate volatility; trade flows; disaggregated data



Citation: Lee, Chien-Hui, Shu-Hui Li, and Jen-Yu Lee. 2022. Examining the Asymmetric Effects of Third Country Exchange Rate Volatility on Trade between the US and the EU. *Journal of Risk and Financial Management* 15: 321. <https://doi.org/10.3390/jrfm15080321>

Academic Editors: Stephen Satchell and Robert Brooks

Received: 23 May 2022

Accepted: 18 July 2022

Published: 22 July 2022

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1. Introduction

Cushman (1986) proposed that bilateral trade is affected by the direct risks of bilateral exchange rates and the indirect risks of third country exchange rates. Empirical findings have shown that ignoring third country effects leads to exchange rate volatility prejudice in bilateral trade. Academia did not pay much attention to this issue until Bahmani-Oskooee and Hegerty (2007) emphasized the third country effects when researching the impact of exchange rate volatility on trade. Then, empirical studies on the third country effect became more widespread.

Following in the footsteps of Cushman (1986), Bahmani-Oskooee et al. (2013) found strong evidence of a “third country” exchange rate volatility effect by comparing the results with and without consideration of a third country. Other studies have revealed that the third country exchange rate risk does indeed play a significant role in bilateral trade (Bahmani-Oskooee et al. 2015). Wang et al. (2016) noticed that exports from China to the US (or Europe or Japan) rise concurrently with an increase in the third country exchange volatility. In similar research, Tunc et al. (2018) found that high volatility in the external exchange rate compared to the volatility of the bilateral exchange rate between the exporting country and the destination country leads the exporting country to shift its exports from the third to a bilateral country.

Moreover, Usman et al. (2021) confirmed that nonlinear models generate more significant results in both the short and long run. Other empirical studies suggest that the asymmetric assumption alone is insufficient (e.g., Choudhry et al. 2014). We also note that it is crucial to consider CNY/USD exchange rate fluctuations while evaluating bilateral trade between the US and EU. In 2020, the EU was the largest trading partner of the United States, followed by China, the largest trading country (647.6 billion USD of trade flows between the US and EU compared to 560.1 billion USD between the US and China; US Census Bureau).

The exchange rate volatility significantly impacts international trade since it influences trade decisions and overall economic performance. Therefore, the prolonged debate on

the influence of exchange rate volatility on international trade continues. Previous studies on this issue have reached different verdicts through different assumptions that cannot be validated in all cases.

The primary rationale for the idea that exchange rate volatility decreases international trade is that volatility increases trade risk between countries. An early study by [Arize \(1997\)](#) argued that an increase in exchange rate volatility would decrease trade with foreign countries. The effect of exchange rate fluctuations on imports and exports is widely accepted. Similar studies, for example, those of [Sukar and Hassan \(2001\)](#), [Ozturk \(2006\)](#), [Hayakawa and Kimura \(2009\)](#), [Ekanayake et al. \(2011\)](#), [Yakub et al. \(2019\)](#), [Dada \(2020\)](#), [Sugiharti et al. \(2020\)](#), and [Bahmani-Oskooee and Karamelikli \(2021\)](#), reinforce this causal relationship.

[McKenzie and Brooks \(1997\)](#) explored the effect of exchange rate volatility on trade between the United States and Germany. Their conclusions differ from those of previous works, as the effects of volatility were positive and statistically significant. [Broil and Eckwert \(1999\)](#) showed a positive impact of exchange rate volatility on export production in countries such as the US, where companies benefit from a large domestic market that permits them to compensate for exchange rate volatility more easily. The empirical findings of [Perée and Steinherr \(1989\)](#) and [McKenzie \(1998\)](#) support exchange rate risks that are positively related to trade.

Earlier research showed either a negative or positive link between exchange rate volatility and international trade; some studies find a biased connection between the two variables (e.g., [Bredin et al. 2003](#); [Wong and Lee 2016](#); [Nyambariga 2017](#); [Bahmani-Oskooee and Gelan 2018](#)). The ambiguous impacts of exchange rate risk expressed in previous studies remain, as in papers by [Serenis and Tsounis \(2013\)](#), [Bahmani-Oskooee and Harvey \(2017\)](#), [Senadza and Diaba \(2017\)](#), [Sharma and Pal \(2018\)](#), and [Bahmani-Oskooee and Noura \(2020\)](#).

Traditionally, studies have used aggregated data in investigating the impact of exchange rate volatility on the trade of one country with the rest of the world or the trade between two countries, including [McKenzie and Brooks \(1997\)](#) for the United States and Germany, [Choudhry \(2003\)](#) for Canada and Japan, and [De Vita and Abbott \(2004\)](#) for the effects of exchange rate volatility on US exports. [Bahmani-Oskooee and Karamelikli \(2019\)](#) employed aggregated data and found that the impact between exchange volatility and exports is insignificant; however, the volatility impact became significant when using disaggregated data, which may be due to aggregation bias. Recently, [Shin et al. \(2014\)](#) argued that traders' reactions to increased volatility could differ with decreased effects. [Chien et al. \(2020\)](#) conducted an asymmetric analysis and found that the long-run asymmetric effect of exchange rate volatility showed far higher impacts on Taiwan's exports to Indonesia than on Taiwan's imports. Similar asymmetric results can be found in [Bahmani-Oskooee and Arize \(2020\)](#) and [Bahmani-Oskooee et al. \(2021\)](#).

The previous studies' disagreement may be due to different methodologies, data sources, and economic development patterns. However, most researchers have used aggregated data and symmetric assumptions, and only a few employed disaggregated data ([Bahmani-Oskooee and Wang 2007](#)) or asymmetric analysis ([Chien et al. 2020](#)). Additionally, most studies have focused on the impact of the exchange rate fluctuations between a country and its relevant trading partners. However, a country has more than one trading partner in the real world, and bilateral trade flows could be interfered with by another important trade partner. Our research attempts to discover the asymmetric effects of real exchange rate volatility on trade flows while accounting for third countries. We employ disaggregated data and focus on the three largest economies in the world: The United States, the European Union (EU)¹, and China—which greatly impact global trade activity. The trends related to both fluctuation measures are shown in Figures 1 and 2.

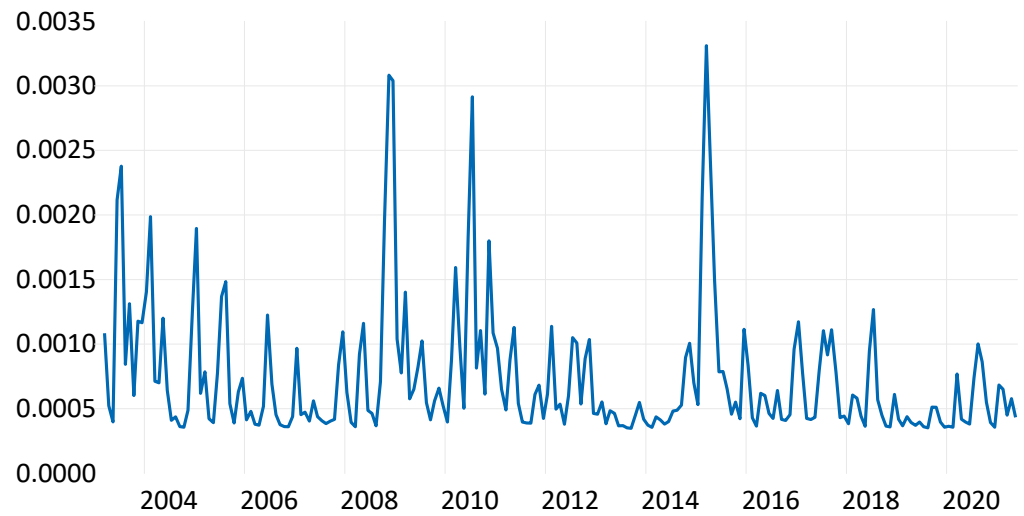


Figure 1. USD/EUR real exchange rate volatility.

