The current issue and full text archive of this journal is available on Emerald Insight at:
https://www.emerald.com/insight/1086-7376.htm

Exchange rate volatility and trade flows in Indonesia and ten main trade partners: asymmetric effects

Unggul Heriqbaldi, Miguel Angel Esquivias,
Rossanto Dwi Handoyo and Alfira Cahyaning Rifami
Department of Economics, Faculty of Economics and Business, Airlangga University, Surabaya, Indonesia, and
Hilda Rohmawati
Department of Economics, Faculty of Economics and Business, Airlangga University, Surabaya and Ministry of National Development Planning, Jakarta, Indonesia

Received 24 October 2021
Revised 3 February 2022
4 June 2022
Accepted 8 September 2022

Abstract

Purpose – This paper aims to examine whether Indonesian cross-border trade responds asymmetrically to exchange rate volatility (ERV).

Design/methodology/approach – An exponential generalized autorgressive conditional heteroscedasticity model is applied to estimate the ERV of Indonesia and ten main trade partners using quarterly data from 2006 to 2020. A nonlinear autoregressive distributed lag estimation is applied to estimate the impact of ERV on cross-border trade. Impacts from the global financial crisis (GFC) of 2008 and the COVID-19 pandemic are covered. Dynamic panel data is used for the robustness test.

Findings – In the short-run, ERV significantly affects exports to most of the top partners (positively, negatively or both). In the long run, asymmetric effects occur in Indonesia's exports to five top destinations. The weakening of the Indonesian Rupiah mainly supports exports in the short term. Imports from top partners are also affected by ERV in both the short run and, to a lesser extent, in the long run. Both the GFC and the COVID-19 pandemic reduced trade: for most cases, in the short run. The dynamic panel model suggests that ERV has asymmetric impact on cross-border trade in the long run.

Practical implications – Exchange rate strategies need to avoid a single-side policy approach and, instead, account for exporter and importer differences in risk behaviour and an asymmetric response to ERV in trade. Policymakers need to consider policies that stabilise the currency.

Originality/value – This study provides evidence that cross-border trade can react asymmetrically to the exchange rate uncertainty and that the impacts of real ERV are asymmetric as well. The authors also apply a dynamic panel that signals that ERV matters in the long run for Indonesian trade with top partners.

Keywords Exchange rate volatility, Asymmetric effects, Nonlinear ARDL, Indonesia, Exchange rates, Global trade

Paper type Research paper

1. Introduction

This study aims to investigate the effects of real exchange rate volatility (ERV) on cross-border trade between Indonesia and its ten largest trade partners and whether the effects of real ERV

JEL classification – C32, F31

This research paper was supported by Universitas Airlangga, Research Grant 2022, and by Bank Indonesia.

Studies in Economics and Finance

1086-7376
DOI 10.1108/SEF-10-2021-0451

are asymmetric. Movements in exchange rate can affect cross-border trade through exchange rate depreciation and through exchange rate uncertainty. Movements in exchange rates expose traders to currency volatility that influences trade flows. Risk averse traders are likely to react negatively to ERV, thereby decreasing trade volume. However, less risk averse exporters may allow some degree of uncertainty in currencies, suggesting that the impact of ERV on cross-border trade is ambiguous rather than necessarily negative. As traders can respond asymmetrically to exchange rates, they may also respond asymmetrically to ERV.

ERV is crucial for traders as deviations in exchange rates from its long-run equilibrium level, may cause negative impacts on trade as deviations in exchange rates lower expectations of profits, alter prices and output and subsequently depress economic growth (Heriqbaldi *et al.*, 2020). The general finding is that ERV harms exports, and this suggests the need for government intervention policies to maintain exchange rate stability (Fang *et al.*, 2009). On the downside, government intervention could exacerbate market uncertainty; and this could impact the currency's long-run equilibrium level, leading to asymmetric responses from the market players (McKenzie, 2002).

Seminal work on ERV by Clark (1973) and Ethier (1973) suggests that uncertainty in exchange rates prompts changes in expected profits and adjustments in demand and supply of goods. Future price uncertainty arising from ERV can lead to inflation uncertainty (*pass-through inflation*), thus affecting final demand for goods (Alexander, 1952). This is due to firms trying to reduce their exposure to risk (Obstfeld and Rogoff, 1998). A perfect forward market can reduce the adverse effect, but will increase costs at the exporters' end (Broll and Eckwert, 1999; Hooper and Kohlhagen, 1978). A risk-neutral exporting firm may respond to a rise in ERV by increasing its trade volume (Franke, 1991; Sercu and Vanhulle, 1992).

Theory and empirical evidence are mixed as to the impact of exchange rate and ERV on cross-border trade. In line with trade theory (Ethier, 1973), some studies find that risk averse traders are likely to react negatively to ERV, thereby decreasing trade volume (Arize *et al.*, 2000; Chit *et al.*, 2010; Hayakawa and Kimura, 2009; Kroner and Lastrapes, 1993). On the other hand, studies such as that by Chi and Cheng (2016) offer evidence of a positive effect from ERV on exports as traders may increase volumes to counterweigh potential losses on revenue. Other studies find no concluding evidence of impacts arising from ERV on exports (Nishimura and Hirayama, 2013). Differences in risk perception and risk attitudes of traders towards ERV suggest that the impact of ERV on cross-border trade is uncertain rather than necessarily negative.

The literature also shows mixed results when it comes to the short- or long-run impacts. Some studies find stronger effects from ERV in the short run (Asteriou *et al.*, 2016; Bahmani-Oskooee and Nouira, 2020), compared to the long run (Chi, 2020; Sharma and Pal, 2018), while few find ERV persistent in both the short and the long run (Bahmani-Oskooee *et al.*, 2017). Ambiguity is also exhibited when studies are restricted to specific commodities or industries (Bahmani-Oskooee and Aftab, 2017); when the levels of economic development among partners differ (Chi and Cheng, 2016); when traders have limited hedging instruments available (Aghion *et al.*, 2009); when partners differ in risk behaviour (De vita and Abbott, 2004); and when it comes to capturing economic shocks (Fitrianti, 2017). In a survey of studies examining ERV impact on trade, McKenzie and Melbourne (1999) argue that existing significant relationships between ERV and trade volume (positive or negative) are seemingly at random, suggesting that ERV exerts specific impacts across markets. Using bilateral or disaggregated data on trade to test the nexus is preferable to pooling data from diverse countries all together.

Considering the disputing evidence of the impact of ERV on trade flows and the different extensions to exchange rate theories, it is reasonable to re-examine how ERV impacts bilateral trade flows at the country level in both the short and long run in the context of a developing country (e.g. Indonesia). Firstly, developing countries often highly rely on exports of natural resources (which are price elastic) and are likely to face challenges in trade balance as a result of uncertainty. Hausmann *et al.* (2006) find that developing countries can face 2 to 2.5 times greater ERV volatility than industrialised countries and that persistency of volatility is longer for them. Secondly, developing countries often lack access to hedging tools, and this increases the risk perception of traders (Fang *et al.*, 2009; Hall *et al.*, 2010), making them more vulnerable to ERV. Furthermore, the link between price volatility and international flows of goods in countries with "commodity currency" (Hegerty, 2016) indicates that for them, cross-border trade is more responsive to adjustments in prices (higher price elasticity) than others, possibly displaying asymmetric effects in trade flows (Belke *et al.*, 2015). In line with findings from previous studies regarding the possible presence of asymmetric effects in ERV (Chi, 2020; Fang *et al.*, 2009; Sharma and Pal, 2018), our study aims to test whether such effects occur for the case of Indonesia and its top ten trade partners, with respect to both exports and imports. The ten largest trade partners of Indonesia are China, the USA, Singapore, South Korea, Japan, India, Malaysia, Australia, Germany, the Netherlands (top only in exports) and Brazil (top only in imports).

From among previous studies involving Indonesia, a large majority focus on the symmetric effects (Asteriou *et al.*, 2016; Bahmani-Oskooee *et al.*, 2015; Doğanlar, 2002; Fang *et al.*, 2006; Fitrianti, 2017; Hayakawa and Kimura, 2009; Sugiharti *et al.*, 2020a, 2020b) and mainly support the claim that exports are significantly affected by currency risk. Most of those earlier studies used moving sample standard deviation or generalized autoregressive conditional heteroskedasticity (GARCH) models to estimate volatility. We opted to use exponential generalized autorgressive conditional heteroscedasticity (EGARCH), which offers some advantages over the GARCH model. EGARCH does not impose restrictions on the parameters of the volatility equation as the model is presented in log variance (fulfilling the positivity of variance) instead of variance itself (Tsay, 2005). Additionally, the EGARCH model can capture negative shocks that may have stronger effects on the variance than positive ones (asymmetry or leverage effect).

Few attempts have questioned the asymmetric effects of ERV on Indonesia's trade, as that of Fang *et al.* (2009), which covered trade between Indonesia and the USA. Sugiharti *et al.* (2020a) covered the total exports of fifteen main commodities from Indonesia to its five largest export markets, finding that top exports to the USA, India and South Korea are mainly negatively affected by ERV. However, little is said of the impact of ERV on imports to Indonesia. Additionally, no previous studies have looked into asymmetric ERV between Indonesia and its large partners in the European Union or Australia (except for Chi and Cheng, 2016), who covered Australian exports to Indonesia using a symmetric approach and found significant ERV effects.

This paper offers several contributions to the literature of ERV. Firstly, our study analyses Indonesia, a country reliant on non-manufactured exports, unlike its ten partners, most of which have not been covered previously. Earlier studies on ERV in Indonesia have suggested that future research might consist of examining country-to-country cases (Bahmani-Oskooee *et al.*, 2015) and the possible asymmetric effect of ERV on Indonesia's trade (Asteriou *et al.*, 2016; Sugiharti *et al.*, 2020a, 2020b). As previous research has not scrutinised both export and import flows in Indonesia, we examine the effect of ERV on both exports and imports at the bilateral level. Secondly, we also consider the impact of ERV using dynamic panel data, thereby pooling together Indonesia's volume of trade with its top partners. Additionally, we incorporate dummy variables into our model to capture the shocks from the global financial crisis (GFC) (2008) and the COVID-19 pandemic. While previous studies have suggested that exchange rate depreciation in Indonesia is detrimental to exports (Asteriou *et al.*, 2016; Fang *et al.*, 2006;

Pino *et al.*, 2016; Sugiharti *et al.*, 2020a, 2020b), which is counterintuitive to what is theoretically expected, proper central bank intervention and fiscal policy may help to reduce uncertainty in the longer time. Exchange rate strategies may need to avoid a single-side policy approach and, instead, account for differences in the risk behaviour (e.g. hedging) of both exporters and importers (Zhang *et al.*, 2018).

As a robustness test, we apply dynamic panel data techniques to test for the impact of asymmetric ERV on cross-border trade. By pooling countries together, we aim at assessing whether common equilibrium association exists within our model. Although models (time series and panel data) are not comparable head-to-head, they can complement each other. For comparison purposes, asymmetric (linear) effects are also estimated. Results using a symmetric model are provided as supplementary materials.

The results show that, in the long run, real ERV has a negative impact on Indonesia's exports to China, Singapore, South Korea and Germany; while in the short run, it significantly affects the case of Indonesia's exports to all top partners, but the USA and Singapore. Moreover, the Wald-test results suggest that the bilateral real ERV has a long-run asymmetric effect on exports to China, Singapore, South Korea, India, Malaysia and the Netherlands. In contrast, the short-run effects are significant on exports to Singapore, South Korea, Japan, India, Australia and the Netherlands.

Moreover, negative ERV has a significant effect on Indonesia's imports from China, Japan, Malaysia, Australia and Germany in the long run. However, we also found that positive ERV influences imports from Japan, Malaysia and Australia. In contrast, in the short run, the results show a dampening effect of real ERV (positive or negative) on Indonesia's imports from the USA, Singapore, South Korea, Japan, Malaysia, Australia and Brazil. Concerning the asymmetric impacts, in the long-run model, we reject the null hypothesis, as in the case of Indonesia's imports from China, Singapore, South Korea, Japan and Germany. In the short-run model, we find that real ERV asymmetrically affects Indonesia's imports from South Korea, Japan and Malaysia.

Finally, the results indicate that in the short run, the COVID-19 pandemic has significant and negative effects on exports and imports to/from most countries. Effects from COVID-19 are larger and more numerous than the impact on trade from the GFC (2008). The nonlinear autoregressive distributed lag (NARDL) indicates that positive and negative ERV vary in terms of their effect on trade volume, providing superior results to autoregressive distributed lag (ARDL) (Supplementary materials). Finally, the dynamic panel data using a Pool Mean Group (PMG) approach indicates that both exports and imports are significantly affected by ERV in the long run, signalling the importance of exchange rate normalisation for Indonesian trade.

2. Methodology

2.1 Specifications

The non-structural equation (reduced form) of Klaassen (2004) and Pozo (1992) offers a framework to empirically analyse the asymmetric effects of ERV in bilateral exports. Although models to estimate trade flows under asymmetric effects can include a substantial number of factors, such structural models can be complex (Koutmos and Martin, 2003). As such, most studies looking into export and import demand models use reduced forms (Arize *et al.*, 2021; Bahmani-Oskooee and Aftab, 2017; Bailey *et al.*, 1987; Fang *et al.*, 2009). Bilateral export flows (EX) can be expressed as a function of real income of foreign partner, real bilateral exchange rate (relative prices) and ERV (exchange rate uncertainty).

To model bilateral trade flows, this article uses the NARDL, beginning with the following standard specification:

$$LnEX_t^{IND} = \beta_0 + \beta_1 LnIIP_t + \beta_2 LnRER_t + \beta_3 LnVOL_t + \varepsilon_t \tag{1}$$

$$LnIM_t^{IND} = \gamma_0 + \gamma_1 LnIIP_t + \gamma_2 LnRER_t + \gamma_3 LnVOL_t + \mu_t \tag{2}$$

We refer to equations (1) and (2) as our standard export and import demand models, where $LnEX_t^{IND}$ and $LnIM_t^{IND}$ represent aggregate real Indonesian exports and imports with respect to ten different partner countries. We use IIP_t *and* IIP_t^* or Index of Industrial Production (IIP) for a proxy of Indonesia and destination countries (*) as a proxy for income and economic activity. Meanwhile, RER_t is the bilateral real exchange rate (RER) between the Indonesian Rupiah (IDR) and the ten currencies of the trading partners; the United States dollar (US$), Renminbi (CNY), Japanese Yen (JPY), European (euro), Indian Rupee (INR), South Korean Won (KRW), Malaysian Ringgit (MYR), Singaporean dollar (SGD), the Australian Dollar (AUD) and the Brazilian Real (BRL). Finally, VOL_t is a measure of the real ERV based on the quarterly conditional variance of a generalised autoregressive conditional heteroskedasticity, GARCH (1,1). It is expected that the estimates coefficient of β_1 and γ_1 are positive. Meanwhile, the estimated coefficient of β_2 is expected to be positive, but that of γ_2 to be negative. The signs of the β_3 and γ_3 coefficients could be positive or negative.

