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THE U.S. TRADE BALANCE WITH PARTNERS FROM DEVELOPING WORLD: AN ASYMMETRY ANALYSIS OF THE J-CURVE EFFECT

MOHSEN BAHMANI-OSKOOEE AND HANAFIAH HARVEY*

University of Wisconsin-Milwaukee, USA
Pennsylvania State University-Mont Alto, USA

We consider the bilateral trade balance of the U.S. with each of her 13 trading partners from the developing world. When we apply the linear ARDL approach of Pesaran et al. (2001), we find support for the J-curve effect with six partners. However, when we apply Shin et al.'s (2014) nonlinear ARDL approach to asymmetry analysis, we find support for the J-curve in the U.S. trade with 10 partners. Additionally, while we find support for the short-run asymmetric effects of exchange rate changes in almost all cases, the short-run effects translate into the long-run significant asymmetric effects in half of the cases.

Keywords: Trade Balance, Exchange Rate, Asymmetry Effects, The U.S. 13 Developing Partners
JEL Classification: F31

1. INTRODUCTION

Although three countries, China, Canada, and Mexico, engage in almost 45% of the U.S. trade, other countries are still important on the ground that their trade with the U.S. adds up to a significant amount. The theoretical literature in international finance continues to emphasize the importance of relative prices, or the real exchange rate, as a main determinant of the trade balance between countries, yet the empirical literature has not been as successful as the theoretical literature. Theoretical developments point at adjustment lags which are said to slow down the response of trade flows to exchange rate changes. So slow that from a position of trade deficit, a currency depreciation may not be effective and the deficit may continue to worsen until adjustment lags are realized, hence the J-curve pattern (Magee, 1973).

Many studies have tried to test the J-curve effect by using trade flows of each country with the rest of the world. Examples include Bahmani-Oskooee (1985, 1989),

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Himarios (1985), Meade (1988), Moffet (1989), Mahdavi and Sohrabian (1993), Felmingham (1988), Brissimis and Leventankis (1989), Noland (1989), Bahmani-Oskooee and Malixi (1992), Bahmani-Oskooee and Alse (1994), Lal and Lowinger (2002), Hacker and Hatemi-J (2003), Akbostabci (2004), De Silva and Zhu (2004), Moura and Da Silva (2005), Duasa (2007), Bahmani-Oskooee and Kutan (2009), and Bahmani-Oskooee and Gelan (2012), The findings are poor and very mixed.[1]

The above studies were criticized by Rose and Yellen (1989) on the ground that they used aggregate trade flows of one country with the rest of the world and the real effective exchange rate. In an effort to discover stronger evidence in support of the J-curve effect, Rose and Yellen (1989) recommended using trade flows at the bilateral level and the real bilateral exchange rate. To that end, they concentrated on the bilateral trade balance between the U.S. and each of her six major partners from the industrial world. After applying Engle and Granger (1987) error-correction and cointegration methods and after defining the J-curve to mean short-run deterioration combined with long-run improvement, they were unsuccessful in establishing any significant short-run or long-run link between the real bilateral exchange rate and the U.S. bilateral trade balances.

Recently the Rose and Yellen (1989) approach was criticized by Bahmani-Oskooee and Fariditavana (2016) on the ground that failure to find a significant relation between the bilateral trade balance and the real bilateral exchange rate could be due to assuming the effects of exchange rate changes to be symmetric. As they argued, traders and market participants could have different expectations and different response to currency depreciations compared to appreciations. If so, the trade balance could respond in an asymmetric manner to changes in the exchange rate. Their conjecture is now even more supported due to evidence of asymmetric response of import and export prices to changes in the exchange rate (Bussiere, 2013). Since asymmetry analysis requires applying nonlinear models, Bahmani-Oskooee and Fariditavana (2016) applied Shin et al.'s (2014) recent nonlinear asymmetry cointegration approach to bilateral trade balance models of the U.S. with Canada, France, Germany, Italy, Japan, and the U.K. They not only found evidence of asymmetric effects of exchange rate changes on the trade balance but also evidence of a significant short-run and long-run link between currency depreciation or appreciation and the trade balance. Overall, more evidence was found from the nonlinear models as compared to the linear models.

