# The J-curve effect in agricultural commodity trade: an empirical study of South East Asian economies



Ivan D. Trofimov
*Victoria University of Wellington, Wellington, New Zealand*

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## Abstract

**Purpose** – In this paper we examine the validity of the J-curve hypothesis in four Southeast Asian economies (Indonesia, Malaysia, the Philippines and Thailand) over the 1980–2017 period.

**Design/methodology/approach** – We employ the linear autoregressive distributed lags (ARDL) model that captures the dynamic relationships between the variables and additionally use the nonlinear ARDL model that considers the asymmetric effects of the real exchange rate changes.

**Findings** – The estimated models were diagnostically sound, and the variables were found to be cointegrated. However, with the exception of Malaysia, the short- and long-run relationships did not attest to the presence of the J-curve effect. The trade flows were affected asymmetrically in Malaysia and the Philippines, suggesting the appropriateness of nonlinear ARDL in these countries.

**Originality/value** – The previous research tended to examine the effects of the real exchange rate changes on the agricultural trade balance and specifically the J-curve effect (deterioration of the trade balance followed by its improvement) in the developed economies and rarely in the developing ones. In this paper, we address this omission.


**Keywords** J-curve, Agriculture, Nonlinear ARDL, Cointegration

**Paper type** Research paper

## Introduction

Agriculture remains a major sector across most of Southeast Asian/SEA economies (with the exception of highly urbanised Singapore and Brunei Darussalam). The Association of Southeast Asian Nations (ASEAN) Secretariat publication (ASEAN Secretariat, 2019) indicates that the agricultural production share of gross domestic product (GDP) ranged from 6.2% in Thailand to 12.5% in Indonesia and 24.6% in Myanmar as of 2018 and for the average of ASEAN economies stood at 12 and 10.3% in 2010 and 2018 respectively. The consideration of the international effects on the agricultural economies of the region is therefore a salient empirical issue. The empirical literature tended to examine the volatility of agricultural commodity prices, the possible deterioration of agricultural terms of trade, the influence of the fluctuations in the exchange rates, amongst other issues. However, the Marshall-Lerner condition and the J-curve effect hypothesis that describe the relationship between the fluctuation of exchange rate and the country's balance in agricultural trade (as opposed to the aggregate trade) received limited consideration in the literature, the works by

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Kim *et al.* (2004), Baek and Koo (2008), Baek *et al.* (2009), Gong and Kinnucan (2015) and Chebbi and Olarreaga (2019) being the notable exceptions.

The purpose of the paper is to address this shortcoming. In contrast to previous studies that focused, due to data availability issues, on a limited number of economies (most commonly the USA), we examine four Southeast Asian economies: Malaysia, Indonesia, the Philippines and Thailand and consider the presence (absence) of the J-curve effect using the annual data for the 1980–2017 period. A multivariate autoregressive distributed lag (ARDL) model is employed to capture delayed effects of the exchange rate on the trade balance and the switch from initially negative to positive effects of the former variable on the latter, as well as to address the possible asymmetric effects (when depreciation and appreciation of the currency deliver changes in the trade balance that are different not only in sign but also in size) [1]. We therefore use linear and nonlinear versions of ARDL for each individual economy, an approach that has been previously taken to examine the J-curve in the aggregate trade of the country (Bahmani-Oskooee and Gelan, 2012; Nusair, 2017). The use of aggregate trade data to establish J-curve effect may result in the rejection of the hypothesis or obfuscate the presence of the curve, particularly when the country in question is industrialised and principally trades in manufactured goods (that tend to have higher import and export elasticities). In this regard, the focus on the developing countries' trade and consideration of the agricultural trade alone (as opposed to the aggregate trade) may be warranted.

The choice of the study period and the sample of economies is warranted due to the following considerations. The economies in question differ in many respects: the countries that were net exporters of agricultural products throughout the period (Indonesia, Malaysia and Thailand) and the economy that switched between two states during the period (Philippines). The economies also differed in terms of foreign exchange management policies and the pace of moving to more flexible exchange rates (independent floating in the Philippines, managed floating in Thailand and Indonesia versus fixed peg in Malaysia; IMF, 2005), the dominance of agricultural exports in the total exports or the proportion of food imports in the total imports, the overall level of development (Malaysia versus Indonesia). The study period includes a number of salient macroeconomic events: Third World debt crisis of the early 1980s, the Asian Financial crisis of 1997–8, the Global Financial crisis of 2008–08, the period of low agricultural commodity prices in the 1980s and conversely the commodity boom of the 2000s. In addition, in line with the remark of Rose and Yellen (1989, p. 58) about the particular equilibrium nature of the J-curve phenomenon that necessitates higher disaggregation of trade data and consideration of the trade of the sector or industry to obtain more reliable results, we focus on a specific economic sector (agriculture) rather than on aggregate trade of the country.

The outline of the paper is as follows. Section 2 reviews the existing literature on the J-curve effect in the aggregate and agricultural trade. Section 3 describes the data, the model and the econometric method. Section 4 presents the empirical results, whilst Section 5 provides the conclusion alongside the discussion of the findings.

## Literature review

The hypothesis that underpins the J-curve effect is the deterioration of the trade balance of the country immediately after depreciation (devaluation) of its currency followed by the balance improvement as economic agents adjust to changed foreign currency conditions. The adjustment processes that constitute the J-curve effect are explained with reference to the Bickerdike-Robinson-Metzler (BRM) and Marshall-Lerner (ML) conditions (the latter as a special case of the former). The exposition of the conditions follows Baek *et al.* (2006, pp. 5–7). The trade balance is defined as:

$$\mathrm{TB} = P_x Q_x - \mathrm{EP}_x Q_m \quad (1)$$

where TB is the trade balance of country A that experiences its currency devaluation (depreciation), $P_x$ is the domestic price of the good in country A bound for exports to country B, $P_x$ is the foreign price of the good in country B bound for exports to country A (or equivalently the good imported by country A), $E$ is a nominal exchange rate, $Q_x$ and $Q_m$ are export and import volumes. The effect of the exchange rate on country A's trade balance is obtained from the differentiation of the above trade balance equation with respect to $E$ (BRM condition):

$$\frac{\mathrm{dTB}}{\mathrm{dE}} = P_x Q_x \left[\frac{(1+\varepsilon)\eta}{(\varepsilon+\eta)}\right] - \mathrm{EP}_x Q_m \left[\frac{(1-\eta)\varepsilon}{(\varepsilon+\eta)}\right] \quad (2)$$

where $\eta$ and $\eta$ are domestic and foreign price elasticities of demand for imports and $\varepsilon$ and $\varepsilon$ domestic and foreign elasticities of the supply of exports. Assuming trade balance in initial equilibrium (TB $= 0$) and perfect supply elasticities in both trading countries ($\varepsilon \to \infty$ and $\varepsilon \to \infty$), the BRM condition transforms to ML condition as:

$$\eta + \eta > 1 \quad (3)$$

Whilst the sum of elasticities in ML exceeds unity and the devaluation is to improve trade balance in the long-run, the interim adjustments in the short-run are more complex, as illustrated by the J-curve effect. Immediately after devaluation (currency-contract period), the prior contracts are executed none-withstanding devaluation, i.e. whilst $E$ rises (or falls depending on whether direct or indirect quotation is used for the exchange rate) and there is deterioration in the trade balance, there is no effect on prices or volumes. In a pass-through period that follows, the prices adjust with no changes in $Q_x$ and $Q_m$. The change in the prices of imported goods initiates the substitution process between imported and domestically-produced goods, but this process is yet incomplete at this stage. The ultimate effect at this stage depends on the values of import demand and export supply elasticity, with multiple outcomes possible (Magee, 1973, p. 317). For instance, the inelastic demand for the exports of the country that devalues its currency (Country A) and the inelastic demand for the imports of that country would result in trade balance deterioration during the pass-through stage, whilst the inelastic supply for the country's exports and imports would bring trade balance improvement. In the quantity adjustment period, the $Q_x$ and $Q_m$ start to adjust, thereby completing the substitution process. With both export and import elasticities increasing compared to previous periods, $Q_x$ rises faster in response to the fall in goods prices (in foreign currency), whilst $Q_m$ falls faster in response to goods price increase (in domestic currency). The trade balance thereby moves into surplus. Given the differential elasticities in currency-contract and pass-through periods and the reactions in the quantity adjustment period, there is no *a priori* assumption that J-curve pattern will eventuate: Magee noted the possibility of I-, L-, M-, N-, V- and W-curves.