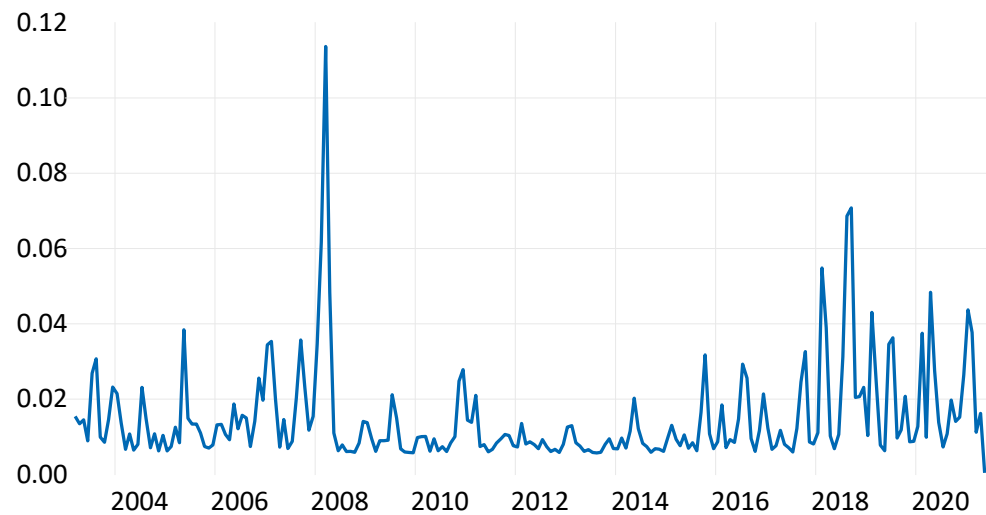


Figure 2. CNY/USD real exchange rate volatility.

Our first contribution in this paper is to add to the literature on the asymmetric effect of exchange rate volatility on international trade and compare the nonlinear autoregressive distributed lag (NARDL) outcomes with linear autoregressive distributed lag (ARDL) outcomes. The second contribution allows traders to gain better insights into the role of third country exchange rate volatility in US–EU bilateral trade. To our knowledge, few studies have used asymmetric analysis to examine the effect of the CNY/USD real exchange rate volatility on trade between the US and EU. Lastly, our paper contributes to using disaggregated data in the NARDL model. We test industries' responses to exchange rate volatility and find that different industries exhibit distinctive behaviors regarding exchange rate risk. In order to achieve our goal, we provide the model specifications in Section 2. Section 3 describes the data sources and empirical results. Finally, we present the major conclusions.

2. Model Specifications

Conventional ARDL models using bilateral trade aggregated data cannot distinguish between the different export and import effects of exchange rate changes across sectors because different industries are subject to different prices and trade contracts. They face different price rigidities and thus reveal relatively asymmetric effects of exchange rate

volatility on trade flows at the industry level. Recently, [Shin et al. \(2014\)](#) modified the ARDL model so that it could be used to assess the possibility of asymmetric effects of the exogenous variables on the dependent variable. A nonlinear ARDL model allows us to decompose the real CNY/USD exchange rate volatility into its positive and negative changes to examine the short-run and long-run asymmetric effects on trade flow between the US and EU.

In the beginning, it is assumed that the exports and imports are autoregressive processes that depend on lagged values of other economic variables. The critical variables are the import country’s income, real exchange rate, real exchange rate volatility, and the lag periods of the export and import volumes. Following [McKenzie and Brooks \(1997\)](#), we use the Industrial Production Index (IPI) as a proxy for income, and we use models by [Bahmani-Oskooee et al. \(2013\)](#) augmented with the volatility measure of a third country. The US export and import demand equations are as follows:

$$\ln EXP_{tj}^{US} = \alpha_{Xj} + \sum_{i=1}^{n1} \alpha_{Xi} \ln EXP_{t-i,j}^{US} + \sum_{i=0}^{n2} \beta_{Xi} \ln IPI_{t-i,j}^{EU} + \sum_{i=0}^{n3} \gamma_{Xi} \ln REX_{t-i,j} + \sum_{i=0}^{n4} \delta_{Xi} \ln V_{t-i,j}^{EU} + \sum_{i=0}^{n5} \zeta_{Xi} \ln V_{t-i,j}^{CN} + \varepsilon_{tj} \quad (1)$$

$$\ln IMP_{tj}^{US} = \alpha_{Mj} + \sum_{i=1}^{n6} \alpha_{Mi} \ln IMP_{t-i,j}^{US} + \sum_{i=0}^{n7} \beta_{Mi} \ln IPI_{t-i,j}^{US} + \sum_{i=0}^{n8} \gamma_{Mi} \ln REX_{t-i,j} + \sum_{i=0}^{n9} \delta_{Mi} \ln V_{t-i,j}^{EU} + \sum_{i=0}^{n10} \zeta_{Mi} \ln V_{t-i,j}^{CN} + v_{tj} \quad (2)$$

where EXP_{tj}^{US} is the export volume of commodity j from the US to the EU at time t . IPI_{t-i}^{EU} is the EU’s income at time $t - i$. IPI_{t-i}^{US} is the US income at time $t - i$. REX_{t-i} is the real USD/EUR rate at time $t - i$. V_{t-i}^{EU} is the volatility of real USD/EUR exchange rate at time $t - i$. V_{t-i}^{CN} is the volatility of the real CNY/USD exchange rate at time $t - i$. ε and v are error terms. $n1$ – $n10$ are the optimum lag periods of variables. IMP_{tj}^{US} is the US import volume of commodity j from the EU at time t .

[Bollerslev \(1986\)](#) proposed a generalized autoregressive conditional heteroskedasticity (GARCH) model, which modifies the ARCH model’s longer lag structure. It is more flexible and reasonable compared to the ARCH model, as the GARCH model can better achieve the principle of the time series model. Hence, we employ the GARCH model to estimate the real exchange volatility. The theoretical specification of a GARCH (p,q) model is as follows:

$$\Delta \ln REX_t = \eta + u_t, \quad u_t \sim N(0, V_t) \quad (3)$$

$$V_t = \alpha_0 + \sum_{i=1}^q \alpha_i u_{t-i}^2 + \sum_{i=1}^p \beta_i V_{t-i} \quad (4)$$

where REX_t follows a first-order autoregressive process. $\Delta \ln REX_t$ represents the change in the logarithm of the real exchange rate between time $t - 1$ and t . η is the mean. u_t is white noise that stands for an error term in period t and follows the normal distribution with volatility V_t . p and q are the optimum lag periods of variables, which have been defined above. Equation (3) is the mean equation. Equation (4) denotes the variation equation and addresses the volatility.

We develop an ARDL model like [Pesaran et al. \(2001\)](#) by extending Equations (1) and (2) to the cointegration model and present them as Equations (5) and (6).

$$\ln EXP_{tj}^{US} = \alpha_{Xj} + \sum_{i=1}^{n1} \alpha'_{Xi} \Delta \ln EXP_{t-i,j}^{US} + \sum_{i=0}^{n2} \beta'_{Xi} \Delta \ln IPI_{t-i,j}^{EU} + \sum_{i=0}^{n3} \gamma'_{Xi} \Delta \ln REX_{t-i,j} + \sum_{i=0}^{n4} \delta'_{Xi} \Delta \ln V_{t-i,j}^{EU} + \sum_{i=0}^{n5} \zeta'_{Xi} \Delta \ln V_{t-i,j}^{CN} + \theta_1 \ln EXP_{t-1,j}^{US} + \theta_2 \ln IPI_{t-1,j}^{EU} + \theta_3 \ln REX_{t-1,j} + \theta_4 \ln V_{t-1,j}^{EU} + \theta_5 \ln V_{t-1,j}^{CN} + \varepsilon_{tj} \quad (5)$$

$$\Delta \ln IMP_{tj}^{US} = \alpha_{Mj} + \sum_{i=1}^{n6} \alpha'_{Mi} \Delta \ln IMP_{t-i,j}^{US} + \sum_{i=0}^{n7} \beta'_{Mi} \Delta \ln IPI_{t-i,j}^{US} + \sum_{i=0}^{n8} \gamma'_{Mi} \Delta \ln REX_{t-i,j} + \sum_{i=0}^{n9} \delta'_{Mi} \Delta \ln V_{t-i,j}^{EU} + \sum_{i=0}^{n10} \zeta'_{Mi} \Delta \ln V_{t-i,j}^{CN} + \theta_6 \ln IMP_{t-1,j}^{US} + \theta_7 \ln IPI_{t-1,j}^{US} + \theta_8 \ln REX_{t-1,j} + \theta_9 \ln V_{t-1,j}^{EU} + \theta_{10} \ln V_{t-1,j}^{CN} + v_{tj} \tag{6}$$

where $\alpha, \beta, \gamma, \delta,$ and ζ are short-run coefficients. θ is the long-run coefficient.