Similar equations are proposed as general models to estimate export and import demand and are further modified to study the effects of asymmetric ERV in bilateral trade in cases like Tunisia, India and Mexico, as in the empirical evidence provided by Arize *et al.* (2017), Bahmani-Oskooee and Nouira (2020), Chi (2020) and Sharma and Pal (2018).

Traders' behaviour may result in asymmetric responses to volatility (EVOL). To capture the asymmetric effects of real ERV, we establish two other variables, namely, *lnVOL_POS*, which represents positive changes in volatility, and *lnVOL_NEG*, including the negative changes. *EVOL* is a measure of the natural logarithm of the real ERV based on a quarterly conditional variance of an exponential GARCH (1,1):

$$\begin{aligned} lnVOL\_POS_t &= \sum\nolimits_{j=1}^{t} \Delta lnVOL_t^{+} = \sum\nolimits_{j=1}^{t} max(\Delta lnVOL_j, 0), \\ lnVOL\_NEG_t &= \sum\nolimits_{j=1}^{t} \Delta lnVOL_t^{-} = \sum\nolimits_{j=1}^{t} min(\Delta lnVOL_j, 0) \end{aligned} \tag{3}$$

The next step is to derive the asymmetric equation for export and import:

$$\begin{aligned} \Delta lnEX_t^{IND} = {} & c_1 + \sum\nolimits_{j=1}^{n1} c_{2j} \Delta lnEX_{t-j}^{IND} + \sum\nolimits_{j=0}^{n2} c_{3j} \Delta lnIIP_{t-j}^{*} + \sum\nolimits_{j=0}^{n3} c_{4j} \Delta lnRER_{t-j} \\ & + \sum\nolimits_{j=0}^{n4} c_{5j} \Delta lnVOL\_POS_{t-j} + \sum\nolimits_{j=0}^{n5} c_{6j} \Delta lnVOL\_NEG_{t-j} + \lambda_1 lnEX_{t-1}^{IND} \\ & + \lambda_2 lnIIP_{t-i}^{*} + \lambda_3 lnRER_{t-1} + \lambda_4 lnVOL\_POS_{t-i} + \lambda_5 lnVOL\_NEG_{t-i} + \varepsilon_t \end{aligned} \tag{4}$$

$$\begin{aligned} \Delta lnIM_t^{IND} = {} & d_1 + \sum\nolimits_{j=1}^{n6} d_{2j} \Delta lnIM_{t-j}^{IND} + \sum\nolimits_{j=0}^{n7} d_{3j} \Delta lnIIP_{t-j}^{*} + \sum\nolimits_{j=0}^{n8} d_{4j} \Delta lnRER_{t-j} \\ & + \sum\nolimits_{j=0}^{n9} d_{5j} \Delta lnVOL\_POS_{t-j} + \sum\nolimits_{j=0}^{n10} d_{6j} \Delta lnVOL\_NEG_{t-j} + \pi_1 lnIM_{t-1}^{IND} + \pi_2 lnIIP_{t-i}^{*} \\ & + \pi_3 lnRER_{t-1} + \pi_4 lnVOL\_POS_{t-1} + \pi_5 lnVOL\_NEG_{t-i} + \varepsilon_t \end{aligned} \tag{5}$$

Equations (4) and (5) were introduced by Shin *et al.* (2014) as nonlinear ARDL (NARDL). For the cointegration test, we use the standard F-test. In equations (4) and (5), Δ indicates the first difference operator. The short-run effects (coefficients attached to Δ) are captured by c_{2j}, c_{3j}, c_{4j}, c_{5j} and c_{6j} in the export demand equation, whereas in the import demand equation, the effects are represented by d_{2j}, d_{3j}, d_{4j}, b_{5j} and d_{6j},. The long-run effects are shown by λ_2, λ_3, λ_4 and $\lambda_5$normalised on λ_1 (for export demand equation) and π_2, π_3, π_4 and π_5 normalised on π_1 (for import demand equation). Pesaran *et al.* (2001) stated that the cointegration condition is required for the long-run outcomes to be valid. Therefore, the F-test will be conducted to establish the joint significance of lagged level variables as a condition for cointegration. Bahmani-Oskooee and Nouira (2020) noted that ARDL models can be used regardless of whether the variables are integrated of order I(0), I(1) or of mixed order I(0) and I(1). We apply a unit root test to verify that the order of integration is not I(2), applying the augmented Dickey–Fuller (ADF) test. Results are offered in Tables A1 and A2 (Appendix).

Using equations (4) and (5), we can establish the analysis of the asymmetric relationship. In the short run, the asymmetric relationship can be represented by the difference in the number of $\Delta ln VOL\_POS$ lags and the number of $\Delta ln VOL\_NEG$ lags. The asymmetric relationship can also be observed by looking at the size or sign of the coefficient estimate of $\Delta ln VOL\_POS_{t-i}$ and $\Delta ln VOL\_NEG_{t-i}$ lags at each lag *i (i = 1, 2, . . .)*. One procedure that can be used to prove the asymmetric effect in the short run is to do an F-test to verify whether $\sum \hat{c}_{5j} \neq \sum \hat{c}_{6j}$ and $\sum \hat{d}_{5j} \neq \sum \hat{d}_{6j}$ in equations (4) and (5). Meanwhile, the long-run asymmetric effects can be identified by conducting an F-test to prove $\hat{\lambda}_4 I - \hat{\lambda}_1 \neq \hat{\lambda}_5 I - \hat{\lambda}_1$ in 4) and, $\hat{\pi}_4 I - \hat{\pi}_1 \neq \hat{\pi}_5 I - \hat{\pi}_1$ in (5).

A number of studies assume symmetric risk behaviour from traders with respect to appreciation or depreciation of exchange rate and ERV (Aftab *et al.*, 2017; Fitrianti, 2017; Gupta and Varshney, 2021). However, we argue that traders are likely to respond asymmetrically to ERV. At least four possible arguments can support the claim that ERV can affect trade asymmetrically, as noted in Fang *et al.* (2009). Firstly, ERV may lead monetary authorities to intervene in the markets, generating uncertainty in the future currency exchange rate (McKenzie, 2002). Central bank intervention may change depending on whether the currency is depreciating or appreciating, adding to the uncertainty in markets (Fratzscher, 2006). Secondly, *pricing-to-market* behaviour of traders often behaves asymmetrically, as prices of export goods are adjusted according to foreign competition rather than to domestic currency alone (Cao *et al.*, 2015; Mahdavi, 2000). Firms oriented towards maintaining or increasing market share may opt not to increase export prices when domestic currency appreciates, to avoid a loss of foreign sales. On the other hand, in an event of currency depreciation, firms targeting an increase in market share may lower foreign prices to increase sales (Cao *et al.*, 2015). Thirdly, asymmetric behaviour of traders may persist after the events that caused the currency changes no longer exist (hysteretic behaviour). Campbell (2020) finds that more open economic sectors experience persistent effects from exchange rate shocks, maintaining asymmetric responses to exchange rate movements to remain competitive. Fourthly, firms may follow asymmetric behaviour by hedging (one side hedge), as they often perceive risk when it entails losses rather than when it involves positive impact on profits (Ali, 2019; Zhang *et al.*, 2018). Considering that emerging Asian countries mainly invoice products in US dollars, a depreciation of domestic currency may increase export revenue, suggesting that exporting firms may hedge against appreciation of the domestic currency, but stand unhedged for depreciation (Fang *et al.*, 2009).

2.2 Data

We obtained the data from quarter 1, 2006 to quarter 4, 2020. The IIP is accessible at the International Financial Statistics site, while quarterly raw data for total bilateral trade (exports and imports) between Indonesia and the leading trade partners are collected from Trade Map.

To find the real value of the exports and imports and the real bilateral exchange rate, we deflated export and import values by the consumer price index (CPI) of the exporter and importer countries, respectively:

$$EX_t = \frac{NEX_t \text{ X } CPI_t^{partner\ countries}}{CPI_t^{Indonesia}} \tag{6}$$

$$IM_t = \frac{NIM_t \text{ X } CPI_t^{partner\ countries}}{CPI_t^{Indonesia}} \tag{7}$$

And:

$$RER_t = \frac{NER_t \text{ X } CPI_t^{partner\ countries}}{CPI_t^{Indonesia}} \tag{8}$$

where EX_t reflects Indonesia's real exports to the destination country; IM_t reflects Indonesia's real imports from the partner countries; NEX_t is a nominal export value; NIM_t is a nominal import value; CPI is the CPI; RER_t is the real bilateral exchange rate between the Rupiah and the partner's currency; and NER_t is the nominal bilateral exchange rate (units of IDR per unit of partner's currency). Thus, an increase in RER indicates a real depreciation of Indonesian currency against a partner's currency.

Because real ERV is not observable, we used the GARCH and EGARCH models to obtain the estimates for ERV. GARCH-based models are appropriate and accurate for capturing uncertainty as they take into account time-varying conditional variance (Sharma and Pal, 2018). GARCH-based models allow that large variance originated in previous periods to result in large variances in future periods, implying a volatility cluster. On the other hand, the main advantage of the EGARCH approach is that it spots the time-varying conditional variance from a time-series model (Choudhry, 2005; Engle and Ng, 1993). Additionally, as noted in Iyke and Ho (2019), in an EGARCH approach, good and bad news can affect volatility (uncertainty) differently, and big news captures greater effect on volatility; however, the same does not happen when using the GARCH model.

Prior to the estimation of ERV using the GARCH-based model, an ADF test (ADF) was conducted on all series to test for a unit root, following Dickey and Fuller (1979). We tested the data of the real bilateral exchange rate between the Rupiah and the partner's currency for stationarity. The test results for all ten exchange rates are displayed in Tables A1 and A2 in the Appendix.

We used GARCH and EGARCH models to determine ERV. We assumed that the logarithm of *RER* mirrors a simple mean equation, as shown by the following equation:

$$lnRER_t = f(lnRER_{t-1}; \gamma) + \varepsilon_t \tag{9}$$

where $lnRER_t$ is the natural logarithm of the real bilateral exchange rate at time t; $f(.)$ is the functional form of RER (assuming linearity in factors); $lnRER_{t-1}$ is a one period lag ($t-1$) of the natural logarithm of real bilateral exchange rate; and γ is a vector of coefficients. ε_t is white noise error with variance, $V(\varepsilon) = h^2$. h_t^2 is conditional variance, which is a measure of real ERV. The simplest GARCH model is GARCH (1,1), which follows the equation:

$$h_t^2 = \overline{\omega} + \beta_1\left(\varepsilon_{t-1}^2 - \overline{\omega}\right) + \emptyset_1\left(h_{t-1}^2 - \overline{\omega}\right) \tag{10}$$

where $\overline{\omega}$, β_1 and $\emptyset_1$ in equation (10) are parameters to be estimated. As noted in Iyke and Ho (2019), "the GARCH (1,1) model is expected to reverse to a constant mean, $\overline{\omega}$".

We could capture asymmetric information in the uncertainty, h_t^2, by modelling the conditional variance as exponential GARCH(1,1) or EGARCH(1,1) of the form:

$$lnh_t^2 = \omega + \beta ln\varepsilon_{t-1}^2 + \emptyset\left|\frac{\varepsilon_{t-1}}{h_{t-1}}\right| + \delta\frac{\varepsilon_{t-1}}{h_{t-1}} \tag{11}$$

where ω, β, $\emptyset$ and δ are parameters to be estimated. Equation (11) shows a relationship between past shocks (t$-$1) and the natural logarithm of the conditional variance, h_t^2, where both positive and negative shocks have an effect of the form of $\emptyset + \delta$ or $\emptyset - \delta$, respectively. The asymmetric effects assume that $\delta \neq 0$. We report the conditional volatility (EVOL) between the IDR and each of the ten currencies using the EGARCH approach in Figure 1(a)–1(i).

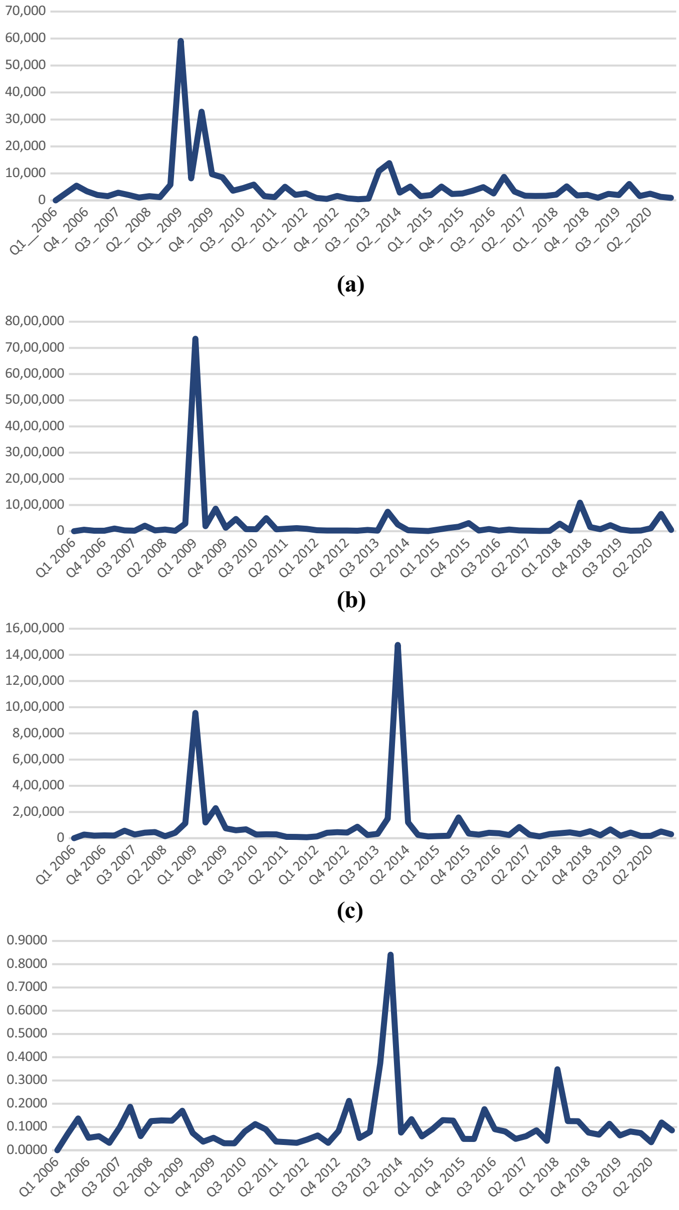
To account for global shocks during the 2008 GFC (*D2008*) and the COVID-19 (*DCOV*) pandemic, two dummy variables were incorporated into the model (Equations 1, 2, 4 and 5). Equations including dummy variables are not displayed due to space limitations. Results for the models without dummy variables were also estimated (see Tables A3 and A4). Results for the linear model are presented in Supplementary Materials (Tables S1 and S2). Quantitatively similar results were obtained when using the models without dummy variables and the baseline model (using dummy variables controlling for GFC and COVID). The GFC period (*D2008*) covers 2008_Q1 to 2009_Q4. The variable for *DCOV* covers 2020_Q1 to 2020_Q4. Earlier studies introduced dummy variables for the GFC period, a period of great volatility and change in global demands for exports (Arize *et al.*, 2021; Bahmani-Oskooee and Nouira, 2020; Iyke and Ho, 2019).