Our goal in this paper is to build upon asymmetry analysis of Bahmani-Oskooee and Fariditavana (2016) by considering bilateral trade balances of the U.S. with her partners from the developing world. The partners we consider are: Argentina, Chile, Ecuador, Indonesia, Israel, India, Korea, Malaysia, Mexico, the Philippines, Singapore, South Africa, and Thailand. They all together engage in almost 30% of the U.S. trade. To that end, we outline the models and methods in Section 2 and present the empirical results in

[1] For the most recent review articles see Bahmani-Oskooee and Ratha (2004) and Bahmani-Oskooee and Hegerty (2010).

Section 3. A summary is provided in Section 4 and data definition and sources are provided in an Appendix.

2. THE MODELS AND METHODS[2]

It is now a well-established fact that in any time-series model, not only we must be concerned with integrating properties of the variables involved, but we also must be concerned with cointegration among those variables. Our approach is no exception and following Bahmani-Oskooee and Fariditavana (2016) we begin with the following long-run bilateral trade balance model of the U.S. with partner i:[3]

$$LnTB_{i,t} = a + bLnY_{US,t} + cLnY_{i,t} + dLnREX_{i,t} + \varepsilon_t, \qquad (1)$$

where TB_i is a measure of the trade balance between the U.S. and partner i. For reasons to be cleared up and explained in great detail later, we define the bilateral trade balance as the ratio of the U.S. imports from partner i over her exports to partner i. As can be seen, three variables are identified to determine the bilateral trade balance. They are the U.S. income (Y_{US}), income of partner i (Y_i), and the real bilateral exchange rate (REX_i). Based on these definitions, we expect an estimate of b to be positive and that of c to be negative.[4] From the Appendix we gather that the REX_i is defined in a way that a decline reflects a depreciation of the U.S. dollar against partner i's currency. Therefore, if dollar depreciation is supposed to to improve the U.S. trade balance with partner i, an estimate of c should be positive. A negative estimate will be an indication of inelastic import demand of the U.S. or partner i.

Specification Eq. (1) lacks the dynamic adjustment process which is needed to assess the short-run effects of exogenous variables. We introduce the dynamic adjustment process by re-writing Eq. (1) as an error-correction model as follows:

$$\Delta LnTB_{i,t} = \alpha + \sum_{j=1}^{n} \beta_{t-j} \Delta LnTB_{i,t-j} + \sum_{j=0}^{n} \delta_{t-j} \Delta LnY_{US,t-j} + \sum_{j=0}^{n} \gamma_{t-j} \Delta LnY_{i,t-j}$$
$$+ \sum_{j=0}^{n} \pi_{t-j} \Delta LnREX_{i,t-j} + \lambda_1 LnTB_{i,t-1} + \lambda_2 LnY_{US,t-1} + \lambda_3 LnY_{i,t-1}$$
$$+ \lambda_4 LnREX_{i,t-1} + \mu_t. \qquad (2)$$

Specification Eq. (2) is due to Pesaran et al. (2001) who suggest applying the F test to establish joint significance of lagged level variables as a sign of cointegration. Since the new critical values that they produce account for integrating properties of variables,

[2] This section closely follows Bahmani-Oskooee and Fariditavana (2016).

[3] For a theoretical derivation of this model see Rose and Yellen (1989).

[4] Of course an estimate of b could be negative and that of c could be positive if increase in income is due to an increase in production of import-substitute goods (Bahmani-Oskooee, 1986).

there is no need for pre-unit-root testing and variables could be a combination of I(0) and I(1).[5] Once cointegration is established, long-run effects of all variables are judged by the estimates of $\lambda_2 - \lambda_4$ normalized on λ_1. Short-run effects are reflected by the estimates of coefficients attached to first-differenced variables. The J-curve effect will be established if estimates of π_{t-j} are insignificant or significant and negative but the normalized estimate of λ_4 is significant and positive.