The empirical research on the Marshall-Lerner condition and J-curve effect has been voluminous and extensive, covering both the aggregate trade of the country with the rest of the world, as well as bilateral trade for a pair of economies, including trade at disaggregated level (Lal and Lowinger, 2002; Nusair, 2017; Bahmani-Oskooee and Harvey, 2016). The findings were generally mixed, conditional on the sample of the economies in the study, the period considered, the econometric methodology used, the specification of the model, the broader political and macroeconomic context as well as definition of the relevant variables.

Some earlier empirical studies of the J-curve effect considered the Southeast Asian economies alongside economies from other regions. This was the case of Lal and Lowinger

(2002) who included Indonesia in a sample of seven economies, applied Johansen cointegration method to the aggregate trade data and established the significant relationship between trade balance and exchange rate in the short- but not the long-run. The study by Onafowora (2003) focused on the bilateral trade relationships of Indonesia and other East Asian economies with the US and Japan respectively. The author applied Johansen cointegration test and impulse-response functions and established improvement in the trade balance of Indonesia with both trading partners in the long run.

More recently, a number of studies with a specific focus on individual South East Asian economies came to the fore. Bahmani-Oskooee and Harvey (2009) considered the trading partners of Indonesia over the 1974Q1–2008Q4 period and confirmed the presence of J-curve in the trade with five out of thirteen trading partners. Bahmani-Oskooee and Kantipong (2001) examined the bilateral trade of Thailand with its five major trading partners (Germany, Japan, Singapore, UK and the US) over the 1973Q1–1997Q4 period and, using cointegration analysis, managed to confirm the J-curve effect in two trading relationship (with Japan and the US). Harvey (2018) applied linear and nonlinear ARDL models to the international trade of Philippines with its major trading partners over the 1981Q1–2015Q4 period and confirmed the significant J-curve effect in the trade with two trading partners, based on a linear model. It additionally established short- and long asymmetry in the trade with three partners. In the case of Malaysian trade, Duasa (2007) and Yusoff (2010) reported no significant link between the exchange rate and Malaysia's aggregate trade balance with the rest of the world. This result echoed the earlier finding that trade balance in Malaysia was affected by real money and less so by the nominal exchange rate (Liew *et al.*, 2003). On the other hand, Bahmani-Oskooee and Aftab (2018) applied nonlinear ARDL model to examine Malaysia–China bilateral trade over the March 2001–December 2015 period and demonstrated significant, yet asymmetric effects of ringgit depreciation in up to one-third of industries, including the largest one that accounted for more than 25% of bilateral trade. In a similar vein, Bahmani-Oskooee and Aftab (2017b), based on monthly Malaysia–Thailand bilateral trade flow data from 61 industries over the April 2000–December 2014 period, indicated the favourable effects of depreciation in the majority of industries and additionally confirmed long-term asymmetric effects in 43% of industries in question. In the case of Malaysia–Singapore bilateral trade during April, 2000–December, 2014 period, Bahmani-Oskooee *et al.* (2016) discovered short-term asymmetric effects of depreciation in almost all industries and showed significant positive influence of depreciation of the trade balance of petroleum and electrical machinery industries that constitute 40% of aggregate trade. In the case of Malaysia-Korea bilateral trade (over the March 2001–December 2015 period), the significant or asymmetric effects of depreciation were found to be limited, despite the application of nonlinear and disaggregated model: significant short-run cumulative, significant adjustment asymmetry and long-run asymmetric effects were discovered only in 13, 21 and 18 out of 55 industries respectively (Bahmani-Oskooee and Aftab, 2017a, b).

The studies focussing on the sectoral trade were limited. Amongst them is Meade (1988) who examined the J-curve in the capital, consumer industrial supplies sectors in the US, Yazici (2010) who considered the effect in services sector in Turkey, Wijeweera and Dollery (2013) and Prakash and Maiti (2016) who compared J-curve effects in the goods and services sectors in Australia and Fiji respectively and Cheng (2020) who estimated J-curve for the aggregate services trade and for the trade in particular services categories in the US. The findings were likewise mixed.

The early research on the linkage between foreign exchange and agricultural trade concerned the estimation of the export demand equations with exchange rates as one the regressors (the overview of the relevant literature contained in Carter and Pick, 1989, pp. 714–715). Konandreas *et al.* (1978) established the sensitivity of the US exports to exchange rate fluctuations during the 1954–72 period, notwithstanding the fact of low statistical significance of the coefficients of the exchange rate. The study of the five export commodities in the US by Chambers and Just (1981) that used US quarterly data generally confirmed the finding, but identified the statistically significant exchange rate coefficients only in the case of corn, but not wheat and soybeans. Batten and Belongia (1986) focused on the exchange rate effects of export volumes and stated that such effects were small and short-lived. On the other hand, Henneberry *et al.* (1987) rejected any significant exchange rate effects, on the grounds of small demand and supply elasticities for agricultural commodities (leading to fluctuation in prices as per cobweb model, but not in export volumes), or higher relative importance of other factors, such as foreign income or terms of trade.

The earliest study of the effects of depreciation on the US agricultural trade balance (as opposed to export prices and volumes alone) by Carter and Pick (1989) established the initially negative effects of depreciation that persisted for a period of nine months and improvement of the trade balance thereafter (hence, the presence of the J-curve effect). The study was based on quarterly data from the 1973–85 period and employed a linear regression model with polynomial distributed lags to capture the pass-through effect of the exchange rate depreciation on trade variables.

Doroodian *et al.* (1999) considered US trade in agricultural and manufacturing goods and estimated the trade balance equation with Shiller lag structure imposed on the exchange rate variable (the preferred approach when the exact functional form of the distributed lag model is unknown). The findings supported the J-curve effect in agricultural trade and failed to support the hypothesis in the manufacturing trade.

Baek *et al.* (2009) examined the relationship between the exchange rate and the agricultural trade balance in the trade of the US with its 15 principal trading partners over 1989Q1–2007Q4 period, using the autoregressive distributed lags (ARDL) methodology and did not identify any J-curve patterns. Their study, however, did not control for the country-specific factors that could affect the relationship, such as national production and trade structure, macroeconomic and exchange rate policies and various structural jolts (spikes in commodity prices, financial crises, political disturbances).

Kim *et al.* (2004, pp. 141–142) examined the bilateral agricultural trade between the USA and Canada using the quarterly time series for the 1983–2000 period and applying the vector error-correction and vector moving average models (VECM and VEM). Exchange rate was found to have significant impact on both USA exports to Canada and US imports from Canada in the short-run, whilst in the long-run the exchange rate effects remained significant only for the US imports, being a major determinant of growing US trade deficit with Canada. On the other hand, the exchange rate was weakly exogenous, I.e. causing deviation of the model from the steady-state, but not affected by other variables (the result consistent with a small size of the agricultural sector relative to the total US economy). The effect of exchange rates on the US agricultural prices and agricultural income in the short- and long-run were also significant, albeit marginal in size.

Chebbi and Olarreaga (2019) considered agricultural trade balance in Tunisia during the 1965–2011 period and employed the Johansen-Juselius test of cointegration and VECM model. In contrast to the majority of the literature, the depreciation was found to have no effect on the trade balance in the short-run and negative effect in the long-run, principally due to the shift in the exchange rate policies that took place at the end of 1980s. Prior to the reform, the devaluations of Tunisian dinar within the fixed exchange rate regime were used to boost the competitiveness of agricultural exports and improve agricultural trade balance, whilst after the policy change a flexible exchange rate regime was introduced, allowing more stable currency but little positive effects on the trade balance.

Yazici (2006) examined the agricultural trade balance in Turkey using similar methodology to Carter and Pick (1989), i.e. linear regression with polynomial distributed lags. The identified movements in the trade balance appeared to contradict J-curve hypothesis (depreciation leading to deterioration of trade balance in the short-term followed by transient improvement and by another deterioration). In addition, the sum of the exchange rate coefficients at different lags was negative, indicating the negative effects of devaluation on the trade balance in the long-run.

The knowledge gap is thus twofold. Firstly, a greater empirical work is needed in the analysis of the Marshall-Lerner condition and J-curve effect in the agricultural trade of the developing economies, including Southeast Asia. The empirical research revealed certain inconsistencies in the analysis of disaggregated trade balance data at the industry level and thus analysis at the level of individual industries and sectors is worthwhile. Secondly, an application of nonlinear models is warranted. As noted by Bussiere (2013), the adjustments by exporters and importers tend to be different during currency appreciation versus depreciation, due to asymmetric reaction of export and import prices to exchange rate changes, implying that trade quantities and trade balance adjust asymmetrically as well. In addition, the asymmetric adjustments in the trade balance may be attributed to the fact that trade balance is a single measure that incorporates exports and imports which originate in countries with different rules, regulations and trade policy context and thus different adjustment dynamics. Overall, as noted by Bahmani-Oskooee and Aftab (2018), the disregard of these nuances in the linear models tends to demonstrate an insignificant link between trade balance and exchange rate in cases where such link exists and is significant.