Equations (5) and (6) are error-correction models. Trade flows are assumed to respond to changes in independent variables in a symmetric manner. Recently, [Shin et al. \(2014\)](#) modified the linear models and assessed the possibility of asymmetric effects of the independent variables on the dependent variable. Our purpose is to assess the asymmetric effects of the third country exchange rate volatility. We follow [Shin et al. \(2014\)](#) and use the following concepts to separate the increased and decreased volatilities:

$$POS_t^{CN} = \sum_{i=1}^t \Delta \ln V_i^{CN+} = \sum_{i=1}^t \max(\Delta \ln V_i^{CN}, 0)$$

$$NEG_t^{CN} = \sum_{i=1}^t \Delta \ln V_i^{CN-} = \sum_{i=1}^t \min(\Delta \ln V_i^{CN}, 0) \tag{7}$$

where POS^{CN} is the partial sum of positive changes in $\Delta \ln V^{CN}$, which reflects only increased volatility. NEG^{CN} is the partial sum of negative changes in $\Delta \ln V^{CN}$ and reflects only decreased volatility. We turn back to Equations (5) and (6) to replace $\ln V_t^{CN}$ with POS_t^{CN} and NEG_t^{CN} , then the new error-correction models can be described as:

$$\Delta \ln EXP_{tj}^{US} = a_{1j} + \sum_{i=1}^{n1} a_{2i} \Delta \ln EXP_{t-i,j}^{US} + \sum_{i=0}^{n2} a_{3i} \Delta \ln IPI_{t-i,j}^{EU} + \sum_{i=0}^{n3} a_{4i} \Delta \ln REX_{t-i,j} + \sum_{i=0}^{n4} a_{5i} \Delta \ln V_{t-i,j}^{EU} + \sum_{i=0}^{n5} a_{6i} \Delta POS_{t-i,j}^{CN} + \sum_{i=0}^{n6} a_{7i} \Delta NEG_{t-i,j}^{CN} + \pi_1 \ln EXP_{t-1,j}^{US} + \pi_2 \ln IPI_{t-1,j}^{EU} + \pi_3 \ln REX_{t-1,j} + \pi_4 \ln V_{t-1,j}^{EU} + \pi_5 POS_{t-1,j}^{CN} + \pi_6 NEG_{t-1,j}^{CN} + \varepsilon_{tj} \tag{8}$$

$$\Delta \ln IMP_{tj}^{US} = b_{1j} + \sum_{i=1}^{n7} b_{2i} \Delta \ln IMP_{t-i,j}^{US} + \sum_{i=0}^{n8} b_{3i} \Delta \ln IPI_{t-i,j}^{US} + \sum_{i=0}^{n9} b_{4i} \Delta \ln REX_{t-i,j} + \sum_{i=0}^{n10} b_{5i} \Delta \ln V_{t-i,j}^{EU} + \sum_{i=0}^{n11} b_{6i} \Delta POS_{t-i,j}^{CN} + \sum_{i=0}^{n12} b_{7i} \Delta NEG_{t-i,j}^{CN} + \rho_1 \ln IMP_{t-1,j}^{US} + \rho_2 \ln IPI_{t-1,j}^{US} + \rho_3 \ln REX_{t-1,j} + \rho_4 \ln V_{t-1,j}^{EU} + \rho_5 POS_{t-1,j}^{CN} + \rho_6 NEG_{t-1,j}^{CN} + v_{tj} \tag{9}$$

where α and b are the short-run coefficients. π and ρ are the long-run coefficients.

3. Data Sources and Empirical Results

3.1. Data Sources

This paper examines the asymmetric effects of the third country, China, on the real exchange rate volatility on commodity trade between the US and EU from January 2003 to March 2021. In order to avoid aggregated bias, the monthly disaggregated data of the top 12 export and import industries are employed². These top 12 export and import industries account for 81.90% of exports and 77.36% of imports. The data sources are reported in Table 1.

Table 1. Data sources.

Variables	Description	Data Sources
<i>EXP</i>	US Export Volume	US Census Bureau
<i>IMP</i>	US Import Volume	US Census Bureau
V^{EU}	The volatility of the real USD/EUR exchange rate	Generated by the GARCH model
V^{CN}	The volatility of the real CNY/USD exchange rate	Generated by the GARCH model
POS^{CN}	The partial sum of positive changes in $\Delta \ln V^{CN}$	Calculated from V^{CN}
NEG^{CN}	The partial sum of negative changes in $\Delta \ln V^{CN}$	Calculated from V^{CN}
IPI^{EU}	Industrial Production Index for the EU	Organization for Economic Cooperation and Development (OECD)
IPI^{US}	Industrial Production Index for the US	OECD
<i>REX</i>	Bilateral Real Exchange Rate between the US and EU	International Financial Statistics (IFS)
22	Beverages, Spirits, and Vinegar	US Census Bureau
27	Mineral Fuel, Oil, etc.; Bitumen Substances; Minerals	US Census Bureau
29	Organic Chemicals	US Census Bureau
30	Pharmaceutical Products	US Census Bureau
38	Miscellaneous Chemical Products	US Census Bureau
39	Plastics and Articles Thereof	US Census Bureau
71	Natural Pearls; Precious Stones; Precious Metals; Coins	US Census Bureau
84	Nuclear Reactors, Boilers, Machinery, etc.; Parts	US Census Bureau
85	Electric Machinery, etc.; Sound Equipment; TV Equipment	US Census Bureau
87	Vehicles, Except Railway or Tramway, and Parts, etc.	US Census Bureau
88	Aircraft, Spacecraft, and Parts Thereof	US Census Bureau
90	Optics, Photography, etc.; Medical or Surgical Instruments, etc.	US Census Bureau
98	Special Classification Provisions; Nesoi	US Census Bureau

3.2. Empirical Results

The advantage of ARDL is that there is no need to consider the order of each variable. Whether variables are I (1) or I (0) does not affect the results; moreover, it is possible to identify which variables are independent and which are dependent (Bahmani-Oskooee and Aftab 2017). However, the presence of I (2) could cause spurious estimates. Therefore, we start with a unit root test and report the results.

3.2.1. The Estimation Results of the Unit Root Tests

There are several types of unit root tests in the literature. This paper applies the augmented Dickey–Fuller (ADF) test with trend and intercept, which evaluates variables as stationary or nonstationary and adopts the Phillips–Perron (PP) robustness test. Table 2 shows that our data are stationary either in level (for example, codes 29) or first difference (for example, *REX*), and there is an absence of I (2). This means that we can apply ARDL in our estimates.

Table 2. Results of unit root tests.

Code	ADF		PP	
	Level	First Difference	Level	First Difference
	EXP (IMP)	EXP (IMP)	EXP (IMP)	EXP (IMP)
22	- (-2.0786)	- (-4.8710 ***)	- (-6.2083 ***)	- (-28.0674 ***)
27	-2.5873 (-3.8459 ***)	-23.0287 *** (-19.2528 ***)	-3.2917 * (-4.5601 ***)	-27.1056 *** (-21.0906 ***)
29	-8.4107 *** (-6.5782 ***)	-11.6154 *** (-10.9815 ***)	-8.8106 *** (-12.3140 ***)	-62.4228 *** (-66.0803 ***)
30	-2.3310 (-0.7706)	-10.1021 *** (-12.6763 ***)	-6.2343 *** (-1.5542)	-45.1170 *** (-97.5126 ***)
38	-4.2044 *** -	-17.7894 *** -	-4.0994 *** -	-20.9102 *** -
39	-4.0034 ** (-1.3543)	-4.3161 *** (-4.3175 ***)	-4.9856 *** (-1.3543)	-20.5492 *** (-4.3175 ***)
71	-3.3483 * (-1.9173)	-14.0898 *** (-4.5924 ***)	-6.9696 *** (-9.4687 ***)	-20.4012 *** (-62.9839 ***)
84	-4.3227 *** (-2.3409)	-3.9509 ** (-2.8710 *)	-7.5628 *** (-2.7223 *)	-47.0528 *** (-25.4081 ***)
85	-3.0295 (-1.5848)	-4.3546 *** (-4.2202 ***)	-7.8802 *** (-3.2098 **)	-37.5643 *** (-21.4753 ***)
87	-2.7866 (-2.1526)	-5.0567 *** (-4.3176 ***)	-6.5573 *** (-6.2841 ***)	-34.2331 *** (-27.4778 ***)
88	-7.2514 *** (-1.8822)	-12.1127 *** (-12.6688 ***)	-7.0820 *** (-8.0740 ***)	-37.3814 *** (-63.7902 ***)
90	-2.4094 (-2.1633)	-4.1103 *** (-4.6780 ***)	-6.8750 *** (-3.0940 **)	-47.5503 *** (-32.4129 ***)
98	-3.7289 ** (-0.6341)	-14.9877 *** (-11.7495 ***)	-5.2079 *** (-2.2660)	-27.0568 *** (-113.9014 ***)
$IP I^{EU}$	-3.0348	-12.2679 ***	-3.2976 *	-12.7832 ***
$IP I^{US}$	-2.0019	-11.2155 ***	-2.1451	-10.6790 ***
REX	-3.0614	-11.5114 ***	-2.9826	-11.4776 ***
V^{EU}	-8.3919 ***	-13.4585 ***	-7.7477 ***	-44.0813 ***
V^{CN}	-8.3689 ***	-10.7328 ***	-7.7496 ***	-46.6138 ***
POS^{CN}	-0.4352	-11.4992 ***	-0.0347	-15.5532 ***
NEG^{CN}	-0.4053	-12.2130 ***	-0.0317	-14.9700 ***