In Eq. (2) it is assumed that exchange rate changes do have a symmetric effect on the bilateral trade balance. In order to introduce asymmetry analysis, we follow Shin et al. (2014) and decompose the real bilateral exchange rate into two new time-series variables, one representing only dollar appreciation and one representing only dollar depreciation. To that end, we form $\Delta LnREX$ which included positive changes (dollar appreciations) and negative changes (dollar depreciations). The two new time-series variables are then generated using the partial sum concept. More precisely, we denote the partial sum of positive changes by POS, a variable that only reflects dollar appreciation and partial sum of negative changes by NEG, a variable that only reflects dollar depreciations. We then replace $LnREX$ in Eq. (2) by POS and NEG variables to arrive at:[6]

$$\Delta LnTB_{i,t} = a' + \sum_{j=1}^{n1} b_j' \Delta LnTB_{i,t-j} + \sum_{j=0}^{n2} c_j' \Delta LnY_{US,t-j} + \sum_{j=0}^{n3} d_j' \Delta LnY_{i,t-j} + \sum_{j=0}^{n4} e_j' \Delta POS_{t-j} + \sum_{j=0}^{n5} f_j' \Delta NEG_{t-j} + \theta_0 LnTB_{i,t-1} + \theta_1 LnY_{US,t-1} + \theta_2 LnY_{i,t-1} + \theta_3 POS_{t-1} + \theta_4 NEG_{t-1} + \xi_t. \quad (3)$$

Shin et al. (2014) label Eq. (3) as a nonlinear ARDL model due to nature of constructing the POS and NEG variables whereas, models like Eq. (2) are labeled as linear ARDL model. They show that Pesaran et al.'s (2001) approach could also be applied to Eq. (3) where POS and NEG should be treated as one variable so that the critical value of the F test stays the same when we move from Eq. (2) to Eq. (3).[7]

Once Eq. (3) is estimated a few assumptions with regards to asymmetric effects of exchange rate changes on the trade balance could be assessed. First, if the number of lags that the ΔPOS variable takes is different than the number of lags that ΔNEG variable takes, then that will be a sign of adjustment asymmetry. This is expected in most cases, because while exports originate in the U.S., imports originate in partner i. Clearly, they are subject to different adjustment lags. Second, by inspection if sign or size of coefficient estimates attached to ΔPOS variable are different than the size or sign of estimates attached to ΔNEG variable, short-run asymmetric effects will be confirmed. However, if $\Sigma \hat{e}_j' \neq \Sigma \hat{f}_j'$, that will be a sign of short-run cumulative or impact

[5] The main assumption is that macro variables are either I(0) or I(1).

[6] For exact formulation of POS and NEG variables refer to Bahmani-Oskooee and Fariditavana (2016). Partial sum of positive (negative) values of a variable at a given time t is the same as cumulative sum up to time t where negative (positive) values are replaced by zeroes.

[7] This is due to dependency between POS and NEG variables.

asymmetry. Finally, long-run asymmetric effects will be established if $-\hat{\theta}_3/\hat{\theta}_0 \neq -\hat{\theta}_4/\hat{\theta}_0$. The Wald test will be used to test these two inequalities.[8]

3. EMPIRICAL RESULTS

As mentioned, we estimate both the linear ARDL model Eq. (2) and the nonlinear ARDL model Eq. (3) using bilateral data between the U.S. and each of her trading partner from the developing world. The list includes Argentina, Chile, Ecuador, Indonesia, Israel, India, Korea, Malaysia, Mexico, the Philippines, Singapore, South Africa, and Thailand. Generally, quarterly data over the period 1993I-2015IV are used to carry out the empirical exercise and exceptions are noted in the Appendix. A maximum of eight lags are imposed on each first-differenced variable in both models and Akaike's Information criterion (AIC) is used to select an optimum model. Results from each model are reported in Table 1. Furthermore, since different statistics have different critical values, we have collected them in the notes of Table, and used them to identify significance by * (**) at the 10% (5%) significance level, respectively. Furthermore, due to the volume of the results, estimates are reported in three panels. While short-run coefficient estimates are reported in Panel A, normalized long-run estimates are reported in Panel B. All diagnostic statistics are reported in Panel C.