## Methodology

### *Model*

In line with previous research, the empirical model is represented by the following equation:

$$\ln \mathrm{TB}_{\mathrm{it}} = \alpha + \beta_{1i} \ln Y_{\mathrm{it}} + \beta_{2i} \ln \mathrm{YROW}_{\mathrm{it}} + \beta_{3i} \ln \mathrm{RER}_{\mathrm{it}} + \varepsilon_t, \tag{4}$$

where $\ln \mathrm{TB}_{\mathrm{it}}$, $\ln Y_{\mathrm{it}}$, $\ln \mathrm{YROW}_{\mathrm{it}}$ and $\ln \mathrm{RER}_{\mathrm{it}}$ indicate respectively the logarithms of the trade balance of country $i$ with the rest of the world, income (GDP) of this country, income (GDP) of the "rest of the world" and real or real effective exchange rate (RER or REER) of the country at period $t$.

Concerning the latter variable, we use two alternative exchange rate measures, obtained respectively from the International Monetary Fund (IMF) and the US Department of Agriculture (USDA) databases. The former measure is the RER of the country, defined as $REER = EP_d/P_f$, where $E$ is the country's nominal exchange rate (units of foreign currency per one unit of domestic currency), $P_d$ and $P_f$ are domestic and foreign prices. The latter is the bilateral RER of the country with the US, defined as $\mathrm{RER} = \mathrm{EP}_f P_d$, where the nominal exchange rate is expressed as units of domestic currency per one unit of foreign currency. In both cases, the foreign currency is the US dollar, the domestic currency is the currency of country in question and domestic and foreign prices are approximated by the respective consumer price indexes. Following Chebbi and Olarreaga (2019), the use of bilateral exchange rate data for the analysis of agricultural trade of the country with rest of the world is justified, given the absence of the IMF REER data for many economies and (as shown further) tight correlation of the two measures (bilateral RER and REER).

With USDA version of RER, we hypothesise the existence of the J-curve if the short-run coefficient of RER is negative but its long-run coefficient is positive, whilst with the IMF version of REER, the relation is the opposite (positive coefficient in the short-run and negative in the long-run). With regard to other variables we expect the negative effect of GDP of the country on the country's trade balance, given that increase in GDP leads to import growth with imports entering the aggregate demand equation with a negative sign. However, the positive effects of domestic GDP on the trade balance may be experienced, if domestic production of importables grows ahead of their consumption, resulting in the decrease in import volume (Magee, 1973). The growth of GDP in the test of the world will stimulate country's export and have positive influence on the trade balance (hence the sign of the respective coefficient is expected to be positive).

*Data*

The data on the value of agricultural exports and imports is taken from the Food and Agriculture Organisation Statistics (*FAOSTAT)* database. The agricultural products traded include crops and livestock, with the primary data collected according to the standard *International Merchandise Trade Statistics (MITS)* methodology. The agricultural trade data covers all major types of transactions, including regular trade, barter trade, goods on consignment, goods on financial lease, goods traded between enterprises under common ownership, goods traded on government account and goods traded with the intent of further processing. The statistical value of exports and imports excludes trade taxes, such customs duties, value added tax, excise duty, levies, export refunds or other taxes with similar effect, but includes expenses associated with bringing the goods to the place of destination (freight and insurance). The quantities of goods exported or imported are measured in tonnes. The value of exports and imports are measured respectively on FOB (free on board) and CIF (cost, insurance and freight) bases. The trade flows cover all crop and livestock products imported and exported during the reference year by country. The complete list of goods at aggregated level includes beverages, cereals and preparations, dairy products and eggs, fats and oils (excluding butter), fodder and feeding stuff, fruit and vegetables, meat and meat preparations, nonedible crude materials, sugar and honey, tobacco, alcoholic beverages, pulses, roots and tubers, other food.

The value of exports or imports is defined as the export or import quantities (tonnes for crops and thousand units for livestock products) multiplied by the per unit export or import values (reported as free-on-board/FOB or cost-insurance-freight/CIF values). The reported data is in nominal terms; it is not deflated to the real terms, since the trade balance in agricultural products is calculated as the ratio of the nominal values of exports to imports. The United States Department of Agriculture Economic Research Service (USDA ERS) agricultural exchange rate data set contains the annual REER for the relevant countries. USDA ERS uses IMF *International Financial Statistics* data on nominal exchange rates alongside the IMF data on the consumer price indexes (with 2010 set as a base year). The trade weights for REER calculation were obtained from the USDA ERS *Global Agricultural Trade System (GATS)* 2014–16 data. As an alternative indicator we also used the IMF REER index derived using Consumer Price Index (CPI) data with 2010 as a base year. The GDP data has likewise been obtained from the USDA ERS *International Macroeconomic Data Set* ("Real GDP, 2010 Dollars, Historical" file). The "rest of the world GDP" for an individual economy is defined as the world GDP in particular year net of GDP of that individual economy. The description of the variables is provided in Table A1 in the Appendix.

*Econometric method*

We used linear and nonlinear version of ARDL model to verify the presence of J-curve effect. In the error-correction form, the linear ARDL is given as:

$$\Delta \ln \mathrm{TB}_{\mathrm{it}} = \alpha_0 + \sum_{k=1}^{n} \beta_{t-k} \Delta \ln \mathrm{TB}_{i,t-k} + \sum_{k=0}^{n} \lambda_{t-k} \Delta \ln Y_{i,t-k} + \sum_{k=0}^{n} \gamma_{t-k} \Delta \ln \mathrm{YROW}_{i,t-k}$$
$$+ \sum_{k=0}^{n} \upsilon_{t-k} \Delta \ln \mathrm{RER}_{i,t-k} + \delta_1 \ln \mathrm{TB}_{i,t-1} + \delta_2 \ln Y_{i,t-1} + \delta_3 \ln \mathrm{YROW}_{i,t-1} + \delta_4 \ln \mathrm{RER}_{i,t-1} + \mu_t \quad (5)$$

Firstly, we conduct Pesaran *et al.* (2001) bounds test (test of the joint significance of the lagged variables in levels) to establish the presence of cointegration amongst the variables. The null hypothesis of the test is of the coefficients of the lagged level variables equal to zero (i.e. $\delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$), whilst the alternative hypothesis is of the absence of such equality ($\delta_1 \neq \delta_2 \neq \delta_3 \neq \delta_4 \neq 0$). The test statistics is compared with the critical values at upper and lower bound, with cointegration present when the statistics exceed the upper bound $I(1)$. The absence of cointegration is indicated when the test statistics is below the lower bound $I(0)$, whilst the indeterminate case when the test statistics is between the bounds. To confirm the finding, the Banerjee-Dolado-Mestre (BDM) cointegration *t*-test is conducted (Banerjee *et al.*, 1998). The null hypothesis of no cointegration ($H_0 : \delta_1 = 0$ and the respective t-statistics is above the test critical value) is contrasted with an alternative hypothesis of the presence of cointegration ($H_1 : \delta_1 < 0$ and the t-statistics is smaller than the critical value). Secondly, we ensure that the ARDL model passed the requisite diagnostic tests (normality, heteroskedasticity, stability, functional form and, importantly, the serial correlation) and that error-correction coefficient is significant and falls within $(0, -1)$ range. Thirdly, the long-run relationships are established by normalising the coefficients of the lagged level regressors on the coefficient of the lagged level dependent variable ($\delta_2$, $\delta_3$ and $\delta_4$ on $\delta_1$). Respectively, the short-run relationships are indicated by the coefficients of the first-differenced variables.

To ensure the appropriateness of ARDL model that does not envisage any variables that are integrated of order two, $I(2)$, we conducted unit root tests of the first differences of the variables. The stationarity of the variables would indicate the absence of $I(2)$ integration order. To address the possible presence of serial correlation in the ARDL model, the sufficient number of lags of the dependent variable and regressors was allowed: given the use of annual data with 38 observations, the maximum lag for the selection procedure was set at four and the optimal number of lags was selected using Akaike Information Criterion (AIC) that, as opposed to Schwarz Information Criterion (SIC), is less restrictive in terms of the lag selection.

Within linear ARDL framework, with the ratio of exports to imports as representation of the trade balance and USDA representation of REER, following depreciation the J-curve effect is indicated if coefficient of the first difference of RER is negative ($\upsilon_t < 0$), but the long-run normalised coefficient of RER is positive ($\delta_4 > 0$). An alternative interpretation, given by Rose and Yellen (1989) is the presence of the J-curve when coefficient of the first difference of RER is negative at lower lags, but positive at higher lags. Conversely, with IMF representation of REER, the J-curve effect holds, when $\upsilon_t > 0$ is positive and $\delta_4 < 0$ is negative, or, in line with Rose-Yellen interpretation, the $\upsilon_t$ at lower lags is positive but at higher lags is negative.