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Our main purpose is to examine the asymmetric effects of the third country exchange rate volatility on trade flows using Equations (8) and (9). However, we also estimate the linear models in Equations (5) and (6) for comparison. Akaike’s information criterion (AIC) is used to select an optimum model.

3.2.2. The Estimation Results of the Linear ARDL Model for Exports

We begin by estimating the linear export demand model (5) and report the results in Tables 3 and 4.

Table 3. Short-run coefficient estimates of volatility for US exports to the EU with third-country effects using the linear ARDL model (5).

Lags on $\Delta \ln V^{EU}$ and V^{CN}						
Code	$\Delta \ln V_t^{EU}$	$\Delta \ln V_{t-1}^{EU}$	$\Delta \ln V_{t-2}^{EU}$	$\Delta \ln V_t^{CN}$	$\Delta \ln V_{t-1}^{CN}$	$\Delta \ln V_{t-2}^{CN}$
84	0.0010			-0.0151	0.0207 **	
88	0.0043	-0.0499	0.0466	-0.0657 ***	0.0419 *	
90	-0.0004			-0.0156	0.0167 *	
30	-0.0067	-0.0509 *	0.0456 **	-0.0028		
85	0.0008	-0.0197 *		-0.0091	0.0131	
27	0.0120			-0.0643 **		
87	-0.0553 **			-0.0268	0.0637 ***	
29	-0.0333			0.0010		
71	-0.0314			0.0096		
39	-0.0117			-0.0074	0.0296 **	-0.0176
98	-0.0069			-0.0210 **	0.0257 **	-0.0264 ***
38	-0.0230 **			0.0082		

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 4. Long-run coefficient estimates of US exports to the EU with third country effects using the linear ARDL model (5).

Long-Run Coefficient Estimates							
Code	$\ln IPI^{EU}$	$\ln REX$	$\ln V^{EU}$	$\ln V^{CN}$	C	F-Statistic	ECM_{t-1}
84	1.6443 ***	-0.1058	0.0020	0.0109	13.9132 ***	10.0130 ***	-0.5175 ***
88	3.8739 **	0.2463	0.0035	-0.0847	3.2568	6.0488 ***	-0.2812 ***
90	1.8363	-0.2534	-0.0058	0.0147	12.8861 *	1.3108	-0.0750 ***
30	2.3209	0.3439	-0.1801	-0.0422	9.1464	1.4321	-0.0665 ***
85	2.2806 ***	-0.0131	-0.0585 *	0.0124	10.2034 ***	9.1998 ***	-0.3235 ***
27	24.0894 **	-4.2737	0.2167	-1.1576 *	-94.1988 **	3.9315 ***	-0.0556 ***
87	3.7619 ***	-0.1732	-0.1202 **	0.0803	2.8649	12.0498 ***	-0.4600 ***
29	0.5379	-0.7666 **	-0.0834	0.0251	17.4808 ***	5.2884 ***	-0.3987 ***
71	1.8282	-1.3964	-0.5106	0.1556	8.4842	1.4490	-0.0614 ***
39	2.5245 ***	0.1223	-0.0922	0.0363	7.9893 **	3.4661 ***	-0.1270 ***
98	6.5076 *	-1.1103	-0.1348	-0.4258 *	-13.0612	2.2320 *	-0.0510 ***
38	3.8603 *	-0.7263	-0.3674	0.1310	-0.0963	2.1723	-0.0625 ***

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 3 reports the short-run coefficient estimates of real exchange rate volatility in the linear ARDL model for USD/EUR and CNY/USD. Of the 12 industries, 4 show at least 1 significant coefficient for USD/EUR exchange rate volatility, while 7 out of 12 sectors show at least 1 significant coefficient for CNY/USD exchange rate volatility. The third country effect generates more significant outcomes than bilateral exchange rate volatility, meaning that this effect plays an important role in US exports to the EU.

Bounds testing is conducted to check whether a cointegration relationship exists. When the F-test is used to investigate cointegration, there are two thresholds. One is the upper-critical bound I (1), and the other is the lower critical bound I (0). If the F-statistic is higher than I (1), it means that the null hypothesis of no cointegration is rejected. If it

is lower than I (0), it means there is no significance, and the null hypothesis cannot be rejected. If the *F-statistic* result falls between the lower and upper critical bounds, then we implement the error correction model (ECM_{t-1}) and re-estimate Equations (5) or (6) and (8) or (9). If the estimated coefficient is significant, we use this as evidence for cointegration (Bahmani-Oskooee et al. 2014).

In Table 4, the ARDL bounds testing results show that the *F-statistics* are significant at the 10% level for industry codes 84, 88, 85, 27, 87, 29, 39, and 98. The results support a cointegration relationship between the variables in the long run, and the volatility of the real CNY/USD exchange rate has implications for trade between the US and EU.

The third country real exchange rate volatility estimation results show that industry codes 27 (Mineral Fuel, Oil, etc.; Bitumen Substances; Minerals) and 98 (Special Classification Provisions, Nesoi) have a significant negative effect with coefficients of 1.16 and 0.43, respectively. This implies that if the CNY/USD volatility increases by 1%, it will cause the US to decrease exports to the EU for codes 27 and 98 by 1.16% and 0.43%, respectively.

The bilateral real exchange rate volatility estimation results show that industry codes 85 (Electric Machinery, etc.; Sound Equipment; TV Equipment) and 87 (Vehicles, Except Railway or Tramway, and Parts, etc.) have a significant negative coefficient. This suggests that higher bilateral exchange rate volatility significantly decreases their export from the US to the EU. Arize (1997) also found that exchange rate risks are negatively related to trade.

The coefficients of EU income in Table 4 show a significantly positive effect in eight industries at the 10% significance level, namely codes 84, 88, 85, 27, 87, 39, 98, and 38. This means that higher income boosts United States exports to the EU.

The real exchange rate could be a price that is level compared to the trading partner, and the local currency depreciation could boost exports. The long-run empirical results indicate that the real exchange rate significantly negatively impacts US exports of code 29.

3.2.3. The Estimation Results of the Linear ARDL Model for Imports

Table 5 reports the short-run coefficient estimates of bilateral volatility for US imports from the EU with third country effects using the linear ARDL model. Six industries coded 87, 90, 98, 88, 22, and 71 carry at least one significant coefficient for $\Delta \ln V^{EU}$. However, eight industries with codes 84, 87, 29, 85, 27, 22, 71, and 39 are affected by third country exchange rate volatility. The estimation outcomes indicate that the third country effect has more significant results than bilateral exchange rate volatility, suggesting that the volatility of the real CNY/USD exchange rate cannot be ignored in trade between the US and EU.

Table 5. Short-run coefficient estimates of volatility effects on US imports from the EU with third country effects using the linear ARDL model (6).