From Panel A and the results from the linear model (L-ARDL column) we gather that exchange rate changes have short-run effects on all bilateral trade balance models except in the models that belongs to U.S.-Ecuador and U.S.-Philippines. However, when we consider the results from the nonlinear model (NL-ARDL column), either ΔPOS or ΔNEG variable has significant short-run effects in all models. Thus, separating appreciations from depreciations and introducing nonlinear adjustment of the real exchange rate yields relatively more short-run effects.[9] Do these short-run effects translate into the long run?

From Panel B we gather that the real bilateral exchange rate carries a positive and significant coefficient in the linear models between the U.S. and Argentina, Chile, Indonesia, Israel, Malaysia and Thailand, supporting Rose and Yellen's (1989) definition of the J-curve.

[8] For some other application of partial sum concept see Apergis and Miller (2006) on the effects of U.S. stock market on consumption, Verheyen (2013) on interest rate pass-through mechanism to deposit rates, and Nusair (2016) on a similar topic as ours. For the application of the linear model see De Vita and Kyaw (2008), Halicioglu (2007, 2008), Tayebi and Yazdani, (2014), Hajilee et al. (2014), and Durmaz (2015).

[9] The U.S. income and partner's income also have short-run effects in almost all models.

Table 1a. Full-Information Estimates of Both Linear ARDL (L-ARDL) and Nonlinear ARDL (NL-ARDL) Models