Additionally, we applied dynamic panel data techniques to provide aggregate results for our sample (pool) as a robustness test. The model proposed for the NARDL (time series) was adjusted into a dynamic panel setting to estimate asymmetric and symmetric results for exports; imports used the same variables. The specific model using panel data is not displayed to save space in the manuscript. Studies such as that by Iyke and Ho (2019) offer detailed explanations of panel dynamic models.

3. Results

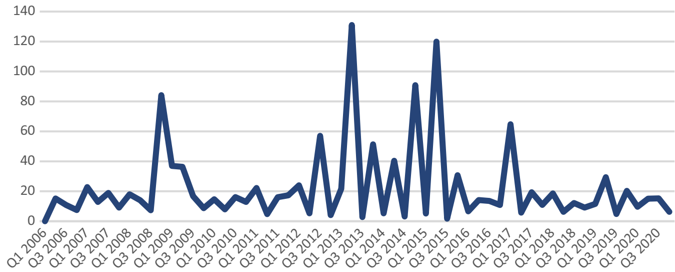
3.1 Asymmetric impacts of exchange rate volatility on exports

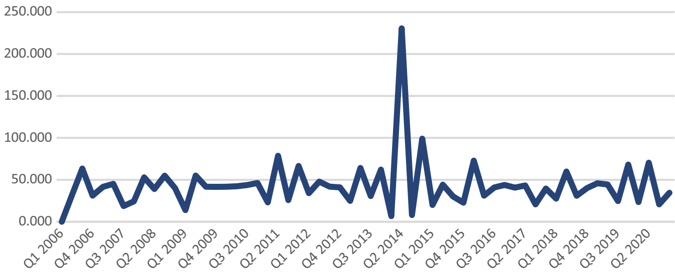
This section focuses on an examination of the impact of ERV on Indonesia's bilateral exports with its ten major trading countries (Tables 1 and 2). Results include dummy variables to consider the effects of the GFC in 2008–2009 and the COVID-19 pandemic (baseline model). Linear estimates of ERV are provided as supplementary files (Tables S1 and S2) as part of the robustness test. We apply dynamic panel data models to provide aggregate results for our sample (pool).



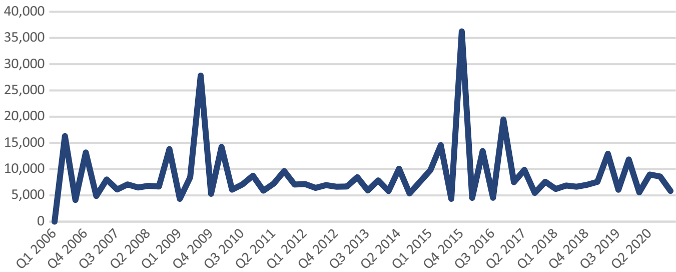
(*continued*)

Figure 1.
Volatility of exchange rate Indonesia (EVOL) and partners EGARCH approach





(e)



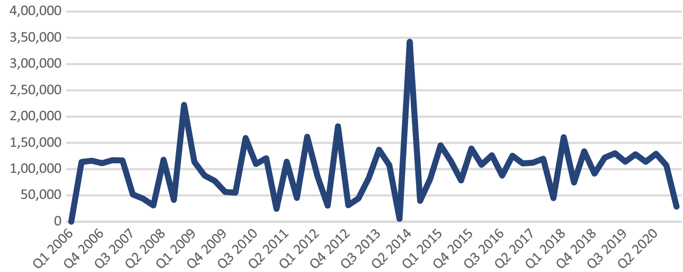
(f)

Figure 1.

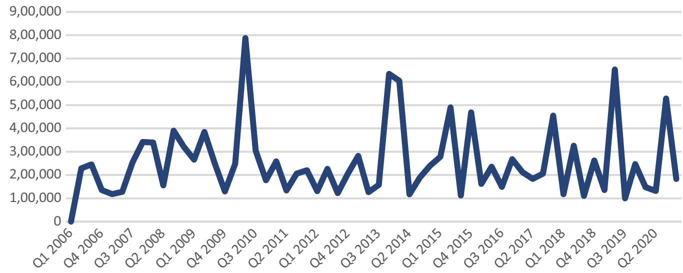
(*continued*)

Panel A of Table 1 presents the normalised cointegrating vector applied to examine the long-run effects of real ERV on Indonesia's bilateral export to China, the USA, Singapore, South Korea, Japan, India, Malaysia, Australia, Germany and the Netherlands, respectively. The positive and negative shocks to a single variable were modelled in the framework as they affect the dependent variable distinctively. This model also allows for the short-run effects to vary from its long-run impacts. Therefore, we test both short- and long-run outcomes. In Panel B, the short-run estimates are reported. In Panel C, we report the diagnostic statistics.

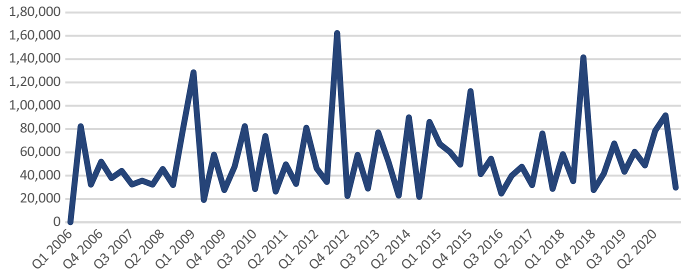
From Table 1, several points can be made. Firstly, the cointegration test results through the bound-tests show cointegration between all existing series. This means that, with this NARDL model specification, we can prove that, in the context of Indonesian exports,



(g)



(h)



(i)

Notes: (a) China EGARCH; (b) United States EGARCH; (c) Singapore EGARCH; (d) South Korea EGARCH; (e) India EGARCH; (f) Malaysia EGARCH; (g) Austraiia EGARCH; (h) Germany – Netherlands EGARCH; (i) Brazil EGARCH

Figure 1.

Indonesia has a long-run relationship with the countries it exports to (see F-stats). The corresponding ECM_{t-1} also shows that the coefficient estimates are significantly negative and less than one. Bahmani-Oskooee and Aftab (2017) stated that when the bound test and ECM_{t-1} results are in different calculations, we might use the ECM_{t-1} result as the reference to decide if there is a long-run relationship among the series in the model. Hence, we can conclude that cointegration exists in all cases because the ECM_{t-1} coefficients are significantly negative and less than one in value. The ECM_{t-1} estimates also show that, on average, 51% of the uncertainty in export demand is offset by short-run correction for each quarterly export.

Table 1.
Results of NARDL for export demand model

Variable	China	USA	Singapore	South Korea	Japan	India	Malaysia	Australia	Germany	The Netherlands
Panel A: Long run										
	(1,2,5,2,1,4,2)	(5,3,2,0,0,4)	(1,4,4,4,2,4,4)	(4,5,0,3,5,5,3)	(2,2,3,3,4,2,2)	(5,0,3,2,3,3,2)	(4,4,3,4,0,2,4)	(2,4,1,0,3,4,1)	(4,4,3,1,2,2,1)	(2,3,0,3,3,1,3)
C	4.977 (3.511)	6.159*** (0.871)	−2.659 (4.270)	−9.509** (3.607)	11.548 (10.968)	6.672 (11.081)	12.684*** (1.586)	−18.709*** (6.823)	12.323*** (2.254)	−4.504* (2.377)
LnIIP	5.872*** (1.334)	2.323*** 0.573	0.223 (1.359)	6.958*** (1.523)	2.323 (2.231)	6.677** (2.606)	−3.793*** (0.445)	−0.798 (1.244)	−2.258* (1.129)	6.128*** (1.014)
LnRER	−3.450*** (0.412)	−1.088*** (0.215)	2.332* (1.185)	2.488** (0.921)	−4.491 (4.523)	−6.112 (4.317)	0.323 (0.363)	6.680*** (1.583)	−0.478 (0.449)	−0.403 (0.252)
LnVOL_POS	−0.055 (0.044)	0.023 (0.014)	−0.261** (0.127)	−0.351** (0.144)	0.240 (0.434)	2.082 (1.679)	0.035 (0.092)	0.125 (0.121)	−0.307* (0.157)	0.026 (0.167)
LnVOL_NEG	−0.180*** (0.044)	0.023 (0.015)	−0.183 (0.129)	−0.190 (0.133)	0.295 (0.446)	2.009 (1.712)	−0.052 (0.094)	0.137 (0.125)	−0.296* (0.169)	0.040 (0.169)
D2008	0.155*** (0.041)	0.042 (0.030)	−0.124 (0.097)	0.439*** (0.131)	−0.028 (0.106)	0.379* (0.197)	0.074 (0.049)	0.214** (0.099)	0.136* (0.070)	−0.014 (0.038)
DCOV	0.006 (0.029)	0.077* (0.039)	0.010 (0.067)	0.072 (0.044)	0.063 (0.119)	−0.370** (0.169)	−0.022 (0.024)	0.173** (0.081)	0.032 (0.047)	−0.017 (0.042)
Panel B: Short run										
Δ (LnEX(−1))		−0.127 (0.115)		0.227* (0.124)	−0.525 (0.100)	−0.351*** (0.101)	0.225* (0.082)	−0.248*** (0.089)	−0.532*** (0.106)	0.189* (0.110)
Δ (LnEX(−2))		−0.048 (0.099)		0.323** (0.128)		−0.210** (0.102)	−0.022 (0.080)		−0.570*** (0.114)	
Δ (LnEX(−3))		0.194** (0.090)		0.203* (0.116)		−0.115 (0.114)	−0.328*** (0.092)		−0.570*** (0.104)	
Δ (LnEX(−4))		0.064 (0.082)				−0.301** (0.109)				
Δ (LnIIP)	2.325*** (0.514)	2.303*** (0.246)	0.827*** (0.258)	0.899** (0.331)	1.143 (0.262)		1.174*** (0.411)	−0.673 (0.416)	0.986*** (0.160)	2.275*** (0.331)
Δ (LnIIP(−1))	−1.460** (0.570)	0.742** (0.284)	1.186*** (0.278)	−2.509*** (0.631)	0.519 (0.245)		4.884*** (0.630)	−0.552 (0.427)	1.507*** (0.279)	−2.989*** (0.526)
Δ (LnIIP(−2))		−1.436*** (0.338)	2.408*** (0.309)	−2.105*** (0.520)			3.472*** (0.608)	1.353*** (0.395)	1.315*** (0.295)	−0.953** (0.388)
Δ (LnIIP(−3))			2.069*** (0.331)	−2.578*** (0.473)			2.796*** (0.581)	1.107** (0.466)	1.064*** (0.274)	
Δ (lLnIIP(−4))				−1.835*** (0.377)						
Δ (LnRER)	−0.512 (0.442)	−0.905*** (0.209)	0.510 (0.453)		−0.135 (0.228)	0.536 (0.520)	0.853** (0.317)	1.639*** (0.477)	0.090 (0.200)	
Δ (LnRER(−1))	0.542 (0.516)		−0.640 (0.568)		0.744 (0.227)	4.484*** (1.231)	−1.356*** (0.363)		−0.048 (0.215)	
Δ (LnRER(−2))	3.337*** (0.487)		−0.051 (0.520)			3.231*** (0.942)	1.628*** (0.405)		0.918*** (0.207)	
Δ (LnRER(−3))	0.233 (0.492)		1.594*** (0.568)							
Δ (LnRER(−4))	1.560*** (0.416)									
Δ (LnVOL_POS)	0.112*** (0.022)		−0.004 (0.029)	−0.091*** (0.025)	−0.075 (0.020)	0.070 (0.080)	−0.141*** (0.034)		−0.054** (0.021)	−0.109*** (0.037)
Δ (LnVOL_POS(−1))	0.141*** (0.030)		0.039 (0.041)	0.174*** (0.029)	−0.075 (0.037)	0.220* (0.117)	−0.200*** (0.035)			0.086* (0.048)
Δ (LnVOL_POS(−2))			0.091** (0.037)	0.170*** (0.032)	−0.171 (0.035)		−0.111** (0.041)			−0.134** (0.049)
Δ (LnVOL_POS(−3))			0.118*** (0.027)				−0.068** (0.033)			
Δ (LnVOL_NEG)	−0.097*** (0.029)		−0.087* (0.043)	0.026 (0.025)	0.022 (0.021)	0.355*** (0.107)		0.004 (0.031)	−0.043*** (0.014)	0.095** (0.046)
Δ (LnVOL_NEG(−1))			0.154*** (0.041)	0.021 (0.024)	−0.143 (0.027)	−0.221*** (0.069)		0.106*** (0.026)	0.040** (0.019)	−0.058 (0.048)
Δ (LnVOL_NEG(−2))				0.022 (0.021)	−0.003 (0.023)	0.166*** (0.056)		0.150*** (0.030)		0.134*** (0.036)
Δ (LnVOL_NEG(−3))				0.083*** (0.022)	−0.052 (0.020)					
Δ (LnVOL_NEG(−4))				0.054** (0.021)						
Δ (D2008)	0.059* (0.032)	−0.335 (0.229)	−0.074* (0.039)	0.088*** (0.027)	−0.101 (0.023)	0.062 (0.040)	0.018 (0.023)	0.079** (0.038)	0.007 (0.019)	−0.100*** (0.034)
Δ (D2008(−1))	−0.064* (0.033)	−0.030** (0.014)	−0.156*** (0.037)	−0.138*** (0.031)	−0.095 (0.022)	−0.119*** (0.039)	−0.074*** (0.023)	−0.096** (0.041)	−0.036** (0.017)	
Δ (D2008(−2))	−0.020 (0.030)	−0.002 (0.014)	−0.069* (0.037)	−0.142*** (0.026)		−0.087** (0.039)		−0.122*** (0.040)		
Δ (D2008(−3))	−0.142*** (0.031)	0.029* (0.014)	0.115*** (0.033)	−0.142*** (0.025)				−0.176*** (0.039)		
Δ (D200801(−4))				−0.138*** (0.025)						
Δ (DCOV)	−0.073* (0.040)	−0.041** (0.020)	−0.010 (0.045)	0.076** (0.034)	−0.067 (0.027)	−0.049 (0.055)	−0.071** (0.029)	−0.118** (0.055)	−0.042* (0.025)	−0.015 (0.042)
Δ (DCOV(−1))	−0.094** (0.042)	−0.133*** (0.026)	0.080* (0.046)	0.020 (0.034)	−0.071 (0.028)	−0.121** (0.054)	0.066** (0.030)			0.034 (0.042)
Δ (DCOV(−2))		−0.071*** (0.025)	0.172*** (0.046)	0.076** (0.034)			0.120*** (0.030)			−0.097** (0.043)
Δ (DCOV(−3))		−0.057** (0.023)	0.129** (0.051)				0.095*** (0.033)			