Code	Lags on $\Delta \ln V^{EU}$ and V^{CN}						
	$\Delta \ln V_t^{EU}$	$\Delta \ln V_{t-1}^{EU}$	$\Delta \ln V_{t-2}^{EU}$	$\Delta \ln V_{t-3}^{EU}$	$\Delta \ln V_t^{CN}$	$\Delta \ln V_{t-1}^{CN}$	$\Delta \ln V_{t-2}^{CN}$
84	0.0027				-0.0282 ***	0.0242**	
30	0.0031				-0.0015		
87	0.0483 **				-0.0438 **		
29	-0.0303				0.0495 **		
90	0.0186 *				-0.0131		
85	-0.0009				-0.0302 ***	0.0274 **	
98	0.0094	0.0018	-0.0437	0.0823 ***	-0.0156		
27	0.0082				-0.0841 ***	0.0673 **	
88	0.0829 *	-0.1643 ***			0.0011		
22	-0.0172	-0.0046	-0.0495 *	0.0386 *	-0.0289 *		
71	0.0140	-0.0080	-0.0932 **	0.0874 **	-0.0180	-0.0155	0.0717 **
39	-0.0059				-0.0214 *	0.0197 *	

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 6 presents the long-run estimates of US imports from the EU with third country effects in the linear ARDL model. The coefficients of the *F*-statistic and ECM_{t-1} are significant and meaningful at the 10% level for seven industries, i.e., 84, 87, 29, 27, 88, 22, and 71. The empirical results confirm a cointegration relationship among these variables in the long run.

Table 6. Long-run coefficient estimates on US imports from the EU with third country effects using the linear ARDL model (6).

Code	Long-Run Coefficient Estimates						
	$\ln IPI^{US}$	$\ln REX$	$\ln V^{EU}$	$\ln V^{CN}$	<i>C</i>	<i>F</i> -Statistic	ECM_{t-1}
84	4.8329 ***	−0.4807	0.0217	−0.0325	0.0455	2.4935*	−0.1254 ***
30	4.1277	1.4665	0.1245	−0.0607	4.6558	1.1487	−0.0246 ***
87	4.0503 ***	0.2681	0.1174 **	−0.1064 **	3.7614	4.8061 ***	−0.4113 ***
29	0.0972	−0.1743	−0.0592	0.0968 **	20.8151 ***	6.2970 ***	−0.5118 ***
90	3.4706 **	0.2072	0.2565	−0.1808	6.6566	1.9074	−0.0726 ***
85	3.5263 ***	0.1871	−0.0083	−0.0256	4.9718	1.3350	−0.1088 ***
98	15.9842	−0.9375	1.5345	−0.4784	−42.9584	1.0348	−0.0325 **
27	4.3212 ***	−3.6576 ***	0.0248	−0.0509	0.0759	6.1983 ***	−0.3308 ***
88	2.6490 *	0.6437	−0.2488	0.0034	6.6605	3.6094 ***	−0.3268 ***
22	2.4198 **	0.5791	−0.1291	−0.1144 *	8.0779 **	3.8661 ***	−0.2528 ***
71	2.6593 **	−0.2762	0.0005	0.1144	8.4176	2.4399 *	−0.3339 ***
39	3.5143 **	0.4291	−0.0758	−0.0220	3.3761	1.0162	−0.0772 **

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

The coefficients of real third country exchange rate volatility are significant in three industries, including one positive effect, coded 29 (Organic Chemicals), and two negative effects, namely coded 87 (Vehicles, Except Railway or Tramway, and Parts, etc.) and 22 (Beverages, Spirits, and Vinegar). This shows that for import industry code 29, a 1% increase in the real CNY/USD exchange rate volatility stimulates the US traders to increase their imports from the EU by 0.10%. However, for industry codes 87 and 22, an increase of 1% of the CNY/USD volatility will induce the US to lower its EU imports by 0.106% and 0.114%, respectively.

The bilateral real exchange rate volatility estimation results show that industry code 87 (Vehicles, Except Railway or Tramway, and Parts, etc.) carries a significant positive coefficient. This implies that higher bilateral exchange rate volatility significantly increases US imports from the EU and that industry code 87 benefits from increasing the volatility.

The income coefficients are significantly positive in 9 out of 12 industries, coded 84, 87, 90, 85, 27, 88, 22, 71, and 39. This means that higher income in the US will raise these industries' imports from the EU.

For the long-run effects of the real exchange rate on the US imports from the EU with third country effects in the linear ARDL model (6), one industry, coded 27, is found to possess a significantly negative effect. This implies that USD depreciation against the EUR increases the import volume. It could be due to price inelasticity or may be because the increase in demand is more than the decrease in price caused by the depreciation, increasing the total import volume from the EU instead.

For comparison, we also conducted long-run export and import estimates on the bilateral trade between the US and EU without including third country exchange rate volatility and found that the results are pretty similar to Tables 4 and 6. Hence, we can infer that the third country effect does not affect USD/EUR volatility.

3.2.4. The Estimation Results of the Nonlinear ARDL Model for Exports

To distinguish between the linear and nonlinear model, we next test the nonlinear ARDL model results with regard to US exports and report the estimation results in Tables 7–9.

Table 7. Short-run coefficient estimates of volatility on US exports to the EU with third country effects using the nonlinear ARDL model (8).

Short-Run Volatility Coefficient Estimates													
Code	$\Delta \ln V_t^{EU}$	$\Delta \ln V_{t-1}^{EU}$	$\Delta \ln V_{t-2}^{EU}$	ΔPOS_t^{CN}	ΔPOS_{t-1}^{CN}	ΔPOS_{t-2}^{CN}	ΔPOS_{t-3}^{CN}	ΔPOS_{t-4}^{CN}	ΔNEG_t^{CN}	ΔNEG_{t-1}^{CN}	ΔNEG_{t-2}^{CN}	ΔNEG_{t-3}^{CN}	ΔNEG_{t-4}^{CN}
84	0.0045			0.0008					-0.0317 **	0.0319 **			
88	0.0183	-0.0497	0.0578 *	-0.0363 *					-0.1052 ***	0.0626 *			
90	0.0072			-0.0093	0.0002	-0.0499 **	0.0075	0.0392 **	-0.0427 **	0.0223	0.0339	0.0264	-0.0579 ***
30	0.0053	-0.0491 *	0.0555 **	-0.0010					-0.0454	0.0382			
85	0.0051	-0.0142		0.0076					-0.0232 *	0.0285 **			
27	0.0335			-0.0863 ***					-0.1015 ***				
87	-0.0490*			-0.0126	0.0866 *	-0.1181 **	0.1322***	-0.0541	-0.0833 *	0.1127 **			
29	-0.0295			-0.0070	0.0686	-0.1410 **	0.0753 *		-0.0060				
71	-0.0169			-0.0071					-0.0173				
39	-0.0130			-0.0003					-0.0210	0.0476 **	-0.0333 **		
98	-0.0007			-0.0047	0.0288	-0.0571 ***			-0.0351 ***				
38	-0.0025			0.0058					-0.0011				

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 8. Long-run coefficient estimates on the US exports to the EU with third country effects in the nonlinear ARDL Model (8).

Code	(Export Share %)	$\ln IPI^{EU}$	$\ln REX$	$\ln V^{EU}$	POS^{CN}	NEG^{CN}	C
84	12.15	1.5775 ***	-0.1957 *	0.0085	0.0015	0.0003	14.1584 ***
88	11.59	2.6110 ***	-0.5170	0.0591	-0.0814 *	-0.0954 *	9.3092 ***
90	10.51	1.1468 ***	-0.6048 ***	0.0165	-0.0281	-0.0414	15.7751 ***
30	8.80	0.2827	-0.8675 *	0.0524	-0.0045	-0.0318	19.3489 ***
85	8.35	1.7541 ***	-0.2410 **	-0.01789	0.0150	0.0104	12.7014 ***
27	7.13	5.3527 ***	-4.9516 ***	0.1633	-0.4200 ***	-0.4943 ***	-5.6105
87	6.22	3.2505 ***	-0.7764 ***	-0.0894 **	0.0620	0.0536	4.7458 **
29	5.51	0.4323	-0.9866 ***	-0.0682	-0.0092	-0.0138	17.8190 ***
71	3.15	2.0227 ***	-1.9288 ***	-0.0505	-0.0212	-0.0519	9.2562 ***
39	3.13	1.3632 ***	-0.5675 ***	-0.0277	-0.0007	-0.0144	13.1713 ***
98	2.79	2.3226 ***	-1.2110 ***	-0.0046	-0.2158 **	-0.2296 **	8.4935 **
38	2.57	2.0686 ***	-1.1541 ***	-0.0072	0.0165	-0.0033	9.6043 ***

Note: ***, **, * denote significance at the 1%, 5%, 10% levels, respectively.

Table 9. Diagnostics in the nonlinear ARDL export demand model (8).