The nonlinear version of ARDL (Shin *et al.*, 2013), as an extension of linear ARDL, likewise disentangles short- and long-run impacts and additionally allows for asymmetric effects of appreciation and depreciation on the trade balance (the necessary feature, if it is assumed that price elasticities and expectations change following exchange rate change). Nonlinear ARDL is given as:

$$\Delta \ln \mathrm{TB}_{\mathrm{it}} = a' + \sum_{k=1}^{n1} b'_k \Delta \ln \mathrm{TB}_{i,t-k} + \sum_{k=0}^{n2} c'_k \Delta \ln Y_{i,t-k} + \sum_{k=0}^{n3} d'_k \Delta \ln \mathrm{YROW}_{i,t-k}$$

$$+ \sum_{k=0}^{n4} e'_k \Delta \ln RER^{+}_{t-k} + \sum_{k=0}^{n5} f'_k \Delta \ln RER^{-}_{t-k} + \delta_0 \ln TB_{i,t-1} + \delta_1 \ln Y_{i,t-1} + \delta_2 \ln YROW_{i,t-1} +$$

$$+\delta_3 \ln RER^{+}_{t-1} + \delta_4 \ln RER^{-}_{t-1} + \upsilon_t \tag{6}$$

where $\mathrm{RER}^{+}_t$ and $\mathrm{RER}^{-}_t$ represent the partial sums of positive and negative changes in the real exchange rate.

The partial sums are calculated as:

$$\mathrm{RER}^{+}_t = \sum_{j=1}^{t} \Delta \ln \mathrm{RER}^{+}_j = \sum_{j=1}^{t} \max(\Delta \ln \mathrm{RER}_j, 0) \tag{7}$$

$$\mathrm{RER}^{-}_t = \sum_{j=1}^{t} \Delta \ln \mathrm{RER}^{-}_j = \sum_{j=1}^{t} \min(\Delta \ln \mathrm{RER}_j, 0) \tag{8}$$

With USDA representation of REER and trade balance as ratio of exports to imports, in the nonlinear ARDL model, the J-curve effect is present following depreciation, when the normalised coefficient of $\ln RER^{-}_{t-1}$ is positive and significant ($\delta_4 > 0$), whilst the coefficients of $\Delta \ln RER^{-}_{t-k}$ are negative and significant. Conversely, with IMF representation of REER, the J-curve effect is established when $\delta_4 < 0$ and the coefficients of $\Delta \ln RER^{-}_{t-k}$ are positive and significant. The Marshall-Lerner condition is satisfied if $\delta_4 > 0$ and $\delta_4 < 0$ in the long-run for the USDA and IMF representations of REER respectively (in both linear and nonlinear ARDL).

The null hypothesis of no cointegration ($H_0 : \delta_0 = \delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$) is contrasted with a cointegration alternative ($H_A : \delta_0 \neq \delta_1 \neq \delta_2 \neq \delta_3 \neq \delta_4 \neq 0$) and the respective statistics is examined with reference to $I(0)$ and $I(1)$ critical bounds. Additionally, the long-term asymmetry test is conducted to ensure the appropriateness of the nonlinear model, assuming the absence of asymmetry under the null, $\gamma^{+} = -\delta_3/_{\delta_0} = \gamma^{-} = -\delta_4/_{\delta_0}$.

## Empirical results

The first step in empirical analysis was examination of descriptive statistics (Table A2 in the Appendix). The means, medians, maximum and minimum values were positive for all variables except the trade balance in the Philippines. The variables were generally symmetric, but exhibited significant skewness in the case of trade balance in the Philippines and Thailand. Excess kurtosis was observed for the REER in Indonesia. In most instances the maximum or minimum values did not exceed (fall below) the mean by more than 20–30% percent, but a number of outliers were observed in the trade balance in each of the four countries (in Malaysia in 2001, in Indonesia in 1981–82, in Thailand in 1982 and in the Philippines in 1980). These outliers correspond to the crises periods – the period immediately after the Asian financial crisis of 1997–98 or the second oil shock and developing countries debt crisis of the early 1980s. On the other hand, the Jarque–Bera test failed to reject the null hypothesis of the normal distribution for all variables and economies in question, except for the logarithm of the trade balance in the Philippines and Thailand. Overall, the descriptive statistics point to the presence of nonlinearities and certain outliers that justify the use of nonlinear models.

To ascertain the appropriateness of applying the ARDL models we performed unit root tests [augmented Dickey–Fuller test (ADF) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS)]

| Country | Test | lnY | lnYROW | lnRER | lnREER | lnTB |
|---|---|---|---|---|---|---|
| Indonesia | ADF (c) | −4.612 | −4.733 | −6.905 | | −6.140 |
| | ADF (ct) | −4.541 | −4.676 | −7.108 | | −6.060 |
| | KPSS (c) | 0.081 | 0.060 | 0.226 | | 0.074 |
| | KPSS (ct) | 0.079 | 0.058 | 0.035 | | 0.064 |
| Malaysia | ADF (c) | −4.937 | −4.752 | −4.820 | −4.612 | −7.831 |
| | ADF (ct) | −4.962 | −4.693 | −4.755 | −4.551 | −7.774 |
| | KPSS (c) | 0.175 | 0.059 | 0.115 | 0.069 | *0.146* |
| | KPSS (ct) | 0.072 | 0.067 | 0.076 | 0.063 | 0.030 |
| Philippines | ADF (c) | *−3.162* | −4.764 | −4.870 | −5.928 | −6.653 |
| | ADF (ct) | −8.541 | −4.708 | −4.869 | −6.047 | −6.806 |
| | KPSS (c) | *0.442* | 0.060 | 0.109 | 0.143 | *0.330* |
| | KPSS (ct) | *0.085* | 0.068 | 0.065 | 0.059 | *0.093* |
| Thailand | ADF (c) | *−3.205* | −4.756 | −4.516 | | −6.396 |
| | ADF (ct) | *−3.571* | −4.696 | −4.561 | | −6.399 |
| | KPSS (c) | 0.320 | 0.058 | 0.159 | | 0.099 |
| | KPSS (ct) | 0.067 | 0.068 | 0.063 | | 0.066 |

**Note(s):** The “c”, “t” and “ct” subscripts indicate test specifications with constant, trend, and constant plus trend
**Source(s):** Table by author

**Table 1.**
Unit root tests’ results

on the first difference of the variables (Table 1). Given that ARDL cannot be used if any of the variables is $I(2)$, the (trend) stationarity of the first difference of the variable indicates the absence of $I(2)$ integration order. Regarding the ADF test (with either constant or constant plus trend deterministic component), the unit root null hypothesis was rejected at 1% significance level for all variables except the logarithms of the GDP in the Philippines and Thailand (where the hypothesis was rejected at the 5% level). For the KPSS test, the null hypothesis of (trend) stationarity was not rejected in the majority of cases. Whilst in most cases, the KPSS statistic was obtained based on automatic selection of the bandwidth, in few instances (logarithm of GDP in the Philippines and the logarithm of the trade balance in Malaysia and the Philippines) the fixed bandwidth was needed to prevent the rejection of the null. In Table 1, these latter KPSS statistics are indicated in italic, whilst the ADF rejection of the unit root null at 5% level is indicated in italics. Overall, we conclude that neither of the variables is $I(2)$.

Tables 2 and 3 contain linear short- and long-run ARDL estimates for the four economies. For Malaysia and the Philippines, the two linear models have been estimated, one with the USDA RER, the other with IMF REER measure (we respectively denoted them as *Models 1 and 2*). The maximum number of lags for the selection purpose varied for each individual economy and, in the case of the second model for Malaysia, the number of lags was set fixed. The lag selection was driven by the need to eliminate serial correlation in ARDL, and to this end, the AIC was used as the one that tends to select a larger number of lags (as opposed to Schwarz Bayesian Criterion that tends to “under-fit” the model).