(*continued*)

Variable	China	USA	Singapore	South Korea	Japan	India	Malaysia	Australia	Germany	The Netherlands
Panel C: Diagnostic tests for NARDL										
F^a	8.768***	4.169***	5.375***	5.509***	4.098***	7.030***	8.816***	5.977***	4.942***	4.415***
ECM_{t-1}	−0.935*** (0.101)	−0.772*** (0.120)	−0.597*** (0.080)	−0.569*** (0.074)	0.196*** (0.031)	−0.257*** (0.031)	−0.970*** (0.102)	−0.557*** (0.073)	−0.367*** (0.053)	−0.761*** (0.116)
Adj. R^2	0.770	0.798	0.691	0.787	0.762	0.672	0.812	0.694	0.735	0.691
LM^b	0.1659	0.0600	0.3413	0.0963	0.7042	0.1122	0.0628	0.6168	0.2965	0.1155
$RESET^c$	0.101	1.706	1.316	0.209	0.251	1.934	3.395	0.808	1.490	0.373
CU_d	S	S	US	S	S	S	S	S	S	S
CUQ	S	S	S	US	US	S	S	S	S	S
Wald test long run[e]	0.438***	1.962	4.803**	10.466***	4.802	5.142***	4.825**	1.404	0.740	8.037***
Wald test short run	2.101		3.634**	7.085***	3.550***	6.242***	0.954	4.929***	0.461	3.952***

Notes: In this table, we display estimates from equations (1) and (4). Volatility takes positive or negative values. a) At the 5% significance level, considering exogenous variables (k = 6), the lower (upper) bound critical value of the F-test is 1.99(3.99); b) Lagrange multiplier (L.M.) test of residual serial correlation with no serial correlation for the null hypothesis – presented by the probability of X^2 with one degree of freedom (first-order); c) RESET is Ramsey's test for misspecification (number of fitted terms = one) and applied by the value of F-statistic (correctly specified because the RESET statistic is insignificant); d) CU means the CUSUM for model stabilisation, S means "Stable" U means "Unstable"; e) asymmetric statistics test using Wald-test with the null hypothesis is symmetric; and f) *, ** and *** denote significance at 1, 5 and 10% levels, respectively

Table 1.

Table 2.
Results of NARDL for import demand model

Variable	China	USA	Singapore	South Korea	Japan	India	Malaysia	Australia	Germany	Brazil
Panel A: Long run										
	(1,3,3,0,1,1,1)	(2,2,1,3,5,0,1)	(2,2,3,2,0,3,2)	(1,3,3,0,3,2,2)	(2,2,3,4,4,2,3)	(2,2,2,0,0,1,2)	(1,3,2,0,1,2,2)	(1,0,0,0,1,3,3)	(1,4,2,2,1,4,4)	(1,4,2,2,1,4,4)
C	13.717*** (4.910)	16.830*** (2.701)	10.806 (10.428)	−18.762 (20.127)	−11.013*** (3.457)	−21.723 (13.656)	13.474** (6.356)	−6.841*** (2.262)	19.313** (7.430)	16.252** (7.794)
LnIIP	0.422 (2.618)	3.321 (2.348)	−3.283 (4.517)	13.997 (10.131)	8.379** (1.829)	9.991*** (3.524)	−5.132 (4.560)	2.552*** (0.735)	−1.037 (3.193)	−7.519** (3.500)
LnRER	−2.609** (1.200)	−4.288*** (1.243)	0.613 (1.894)	−2.231* (1.237)	0.594 (0.730)	3.583 (4.239)	0.761 (2.827)	1.988*** (0.345)	−2.723*** (0.980)	1.005 (0.903)
LnVOL_POS	−0.113 (0.090)	−0.022 (0.072)	0.091 (0.201)	−0.191 (0.251)	−0.455** (0.168)	0.096 (0.227)	−0.741** (0.336)	0.092* (0.051)	0.489 (0.301)	0.237 (0.298)
LnVOL_NEG	−0.214** (0.095)	0.013 (0.083)	0.028 (0.191)	0.260 (0.212)	−0.342* (0.179)	0.366 (0.248)	−0.897** (0.365)	0.140** (0.059)	0.533* (0.282)	−0.032 (0.296)
D2008	0.237** (0.098)	0.202* (0.109)	−0.121 (0.204)	0.089 (0.194)	−0.003 (0.057)	−0.129 (0.168)	−0.238 (0.152)	0.034 (0.026)	−0.095 (0.097)	0.074 (0.082)
DCOV	−0.112 (0.101)	0.095 (0.075)	−0.122 (0.232)	0.213 (0.285)	−0.284*** (0.069)	0.412 (0.299)	0.265 (0.224)	−0.013 (0.063)	0.210 (0.152)	0.072 (0.135)
Panel B: Short run										
Δ (LnIM(−1))		−0.338*** (0.094)	0.329*** (0.119)		0.144* (0.078)	0.020 (0.120)				
Δ (LnIIP)	2.156*** (0.534)	1.545* (0.808)	0.272 (0.889)	4.208*** (0.983)	0.962* (0.551)	0.228 (0.764)	1.291* (0.737)		2.145 (0.564)	−2.142 (1.900)
Δ (LnIIP(−1))	−0.204 (0.508)	−1.846** (0.750)	−0.952 (0.827)	−4.923*** (1.017)	−4.373*** (0.705)	−2.793*** (0.735)	0.430 (0.704)		1.086 (0.629)	6.454*** (1.974)
Δ (LnIIP(−2))	1.280** (0.526)			−2.146** (0.883)			1.915*** (0.688)		1.012 (0.496)	9.036*** (1.830)
Δ (LnIIP(−3))									−0.668 (0.461)	
Δ (LnRER)	−0.939** (0.358)	−0.963* (0.531)	1.540** (0.719)	−1.466** (0.589)	0.790*** (0.208)	−1.270** (0.615)	−1.181** (0.561)		−0.789 (0.316)	
Δ (LnRER(−1))	0.067 (0.387)		0.936 (0.717)	1.815*** (0.633)	−1.303*** (0.241)	0.116 (0.617)	−2.705*** (0.594)		0.913 (0.310)	
Δ (LnRER(−2))	1.005*** (0.364)		2.332*** (0.728)	1.848*** (0.670)	0.569** (0.253)					
Δ (LnVOL_POS)		−0.020 (0.019)	−0.015 (0.032)		−0.104*** (0.023)				0.032 (0.027)	0.154* (0.084)
Δ (LnVOL_POS(−1))		0.054** (0.022)	−0.101*** (0.037)		−0.060 (0.044)				−0.120 (0.037)	−0.254** (0.117)
Δ (LnVOL_POS(−2))		0.077*** (0.026)			0.038 (0.035)					−0.314*** (0.087)
Δ (LnVOL_POS(−3))					−0.051** (0.020)					
Δ (LnVOL_NEG)	−0.003 (0.023)	0.037 (0.025)		0.029 (0.042)	−0.102*** (0.026)		−0.048 (0.038)	0.054** (0.021)	−0.011 (0.031)	−0.288*** (0.106)
Δ (LnVOL_NEG(−1))		−0.010 (0.025)		−0.081* (0.043)	0.015 (0.031)					
Δ (LnVOL_NEG(−2))		−0.044* (0.023)		−0.178*** (0.041)	−0.065* (0.034)					
Δ (LnVOL_NEG(−3))		0.019 (0.019)			0.064*** (0.022)					
Δ (LnVOL_NEG(−4))		0.067*** (0.018)								
Δ (D2008)	0.000 (0.031)		−0.115** (0.049)	−0.129*** (0.047)	−0.032 (0.024)	0.002 (0.045)	−0.085** (0.041)	−0.037 (0.033)	0.073 (0.028)	0.089 (0.095)
Δ (D2008(−1))			0.030 (0.049)	0.081* (0.047)	0.074*** (0.026)		0.180*** (0.041)	−0.026 (0.033)	0.046 (0.029)	−0.219** (0.091)
Δ (D2008(−2))			−0.148*** (0.053)					−0.071** (0.034)	0.035 (0.028)	
Δ (D2008(−3))									−0.071 (0.029)	
Δ (DCOV)	−0.113*** (0.041)	−0.110* (0.055)	−0.044855 (0.062)	−0.043 (0.062)	−0.141*** (0.033)	−0.028 (0.060)	−0.047 (0.058)	0.051 (0.047)	0.035 (0.039)	−0.137 (0.114)
Δ (DCOV(−1))			−0.167*** (0.061)	−0.156** (0.063)	0.029 (0.033)	−0.175*** (0.064)	−0.336*** (0.059)	−0.111** (0.047)	0.011 (0.037)	
Δ (DCOV(−2))					0.090*** (0.033)			−0.146*** (0.048)	0.055 (0.039)	
Δ (DCOV(−3))									−0.090 (0.038)	

(*continued*)

Table 2.

Variable	China	USA	Singapore	South Korea	Japan	India	Malaysia	Australia	Germany	Brazil
Panel C: Diagnostic tests for NARDL										
F[a]	3.754**	2.952*	3.036**	2.963*	6.798***	2.954*	5.535***	13.001***	2.968*	4.556***
ECM_{t-1}	−0.247*** (0.049)	−0.495*** (0.093)	−0.271*** (0.061)	−0.309*** (0.061)	−0.498*** (0.060)	−0.223*** (0.045)	−0.285*** (0.039)	−0.935*** (0.085)	−0.270*** (0.050)	−0.828 (0.126)
Adj. R^2	0.547	0.604	0.481	0.650	0.824	0.354	0.633	0.696	0.641	0.625
LM[b]	0.3805	0.5090	0.0590	0.7579	0.1117	0.4309	0.8494	0.1498	0.5007	0.3040
RESET[c]	1.927	1.117	1.706	0.917	0.386	1.787	0.758	1.879	0.269	0.563
CU[d]	S	S	S	S	S	S	S	S	S	S
CUQ	US	S	S	S	S	S	S	US	US	S
Wald test long run[e]	40.187***	1.308	27.090***	14.718***	21.270***	0.441	2.298	0.225	45.627***	2.284
Wald test short run	2.298	2.422	2.117	16.205***	8.839***		8.528***	0.009	2.377	1.893

Notes: In this table, we display estimates from equations (2) and (5). Volatility takes positive or negative values. a) At the 5% significance level, considering exogenous variables (k = 6), the lower (upper) bound critical value of the F-test is 1.99(3.99); b) Lagrange multiplier (L.M.) test of residual serial correlation with no serial correlation for the null hypothesis – presented by the probability of X^2 with one degree of freedom (first-order). c) RESET is Ramsey's test for misspecification (number of fitted terms = one) and applied by the value of F-statistic (correctly specified because the RESET statistic is insignificant); d) CU means the CUSUM for model stabilisation, S means "Stable" U means "Unstable"; e) asymmetric statistics test using Wald-test with the null hypothesis is symmetric; and f) *, ** and *** denote significance at 1, 5 and 10% levels, respectively

Secondly, in the short run, the estimation outcomes show that the IIP of trading partners (indicated by at least one significant coefficient) affects export demand in all cases, but for Japan and India. In the long run, the significant effects can be found in all cases but for Singapore, Japan and Australia. As noted in earlier studies, industrial activities are an important determinant of exports (Chi and Cheng, 2016; Nishimura and Hirayama, 2013). Previous studies on Indonesia suggest that exports are propelled by the partners' industrial activities (Asteriou *et al.*, 2016; Sugiharti *et al.*, 2020a, 2020b). In the long-run, the IIP of Malaysia and Germany are negative, suggesting that an expansion in industrial activity in those partners negatively relates to exports from Indonesia to those destinations.

Thirdly, significant positive short-run effects are found in RERs in the cases of China, Singapore, Malaysia, India, Australia and Germany. Significant negative short-run effects from RER are found in the case of the USA. Meanwhile, the long-run significant positive impact of the RER on exports only occurs in shipments to Singapore, South Korea and Australia. Significant and negative impact from RER in the long-run appears in exports to China and the USA. The negative estimates indicate that the IDR's depreciation will worsen Indonesia's exports (i.e. a depreciation of the IDR will worsen exports to China and the USA), contrary to what the theory generally suggests (Alexander, 1952). This could be due to the inelastic export demand in those countries. Several studies have found that periods of domestic currency depreciation could be linked to decreased exports (Bahmani-Oskooee *et al.*, 2016). The link is open to debate, as it depends on, among other things, the characteristics of bilateral trade flows, income and traders' price elasticity (Hegerty, 2016; Sharma and Pal, 2018; Sugiharti *et al.*, 2020a, 2020b). Another possibility is that the impact of exchange rate on exports is asymmetric, as Chi (2020) noted in the case of bilateral flows between Canada and the USA.

Fourthly, the estimation results using NARDL also provide evidence of asymmetric effects in the long-run of real ERV in the case of bilateral Indonesian exports to partners (i.e. China, Singapore, South Korea and Germany), as shown by the F-test that rejects the no-asymmetry hypothesis. In the case of exports to Singapore, South Korea and Germany, the coefficients support a negative ERV. In the case of exports to China, shipments are affected by negative volatility in the long run. In the case of exports to Germany, both positive and negative volatility in the ERV had a dampening impact on Indonesia's shipments in the long run.

In the short run, the F-test estimates show that asymmetric impacts are significant in exports to all countries but the USA and Japan. Significant effects for both positive and negative ERV are found for shipments to China, Singapore, South Korea, India, Germany and the Netherlands. As such, ERV has a dampening impact on Indonesia's exports in the short run for most destination countries. These findings were just as expected and suggest that ERV adversely affects exports. However, the impacts of ERV are mainly within the short run. It should be noted that positive effects that might originate from depreciation of the currency or could be offset by the impact of volatility.

Furthermore, we tested whether the GFC of 2008–2009 and the COVID-19 pandemic have had a significant effect on Indonesia's exports to its main partners. Two dummy variables were introduced in the model to estimate the potential effects from such shocks. First, in the short run, the GFC impact on exports varied across partners. Significant and negative effects were found on exports to the USA, Singapore, India, Malaysia, Germany and the Netherlands. Positive effects from the GFC were found on exports to China, South Korea and Australia. However, in the long run, the effects from the GFC on exports were positive for the case of shipments to China, South Korea, India, Australia and Germany. No negative or significant impacts from the GFC were found in the long run.