Code	Diagnostics						
	F-Statistic	ECM_{t-1}	LM	RESET	CSM(SQ)	Wald-SR	Wald-LR
84	8.8694 ***	-0.5287 ***	1.9227	1.7533	S(S)	5.2690 *	1.9504
88	10.8440 ***	-0.4457 ***	1.0139	3.5968 *	S(S)	3.4055 *	24.4481 ***
90	7.6505 ***	-0.4362 ***	1.9334	8.1295 ***	US(S)	11.1704 ***	145.0658 ***
30	2.0472	-0.2238 ***	6.0101 ***	4.0502 **	S(S)	2.5870	39.8624 ***
85	10.8691 ***	-0.5092 ***	0.8713	0.2715	S(S)	5.1811 **	26.4939 ***
27	5.4159 ***	-0.2054 ***	0.8793	0.1313	S(S)	15.2927 ***	78.1205 ***
87	12.0988 ***	-0.5477 ***	0.8061	0.0537	S(S)	0.7664	15.0903 ***
29	4.5388 ***	-0.4328 ***	0.9980	0.5442	S(US)	0.0000	2.0511
71	3.9798 ***	-0.3348 ***	4.5988 **	1.2110	S(S)	18.4150 ***	92.273 ***
39	10.5797 ***	-0.4699 ***	1.8526	2.2994	S(S)	0.3027	176.4649 ***
98	3.0985 **	-0.1531 ***	1.0693	0.0978	S(S)	10.8352 ***	19.1761 ***
38	6.3268 ***	-0.3501 ***	1.4506	0.0009	S(S)	31.0794 ***	168.5321 ***

Note: ***, **, * denote significance at the 1%, 5%, 10% levels, respectively. Wald-SR refers to the Wald test for short-run symmetry, while Wald-LR denotes the Wald test for long-run symmetry. S stands for stable at 5% significance, and US stands for unstable.

Table 7 reports the short-run coefficient estimates of volatility on US exports to the EU with third country effects using the nonlinear ARDL model. At least one significant short-run estimate related to increased or decreased CNY/USD volatility measures is found in nine industries with codes 84, 88, 90, 85, 27, 87, 29, 39, and 98. This is two more industries than were found to be affected when using the previous estimation of the linear export model. This significant increase in the number of industries affected may be due to the separation of positive and negative changes in the nonlinear model leading to capturing the positive and negative exchange rate fluctuations simultaneously. Moreover, only three industries, coded 88, 30, and 87, are affected by USD/EUR volatility when using this model. This means that the volatility of the real CNY/USD exchange rate plays an important role in trade between the US and EU.

Table 8 reports the long-run coefficient estimates on US exports to the EU with third country effects using the nonlinear ARDL model.

The positive and negative volatilities of the CNY/USD rate have a significant effect on three industries, which is one more than when using the linear model. The second-largest industry is code 88 (Aircraft, Spacecraft, and Parts Thereof), with an export share of 11.59%. Its POS^{CN} coefficient is -0.0814. This result confirms that a 1% increase in the CNY/USD

volatility hurts the exports of bilateral trade by 0.08%. Moreover, the NEG^{CN} coefficient is -0.0954 , implying that a 1% decrease in the CNY/USD volatility increases bilateral trade export volumes by 0.10%; for a similar study, see [Usman et al. \(2021\)](#). As for codes 27 (Mineral Fuel, Oil, etc.; Bitumen Substances; Minerals) and 98 (Special Classification Provisions, Nesoi), the POS^{CN} coefficients are -0.4200 and -0.2158 , which means that a 1% increase in the CNY/USD volatility hurts the exports of bilateral trade by 0.42% and 0.22%, respectively. The NEG^{CN} coefficients are -0.4943 and -0.2296 , which suggests that a 1% decrease in the CNY/USD volatility benefits the exports of bilateral trade by 0.49% and 0.23%, respectively. This implies that the third country exchange rate volatility positively correlates with the US exports to the EU. If CNY/USD exchange rate fluctuations increase, US exports to the EU will be boosted in the long run. Similar results can be found in [Tunc et al. \(2018\)](#). The effect of the volatility of the USD/EUR rate on industry 87 (Vehicles, Except Railway or Tramway, and Parts, etc.) has a significant negative effect (coefficient = -0.0894). This means that a 1% increase in bilateral volatility leads to a 0.09% decrease in US exports to the EU.

The expected sign of $\ln IPI^{EU}$ is positive, i.e., increasing EU incomes leads the US to export more goods to the EU. Our empirical analysis shows significantly positive effects in ten industries, namely codes 84 (Nuclear Reactors, Boilers, Machinery, etc.; Parts), 88 (Aircraft, Spacecraft, and Parts Thereof), 90 (Optical, Photography, etc.; Medical or Surgical Instruments, etc.), 85 (Electric Machinery, etc.; Sound Equipment; TV Equipment), 27 (Mineral Fuel, Oil, etc.; Bitumen Substances; Minerals), 87 (Vehicles, Except Railway or Tramway, and Parts, etc.), 71 (Natural Pearls, Precious Stones, Precious Metals; Coins, etc.), 39 (Plastics and Articles Thereof), 98 (Special Classification Provisions, Nesoi), and 38 (Miscellaneous Chemical Products).

The local currency depreciation will cause domestic goods to be cheaper than foreign goods, thus pushing up exports. We expect the coefficient of $\ln REX$ to be negative. Our empirical results support the expectation, and significantly negative effects are found in eleven industries, namely codes 84 (Nuclear Reactors, Boilers, Machinery, etc.; Parts), 90 (Optic, Photo, etc., Medic or Surgical Instruments, etc.), 30 (Pharmaceutical Products), 85 (Electric Machinery, etc.; Sound Equip; Tv Equip), 27 (Mineral Fuel, Oil, etc.; Bitumin Subst; Mineral), 87 (Vehicles, Except Railway or Tramway, and Parts, etc.), 29 (Organic Chemicals), 71 (Natural, etc. Pearls, Precious, etc. Stones, Precious Metals, etc.; Coin), 39 (Plastics and Articles Thereof), 98 (Special Classification Provisions, Nesoi), and 38 (Miscellaneous Chemical Products).

3.2.5. Diagnostic Test Results in the Nonlinear ARDL Model for Exports

Table 9 reports the diagnostics in the nonlinear ARDL export demand model. The empirical bounds testing and error correction model results reveal a cointegration relationship among the variables. Out of 12 industries, 11 are cointegrated and meaningful. LM is the Lagrange Multiplier statistic to test autocorrelation, and a maximum of 12 lags is adopted. Most LM statistics are insignificant, implying that most models are autocorrelation-free. Most $RESET$ statistics are insignificant, which means that most models are correctly specified. The $CUSUM$ and $CUSUMSQ$ tests are significant at the 5% critical bound and reject the structural instability of estimates in most industries. Furthermore, the strong evidence of asymmetric by the Wald test rejects the null hypothesis $\sum \hat{a}_{6i} = \sum \hat{a}_{7i}$ for the short-run, $\hat{\pi}_5 = \hat{\pi}_6$ for the long-run model (8) and $\sum \hat{b}_{6i} = \sum \hat{b}_{7i}$ for the short-run, $\hat{\rho}_5 = \hat{\rho}_6$ for the long-run model (9). The outcomes of Table 9 indicate that both short-run and long-run are significant, at least at the 10% level, and the symmetry assumption can be rejected by the Wald test in eight and ten industries, respectively.

3.2.6. The Estimation Results of the Nonlinear ARDL Model for Imports

In this section, we present the results of the nonlinear ARDL for the import model (9) subsequent to the nonlinear ARDL for export earlier and report the results in Tables 10–12.

Table 10. Short-run coefficient estimates of volatility on the US imports from the EU with third country effects in the nonlinear ARDL model (9).