All four models have passed the requisite diagnostic tests: the residuals are normally distributed, as indicated by the Jarque-Bera test and contain no serial correlation, as demonstrated by the Breusch–Pagan Lagrange multiplier (LM) test (in the case of Indonesia, the LM statistics only barely exceeds the 5% critical level, however, according to the $F$-test version of the Breusch–Godfrey test the null hypothesis of no serial correlation is not rejected). All models have also passed Ramsey (RESET) functional misspecification test (with null hypothesis of correct specification not rejected in all models) and the Breusch–Pagan–Godfrey LM test of heteroskedasticity (with the null hypothesis of no heteroskedasticity likewise not rejected). All models were stable, as indicated by the cumulative sum of recursive

**Table 2.** Linear ARDL results - short-run results

| Variable | Indonesia Coeff | Indonesia $p$-value | Thailand Coeff | Thailand $p$-value | Malaysia (1) Coeff | Malaysia (1) $p$-value | Malaysia (2) Coeff | Malaysia (2) $p$-value | Philippines (1) Coeff | Philippines (1) $p$-value | Philippines (2) Coeff | Philippines (2) $p$-value |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| D(lnTB) | 0.190 | 0.173 | | | 0.222 | 0.057 | | | | | | |
| D(lnTB) | | | | | −0.164 | 0.142 | | | | | | |
| D(lnY) | −0.526 | 0.469 | −2.925 | 0.000 | −1.224 | 0.020 | −1.184 | 0.094 | −1.527 | 0.089 | −1.589 | 0.053 |
| D(lnY) | | | | | 0.080 | 0.875 | −0.452 | 0.520 | 0.922 | 0.352 | 0.480 | 0.618 |
| D(lnY) | | | | | −1.023 | 0.039 | −0.603 | 0.323 | −2.774 | 0.002 | −2.589 | 0.005 |
| D(lnY) | | | | | 0.267 | 0.614 | | | | | | |
| D(lnY) | | | | | 0.872 | 0.064 | | | | | | |
| D(lnYROW) | 4.396 | 0.006 | 1.391 | 0.112 | 4.036 | 0.000 | 3.539 | 0.016 | 1.228 | 0.310 | 2.078 | 0.093 |
| D(lnYROW) | | | | | | | 0.258 | 0.881 | | | | |
| D(lnYROW) | | | | | | | −1.872 | 0.198 | | | | |
| D(lnRER) | −0.992 | 0.507 | −1.274 | 0.000 | 0.062 | 0.798 | −0.216 | 0.541 | −0.123 | 0.675 | −0.374 | 0.220 |
| D(lnRER) | | | | | −0.294 | 0.269 | 0.405 | 0.274 | −0.471 | 0.151 | | |
| D(lnRER) | | | | | −0.929 | 0.002 | 1.171 | 0.004 | | | | |
| D(lnRER) | | | | | −1.105 | 0.004 | | | | | | |
| ECT | −0.426 | 0.000 | −0.658 | 0.000 | −0.908 | 0.000 | −0.741 | 0.000 | −0.592 | 0.000 | −0.520 | 0.000 |

**Note(s):** ECT indicates the error-correction term, $p$-values are put in the parentheses. Formula-wise DlnTB is an equivalent representation of $\Delta \ln TB$ in text (likewise for other variables)
**Source(s):** Table by author

**Table 3.**
Linear ARDL results - long-run results

| Variable | Indonesia Coeff | Indonesia $p$-value | Thailand Coeff | Thailand $p$-value | Malaysia (1) Coeff | Malaysia (1) $p$-value | Malaysia (2) Coeff | Malaysia (2) $p$-value | Philippines (1) Coeff | Philippines (1) $p$-value | Philippines (2) Coeff | Philippines (2) $p$-value |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Constant | −12.806 | 0.215 | −7.663 | 0.011 | −6.382 | 0.145 | −0.612 | 0.914 | 33.219 | 0.000 | 40.711 | 0.000 |
| lnY | −2.060 | 0.038 | −1.479 | 0.000 | −0.924 | 0.002 | −0.914 | 0.030 | 2.896 | 0.000 | 3.013 | 0.001 |
| lnYROW | 4.048 | 0.018 | 1.665 | 0.000 | 1.027 | 0.052 | 0.835 | 0.221 | −4.654 | 0.000 | −4.669 | 0.000 |
| lnRER | −8.030 | 0.002 | −0.292 | 0.105 | 0.449 | 0.077 | −0.651 | 0.162 | 0.703 | 0.065 | −1.132 | 0.065 |
| DUM | 1996 | | | | | | 2001 | | | | | |
| Model | (2, 0, 1, 1) | | (1, 1, 0, 1) | | (3, 5, 1, 4) | | (1, 3, 3, 3) | | (1, 3, 1, 2) | | (1, 3, 1, 0) | |
| $R^2$ adj | 0.301 | | 0.390 | | 0.689 | | 0.375 | | 0.356 | | 0.392 | |
| $F$-test | 4.709 | | 5.588 | | 6.464 | | 3.788 | | 5.684 | | 6.212 | |
| J-B | 0.263 | 0.877 | 0.124 | 0.940 | 1.221 | 0.543 | 1.884 | 0.390 | 1.346 | 0.510 | 0.823 | 0.663 |
| LM ($F$) | 2.368 | 0.114 | 0.889 | 0.422 | 0.190 | 0.829 | 0.332 | 0.722 | 0.052 | 0.949 | 0.190 | 0.828 |
| LM | 5.734 | 0.057 | 2.210 | 0.331 | 0.873 | 0.646 | 1.245 | 0.537 | 0.166 | 0.921 | 0.545 | 0.762 |
| Heterosk | 5.739 | 0.677 | 11.565 | 0.072 | 13.830 | 0.611 | 15.114 | 0.370 | 6.107 | 0.806 | 4.153 | 0.843 |
| RESET | 0.129 | 0.898 | 0.643 | 0.525 | 0.379 | 0.710 | 1.586 | 0.129 | 0.068 | 0.947 | 0.213 | 0.833 |
| CUSUM | S | | S | | S | | S | | S | | S | |
| CUSUMSQ | S | | S | | S | | S | | S | | S | |

**Note(s):** $F$-test, J-B, LM ($F$), LM, Heterosk., RESET, CUSUM and CUSUMSQ represent respectively bounds $F$-test, Jarque-Bera normality test, $F$-test version of the Breusch-Godfrey test of serial correlation, Breusch-Godfrey LM test of serial correlation, Breusch-Pagan-Godfrey LM test of heteroskedasticity, Ramsey regression equation specification error test, cumulative sum and cumulative sum squared of recursive residuals tests of stability. DUM indicates time dummy variables and $p$-values are put in the parentheses. “S” indicates stability of the model
**Source(s):** Table by author

residuals (squared) tests (CUSUM and CUSUMSQ), however, in the case of Indonesia and in *Model 2* for Malaysia the introduction of dummy variables was required to achieve stability. The overall significance of the models was sufficient with adjusted $R^2$ ranging from 0.301 to 0.689. According to the bounds $F$-test, the null hypothesis of no cointegration was rejected at the 5% significance level in Indonesia and Thailand and at the 1% level in *Model 1* for Malaysia and in both models for the Philippines. In the *Model 2* for Malaysia, the $F$-test statistic fell within $I(0)$ and $I(1)$ bounds, however, with reference to the error-correction term (ECT), that is negative and highly significant, we conclude that there was cointegration amongst the variables in this case as well. In other economies the ECT was likewise negative and the speed of adjustment to the long-run equilibrium was substantially high (ranging from approximately 0.426 in Indonesia to as high as 0.908 in Malaysia's *Model 1*, i.e. up to 91% of disequilibrium was corrected in the following period).

In line with theoretical predictions, in Indonesia, Malaysia and Thailand, the long-run coefficient of domestic GDP was significant and negative, whilst the coefficient of the "rest of the world" GDP was positive. Domestic economic growth of the three countries was accompanied by the growth of agricultural imports, whilst economic growth overseas stimulated agricultural exports (manifested, for instance, in a number of vibrant export industries, such as rubber in Thailand, palm oil in Malaysia and cocoa and palm oil in Indonesia (David *et al.*, 2007, p. 4).

In the Philippines the signs of the effects were the opposite: positive effects of domestic GDP and the negative effects of the "rest of the world" GDP. In the post-1980 period the Philippines had the lowest GDP growth rate in South East Asia, whilst the deterioration of the agricultural trade balance (that was particularly drastic in the early 1980s) slowed down to the extent that the balance was rather stable since the mid-1990s. On the other hand, the trade liberalisation initiatives of the early 1980s were reversed a number of times and, since the end of 1980s, have been losing momentum, with import-substitution and protectionist practices and policies, including the most distorting ones, persistent in a large number of agricultural products (David *et al.*, 2007, pp. 8–9). The combined effect of the two developments was a positive effect of domestic GDP on the agricultural trade balance. The "rest of the world GDP" exercised a negative effect on Philippines agricultural trade balance. Whilst the upswings in international commodity prices were experienced during the 1980–2013 period (e.g. "commodity super cycle" of the 2000s), the Philippines' agricultural exports were sluggish. The problem was manifested in the slowdown of traditional commodities' exports (coconut, tobacco, sugar) and the insufficient growth in the high-value-added agricultural exports, whilst the sector as a whole was constrained by institutional and governance weaknesses, rent seeking and infrastructural and research bottlenecks (David, 2003).