As for the COVID-19 pandemic, in the short run, exports to most countries were significantly affected by the pandemic (China, the USA, India, Malaysia, Australia, Germany and the Netherlands). Only exports to Singapore and South Korea reported positive impacts in the short run amid the COVID-19 pandemic. In the long run, only exports to India were significantly affected, while those to the USA and Australia reported significant and positive coefficients.

As for the diagnostic test, we found several important points. *First*, to test for autocorrelation, the Lagrange multiplier statistic was applied. The test results in all cases showed that the residuals do not suffer from autocorrelation. We also ran Ramsey's RESET statistics to check for misspecification. This test was distributed as $\chi 2$ with one degree of freedom. The results show that the statistics for all cases are not significant, except for India, Malaysia and Germany, indicating that most of the models are well identified. We also performed the cumulative sum (CUSUM) and cumulative sum of squares of recursive residuals tests to test the stability of all short- and long-run coefficient estimates. These two are reported as cumulative (CU) and cumulative squares (CUQ) in panel C, Table 1. The CUSUM test was also used to test for stability in the applied models. Finally, we observed the size of adjusted R^2 to measure the goodness of fit.

3.2 Asymmetric effects of real exchange rate volatility on import

Table 2 displays the results of NARDL estimation for the import demand model. Panel A presents the long-run impacts of the IIP, RER and asymmetric effects of real ERV on Indonesia's imports from China, the USA, Singapore, South Korea, Japan, India, Malaysia, Australia, Germany and Brazil, respectively. The short-run estimations are reported in Panel B, and the diagnostic statistics are presented in Panel C.

From Table 2, we gathered that the import demand model is significant in all the studied cases, based on the bound test. The ECM_{t-1} coefficient estimates show long-term relationships or cointegration in all import demand models except for the case of Brazil. The ECM_{t-1} also indicates that, on average, 38% of the ERV in import demand is offset by short-run correction for each quarterly import. In the short run, the estimates show that the IIP of Indonesia has an effect (indicated by at least one significant coefficient estimate) on import demand in the cases of China, the USA, Japan, South Korea, India, Malaysia and Brazil. In the long run, the significant effects only occur for Indonesia's imports from Japan, India, Australia and Brazil (negative). A positive impact on IIP suggests that improvements in industrial activity in Indonesia are linked to an increase in imports. In terms of RER, the short-run outcomes indicate significant negative effects on Indonesia's imports from China, the USA, South Korea, India and Malaysia, and positive effects on imports from Singapore and Japan. Meanwhile, the long run, negative impacts of the RER occur in the cases of imports from China, the USA, South Korea and Germany (negative). Four of the five significant cases show a negative coefficient, which means that as the IDR depreciates, Indonesia's imports from the respective three countries will decline. Only imports from Australia show a positive coefficient, suggesting that imports may be price inelastic, as depreciation leads to higher value in imports.

Regarding ERV, the F-test confirms the existence of the long- and short-run asymmetric effect for our estimates. We find a significant impact from ERV in the long run, on imports from China, Japan, Malaysia, Australia and Germany. Japan, Malaysia and Germany have both significant *LnVOL_POS* and *LnVOL_NEG* coefficients. In the short run, we observed that significant lag coefficients in *LnVOL_POS* or *LnVOL_NEG* are found in all cases but China, India and Germany. As such, imports to Indonesia are substantially affected by ERV, in both positive and negative ERV.

In the short-run, negative and significant impacts from the GFC were found on imports from Singapore, South Korea, Malaysia, Australia and Brazil. The impact of the GFC was positive only for the case of imports from Japan. However, in the long run, only positive and significant effects on imports were identified for imports from China and the USA. Regarding the COVID-19 pandemic, in the short run, the impact was detrimental (significant negative) on imports from all top partners but Germany and Brazil. On the other hand, the COVID-19 outbreak resulted in a negative and significant effect on imports from South Korea.

As for the diagnostic test, we found that the Lagrange multiplier statistics for all cases are not significant, except for Brazil. We also ran Ramsey's RESET statistics to test for misspecification. The test results show that the statistics for all cases are not significant, implying that most of the models are well specified. The CUSUM test shows that the models are stable for all cases, except for China, Australia and Germany. Finally, we used the adjusted R^2 to measure the goodness of fit.

3.3 Discussion

This study suggests that economic policies to support a stable exchange rate are important. Although in the short run, a weak Rupiah can support exports, in the long run, Indonesia's exports do not seem to be supported by the weakening of the IDR. Besides, the large exposure of exports to ERV may offset the positive effects on shipments from a weak currency in the short run. As noted in previous studies on Indonesia's trade (Hegerty, 2016; Pino *et al.*, 2016), depreciation of the IDR may be detrimental to exports from Indonesia as the currency could change global prices for those commodities or spur anxiety among traders (Fang *et al.*, 2006). For instance, export promotion policies through IDR depreciation require stability to be effective (Arize *et al.*, 2021; Fang *et al.*, 2009).

Risk perception among exporters–importers and the government's hedging strategies may contribute to the detrimental effect of ERV on trade. As noted in our findings, asymmetric effects are both positive and negative, suggesting that hedging strategies may need to be holistic and not just one-sided. In some cases, traders in a country with a weakening currency may exercise their hedging strategies to protect themselves from currency appreciation and the expectation that exports will increase. In the case of Indonesia, although the weakening of the IDR may support exports in the short-run, the effects of ERV are significant for both positive and negative volatility in six cases (out of ten). As such, hedging becomes complex and most likely, costly. Such asymmetric hedging may lead to subsequent asymmetric effects on the exchange rate as traders try to recover from high risk. Our findings suggest the need to look for more innovative instruments to hedge and to look at specific market behaviour. Besides, as noted by Bahmani-Oskooee and Fariditavana (2016), real depreciation of currencies is more likely to stem from changes in traders' behaviour than appreciation. Furthermore, due to rigidity of prices, export and import prices may respond asymmetrically to exchange rate movements (Arize *et al.*, 2021).

On the other hand, economic activity in partner countries is a crucial driver of their export demands. A decrease in industrial activity (global economic slowdown) may have significant negative impacts on demand for exports to those countries (i.e. China, the USA, South Korea, India and the Netherlands). Finally, Indonesia's input dependency on its main trade partners is often highlighted in the literature of regional trade (Purwono *et al.*, 2020). It is worth noting that imports from China, Singapore, South Korea and Japan are strategic inputs. This may explain the strong effects of the IIP on imports and the insignificant impact of ERV on imports from China, USA, Singapore, South Korea and Japan.

The short-term positive impact of RER in the case of imports from Singapore and Japan could also be attributed to Indonesia's input dependency.

3.4 Robustness test

As a complementary approach to the estimation of error-correction using time series, we applied three alternative estimators using dynamic panel data. The results are not directly comparable, as time series (ARDL and NARDL) focus on bilateral trade compared to panel data, which pools the sample of countries together. Still, panel data provide some advantages as they allow verification as to whether or not a common equilibrium association exists in the model for ERV and the other variables. Following Iyke and Ho (2019), who estimate the effects of ERV on real consumption in Asian countries, if the parameters on our main equations are assumed to be heterogeneous across the countries included, then a mean group (MG) approach can be used. The MG estimates the parameters for each country and provides a group estimate through the average. Alternatively, if only the intercepts are assumed to be heterogenous across countries, then we can estimate our model by applying the dynamic fixed-effects (DEF). Finally, if we assume that in the short-run, the intercept and the error terms can be heterogeneous across partners, but restrict the long-run estimates to be uniform, then we can us the PMG. The PMG of Pesaran *et al.* (1999) somehow blend the MG and DEF advantages (averaging and pooling, respectively), suggesting it can be advantageous as it is more flexible, as noted in Iyke and Ho (2019). We tested the suitability of the models using the Hausman test, to find whether MG, DEF or PMG estimators fit the data and model best. Results for symmetric effects are also estimated for comparison with asymmetric ones (results in the Supplementary file, Tables S3 and S4).

For both the export and import models, the Hausman test suggests that PMG is more efficient and consistent compared to MG and DEF (Tables 3 and 4). We estimated the results considering the possible effects from GFC and COVID-19 (including a dummy variable), as well as a model without that consideration. The results indicate that in the short-run (Table 3), exports are significantly related to industrial activity (IIP). RER is positively related to exports. However, the results are not significant. Related to ERV, in the short run, the PMG model does not capture significant effects; only the DFE (model without shocks for GFC and COVID) and the MG find that VOL_NEG (negative ERV) has an impact on exports. As for the GFC shock, the impact in the short run is positive, and the effect of COVID-19 on exports is negative. Both the GFC shock and the COVID-19 effects on trade are statistically significant.

In the long-run, the dynamic model finds a negative and significant impact from IIP and RER on exports, contrary to what is expected. However, the results for ERV in the long run (model without shocks) are positive and significant, suggesting that exports are affected by ERV. In the short run, the effects from ERV are not significant. By using a flexible dynamic panel approach, the long-run effects are homogenous, and we allow the short-run effects to be heterogenous. That may be a reason why the results using panel data differ from those for the time series.

Regarding imports (Table 4), we find more consistent results regarding the role of IIP and RER for the short-run, as they indicate that imports are driven by industrial activity and that a depreciation of the IDR may discourage imports. However, in the long run, the estimates suggest that IIP is negatively associated with imports and that exchange rate appreciation may discourage imports. Furthermore, we find evidence that in the long run, imports significantly respond to both positive and negative ERV. As such, both models indicate that ERV (positive and negative) significantly affects exports and imports in the long run. This result suggests that policymakers in Indonesia should consider long-run

Table 3.
Panel data estimates for exports demand model – asymmetric

Variables	PMG	MG	DFE	PMG	MG	DFE
	Coefficient					
Panel A: Long run						
	(2,2,2,4,4)	(2,1,1,4,1)	(1,2,1,3,1)	(2,1,1,3,3,3,3)	(2,1,1,2,2,3,2)	(1,1,1,3,3,3,3)
C	0.735** (0.362)	0.362 (0.417)	0.989*** (0.294)	0.724*** (0.245)	−0.255 (0.572)	1.142*** (0.301)
LnIIP	−0.971** (0.480)	15.609** (7.512)	−0.403 (0.595)	−3.492 (2.253)	−0.027 (3.659)	−0.746 (0.604)
LnRER	−1.379* (0.476)	1.112 (2.839)	−0.011 (0.524)	−6.327 (4.025)	1.352 (2.075)	−0.090 (0.546)
LnVOL_POS	0.275** (0.121)	0.643 (1.221)	−0.120 (0.116)	−0.434 (0.436)	0.215 (0.340)	−0.041 (0.128)
LnVOL_NEG	0.337* (0.123)	0.750 (1.248)	−0.102 (0.116)	−0.463 (0.451)	0.175 (0.303)	−0.017 (0.127)
D2008				−0.030 (0.185)	0.005 (0.121)	−0.134* (0.078)
DCOV				2.586* (1.431)	−0.364 (0.276)	0.379** (0.163)
Panel B: Short run						
Δ (LnEX(−1))	0.462*** (0.017)	0.429*** (0.031)		0.459*** (0.017)	0.439*** (0.024)	
Δ (LnIIP)	1.522*** (0.440)	0.230 (0.201)	1.148*** (0.220)	0.692*** (0.151)	0.133 (0.234)	0.863*** (0.127)
Δ (LnIIP(−1))	−0.810*** (0.248)		−0.144 (0.129)			
Δ (LnRER)	0.168 (0.259)	0.117 (0.195)	0.037 (0.163)	−0.037 (0.091)	−0.090 (0.105)	−0.049 (0.165)
Δ (LnRER(−1))	0.010 (0.186)					
Δ (LnVOLPOS)	−0.056 (0.037)	0.011 (0.037)	0.030 (0.032)	−0.003 (0.039)	0.039 (0.025)	−0.009 (0.037)
Δ (LnVOL_POS(−1))	0.079 (0.058)	0.017 (0.053)	−0.028 (0.030)	0.015 (0.033)	−0.009 (0.011)	0.011 (0.036)
Δ (LnVOL_POS(−2))	−0.049 (0.047)	−0.022 (0.044)	0.003 (0.001)	−0.004 (0.012)		−0.017 (0.013)
Δ (LnVOL_POS(−3))	0.014 (0.015)	0.006 (0.013)				
Δ (LnVOL_NEG)	−0.028 (0.026)	0.016 (0.012)	0.028* (0.016)	0.022 (0.019)	0.042** (0.021)	0.008 (0.032)
Δ (LnVOL_NEG(−1))	0.039 (0.050)			−0.007 (0.020)	−0.015 (0.011)	0.009 (0.028)
Δ (LnVOL_NEG(−2))	−0.026 (0.041)			−0.002 (0.001)		0.010 (0.011)
Δ (LnVOL_NEG(−3))	0.003 (0.012)					
Δ (D2008)				0.043*** (0.008)	−0.005 (0.027)	0.059** (0.030)
Δ (D2008(−1))				−0.037** (0.017)	−0.004 (0.037)	−0.009 (0.036)
Δ (D2008(−2))				0.005 (0.001)	−0.005 (0.013)	−0.007 (0.015)
Δ (DCOV)				−0.224*** (0.073)	−0.080* (0.045)	−0.192*** (0.068)
Δ (DCOV(−1))				0.241*** (0.074)	0.045 (0.029)	0.161** (0.074)
Δ (DCOV(−2))				−0.081*** (0.021)		−0.032 (0.028)
ECT	−0.055** (0.028)	−0.095*** (0.031)	−0.137*** (0.021)	−0.021*** (0.007)	−0.091** (0.038)	−0.138*** (0.022)
Hausman test		0.0718	1.0000		0.1088	1.0000

Notes: *, ** and *** indicate 10, 5 and 1% (significance level), respectively. The Hausman test estimates accept the null hypothesis of homogeneity restrictions on the long-run regressors, indicating that PMG is a more efficient and consistent estimator than MG or DFE