Short-Run Volatility Coefficient Estimates													
Code	$\Delta \ln V_t^{EU}$	$\Delta \ln V_{t-1}^{EU}$	$\Delta \ln V_{t-2}^{EU}$	$\Delta \ln V_{t-3}^{EU}$	$\Delta \ln V_{t-4}^{EU}$	ΔPOS_t^{CN}	ΔPOS_{t-1}^{CN}	ΔPOS_{t-2}^{CN}	ΔNEG_t^{CN}	ΔNEG_{t-1}^{CN}	ΔNEG_{t-2}^{CN}	ΔNEG_{t-3}^{CN}	ΔNEG_{t-4}^{CN}
84	0.0021					-0.0158	0.0244	-0.0441 **	-0.0612 ***	-0.0058	0.0410 **	0.0162	-0.0307 **
30	0.0067					0.0076			-0.0120				
87	0.0492 **					-0.0433 **			-0.0436 **				
29	-0.0296					0.0565 ***			0.0013	0.0539			
90	0.0165					-0.0167 *			-0.0254 ***				
85	0.0098					-0.0098			-0.0586 ***	0.0451 ***			
98	0.0202	0.0039	-0.0385	0.0910 ***		0.0005			-0.0215				
27	0.0003					-0.1319 ***	0.1239 **		-0.0068				
88	0.0869 *	-0.1175 **				0.0159	0.0043	-0.1483 *	-0.0456	-0.0760	0.1067	-0.1281 **	
22	-0.0097					-0.0228			-0.0333 **				
71	0.0284	0.0001	-0.0642	0.0583	0.0580	-0.0262			-0.0364				
39	-0.0014					-0.0166 *			-0.0233 **				

Note: ***, **, * denote significance at the 1%, 5%, 10% levels, respectively.

Table 11. Long-run coefficient estimates on US imports from the EU with third country effects using the nonlinear ARDL model (9).

Code	Import Share %	$\ln IPI^{US}$	$\ln REX$	$\ln V^{EU}$	POS^{CN}	NEG^{CN}	C
84	16.28	3.0570 ***	−0.5447 ***	0.0046	−0.0777 ***	−0.0885 ***	7.8334 ***
30	12.53	0.5397 *	0.2039	0.0119	0.01342	−0.0212	18.7473 ***
87	11.58	3.9741 ***	0.2239	0.1217 *	−0.1069 **	−0.1078 **	4.5540
29	6.67	−0.2081	−0.2788	−0.0493	0.0942 **	0.0919 **	21.7854 ***
90	6.60	1.4762 ***	−0.3290 ***	0.0274	−0.0277 *	−0.0420 ***	14.3434 ***
85	6.03	1.9003 ***	−0.1943	0.0303	−0.0304	−0.0419	12.3994 ***
98	4.19	0.6495	−0.3005	0.1291 **	0.0009	−0.0363	17.8343 ***
27	3.89	4.5172 ***	−3.3241 ***	0.0008	−0.0233	−0.0199	−0.5435
88	2.90	2.2953 ***	−0.2789	−0.0373	−0.1565 **	−0.1745 ***	9.1278 ***
22	2.88	1.5188 ***	−0.1789	−0.0123	−0.0290	−0.0423 **	13.0726 ***
71	2.08	2.3121 ***	−0.9479 ***	0.1056	−0.0343	−0.0477	9.7998 ***
39	1.73	2.0333 ***	0.0287	−0.0033	−0.0382 *	−0.0537 **	10.2658 ***

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 12. Diagnostics in the nonlinear ARDL import demand model (9).

Diagnostics							
Code	F-Statistic	ECM_{t-1}	LM	RESET	CSM(SQ)	Wald-SR	Wald-LR
84	8.9789 ***	−0.4568 ***	1.7695	1.3872	S(S)	2.4036	99.6392 ***
30	5.3399 ***	−0.5656 ***	1.0763	0.3644	S(S)	30.8561 ***	685.3393 ***
87	4.1153 ***	−0.4047 ***	4.6378 **	7.8975 ***	S(S)	0.1001	0.0965
29	2.6912 **	−0.1113 ***	1.5402	6.3643 **	S(S)	1.0523	1.0705
90	6.9633 ***	−0.6042 ***	0.1190	2.7887 *	S(S)	37.3727 ***	277.7342 ***
85	4.5122 ***	−0.3233 ***	0.6873	1.3849	S(US)	9.0488 ***	42.7281 ***
98	5.8005 ***	−0.5935 ***	0.8053	0.0187	S(US)	34.2494 ***	376.5459 ***
27	5.8955 ***	−0.3420 ***	1.4678	1.3799	S(S)	4.4972 **	0.4772
88	22.2503 ***	−0.8192 ***	1.1902	0.7947	S(S)	0.0458	56.1025 ***
22	13.7898 ***	−0.7855 ***	1.2310	12.0802 ***	US(S)	52.5791 ***	150.6753 ***
71	6.0020 ***	−0.7641 ***	2.6424 **	0.2891	S(S)	30.3843 ***	43.5830 ***
39	8.5275 ***	−0.4348 ***	0.8443	1.8100	S(S)	42.4359 ***	129.4087 ***

Notes: ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. Wald-SR refers to the Wald test for the short-run symmetry. Wald-LR denotes the Wald test for long-run symmetry.

Table 10 shows short-run coefficient estimates of volatility on the US imports from the EU with third country effects in the nonlinear ARDL model. At least one significant short-run estimate, attached with increased or decreased CNY/USD volatility measures, is found in nine industries coded 84, 87, 29, 90, 85, 27, 88, 22, and 39, which is one more than the previous estimate of the linear import model. Furthermore, three industries coded 87, 98, and 88 are affected by USD/EUR volatility. It is clear that the impact of the CNY/USD volatility on US imports from the EU is considerable.

The long-run coefficient estimates on the US imports from the EU with third country effects in the Nonlinear ARDL model are reported in Table 11. We notice that increased or decreased third country volatility of CNY/USD rate significantly affects imports in seven industries, namely coded 84 (Nuclear Reactors, Boilers, Machinery, etc.; Parts), 87 (Vehicles,

Except Railway or Tramway, and Parts, etc.), 29 (Organic Chemicals), 90 (Optic, Photo, etc., Medic or Surgical Instruments, etc.), 88 (Aircraft, Spacecraft, Additionally, Parts Thereof), 22 (Beverages, Spirits and Vinegar), and 39 (Plastics and Articles Thereof). However, only three industries carry a significant estimate in the linear model. The asymmetric estimate results show that significant coefficients are negative in 84, 87, 90, 88, 22, and 39. We could explain them by the sample of the largest industry code 84, with an import share of 16.28%. The estimated POS^{CN} coefficient is -0.0777 , which implies that a 1% increase in the CNY/USD volatility hurts the exports of bilateral trade by 0.08%. Furthermore, the estimated NEG^{CN} coefficient is -0.0885 , suggesting a 1% decrease in the CNY/USD volatility increases bilateral trade export volumes by 0.09%. For industry 29, both significant coefficients are positive. The POS^{CN} coefficient is 0.0942. It means that a 1% increase in the CNY/USD volatility boosts the imports of bilateral trade by 0.09%. Moreover, the estimated NEG^{CN} coefficient is 0.0919, representing that a 1% decrease in the CNY/USD volatility increases the import volumes of bilateral trade by 0.09%. These estimates appear that decreasing CNY–USD exchange rate volatility will cause the US to reduce imports from the EU. Wang et al. (2016) showed a similar result. The bilateral volatility coefficients of USD/EUR are significant in two industries, namely 87 (Vehicles, Except Railway or Tramway, and Parts, etc.), and 98 (Special Classification Provisions, Nesoi). Both significant coefficients are positive. This suggests that increased bilateral exchange rate volatility will boost US imports from the EU.

It is expected that the coefficient of the US income is positive. The empirical analysis shows that the coefficient is significantly positive in ten industries, namely 84 (Nuclear Reactors, Boilers, Machinery; Parts), 30 (Pharmaceutical Products), 87 (Vehicles, Except Railway or Tramway, and Parts, etc.), 90 (Optics, Photography, etc., Medical or Surgical Instruments, etc.), 85 (Electric Machinery, etc.; Sound Equipment; TV Equipment), 27 (Mineral Fuel, Oil, etc.; Bitumen Substances; Minerals), 88 (Aircraft, Spacecraft, and Parts Thereof), 22 (Beverages, Spirits, and Vinegar), 71 (Natural Pearls, Precious Stones, Precious Metals, Coins), and 39 (Plastics and Articles Thereof). This means that higher US income leads the US to increase imports from the EU.

The coefficients of real exchange rates show significant negative effects on US imports from the EU in four industries, namely 84 (Nuclear Reactors, Boilers, Machinery, etc.; Parts), 90 (Optics, Photography, Medical or Surgical Instruments, etc.), 27 (Mineral Fuel, Oil, etc.; Bitumen Substances; Minerals), and 71 (Natural Pearls, Precious Stones, Precious Metals, Coins). This means that the USD's depreciation against the EUR increases the import volume. We can infer that due to price inelasticity, the increase in demand is more significant than the decrease in the price caused by depreciation.