The definition of the J-curve that relies on the comparison of the long- and short-run RER coefficients, indicates its presence when the long-run normalised coefficient of RER is positive but the short-run coefficient is negative (if USDA measure of RER is used) and conversely positive short-run and negative long-run coefficients (if IMF measure of RER is applied). In the case of Indonesia and Thailand both types of coefficients were negative. In Malaysia, the J-curve hypothesis was supported in the model with USDA measure of RER, as well as with IMF measure, albeit in the latter case the long-run coefficient, whilst having correct sign, was not significant. In the Philippines, the *Model 1* estimates suggest the presence of J-curve, given that the long-run coefficient is positive but short-run coefficient is negative (and not significant). The *Model 2* rejects the J-curve hypothesis in the Philippines, given that both types of coefficients are negative. The definition of the J-curve adopted by Rose and Yellen (1989) examines the change in the signs of RER short-run coefficients. In this paper, this version of the hypothesis could only be verified in Malaysia, where the ARDL model selected a sufficient number of RER first-difference terms. In *Model 1*, the first lag of the differenced RER was positive and insignificant, whilst further lags were negative and significant. In

*Model 2*, the change of the sign was the opposite (from negative and insignificant to positive and significant). Overall, there is certain evidence supporting the J-curve hypothesis in Malaysia, but not the other three economies.

In Indonesia and Thailand, depreciation did not improve agricultural trade balance in the long-run, whilst in Malaysia and the Philippines the improvement was witnessed in the models with USDA based RER, but the size of the trade balance elasticities was smaller than one. This empirical result is in line with studies that indicated limited effectiveness of exchange rate adjustments in fixing the external imbalances (Mundell, 1988).

Tables 4 and 5 present the short- and long-run findings from the nonlinear ARDL model. Similarly to the linear ARDL, the two alternative models were estimated for Malaysia and the Philippines (based on the USDA and IMF alternative RER indicators). Additionally, due to inconsistent signs of coefficients and specification issues, we tried for the Philippines a model with a different lag structure (the models are respectively called *Models 1, 2 and 3*). The maximum number of lags varied in each case, the AIC was used for selection purposes; however, no fixed lags were imposed in any of the models. The normality and heteroskedasticity tests were passed in all models (albeit in *Model 2* for the Philippines, the evidence of no heteroskedasticity was somewhat weaker than in other models). The Breusch-Pagan LM test did not reject the null hypothesis of no serial correlation in Indonesia, Thailand and the Philippines (*Models 2 and 3*), but indicated the presence of serial correlation in Malaysia (both models) and the Philippines (*Model 1*). However, the $F$-test version of the test (as well as the correlogram Q-statistic, not presented here to conserve space), unambiguously gave support to the null hypothesis. All models in question were correctly specified and were stable (in the case of Malaysia and Philippines, following the introduction of time dummies). Wald long-run asymmetry statistic was significant only in the models for Malaysia and the Philippines, hence for interpretation purposes the nonlinear ARDL results are more informative for these two economies. The overall significance of the models was adequate, with adjusted $R^2$ ranging from 0.236 to 0.630. The bounds $F$-test statistic ranged from 3.126 to 9.177, exceeding the $I(1)$ critical bound in all economies except Malaysia *Model 1* and the Philippines *Model 3*, where it fell within $I(0)$ and $I(1)$ bounds. We therefore conclude that the null hypothesis of no cointegration was rejected in all models (in the latter two cases based on the significance of the ECT). The speed of adjustment to equilibrium was sufficient to correct between 39.3 and 92.2% of disequilibrium in the period following the shock.

In the nonlinear ARDL, the J-curve hypothesis requires that in the model with USDA measure of RER the long-run normalised coefficient of $\ln RER^{-}$ is positive whilst the short-run coefficient $\Delta \ln RER^{-}$ is negative, and conversely, in the model with IMF measure of REER, the long- and short-run coefficients are negative and positive.

The estimates are largely in line with the ones from the linear ARDL. In Thailand and in *Model 1* for the Philippines, both $\ln RER^{-}$ and $\Delta \ln RER^{-}$ are negative, and the J-curve hypothesis is not supported. In Indonesia, the negative long-run coefficient is coupled with the short-run coefficient that becomes positive at the second and third lag, i.e. the pattern that is inverse to J-curve is observed. In Malaysia, the long- and short-run coefficients of $\ln RER^{-}$ are respectively positive and negative, and the J-curve is likely to be present (however, in *Model 2* we observe the variation in the short-run coefficient sign, that was initially positive, turning negative at the second and third lag and finally becoming positive at the fourth lag). The estimates of *Model 2* in the Philippines arguably indicate the J-curve as well (with the short-run $\Delta \ln RER^{-}$ becoming positive at lags two and three, whilst $\ln RER^{-}$ is negative albeit insignificant). With regard to *Model 3* in the Philippines, both types of coefficients were positive and insignificant, hence no J-curve. Overall, only in two cases (*Model 1* in Malaysia and possibly *Model 2* in the Philippines) the depreciation improves trade balance in the long-run following temporal deterioration in the short-run. In all other cases (where coefficients are

**Table 4.** Nonlinear ARDL results - short-run results

| Variable | Indonesia Coeff | Indonesia *p*-value | Thailand Coeff | Thailand *p*-value | Malaysia (1) Coeff | Malaysia (1) *p*-value | Malaysia (2) Coeff | Malaysia (2) *p*-value | Philippines (1) Coeff | Philippines (1) *p*-value | Philippines (2) Coeff | Philippines (2) *p*-value | Philippines (3) Coeff | Philippines (3) *p*-value |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| D(lnTB) | | | | | 0.367 | 0.006 | 0.193 | 0.011 | | | | | | |
| D(lnTB) | | | | | −0.087 | 0.491 | | | | | | | | |
| D(lnY) | −2.287 | 0.004 | −1.988 | 0.002 | −2.080 | 0.000 | −1.607 | 0.000 | −1.643 | 0.046 | −1.618 | 0.035 | −1.068 | 0.118 |
| D(lnY) | | | 1.143 | 0.016 | | | 1.329 | 0.003 | 0.212 | 0.820 | −0.652 | 0.476 | | |
| D(lnY) | | | | | | | 1.765 | 0.000 | −2.430 | 0.005 | −2.650 | 0.005 | | |
| D(lnYROW) | 3.710 | 0.033 | −0.273 | 0.832 | 4.799 | 0.000 | 4.684 | 0.000 | 1.434 | 0.382 | 3.560 | 0.061 | 5.047 | 0.003 |
| D(lnYROW) | | | | | | | 0.635 | 0.487 | | | | | | |
| D(lnYROW) | | | | | | | −1.282 | 0.176 | | | | | | |
| D(lnYROW) | | | | | | | 7.227 | 0.000 | | | | | | |
| D(lnRERPOS) | −2.347 | 0.237 | 0.118 | 0.733 | −0.579 | 0.042 | 2.721 | 0.000 | −0.139 | 0.740 | −1.739 | 0.041 | −3.186 | 0.000 |
| D(lnRERPOS) | | | | | | | −0.645 | 0.168 | −1.390 | 0.004 | | | | |
| D(lnRERPOS) | | | | | | | 3.217 | 0.000 | −1.057 | 0.020 | | | | |
| D(lnRERPOS) | | | | | | | −3.508 | 0.000 | | | | | | |
| D(lnRERNEG) | −1.655 | 0.480 | −3.691 | 0.000 | −0.995 | 0.068 | 0.049 | 0.809 | −0.022 | 0.972 | −0.090 | 0.842 | 0.495 | 0.260 |
| D(lnRERNEG) | 6.689 | 0.007 | | | | | −0.777 | 0.002 | | | 1.240 | 0.004 | | |
| D(lnRERNEG) | 5.729 | 0.013 | | | | | −0.283 | 0.208 | | | 0.828 | 0.031 | | |
| D(lnRERNEG) | | | | | | | 1.277 | 0.000 | | | | | | |
| ECT | −0.840 | 0.000 | −0.706 | 0.000 | −0.922 | 0.000 | −0.578 | 0.000 | −0.749 | 0.000 | −0.518 | 0.000 | −0.392 | 0.002 |