Table 4. Panel data estimates for imports demand model – asymmetric

Variables	Coefficient PMG	MG	DFE	PMG	MG	DFE
Panel A: Long run						
	(1,3,3,2,2)	(1,3,3,2,2)	(2,3,3,4,3)	(1,3,3,2,2,2,2)	(2,2,2,3,3,2,2)	(2,2,2,3,3,2,2)
C	3.352*** (0.772)	2.491** (1.196)	0.090*** (0.309)	3.017*** (0.575)	1.708* (0.998)	0.936*** (0.292)
LnIIP	−1.610*** (0.348)	0.704 (1.332)	−2.885 (2.819)	−1.494*** (0.354)	14.230 (9.063)	−5.629 (4.415)
LnRER	−1.577*** (0.305)	−1.233 (0.646)	−3.272 (2.227)	−1.296*** (0.366)	−0.578 (4.719)	−3.619 (2.840)
LnVOL_POS	0.068 (0.050)	0.090 (0.092)	0.523 (0.490)	0.071 (0.071)	0.604 (0.665)	0.507 (0.536)
LnVOL_NEG	0.042 (0.049)	0.086 (0.102)	0.534 (0.486)	0.126* (0.070)	1.035 (0.733)	0.488 (0.525)
D2008				−0.091** (0.042)	0.043 (0.216)	−0.484 (0.357)
DCOV				0.059 (0.065)	0.943 (0.643)	0.833 (0.690)
Panel B: Short run						
Δ (LnIM(−1))			0.480*** (0.019)		0.450*** (0.016)	0.482*** (0.018)
Δ (LnIM(−2))						
Δ (LnIM(−3))						
Δ (LnIIP)	3.836*** (1.464)	2.978*** (0.832)	2.035*** (0.616)	3.977** (1.760)	1.638** (0.690)	2.567*** (0.433)
Δ (LnIIP(−1))	−1.872 (1.794)	−1.406 (1.270)	−0.721 (0.687)	−2.113 (1.950)	−0.635** (0.334)	−1.162*** (0.267)
Δ (LnIIP(−2))	0.850 (0.580)	0.717 (0.450)	−0.079 (0.269)	0.913 (0.614)		
Δ (LnIIP(−3))						
Δ (LnRER)	0.505 (0.332)	0.413 (0.421)	−0.075 (0.245)	−0.033 (0.332)	−0.400 (0.327)	−0.453** (0.187)
Δ (LnRER(−1))	−1.314** (0.522)	−1.341** (0.589)	−0.155 (0.305)	−0.736 (0.513)	0.138 (0.220)	0.243* (0.139)
Δ (LnRER(−2))	0.577** (0.235)	0.607** (0.235)	0.158 (0.146)	0.382* (0.229)		
Δ (LnVOLPOS)	0.027 (0.017)	0.051 (0.070)	−0.038 (0.048)	−0.025 (0.021)	−0.005 (0.042)	−0.023 (0.032)
Δ (LnVOL_POS(−1))	0.005 (0.028)	−0.008 (0.025)	0.086 (0.065)	0.025 (0.025)	0.059 (0.055)	0.036 (0.031)
Δ (LnVOL_POS(−2))			−0.059 (0.043)		−0.034 (0.027)	−0.017 (0.012)
Δ (LnVOL_POS(−3))			0.012 (0.012)			
Δ (LnVOL_NEG)	0.020 (0.014)	0.043 (0.050)	−0.050 (0.040)	−0.006 (0.024)	0.068 (0.077)	−0.006 (0.028)
Δ (LnVOL_NEG(−1))	−0.030 (0.022)	−0.040 (0.031)	0.003 (0.037)	−0.014 (0.014)	−0.097* (0.054)	−0.032 (0.025)
Δ (LnVOL_NEG(−2))			0.007 (0.014)		0.043** (0.019)	0.018 (0.010)
Δ (LnVOL_NEG(−3))						
Δ (D2008)				0.136*** (0.020)	0.121*** (0.019)	0.116*** (0.020)
Δ (D2008(−1))				−0.060*** (0.019)	−0.065*** (0.011)	−0.064*** (0.013)
Δ (D2008(−2))						
Δ (DCOV)				−0.132** (0.056)	−0.109** (0.049)	−0.097*** (0.034)
Δ (DCOV(−1))				0.075* (0.045)	0.067** (0.032)	0.064*** (0.021)
ECT	−0.227*** (0.049)	−0.378*** (0.077)	−0.040** (0.020)	−0.224*** (0.041)	−0.073* (0.041)	−0.032* (0.019)
Hausman test		0.0669	0.8926		0.9897	0.9887

Notes: *, ** and *** indicate 10, 5 and 1% (significance level), respectively. The Hausman test estimates accept the null hypothesis of homogeneity restrictions on the long-run regressors, indicating that PMG is a more efficient and consistent estimator than MG or DFE (more permanent) volatility (uncertainty) as it can disrupt both exports and imports. At the bilateral trade level, the NARDL model can provide country specific insights to handle uncertainty, mostly in the short run.

4. Conclusion

This paper explores the impact of exchange rate volatility (ERV) on Indonesia's bilateral trade with ten main trading partners using a nonlinear ARDL approach. We separate the effects of real ERV on the export and import sides. An exponential generalised autoregressive conditional heteroscedasticity model is used to estimate the ERV. Using quarterly data over 2006 to 2020, our findings can be summarised as follows.

Concerning Indonesia's exports, a significant error-correction term shows that a long-run cointegrated relationship exists between aggregate exports, the IIP, RERs and real ERV. In the long run, real ERV has a negative effect on Indonesia's shipments to China and Germany, while in the short run, negative effects exist for Indonesia's exports to Australia, China, Germany, India, Malaysia, the Netherlands, Singapore and South Korea. Moreover, the Wald-test results suggest that the bilateral real ERV has a long-run asymmetric impact on exports to China, Singapore, South Korea, India, Malaysia and the Netherlands. In contrast, the short-run effects are significant for exports to six out of ten countries.

As for imports, we can conclude that in the long-run model, uncertainty (ERV) has a significant negative impact on Indonesia's imports from China, Japan, Australia, Malaysia and Germany. However, we also find a positive effect in the case of imports from Australia. In contrast, in the short-run, the results also show a dampening effect of real ERV on Indonesia's imports from the USA, Singapore, South Korea, Japan, Australia and Brazil. Concerning the asymmetric effects, in the long-run model, we reject the null hypothesis in the case of Indonesia's imports from China, Singapore, South Korea, Japan and Germany. In the short-run model, we find that real ERV asymmetrically affects Indonesia's imports from South Korea and Japan.

For a robustness test, we apply a dynamic panel finding that in the long run, asymmetric uncertainty (ERV) plays an important role in Indonesian trade (exports and imports) with top partners. The detrimental effects from depreciation of the IDR on exports and the significant impacts of asymmetric volatility, suggest that policymakers need to consider policies to stabilise IDR, pay more attention to traders' risk behaviour and consider the availability of hedging tools. A weakening currency and the available hedging tools may account for the IDR's asymmetric movements and the possible asymmetry between currency and prices. The asymmetric model seems to be superior to the linear model as it signals that positive and negative effects from ERV differ in their impact on export–import volumes. Further research may look into asymmetric links between the exchange rate and export prices.

References

Aftab, M., Syed, K.B.S. and Katper, N.A. (2017), "Exchange-rate volatility and Malaysian-Thai bilateral industry trade flows", *Journal of Economic Studies*, Vol. 44 No. 1, pp. 99-114, doi: 10.1108/JES-05-2015-0091.

Aghion, P., Bacchetta, P., Ranciere, R. and Rogoff, K. (2009), "Exchange rate volatility and productivity growth: the role of financial development", *Journal of Monetary Economics*, Vol. 56 No. 4, pp. 494-513, doi: 10.1016/j.jmoneco.2009.03.015.

Alexander, S.S. (1952), "Effects of a devaluation on a trade balance", *Staff Papers – International Monetary Fund*, Vol. 2 No. 2, pp. 263-278, doi: 10.2307/3866218.

Ali, H. (2019), "Does downside risk matter more in asset pricing? Evidence from China", *Emerging Markets Review*, Vol. 39, pp. 154-174, doi: 10.1016/j.ememar.2019.05.001.

Arize, A.C., Malindretos, J. and Igwe, E.U. (2017), "Do exchange rate changes improve the trade balance: an asymmetric nonlinear cointegration approach", *International Review of Economics and Finance*, Vol. 49, pp. 313-326, doi: 10.1016/j.iref.2017.02.007.

Arize, A.C., Osang, T. and Slottje, D.J. (2000), "Exchange-rate volatility and foreign trade: evidence from thirteen LDC's", *Journal of Business and Economic Statistics*, Vol. 18 No. 1, pp. 10-17, doi: 10.1080/07350015.2000.10524843.

Arize, A.C., Ogunc, A., Kalu, E.U. and Malindretos, J. (2021), "New evidence on exchange-rate volatility and export flows in Thailand: nonlinearity and asymmetric ARDL investigation", *The International Trade Journal*, Vol. 35 No. 2, pp. 194-218, doi: 10.1080/08853908.2020.1799886.

Asteriou, D., Masatci, K. and Pılbeam, K. (2016), "Exchange rate volatility and international trade: international evidence from the MINT countries", *Economic Modelling*, Vol. 58, pp. 133-140, doi: 10.1016/j.econmod.2016.05.006.

Bahmani-Oskooee, M. and Aftab, M. (2017), "On the asymmetric effects of exchange rate volatility on trade flows: new evidence from US-Malaysia trade at the industry level", *Economic Modelling*, Vol. 63, pp. 86-103, doi: 10.1016/j.econmod.2017.02.004.

Bahmani-Oskooee, M. and Fariditavana, H. (2016), "Nonlinear ARDL approach and the J-curve phenomenon", *Open Economies Review*, Vol. 27 No. 1, pp. 51-70, doi: 10.1007/s11079-015-9369-5.

Bahmani-Oskooee, M. and Nouira, R. (2020), "On the impact of exchange rate volatility on Tunisia's trade with 16 partners: an asymmetry analysis", *Economic Change and Restructuring*, Vol. 53 No. 3, pp. 357-378, doi: 10.1007/s10644-019-09250-y.

Bahmani-Oskooee, M., Ghodsi, S.H. and Halicioglu, F. (2017), "UK trade balance with its trading partners: an asymmetry analysis", *Economic Analysis and Policy*, Vol. 56, pp. 188-199, doi: 10.1016/j.eap.2017.09.005.

Bahmani-Oskooee, M., Harvey, H. and Hegerty, S.W. (2015), "Exchange-rate volatility and commodity trade between the USA and Indonesia", *New Zealand Economic Papers*, Vol. 49 No. 1, pp. 78-102, doi: 10.1080/00779954.2014.901136.

Bahmani-Oskooee, M., Iqbal, J. and Salam, M. (2016), "Short run and long run effects of exchange rate volatility on commodity trade between Pakistan and Japan", *Economic Analysis and Policy*, Vol. 52, pp. 131-142, doi: 10.1016/j.eap.2016.09.002.

Bailey, M.J., Tavlas, G.S. and Ulan, M. (1987), "The impact of exchange-rate volatility on export growth: Some theoretical considerations and empirical results", *Journal of Policy Modeling*, Vol. 9 No. 1, pp. 225-243, doi: 10.1016/0161-8938(87)90010-X.

Belke, A.H., Göcke, M. and Werner, L. (2015), "Exchange rate volatility and other determinants of hysteresis in exports-empirical evidence for the euro area", *Review of Economic Analysis*, Vol. 7 No. 1, pp. 24-53, doi: 10.15353/rea.v7i1.1484.

Broll, U. and Eckwert, B. (1999), "Exchange rate volatility and international trade", *Southern Economic Journal*, Vol. 66 No. 1, pp. 178-185, doi: 10.2307/1060843.

Campbell, D.L. (2020), "Relative prices and hysteresis: evidence from US manufacturing", *European Economic Review*, Vol. 129, p. 103474, doi: 10.1016/j.euroecorev.2020.103474.

Cao, S., Dong, W. and Tomlin, B. (2015), "Pricing-to-market, currency invoicing and exchange rate pass-through to producer prices", *Journal of International Money and Finance*, Vol. 58, pp. 128-149, doi: 10.1016/j.jimonfin.2015.08.004.

Chi, J. (2020), "Asymmetric exchange rate effects on cross-border freight flows between the United States and Canada", *Transportation Research Record: Journal of the Transportation Research Board*, Vol. 2674 No. 4, pp. 348-361, doi: 10.1177/0361198120912242.

Chi, J. and Cheng, S.K. (2016), "Do exchange rate volatility and income affect Australia's Maritime export flows to Asia?", *Transport Policy*, Vol. 47, pp. 13-21, doi: 10.1016/j.tranpol.2015.12.003.

Chit, M.M., Rizov, M. and Willenbockel, D. (2010), "Exchange rate volatility and exports: new empirical evidence from the emerging east Asian economies", *World Economy*, Vol. 33 No. 2, pp. 239-263, doi: 10.1111/j.1467-9701.2009.01230.x.

Choudhry, T. (2005), "Exchange rate volatility and the United States exports: evidence from Canada and Japan", *Journal of the Japanese and International Economies*, Vol. 19 No. 1, pp. 51-71, doi: 10.1016/j.jjie.2003.11.002.

Clark, P.B. (1973), "Uncertainty, exchange risk, and the level of international trade", *Economic Inquiry*, Vol. 11 No. 3, pp. 302-313, doi: 10.1111/j.1465-7295.1973.tb01063.x.

De vita, G. and Abbott, A. (2004), "The impact of exchange rate volatility on UK exports to EU countries", *Scottish Journal of Political Economy*, Vol. 51 No. 1, pp. 62-81, doi: 10.1111/j.0036-9292.2004.05101004.x.

Dickey, D.A. and Fuller, W.A. (1979), "Distribution of the estimators for autoregressive time series with a unit root", *Journal of the American Statistical Association*, Vol. 74 No. 366a, pp. 427-431, doi: 10.1080/01621459.1979.10482531.

Doğanlar, M. (2002), "Estimating the impact of exchange rate volatility on exports: evidence from Asian countries", *Applied Economics Letters*, Vol. 9 No. 13, pp. 859-863, doi: 10.1080/13504850210150906.

Engle, R.F. and Ng, V.K. (1993), "Measuring and testing the impact of news on volatility", *The Journal of Finance*, Vol. 48 No. 5, pp. 1749-1778, doi: 10.1111/j.1540-6261.1993.tb05127.x.

Ethier, W. (1973), "International trade and the forward exchange market", *The American Economic Review*, Vol. 63 No. 3, pp. 494-503.

Fang, W., Lai, Y. and Miller, S.M. (2006), "Export promotion through exchange rate changes: exchange rate depreciation or stabilization?", *Southern Economic Journal*, Vol. 72 No. 3, p. 611, doi: 10.2307/20111836.

Fang, W., Lai, Y. and Miller, S.M. (2009), "Does exchange rate risk affect exports asymmetrically? Asian evidence", *Journal of International Money and Finance*, Vol. 28 No. 2, pp. 215-239, doi: 10.1016/j.jimonfin.2008.11.002.

Fitrianti, S. (2017), "The exchange rate volatility and export performance: the case of Indonesia's exports to Japan and US", *Buletin Ekonomi Moneter Dan Perbankan*, Vol. 20 No. 1, pp. 1-22, doi: 10.21098/bemp.v20i1.724.

Franke, G. (1991), "Exchange rate volatility and international trading strategy", *Journal of International Money and Finance*, Vol. 10 No. 2, pp. 292-307, doi: 10.1016/0261-5606(91)90041-H.