3.2.7. Diagnostic Test Results in the Nonlinear ARDL Model for Imports

Table 12 reports the diagnostics in the nonlinear ARDL import demand model. In order to avoid spurious estimates, we perform bounds testing and use an error correction model. According to the F-statistics and coefficients of ECM, the results reveal that a meaningful cointegration relationship between these variables is found for all twelve industries at least at the 5% level of significance. Most LM and RESET statistics are insignificant, meaning that most models are autocorrelation-free and correctly specified. The CUSUM and CUSUMSQ tests reject the structural instability of estimates in most industries. Our empirical symmetry test results show that the Wald tests are significant, at least at the 10% level. Moreover, rejecting the long-run symmetry in nine sectors and the short-run symmetry in eight. To date, most empirical analyses have provided strong evidence of asymmetric assumptions.

Our sample is taken from January 2003 through March 2021. The period included the Global Financial Crisis. In order to investigate the effect of the financial crisis, we added a dummy variable (D08) during 2008 in our estimates. We found that the estimated results with and without the dummy did not exhibit a significant difference in exchange rate volatility. Appendices A and B are the long-run coefficient estimates of the US exports to and imports from the EU with dummy variables in the nonlinear ARDL models.

4. Conclusions

This paper demonstrates the symmetric and asymmetric effects of third country real exchange rate (CNY/USD) volatility on commodity trade between the US and EU. The nonlinear ARDL models are employed in order to compare them with the linear models.

Our research reveals that the third country exchange rate risk plays a significant role in bilateral trade between the US and EU. The bounds-testing approach to cointegration and error correction terms are significant, and the Ramsey statistics are insignificant in most industries, which provides robust evidence. This implies that bilateral trade has a long-run equilibrium relationship with income, real exchange rate, and bilateral and third country exchange rate volatility and that most models are specified correctly. Furthermore, substantial evidence of asymmetry by the Wald test is identified in most industries.

When comparing the linear and nonlinear ARDL models, our results show that the inclusion of the third country exchange rate effect does not disturb the results in the case of bilateral trade, and the nonlinear adjustment of the volatility has a more significant outcome than the ARDL model, which may prevent positive changes from being canceled out by negative changes in the linear model.

Different industries exhibit distinctive behaviors regarding exchange rate risk. However, in the export sector, US exports to the EU are hurt by an increase in bilateral real exchange rate volatility, in which case the US exports less to the EU. Third country exchange rate volatility has significant negative impacts, and the coefficient of negative changes is greater than that of positive changes. This implies that an increase in the real exchange rate (CNY/USD) volatility will boost US exports to the EU in the long run. In the import sectors, US imports from the EU will benefit from increasing bilateral exchange rate volatility, and the US will import more goods from the EU to create more profit in the future. In addition, for most industries, increased CNY/USD exchange rate volatility will raise the trade volume between the US and EU. The nonlinear ARDL model provides more significant results than the linear ARDL model since the increased and decreased volatility may cancel each other out in the symmetric ARDL model. We decompose the volatility into positive and negative changes and present more detailed results to reflect the actual circumstances.

Our empirical analysis provides valuable advice to policymakers who should not ignore the third country exchange rate fluctuations when dealing with US–EU trade friction. US traders who manage potential risks in global trade should pay attention to both the bilateral and third country exchange rate risks simultaneously.

The limitation of this paper is the trade flows include all of the EU, but the volatility measures are based on the USD–EUR exchange rate only. Thus, it seems to be using an exchange rate volatility measure that may not accurately reflect the currencies of all the EU countries. In future and extended research: (1) We may consider the impact of exchange rate volatility on the trade flows of the EUR-originating countries; (2) We should try to aggregate those predictions over all the industries and generate a predicted change in overall US–EU trade.

Author Contributions: Conceptualization, Methodology, Supervision. (C.-H.L.); Data curation, Software, Formal analysis, Writing—Original draft preparation. (S.-H.L.); Validation, Investigation, Writing—Reviewing and Editing, Supervision (J.-Y.L.). All authors have read and agreed to the published version of the manuscript.

Funding: This research received no external funding.

Conflicts of Interest: The authors declare no conflict of interest.

Appendix A

Table A1. Long-run coefficient estimates on the US exports to the EU with dummy and third country effects in the nonlinear ARDL model (8).

Code	(Export Share %)	$\ln IPI^{EU}$	$\ln REX$	$D08$	$\ln V^{EU}$	POS^{CN}	NEG^{CN}	C
84	12.15	1.5489 ***	-0.1763	0.0148	0.0073	0.0011	-0.0002	14.8247 ***
88	11.59	2.6791 ***	-0.5738	-0.0487	0.0656	-0.0795 *	-0.0935 *	9.0323 ***
90	10.51	1.1464 ***	-0.6044 ***	0.0003	0.0164	-0.0281	-0.0414	15.7767 ***
30	8.80	0.1303	-0.3858	0.3761*	0.0974	-0.4234	-0.0707	20.5243 ***
85	8.35	1.6925 ***	-0.2056*	0.0301	-0.0202	0.0143	0.0098	12.9735 ***
27	7.13	5.5512 **	-5.0997 ***	-0.1345	0.1761	-0.4166 ***	-0.4909 ***	-6.4550
87	6.22	2.7232 ***	-0.2586	0.3654 ***	-0.1100 ***	0.0433	0.0350	7.1458 **
29	5.51	1.0004	-1.3928 ***	-0.2964 **	0.0070	-0.0216	-0.0272	15.6448 ***
71	3.15	2.5084 ***	-2.1780 ***	-0.2214	-0.0241	-0.0195	-0.0502	7.1744 **
39	3.13	1.4402 ***	-0.7199 ***	-0.1114 *	-0.0213	0.0102	-0.0036	12.8308 ***
98	2.79	2.4353 ***	-1.2830 ***	-0.0626	0.0018	-0.2132 **	-0.2270 **	8.0108 *
38	2.57	1.4235 ***	-0.7876 ***	0.2774 ***	-0.0181	0.0044	-0.0154	12.5460 ***

Note: ***, **, and * denote significance at the 1%, 5%, 10% levels, respectively.

Appendix B

Table A2. Long-run coefficient estimates on the US imports from the EU with dummy and third Country effects in the nonlinear ARDL model (9).

Code	Import Share %	$\ln IPI^{US}$	$\ln REX$	$D08$	$\ln V^{EU}$	POS^{CN}	NEG^{CN}	C
84	16.28	3.1616 ***	-0.6634 ***	-0.0780	0.0238	-0.0614 **	-0.0719 **	7.4671 ***
30	12.53	0.5941 **	0.1349	-0.0619	0.0172	0.0166	-0.0181	18.5248 ***
87	11.58	3.7940 ***	0.3580	0.1054	0.1139 *	-0.1119 **	-0.1130 **	5.3396
29	6.67	-0.3162	-0.0515	0.1705 *	-0.0679	0.0721 **	0.0701 **	22.2023 ***
90	6.60	1.6544 ***	-0.4008 ***	-0.0417	0.0296 *	-0.0284 **	-0.0426 ***	13.5343 ***
85	6.03	1.7252 ***	-0.0309	0.1207	0.0180	-0.0339	-0.0453	13.1437 ***
98	4.19	0.5564	-0.2273	0.0561	0.1221 **	-0.0009	-0.0380	18.2213 ***
27	3.89	4.4688 ***	-3.2745 ***	0.0413	-0.0029	-0.0026	-0.0223	-0.3420
88	2.90	2.4327 ***	-0.6339 *	-0.2139 *	-0.0111	-0.1132 *	-0.1311 **	8.5864 ***
22	2.88	1.4878 ***	-0.1397	0.0345	-0.0153	-0.0308 *	-0.0441 **	13.1998 ***
71	2.08	2.2499 ***	-0.9009 ***	0.0352	0.1004	-0.0352	-0.0486	10.0547 ***
39	1.73	2.0259 ***	-0.0778	-0.1807 ***	0.0135	-0.0579 ***	-0.0737 ***	10.3971 ***

Note: ***, **, and * denote significance at the 1%, 5%, 10% levels, respectively.

Notes

- ¹ The EU consists of 27 members: on/before 2003, there were 15 members, namely Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Spain, Sweden, and the United Kingdom. In 2004, Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia, and Slovenia joined. In 2007, Bulgaria and Romania joined. In 2013, Croatia joined, but the United Kingdom withdrew in 2020.
- ² This list comprises the 12 largest industries by trade market share, representing the trade flows of an industry as a fraction of the exports or imports of the United States and the EU.

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