**Note(s):** As per Table 1
**Source(s):** Table by author

**Table 5.** Nonlinear ARDL results - long-run results

| Variable | Indonesia Coeff | Indonesia $p$-value | Thailand Coeff | Thailand $p$-value | Malaysia (1) Coeff | Malaysia (1) $p$-value | Malaysia (2) Coeff | Malaysia (2) $p$-value | Philippines (1) Coeff | Philippines (1) $p$-value | Philippines (2) Coeff | Philippines (2) $p$-value | Philippines (3) Coeff | Philippines (3) $p$-value |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Constant | −26.284 | 0.108 | 9.000 | 0.503 | −12.595 | 0.012 | 13.269 | 0.124 | 58.050 | 0.022 | 69.164 | 0.364 | −65.643 | 0.208 |
| lnY | −2.515 | 0.001 | −1.096 | 0.005 | −0.541 | 0.011 | −1.909 | 0.000 | 2.978 | 0.001 | 4.276 | 0.077 | 0.024 | 0.987 |
| lnYROW | 3.980 | 0.042 | −0.235 | 0.871 | 1.526 | 0.005 | −0.635 | 0.441 | −6.840 | 0.013 | −8.500 | 0.313 | 6.480 | 0.251 |
| lnRERPOS | −7.494 | 0.009 | 0.202 | 0.628 | −0.718 | 0.009 | 8.893 | 0.005 | 1.429 | 0.051 | −1.202 | 0.668 | −5.643 | 0.025 |
| lnRERNEG | −11.128 | 0.000 | −1.293 | 0.046 | 0.455 | 0.235 | −0.729 | 0.063 | −0.122 | 0.875 | −3.106 | 0.290 | 2.212 | 0.234 |
| DUM | | | | | 2003 | | 1995 | | 2012 | | 2013 | | 2013 | |
| Model | (1, 0, 0, 1, 3) | | (1, 2, 0, 0, 1) | | (3, 1, 1, 0, 1) | | (2, 3, 4, 4, 4) | | (1, 3, 1, 3, 0) | | (1, 3, 1, 1, 3) | | (1, 0, 0, 0, 1) | |
| $R^2$ adj | 0.532 | | 0.472 | | 0.546 | | 0.630 | | 0.471 | | 0.415 | | 0.236 | |
| $F$-test | 9.177 | | 4.640 | | 3.263 | | 5.486 | | 7.040 | | 5.311 | | 3.162 | |
| J-B | 1.090 | 0.580 | 0.078 | 0.962 | 0.456 | 0.796 | 0.278 | 0.870 | 2.533 | 0.282 | 0.941 | 0.625 | 2.225 | 0.329 |
| LM ($F$) | 1.039 | 0.371 | 0.542 | 0.588 | 1.767 | 0.195 | 1.301 | 0.324 | 0.481 | 0.626 | 1.452 | 0.262 | 0.327 | 0.724 |
| LM | 2.935 | 0.231 | 1.497 | 0.473 | 5.041 | 0.080 | 8.100 | 0.017 | 1.725 | 0.422 | 4.961 | 0.084 | 0.883 | 0.643 |
| Heterosk | 10.996 | 0.276 | 9.153 | 0.330 | 11.438 | 0.407 | 26.990 | 0.212 | 9.441 | 0.739 | 22.121 | 0.076 | 4.885 | 0.674 |
| RESET | 0.407 | 0.688 | 0.947 | 0.353 | 1.361 | 0.187 | 0.812 | 0.438 | 0.527 | 0.605 | 0.780 | 0.445 | 1.690 | 0.103 |
| Wald asymm | 0.122 | 0.727 | 2.336 | 0.126 | 12.366 | 0.000 | 4.356 | 0.037 | 6.073 | 0.014 | 0.007 | 0.935 | 5.068 | 0.024 |
| CUSUM | S | | S | | S | | S | | S | | S | | S | |
| CUSUMSQ | S | | S | | S | | S | | S | | S | | S | |

**Note(s):** As per Table 2
**Source(s):** Table by author

significant), the depreciation either exercises a negative effect on the trade balance all the way through, or results in the inversion of the J-curve.

As far as the comparative significance of regressors is concerned, in a total of 13 linear or nonlinear ARDL models, the GDP variable was significant in all but one case. The "rest of the world" GDP and exchange rate were insignificant in respectively five and six cases. Whilst the long-run asymmetry was confirmed in four models, the estimation of nonlinear ARDL did not deliver greater significance of the regressors (in fact, in the nonlinear models, "rest of the world" GDP and exchange rate were insignificant in four cases each).

## Conclusion and discussion

The empirical findings demonstrated negative effects of the domestic GDP and positive effects of the "rest of the world" GDP in all countries except the Philippines, where the signs of the effects were the opposite (attributed to the persistence of protectionist agricultural policies, difficulties in expanding agricultural exports and certain slowdown in the growth of agricultural trade deficit). The J-curve effect was only observed in Malaysia, but the size of the trade balance elasticity coefficient (in those cases where long-run positive effect of depreciation on the trade balance was observed) was smaller than unity. This result is in line with previous research of the J-curve in agricultural and primary products' trade: the study by Yazici (2006) of Turkish agricultural trade balance and the analyses by Baek *et al.* (2009) and Gong and Kinnucan (2015) of the J-curve effect in the US agricultural trade. The absence of the J-curve was also documented on numerous occasions in the studies of the nonagricultural trade.

The output variables generally exercised a more significant effect on the agricultural trade balance than the real (effective) exchange rate. This finding confirms certain previous studies (e.g. Batten and Belongia, 1986, in the US context and Chebbi and Olarreaga, 2019, in Tunisian context, amongst others) and contravenes those analyses that establish statistically significant effects of the exchange rate on exports and trade balance (e.g. Gardner, 1981; Tweeten, 1989; Baek and Koo, 2008; Gong and Kinnucan, 2015, in the US context). Policy-wise, this finding suggests that fiscal and monetary measures that affect output levels may be more efficient in correction of external imbalances than targeting or tinkering with exchange rates (Noland, 1989, p. 178).

The findings (the absence of J-curve in the majority of the economies in question) are noteworthy, given that J-curve is more likely to exist in agricultural commodities as opposed to industrial products trade (low export and import elasticities and slow adjustment of quantities to changes in relative prices, due to substantial lags in production, consumption and transaction and the payments that are made after the delivery) (Doroodian *et al.*, 1999, p. 687; Junz and Rhomberg, 1973).

The absence of the J-curve and violation of Marshall-Lerner condition may be attributed to several factors. From methodological point of view, the J-curve could be identified, if the bilateral trade is considered (e.g. J-curve in the trade between two countries and the curve absence in the trade of the country with the rest of the world, as is the case in this paper), or if reactions of exports and imports to exchange rate fluctuations are examined separately. The use of a partial equilibrium framework that does not consider a complete set of the relations in the system (e.g. induced effects of depreciation of domestic output, or an interaction between domestic agricultural and foreign exchange policies) could likewise distort the results. Also, the results could be sensitive to the degree of aggregation and composition of agricultural exports [as noted by Gong and Kinnucan (2015), the sensitivity of exports and imports of bulk versus high-value-added consumer commodities may differ] and to the degree of currency misalignment prior to depreciation (Lal and Lowinger, 2002, pp. 412–413).

The theoretical explanations of missing J-curve are as follows. Firstly, at the currency-contract stage, the requisite assumptions behind the J-curve effect may be violated. For instance, the domestic exports (i.e. exports from the respective Southeast Asian economy) may be denominated in foreign currency (e.g. not baths or ringgits), whilst the imports are denominated in the foreign currency (US dollars), hence unexpected signs of the short-run coefficients of REER or RER (Baek *et al.*, 2009, p. 222). This explanation may nonetheless be implausible, if such effect is mitigated by the existence of futures markets for the exported commodities (e.g. palm oil futures in Bursa Malaysia, sugar futures on various commodity exchanges etc.). Secondly, the incomplete pass-through may be observed, where prices of exports in foreign currency rise proportionately to devaluation of domestic currency, whilst prices of exports in domestic currency remain the same, thus improving trade balance at the pass-through stage instead of deteriorating it (Baek *et al.*, 2009, p. 217). In a related vein, imperfectly competitive markets may cause firms to engage in oligopolistic behaviour (altering profit margins to compensate for exchange rate changes). Whilst such behaviour is more typical for production and exports of differentiated manufactured products, the extent of oligopolistic behaviour in international agricultural markets may need to be investigated (Noland, 1989, p. 177).

Thirdly, other factors that affect J-curve and Marshall-Lerner condition include agricultural trade restrictions and protectionism (particularly prior to World Trade Organisation (WTO) Uruguay Round) that potentially create downward bias in price elasticities; the trading decisions that are based on the expected future exchange rates rather than actual ones, resulting in low elasticities contra Marshall-Lerner condition; and the relative size of agricultural exporting economy. The latter factor implies that “small sellers” face perfectly elastic export demand in international markets. Thus, depreciation of domestic currency will lead to improvement of the domestic trade balance (since export values rise faster than import values). Arguably, this is not the case of the economies in this paper: in the specific export commodities none of them are small exporters, hence trade balance may not improve following depreciation.