Fratzscher, M. (2006), "On the long-term effectiveness of exchange rate communication and interventions", *Journal of International Money and Finance*, Vol. 25 No. 1, pp. 146-167, doi: 10.1016/j.jimonfin.2005.10.007.

Gupta, M. and Varshney, S. (2021), "Does exchange rate volatility influence import commodities of India-US? Evidence from ARDL approach", *South Asian Journal of Business Studies, Ahead-of-Print*, doi: 10.1108/SAJBS-02-2021-0056.

Hall, S., Hondroyiannis, G., Swamy, P., Tavlas, G. and Ulan, M. (2010), "Exchange-rate volatility and export performance: do emerging market economies resemble industrial countries or other developing countries?", *Economic Modelling*, Vol. 27 No. 6, pp. 1514-1521, doi: 10.1016/j.econmod.2010.01.014.

Hausmann, R., Panizza, U. and Rigobon, R. (2006), "The long-run volatility puzzle of the real exchange rate", *Journal of International Money and Finance*, Vol. 25 No. 1, pp. 93-124, doi: 10.1016/j.jimonfin.2005.10.006.

Hayakawa, K. and Kimura, F. (2009), "The effect of exchange rate volatility on international trade in east Asia", *Journal of the Japanese and International Economies*, Vol. 23 No. 4, pp. 395-406, doi: 10.1016/j.jjie.2009.07.001.

Hegerty, S.W. (2016), "Commodity-price volatility and macroeconomic spillovers: evidence from nine emerging markets", *The North American Journal of Economics and Finance*, Vol. 35, pp. 23-37, doi: 10.1016/j.najef.2015.10.014.

Heriqbaldi, U., Widodo, W. and Ekowati, D. (2020), "Real exchange rate misalignment and currency crises", *Bulletin of Indonesian Economic Studies*, Vol. 56 No. 3, pp. 345-362, doi: 10.1080/00074918.2019.1662885.

Hooper, P. and Kohlhagen, S.W. (1978), "The effect of exchange rate uncertainty on the prices and volume of international trade", *Journal of International Economics*, Vol. 8 No. 4, pp. 483-511, doi: 10.1016/0022-1996(87)90001-8.

Iyke, B.N. and Ho, S.Y. (2020), "Consumption and exchange rate uncertainty: evidence from selected Asian countries", *The World Economy*, Vol. 43 No. 9, pp. 2437-2462.

Klaassen, F. (2004), "Why is it so difficult to find an effect of exchange rate risk on trade?", *Journal of International Money and Finance*, Vol. 23 No. 5, doi: 10.1016/j.jimonfin.2004.03.009.

Koutmos, G. and Martin, A.D. (2003), "Asymmetric exchange rate exposure: theory and evidence", *Journal of International Money and Finance*, Vol. 22 No. 3, pp. 365-383, doi: 10.1016/S0261-5606(03)00012-3.

Kroner, K.F. and Lastrapes, W.D. (1993), "The impact of exchange rate volatility on international trade: reduced form estimates using the GARCH-in-mean model", *Journal of International Money and Finance*, Vol. 12 No. 3, pp. 298-318, doi: 10.1016/0261-5606(93)90016-5.

McKenzie, M. (2002), "The economics of exchange rate volatility asymmetry", *International Journal of Finance and Economics*, Vol. 7 No. 3, pp. 247-260, doi: 10.1002/ijfe.189.

McKenzie, M.D. and Melbourne, R.M.IT. (1999), "The impact of exchange rate volatility on international trade flows", *Journal of Economic Surveys*, Vol. 13 No. 1, pp. 71-106, doi: 10.1111/1467-6419.00075.

Mahdavi, S. (2000), "Do German, Japanese, and US export prices asymmetrically respond to exchange rate changes? Evidence from aggregate data", *Contemporary Economic Policy*, Vol. 18 No. 1, pp. 70-81, doi: 10.1111/j.1465-7287.2000.tb00007.x.

Nishimura, Y. and Hirayama, K. (2013), "Does exchange rate volatility deter Japan-China trade? Evidence from pre- and post-exchange rate reform in China", *Japan and the World Economy*, Vol. 25-26, pp. 90-101, doi: 10.1016/j.japwor.2013.03.002.

Obstfeld, M. and Rogoff, K. (1998), "Risk and exchange rates", *National Bureau of Economic Research*, doi: 10.3386/w6694.

Pesaran, M.H., Shin, Y. and Smith, R.J. (2001), "Bounds testing approaches to the analysis of level relationships", *Journal of Applied Econometrics*, Vol. 16 No. 3, pp. 289-326, doi: 10.1002/jae.616.

Pesaran, M.H., Shin, Y. and Smith, R.P. (1999), "Pooled mean group estimation of dynamic heterogeneous panels", *Journal of the American Statistical Association*, Vol. 94 No. 446, pp. 621-634, doi: 10.1080/01621459.1999.10474156.

Pino, G., Tas, D. and Sharma, S.C. (2016), "An investigation of the effects of exchange rate volatility on exports in east Asia", *Applied Economics*, Vol. 48 No. 26, pp. 2397-2411, doi: 10.1080/00036846.2015.1122730.

Pozo, S. (1992), "Conditional Exchange-Rate volatility and the volume of international trade: evidence from the early 1900s", *The Review of Economics and Statistics*, Vol. 74 No. 2, p. 325, doi: 10.2307/2109665.

Purwono, R., Esquivias, M.A., Sugiharti, L., Heriqbaldi, U. and Handoyo, R.D. (2020), "Value chains, production networks and regional integration: the case of Indonesia", *Jurnal Ekonomi Malaysia*, Vol. 54 No. 1, pp. 135-151, doi: 10.17576/JEM-2020-5401-10.

Sercu, P. and Vanhulle, C. (1992), "Exchange rate volatility, international trade, and the value of exporting firms", *Journal of Banking and Finance*, Vol. 16 No. 1, pp. 155-182, doi: 10.1016/0378-4266(92)90083-C.

Sharma, C. and Pal, D. (2018), "Exchange rate volatility and India's cross-border trade: a pooled mean group and nonlinear cointegration approach", *Economic Modelling*, Vol. 74, pp. 230-246, doi: 10.1016/j.econmod.2018.05.016.

Shin, Y., Yu, B. and Greenwood-Nimmo, M. (2014), "Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework", in Sickles, R.C. and Horrace, W.C. (Eds), *Festschrift in Honor of Peter Schmidt*, Springer New York, NY, pp. 281-314, doi: 10.1007/978-1-4899-8008-3_9.

Sugiharti, L., Esquivias, M.A. and Setyorani, B. (2020a), "The impact of exchange rate volatility on Indonesia's top exports to the five main export markets", *Heliyon*, Vol. 6 No. 1, p. e03141, doi: 10.1016/j.heliyon.2019.e03141.

Sugiharti, L., Purwono, R. and Esquivias Padilla, M.A. (2020b), "Analysis of determinants of Indonesian agricultural exports", *Entrepreneurship and Sustainability Issues*, Vol. 7 No. 4, pp. 2676-2695, doi: 10.9770/jesi.2020.7.4(8).

Tsay, R.S. (2005), *Analysis of Financial Time Series*, John Wiley and Sons, Vol. 543, doi: 10.1002/0471746193.

Zhang, W.G., Yu, X. and Liu, Y.J. (2018), "Trade and currency options hedging model", *Journal of Computational and Applied Mathematics*, Vol. 343, pp. 328-340, doi: 10.1016/j.cam.2018.04.059.

Further reading

Bahmani-Oskooee, M. and Gelan, A. (2018), "Exchange-rate volatility and international trade performance: Evidence from 12 African countries", *Economic Analysis and Policy*, Vol. 58, pp. 14-21, doi: 10.1016/j.eap.2017.12.005.

Dellas, H. and Zilberfarb, B.-Z. (1993), "Real exchange rate volatility and international trade: a reexamination of the theory", *Southern Economic Journal*, Vol. 59 No. 4, p. 641, doi: 10.2307/1059729.

Grauwe, P.D. (1988), "Exchange rate variability and the slowdown in growth of international trade", *Staff Papers – International Monetary Fund*, Vol. 35 No. 1, p. 63, doi: 10.2307/3867277.

Supplementary material

The supplementary material for this article can be found online

Corresponding author

Miguel Angel Esquivias can be contacted at: miguel@feb.unair.ac.id

Appendix

Country	Test for unit root in	LnEX	LnIIP	LnRER	LnVOL_POS	LnVOL_NEG
China	I(0)	−2.121	−2.198	−1.962	−7.846	−0.733
	I(I)	−9.750***	−10.942***	−5.676***	−2.336***	−11.114***
USA	I(0)	−3.969***	−2.409	−2.029	−0.651	−0.357
	I(I)	−8.635***	−3.869***	−5.465***	−8.769***	−10.264***
Singapore	I(0)	−2.177	−0.378	−1.766	−0.428	−0.076
	I(I)	−9.213***	−4.490***	−6.577***	−7.494***	−8.014***
South Korea	I(0)	−1.827	−1.816	−1.699	−0.131	−0.025
	I(I)	−5.589***	−3.769***	−5.402***	−8.080***	−10.489***
Japan	I(0)	−0.781	−3.309**	−2.373	−0.286	−0.066
	I(I)	−7.268***	−6.535***	−7.420***	−3.328**	−4.038***
India	I(0)	−3.089**	−1.879	−1.378	−0.219	−0.032
	I(I)	−11.012***	−8.960***	−7.846***	−3.650***	−3.026**
Malaysia	I(0)	−2.412	−0.467	−1.879	−0.327	−0.059
	I(I)	−7.017***	−3.556**	−6.612***	−3.589***	−4.002***
Australia	I(0)	−2.220	−1.108	−2.390	−0.791	−0.585
	I(I)	−7.373***	−4.281***	−6.897***	−9.072***	−4.248***
Germany	I(0)	−2.056	−2.958**	−2.264	2.052	1.892
	I(I)	−8.858***	−7.826***	−5.334***	−11.094***	−10.659***
The Netherlands	I(0)	−2.690*	−3.297**	−2.393**	1.348	1.314
	I(I)	−10.188***	−6.128***	−5.377***	−11.581***	−11.135***

Notes: *, ** and *** indicate a 10, 5 and 1% significance level, respectively. Unit root test of variables using the ADF. All variables are free from unit root problem. All variables are stationary at the first difference (I[I]). Several variables are also stationary at level (I[0]), namely, LnEX (the USA, India and the Netherlands), LnIIP (Japan, Germany and the Netherlands) and LnRER (the Netherlands) in the export model. LnIM is also stationary at level (I[0]) for the import model for the cases of China, South Korea, Malaysia and Australia. Unit root tests for the data, including the dummy variables, were also carried out but are not displayed due to space limitations (available upon request). The variables have no unit root problems

Table A1. Unit root test (dependent variable: export)

Country	Test for unit root in	LnIM	LnIIP	LnRER	LnVOL_POS	LnVOL_NEG
China	I(0)	−2.754*	−1.276	−1.962	−0.794	−0.733
	I(I)	−3.092**	−9.179***	−5.676***	−7.846***	−11.114***
USA	I(0)	−2.203	−1.276	−2.029	−0.651	−0.357
	I(I)	−11.521***	−9.179***	−5.465***	−8.769***	−10.264***
Singapore	I(0)	−2.168	−1.276	−1.766	−0.428	−0.076
	I(I)	−6.940***	−9.179***	−6.577***	−7.494***	−8.014***
South Korea	I(0)	−3.449**	−1.276	−1.699	−0.131	−0.025
	I(I)	−8.456***	−9.179***	−5.402***	−8.080***	−10.489***
Japan	I(0)	−1.799	−1.276	−2.373	−0.286	−0.066
	I(I)	−6.992***	−9.179***	−7.420***	−3.328**	−4.038***
India	I(0)	−2.339	−1.276	−1.378	−0.219	−0.032
	I(I)	−7.681***	−9.179***	−7.846***	−3.650***	−3.026**
Malaysia	I(0)	−3.737***	−1.276	−1.879	−0.327	−0.059
	I(I)	−6.466***	−9.179***	−6.612***	−3.589***	−4.002***
Australia	I(0)	−4.728***	−1.276	−2.390	−0.791	−0.585
	I(I)	−11.979***	−9.179***	−6.897***	−9.072***	−4.248***
Germany	I(0)	−2.581	−1.276	−2.264	2.052	1.892
	I(I)	−8.441***	−9.179***	−5.334***	−11.094***	−10.659***
Brazil	I(0)	−2.455	−1.276	0.272	−0.114	−0.034
	I(I)	−8.536***	−9.179***	−6.151***	−12.974***	−11.555***

Notes: *, ** and *** indicate a 10, 5 and 1% significance level, respectively. Unit root test of variables using the ADF. All variables are free from unit root problem. All variables are stationary at the first difference (I[I]). Several variables are also stationary at level (I[0]), namely, LnEX (the USA, India and the Netherlands), LnIIP (Japan, Germany and the Netherlands) and LnRER (the Netherlands) in the export model. LnIM is also stationary at level (I[0]) for the import model for the cases of China, South Korea, Malaysia and Australia. Unit root tests for the data, including the dummy variables, were also carried out but are not displayed due to space limitations (available upon request). The variables have no unit root problems

Table A2.
Unit root test (dependent variable: import)