The absence of the J-curve effect may indicate the importance of other mechanisms of adjustments to currency depreciation, e.g. the industries that do not benefit from depreciation may have to adjust their profit margins in order to maintain their market share (Bahmani-Oskooee and Aftab, 2017a). The realisation of the benefits from depreciation also requires high price elasticity of the exported goods. This in turn necessitates that countries in question continue to expand their export base by moving towards the production of manufactured goods or higher-value-added agricultural products with higher price elasticity. Additionally, the improvements in infrastructure would facilitate faster response in the production and supply to the favourable opportunities arising from currency depreciation (Begum and Alhelal, 2016). In terms of macroeconomic policy, the weak or absent link between exchange rate and trade balance (e.g. due to frequent intervention in the foreign exchange market as was the case of Malaysia in the 1990s) may imply that correction of trade balance and balance of payments may need to include prior changes to the level of money supply or to the aggregate income (Duasa, 2007).

The future empirical study of the agricultural trade J-curve effect may need to differentiate the behaviour of the trade balance in the economies with flexible or fixed exchange rate systems. Furthermore, in order to get a more complete understanding of the agricultural trade balance dynamics, an analysis at disaggregated level may be required (i.e. consideration of exchange rate effects on individual agricultural commodities or groups of commodities). As noted by Bahmani-Oskooee and Aftab (2018), the disaggregation tends to increase the number of significant effects and additionally allows distinction between the commodities that benefit from a currency depreciation, those that do not and those that are not affected by depreciation. In addition, the future research may benefit from incorporating policy factors in the analysis. The international trade in agricultural commodities has historically been regulated more extensively and profoundly than trade in manufacturing goods and has been subject to various nontariff barriers (Chebbi and Olarreaga, 2019) [2]. Lastly, the future research may reformulate the model on a bilateral or regional basis, i.e. consider the effects of exchange rate changes on the agricultural trade (aggregate or disaggregated) in a pair or group of economies.

## Notes

1. The ARDL class models serve this purpose well, given the flexible lag structure across dependent and independent variables. This feature is advantageous, compared to VAR and VECM models that impose uniform lags for all variables.
2. Such analytical extension may include trade barrier measures as regressor in the model, or may utilise some form of trade policy indicator (which itself is a result of operation of domestic and international political-economic forces).

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**Corresponding author**
Ivan D. Trofimov can be contacted at: ivan.trofimov1@gmail.com

## Appendix

| Variable | Measurement | Description/source |
|---|---|---|
| Export and import values | Thousands of $US | Based on export (import) quantities multiplied by per unit export (import) values. Obtained from FAO "Crops and Livestock Products" dataset, sub-heading "Items Aggregated: Total Agricultural Products". Accessed at http://www.fao.org/faostat/en/#data/TP |
| Real exchange rate | Local currency per $US | Obtained from USDA "Real Annual Country Average Exchange Rates, Local Currency per USD" dataset. Accessed at https://www.ers.usda.gov/data-products/agricultural-exchange-rate-data-set/ |
| Real effective exchange rate | Units | Obtained from IMF *International Financial Statistics* "Real Effective Exchange Rate Based on the CPI" dataset. CPI index has base in 2010. Accessed at https://data.imf.org/regular.aspx?key=61545862 |
| GDP | Billions of 2010 $US | Obtained from USDA "Historical Real Gross Domestic Product (GDP) and Growth Rates of GDP for Baseline Countries/Regions" dataset. Accessed at https://www.ers.usda.gov/data-products/international-macroeconomic-data-set/ |

**Source(s):** Table by author

**Table A1.**
Description of the variables

**Table A2.** Descriptive statistics

Indonesia

| LGDP | | LGDPROW | | LREER | | LTB | |
|---|---|---|---|---|---|---|---|
| Mean | 6.105 | Mean | 10.754 | Mean | 9.128 | Mean | 0.560 |
| Median | 6.135 | Median | 10.750 | Median | 9.110 | Median | 0.599 |
| Maximum | 6.994 | Maximum | 11.271 | Maximum | 9.988 | Maximum | 1.008 |
| Minimum | 5.201 | Minimum | 10.233 | Minimum | 8.349 | Minimum | 0.049 |
| Std. Dev | 0.522 | Std. Dev | 0.319 | Std. Dev | 0.359 | Std. Dev | 0.295 |
| Skewness | −0.047 | Skewness | −0.038 | Skewness | −0.277 | Skewness | −0.289 |
| Kurtosis | 1.961 | Kurtosis | 1.770 | Kurtosis | 3.424 | Kurtosis | 1.910 |
| J-B | 1.723 | J-B | 2.404 | J-B | 0.772 | J-B | 2.409 |
| J-B prob | 0.423 | J-B prob | 0.301 | J-B prob | 0.680 | J-B prob | 0.300 |

Malaysia

| LGDP | | LGDPROW | | LREER | | LREERIMF | | LTB | |
|---|---|---|---|---|---|---|---|---|---|
| Mean | 4.920 | Mean | 10.761 | Mean | 1.102 | Mean | 4.740 | Mean | 0.668 |
| Median | 5.014 | Median | 10.756 | Median | 1.125 | Median | 4.639 | Median | 0.646 |
| Maximum | 5.899 | Maximum | 11.281 | Maximum | 1.397 | Maximum | 5.214 | Maximum | 1.066 |
| Minimum | 3.824 | Minimum | 10.238 | Minimum | 0.687 | Minimum | 4.444 | Minimum | 0.316 |
| Std. Dev | 0.638 | Std. Dev | 0.320 | Std. Dev | 0.212 | Std. Dev | 0.217 | Std. Dev | 0.205 |
| Skewness | −0.217 | Skewness | −0.036 | Skewness | −0.298 | Skewness | 0.770 | Skewness | 0.220 |
| Kurtosis | 1.764 | Kurtosis | 1.774 | Kurtosis | 1.857 | Kurtosis | 2.470 | Kurtosis | 2.333 |
| J-B | 2.718 | J-B | 2.389 | J-B | 2.631 | J-B | 4.202 | J-B | 1.012 |
| J-B prob | 0.257 | J-B prob | 0.303 | J-B prob | 0.268 | J-B prob | 0.122 | J-B prob | 0.603 |

Philippines

| LGDP | | LGDPROW | | LREER | | LREERIMF | | LTB | |
|---|---|---|---|---|---|---|---|---|---|
| Mean | 4.866 | Mean | 10.761 | Mean | 3.872 | Mean | 4.605 | Mean | −0.129 |
| Median | 4.776 | Median | 10.757 | Median | 3.881 | Median | 4.606 | Median | −0.340 |
| Maximum | 5.715 | Maximum | 11.281 | Maximum | 4.186 | Maximum | 4.885 | Maximum | 1.137 |
| Minimum | 4.318 | Minimum | 10.237 | Minimum | 3.628 | Minimum | 4.331 | Minimum | −0.810 |

(*continued*)

**Table A2.**

Philippines

| LGDP | | LGDPROW | | LREER | | LREERIMF | | LTB | |
|---|---|---|---|---|---|---|---|---|---|
| Std. Dev | 0.416 | Std. Dev | 0.321 | Std. Dev | 0.154 | Std. Dev | 0.130 | Std. Dev | 0.542 |
| Skewness | 0.498 | Skewness | −0.039 | Skewness | 0.350 | Skewness | 0.040 | Skewness | 0.945 |
| Kurtosis | 2.018 | Kurtosis | 1.774 | Kurtosis | 2.308 | Kurtosis | 2.611 | Kurtosis | 2.656 |
| J-B | 3.101 | J-B | 2.389 | J-B | 1.533 | J-B | 0.249 | J-B | 5.844 |
| J-B prob | 0.212 | J-B prob | 0.303 | J-B prob | 0.465 | J-B prob | 0.883 | J-B prob | 0.054 |

Thailand

| LGDP | | LGDPROW | | LREER | | LTB | |
|---|---|---|---|---|---|---|---|
| Mean | 5.278 | Mean | 10.760 | Mean | 3.489 | Mean | 1.271 |
| Median | 5.379 | Median | 10.755 | Median | 3.453 | Median | 1.141 |
| Maximum | 6.047 | Maximum | 11.280 | Maximum | 3.826 | Maximum | 2.046 |
| Minimum | 4.197 | Minimum | 10.237 | Minimum | 3.174 | Minimum | 0.901 |
| Std. Dev | 0.568 | Std. Dev | 0.320 | Std. Dev | 0.165 | Std. Dev | 0.313 |
| Skewness | −0.501 | Skewness | −0.034 | Skewness | 0.362 | Skewness | 0.993 |
| Kurtosis | 2.024 | Kurtosis | 1.773 | Kurtosis | 2.448 | Kurtosis | 2.689 |
| J-B | 3.100 | J-B | 2.392 | J-B | 1.312 | J-B | 6.392 |
| J-B prob | 0.212 | J-B prob | 0.302 | J-B prob | 0.519 | J-B prob | 0.041 |

**Source(s):** Table by author