Table A3. Results of NARDL for export demand model

Variable	China	USA	Singapore	South Korea	Japan	India
Panel A: Long run						
	(4, 0, 3, 2, 1)	(4, 3, 0, 0, 0)	(1, 2, 3, 0, 1)	(1, 3, 1, 0, 2)	(2, 2, 3, 3, 2)	(3, 0, 4, 4, 0)
Constant	6.935* (8.078)	6.695* (5.637)	−1.626* (−5.248)	1.094* (3.648)	5.363* (4.433)	−5.733* (−5.763)
$LnIIP^*_{t\ t}$	3.562** (2.537)	1.640* (5.206)	3.253* (4.656)	1.821 (1.218)	−4.227 (−0.975)	19.619 (0.670)
$LnRER_{t\ t}$	−3.958* (−5.593)	−1.096* (−6.187)	1.017 (1.220)	−1.372 (−1.459)	−9.369 (−1.260)	25.309 (0.518)
$LnVOL\_POS_t$	−0.032 (−0.519)	0.014 (1.636)	0.002 (0.036)	0.053 (0.461)	−0.230 (−0.663)	7.272 (0.646)
$LnVOL\_NEG_t$	−0.138** (−2.425)	0.013 (1.473)	0.114*** (1.925)	0.127 (1.234)	−0.147 (−0.421)	7.814 (0.640)
Panel B: Short run						
Δ(LnEX(t−1))	−0.204** (−2.073)	0.334** (2.331)			−0.303* (−2.815)	−0.589* (−5.119)
Δ(LnEX(t−2))	−0.063 (−0.626)	0.270** (2.132)				−0.386* (−3.325)
Δ(LnEX(t−3))	−0.264* (−2.880)	0.307* (2.709)				
Δ(LnIIP*)		1.790* (6.363)	0.901* (3.108)	1.349* (4.460)	−0.091 (−0.304)	
Δ(LnIIP*(t−1))		−0.618 (−1.518)	−0.969* (−3.117)	1.834* (5.709)	0.573** (2.151)	
Δ(LnIIP*(t−2))		−2.515* (−5.706)		0.550*** (1.743)		
Δ(LnIIP*(t−3))						
Δ(LnRER)	−0.827*** (−1.677)		0.049 (0.079)	0.685*** (1.770)	−1.143* (−4.648)	1.273** (2.085)
Δ(LnRER(t−1))	−0.349 (−0.640) (−0.640)		−0.823 (−1.289) (−1.289)		0.093 (0.283)	3.205* (3.609)
Δ(LnRER(t−2))	2.106* (4.019)		−1.745* (−2.835)		0.731* (2.816)	2.051* (2.077)
Δ(LnRER(t−3))						−3.822* (−3.723)
Δ(LnVOL_POS)	0.071** (2.143)				−0.091* (−3.661)	0.098 (1.396)
Δ(LnVOL_POS(t−1)) Δ(LnVOL_POS(t−1))	0.1523* (4.052)				−0.081* (−3.706)	0.357* (5.081)
Δ(LnVOL_POS(t−2))					−0.103* (−4.665)	0.109** (2.599)
Δ(LnVOL_POS(t−3))						−0.182* (−4.234)
Δ(LnVOL_NEG)	−0.009 (−0.224)		−0.035 (−1.009)	−0.029 (−0.969)	−0.017 (−0.901)	
Δ(LnVOL_NEG(t−1))				−0.064** (−2.144)	−0.079* (−2.765)	
Δ(LnVOL_NEG(t−2))						
Panel C: Diagnostic tests for NARDL						
F^a	11.811*	5.823*	4.966**	2.555	3.527	6.083**
ECM_{t-1}	−0.606** (−8.051)	−0.882** (−5.635)	−0.513** (−5.200)	−0.264** (−3.733)	−0.155** (−4.415)	−0.064** (−5.798)
Adj. R^2	0.619	0.660	0.360	0.531	0.639	0.526
LM^b	0.036	0.264	0.514	0.861	0.505	0.063
$RESET^c$	0.314	2.560	0.124	3.884	0.489	7.89**
CU^d	S	S	US	S	S	S
CUQ	S	S	S	S	S	S
Wald test long run	35.558**	0.035	24.963**	9.248**	2.336	1.816
Wald test short run	12.113**		1.068	2.637	3.601	2.977

Notes: In this table, we display estimates from equations (1) and (4). Volatility takes positive or negative values. a) At the 5% significance level, considering exogenous variables (k = 4), the lower (upper) bound critical value of the F-test is 2.86(4.01); b) Lagrange multiplier (L.M.) test of residual serial correlation with no serial correlation for the null hypothesis – presented by the probability of X^2 with one degree of freedom (first-order). c) RESET is Ramsey's test for misspecification (number of fitted terms = one) and applied by the value of F-statistic (correctly specified because the RESET statistic is insignificant); d) CU means the CUSUM for model stabilisation, S means "Stable" U means "Unstable"; e) asymmetric statistics test using Wald-test with the null hypothesis is symmetric; and f) *, ** and *** denote significance at 1, 5 and 10% levels, respectively

(*continued*)

Table A3.

Variable	Malaysia	Australia	Germany	The Netherlands
Panel A: Long run				
	(4, 4, 4, 4, 0)	(4, 4, 2, 3, 3)	(4, 4, 3, 1, 0)	(2, 3, 3, 4, 2)
Constant	10.750* (6.167)	8.030* (4.095)	5.170* (6.251)	−1.815* (−6.818)
$LnIIP*_{t\ t}$	−3.451* (−11.128)	−4.413 (−0.660)	−2.824** (−2.621)	5.900* (5.795)
$LnRER_{t\ t}$	0.623 (1.388)	−9.170 (−0.571)	−0.393 (−1.060)	−0.855* (−3.202)
$LnVOL\_POS_t$	0.146 (1.109)	−3.885 (−0.742)	−0.125*** (−1.689)	0.265** (2.253)
$LnVOL\_NEG_t$	0.070 (0.530)	−3.919 (−0.737)	−0.104 (−1.375)	0.281** (2.470)
Panel B: Short run				
Δ(LnEX(t−1))	0.421* (3.356)	−0.532* (−4.534)	−0.426* (−3.940)	0.153 (1.286)
Δ(LnEX(t−2))	0.181 (1.600)	−0.410* (−3.620)	−0.489* (−4.302)	
Δ(LnEX(t−3))	−0.221*** (−1.909)	−0.416* (−3.748)	−0.459* (−4.248)	
Δ(LnIIP*)	0.872 (1.614)	−0.00373 (−0.006)	0.759* (4.098)	1.951* (5.765)
Δ(LnIIP*(t−1))	3.466* (4.530)	0.896 (1.495)	1.407* (4.966)	−2.937* (−5.714)
Δ(LnIIP*(t−2))	1.637** (2.292)	1.895* (3.497)	1.302* (4.163)	−1.222* (−3.275)
Δ(LnIIP*(t−3))	2.020* (2.792)	1.280** (2.219)	0.921* (3.131)	
Δ(LnRER)	1.058** (2.412)	1.418** (2.434)	0.160 (0.725)	0.041 (0.110)
Δ(LnRER(t−1))	−1.543* (−2.907)	1.394 (1.440)	−0.156 (−0.663)	−0.169 (−0.396)
Δ(LnRER(t−2))	1.232** (2.062)		0.760* (3.343)	0.983** (2.447)
Δ(LnRER(t−3))	−1.171*** (−1.891)			
Δ(LnVOL_POS)	−0.110** (−2.246)	−0.042 (−0.513)	−0.018 (−0.885)	−0.091** (−2.122)
Δ(LnVOL_POS(t−1)) Δ(LnVOL_POS(t−1))	−0.140** (−2.481)	0.331* (4.115)		−0.134** (−2.267)
Δ(LnVOL_POS(t−2))	−0.139** (−2.279)	0.172* (3.080)		−0.226* (−3.749)
Δ(LnVOL_POS(t−3))	−0.113** (−2.242)			−0.108** (−2.543)
Δ(LnVOL_NEG)		−0.178** (−2.297)		0.146* (2.966)
Δ(LnVOL_NEG(t−1))		0.419* (3.813)		−0.191* (−3.431)
Δ(LnVOL_NEG(t−2))		0.199* (3.300)		
Panel C: Diagnostic tests for NARDL				
F^a	6.785**	3.013	7.099**	8.388**
ECM_{t-1}	−0.972** (−6.167)	−0.160** (−4.104)	−0.393** (−6.256)	−0.836** (−6.836)
Adj. R^2	0.648	0.579	0.661	0.685
LM^b	0.031	0.182	0.739	0.299
$RESET^c$	6.068**	1.361	6.063**	0.057
CU^d	S	US	C	S
CUQ	US	S	C	S
Wald test long run	1.326	1.148	3.262	14.803**
Wald test short run	1.092	0.279	0.604	7.525**

Table A4. Results of NARDL for import demand model

Variable	China	USA	Singapore	South Korea	Japan	India
Panel A: Long run						
	(4, 2, 0, 0, 0)	(2, 1, 0, 0, 0)	(4, 4, 3, 4, 0)	(1, 4, 2, 4, 3)	(1, 4, 3, 4, 3)	(1, 2, 1, 0, 0)
Constant	0.548* (5.106)	7.483* (4.424)	3.359* (4.925)	1.623* (5.175)	0.546* (4.244)	−4.162* (−3.610)
$LnIIP_t$	9.620 (0.782)	−0.757 (−0.565)	4.164 (0.539)	1.758 (0.283)	5.132 (1.212)	9.293* (2.889)
LnRER	−5.309 (−1.059)	−2.187* (−3.676)	−4.849) (−1.022)	−2.461** (−2.525)	−2.871 (−1.171)	2.932 (0.860)
LnVOL_POSt	−0.245 (−0.889)	0.044 (1.353)	0.093 (0.385)	−0.478 (−1.575)	−0.065 (−0.161)	0.030 (0.171)
LnVOL_NEGt	−0.060 (−0.320)	0.027 (0.629)	0.209 (0.714)	−0.426 (−1.278)	0.036 (0.085)	0.267 (1.335)
Panel B: Short run						
Δ(LnIM(−1))	−0.265** (−2.237)	−0.239** (−2.183)	0.038 (0.347)			
Δ(LnIM(−2))	−0.245** (−2.303)		−0.027 (−0.245)			
Δ(LnIM(−3))	−0.327* (−3.050)		−0.464* (−3.918)			
Δ(LnIIP)	1.819* (2.908)	1.556** (2.242)	2.689* (2.724)	4.661* (4.199)	1.609*** (1.974)	0.285 (0.339)
Δ(LnIIP(−1))	−1.018 (−1.576)		−1.052 (−1.052)	−1.851 (−1.581)	−1.062 (−1.188)	−2.886* (−3.732)
Δ(LnIIP(−2))			0.129 (0.128)	−0.421 (−0.401)	1.959** (2.235)	
Δ(LnIIP(−3))			2.737* (2.803)	2.769* (2.728)	2.502* (3.340)	
Δ(LnRER)			0.120 (0.181)	−2.541* (−3.787)	0.525*** (1.734)	−1.286** (−2.157)
Δ(LnRER(−1))			0.903 (1.382)	2.213* (3.126)	−0.954** (−2.657)	
Δ(LnRER(−2))			2.110* (3.142)		1.001* (2.999)	
Δ(LnRER(−3)) Δ(LnVOL_POS)			0.002	0.045	−0.097*	
Δ(LnVOL_POS(−1))			(0.065) –0.105*	(0.863) 0.106***	(−3.070) –0.180*	
Δ(LnVOL_POS(−2))			(−2.915) –0.009	(1.673) 0.199*	(−3.098) –0.112*	
Δ(LnVOL_POS(−3))			(−0.259) 0.063***	(3.195) 0.122**	(−3.849) –0.058**	
Δ(LnVOL_NEG)			(1.876)	(2.021) –0.097***	(−2.061) –0.057	
Δ(LnVOL_NEG(−1))				(−1.688) 0.086	(−1.657) –0.142*	
Δ(LnVOL_NEG(−2))				(1.382) –0.132**	(−3.64) –0.108**	
Δ(LnVOL_NEG(−3))				(−2.353)	(−2.378)	
Panel C: Diagnostic tests for NARDL						
F^a	4.592**	3.63	4.331**	5.032**	3.154	2.439
ECM_{t-1}	−0.1088** (−5.000)	−0.455** (−4.427)	−0.193** (−4.920)	−0.330** (−5.295)	−0.221** (−4.198)	−0.220** (−3.632)
Adj. R^2	0.449	0.432	0.525	0.582	0.640	0.309
LM^b	0.154	0.348	0.092	0.452	0.239	0.510
$RESET^c$	16.271**	4.336**	0.372	2.822	0.007	2.141
CU^d	US	US	S	S	S	S
CUQ	S	US	S	US	US	S
Wald test long run	17.178**	0.724	14.165**	13.750**	12.529**	2.410
Wald test short run			0.835	11.331**	6.563**	

Notes: In this table, we display estimates from equations (2) and (5). Volatility takes positive or negative values. a) At the 5% significance level, considering exogenous variables (k = 4), the lower (upper) bound critical value of the F-test is 2.86(4.01); b) Lagrange multiplier (L.M.) test of residual serial correlation with no serial correlation for the null hypothesis – presented by the probability of X^2 with one degree of freedom (first-order). c) RESET is Ramsey's test for misspecification (number of fitted terms = one) and applied by the value of F-statistic (correctly specified because the RESET statistic is insignificant); d) CU means the CUSUM for model stabilisation, S means "Stable" U means "Unstable"; e) asymmetric statistics test using Wald-test with the null hypothesis is symmetric; and f) *, ** and *** denote significance at 1, 5 and 10% levels, respectively

(*continued*)

Table A4.

Variable	Malaysia	Australia	Germany	Brazil
Panel A: Long run				
	(1, 0, 2, 0, 1)	(1, 1, 0, 0, 1)	(1, 1, 2, 0, 1)	(1, 3, 1, 0, 4)
Constant	4.384* (4.903)	−4.246* (−6.680)	3.772* (5.103)	19.934* (7.676)
$LnIIP_t$	1.693 (0.774)	2.507** (2.401)	−0.058 (−0.036)	−9.909* (−3.404)
LnRER	−2.419 (−1.237)	1.811* (4.165)	−1.712** (−2.465)	0.868 (1.555)
LnVOL_POSt	−0.806* (−2.701)	0.111 (1.650)	0.144 (1.088)	0.075 (0.634)
LnVOL_NEGt	−0.700** (−2.238)	0.162** (2.085)	0.183 (1.425)	−0.243 (−1.659)
Panel B: Short run				
Δ(LnIM(−1))				
Δ(LnIM(−2))				
Δ(LnIM(−3))				
Δ(LnIIP)		3.648* (5.134)	2.567* (4.453)	−1.852 (−0.946)
Δ(LnIIP(−1))				6.084* (3.597)
Δ(LnIIP(−2))				7.496* (4.432)
Δ(LnIIP(−3))				
Δ(LnRER)	−1.259*** (−1.722)		−0.378 (−1.062)	−0.658 (−0.908)
Δ(LnRER(−1))	−2.505* (−3.272)		0.922** (2.528)	
Δ(LnRER(−2))				
Δ(LnRER(−3)) Δ(LnVOL_POS)				
Δ(LnVOL_POS(−1))				
Δ(LnVOL_POS(−2))				
Δ(LnVOL_POS(−3))				
Δ(LnVOL_NEG)	0.011	0.039	−0.0463	−0.202**
Δ(LnVOL_NEG(−1))	(0.212)	(1.469)	(−1.654)	(−2.599) 0.135
Δ(LnVOL_NEG(−2))				(1.560) 0.163***
Δ(LnVOL_NEG(−3))				(1.964) 0.275*
				(3.506)
Panel C: Diagnostic tests for NARDL				
F^a	4.414**	8.210*	4.835**	10.699**
ECM_{t-1}	−0.384** (−4.890)	−0.703** (−6.663)	−0.289** (−5.122)	−0.928** (−7.671)
Adj. R^2	0.336	0.617	0.430	0.637
LM^b	0.584	0.255	0.912	0.048**
$RESET^c$	0.040	1.078	0.126	0.706
CU^d	S	S	S	S
CUQ	S	S	S	US
Wald test long run	0.351	0.541	5.365**	1.004
Wald test short run			1.800	0.027