	i=Argentina		i=Chile		i=Ecuador		i=India	
	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL
Panel A: Short–Run Estimates								
$\Delta lnY_{US,t}$	0.34(0.59)	-1.32(0.45)	5.0(3.07)**	-0.58(0.12)	1.12(2.1)**	4.21(0.74)	-11.7(2.24)**	-8.23(1.75)
$\Delta lnY_{US,t-1}$		-4.45(1.61)		-3.60(0.96)		3.71(0.70)	13.97(2.75)**	11.93(2.68)**
$\Delta lnY_{US,t-2}$		-6.14(2.16)**		7.62(2.09)**		12.34(2.48)**		9.01(2.36)**
$\Delta lnY_{US,t-3}$						21.41(3.63)**		
$\Delta lnY_{i,t}$	-2.95(7.22)**	-2.69(4.55)**	0.19(0.24)	0.19(3.05)**	0.11(1.23)	0.56(0.85)	-0.24(0.72)	-0.26(1.13)
$\Delta lnY_{i,t-1}$	-1.95(3.93)**	-1.19(1.85)*	2.11(2.41)**	1.48(1.62)		-1.29(1.94)*	-0.34(0.98)	
$\Delta lnY_{i,t-2}$	-0.96(2.09)**	-0.90(1.63)	1.17(1.30)	3.13(3.41)**		-0.88(1.39)	-0.92(2.78)**	
$\Delta lnY_{i,t-3}$	1.36(3.29)**	-1.26(2.39)**	2.62(2.98)**	2.61(2.74)**			-0.46(1.76)*	
$\Delta lnY_{i,t-4}$		-0.19(0.38)	1.86(2.07)**	2.01(2.24)**				
$\Delta lnY_{i,t-5}$		-0.84(1.88)*	1.56(1.84)*	2.46(2.76)**				
$\Delta lnY_{i,t-6}$			2.29(2.56)**					
$\Delta lnREX_{i,t}$	0.92(2.88)**		0.83(3.84)**		0.28(1.31)		-0.73(0.88)	
$\Delta lnREX_{i,t-1}$	-1.85(3.59)**						2.01(2.61)**	
$\Delta lnREX_{i,t-2}$	-1.18(2.03)**						0.15(0.16)	
$\Delta lnREX_{i,t-3}$	-1.64(3.05)**						1.44(1.62)	
$\Delta lnREX_{i,t-4}$	-0.95(2.09)**							
$\Delta lnREX_{i,t-5}$	-0.33(0.85)							
$\Delta lnREX_{i,t-6}$	-0.60(1.69)*							
$\Delta lnREX_{i,t-7}$	-0.60(1.89)*							
ΔPOS_t		1.97(2.28)**		0.04(0.03)		6.21(2.49)**		0.04(0.01)
ΔPOS_{t-1}		-4.59(3.35)**		-2.84(1.73)*		3.92(1.24)		
ΔPOS_{t-2}		-3.75(2.27)**				7.50(3.00)**		
ΔPOS_{t-3}		-5.57(3.24)**				2.81(1.39)		
ΔPOS_{t-4}		-3.57(2.46)**						
ΔPOS_{t-5}		-2.95(2.12)**						
ΔPOS_{t-6}		-2.59(2.58)**						
ΔPOS_{t-7}								
ΔNEG_t		2.08(0.62)		2.29(2.36)**		0.23(0.06)		2.49(0.69)
ΔNEG_{t-1}		-11.6(2.98)**				-2.93(1.05)		7.72(2.75)**
ΔNEG_{t-2}						-7.40(3.22)**		
Panel B: Long–Run Estimates								
lnY_{US}	0.48(0.60)	-0.24(0.25)	6.67(6.65)**	6.28(7.26)**	2.91(2.68)**	-5.46(1.51)	-29.03(0.55)	-14.59(1.32)
lnY_i	-0.81(2.23)**	-0.57(2.21)**	-2.48(9.29)**	-2.65(2.92)**	0.28(1.34)	1.22(2.76)**	2.38(0.55)	-0.81(0.96)
$lnREX_i$	1.38(2.84)**		1.11(5.61)**		0.72(1.41)		-1.51(0.34)	
POS		4.99(3.89)**		2.94(3.81)**		-1.34(0.30)		16.16(1.85)*
NEG		5.53(3.85)**		2.52(3.22)**		-15.02(1.78)*		0.47(0.17)
Constant	5.94(0.35)	22.71(0.92)	-133.2(5.8)**	-122.9(3.8)**	-101.0(2.5)**	123.01(1.27)	815.55(0.55)	451.94(1.34)
Panel C: Diagnostic Statistics								
F	9.32**	9.37**	7.58**	5.68**	2.84	6.72**	4.33*	6.90**
ECM_{t-1}	-0.71(5.81)**	-0.76(5.82)**	-0.75(4.44)**	-0.91(4.52)*	-0.38(3.56)*	-0.44(2.72)	-0.15(0.66)	-0.31(1.75)
LM	5.92	3.82	8.79*	5.91	6.96	10.04**	9.46*	13.00**
RESET	2.99*	3.71*	4.09**	1.50	0.23	2.29	1.73	0.06
AdjustedR2	0.49	0.48	0.69	0.74	-0.04	0.02	0.33	0.42
CS(CS2)	S(UNS)	S(S)	S(S)	S(S)	S(S)	S(S)	S(S)	S(S)
WALD-S		3.61*		2.61		5.33**		1.65
WALD-L		1.37		0.38		4.82**		1.27

Notes: See notes at the end.

Table 1b. Full-Information Estimates of Both Linear ARDL (L-ARDL) and Nonlinear ARDL (NL-ARDL) Models

	i=Indonesia		i=Israel		i=South Korea		i=Mexico	
	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL
Panel A: Short–Run Estimates								
$\Delta lnY_{US,t}$	0.50(0.17)	3.87(1.24)	4.28(2.22)**	4.17(2.09)**	-1.41(0.75)	0.04(0.02)		1.42(1.67)
$\Delta lnY_{US,t-1}$	-0.67(0.22)	-1.95(0.62)	-0.93(0.46)	-0.89(0.44)	0.76(0.43)	-0.22(0.13)		-1.16(1.11)
$\Delta lnY_{US,t-2}$	-2.11(0.68)	1.02(0.32)	-5.92(2.98)**	-5.85(2.89)**	3.71(2.02)**	3.36(1.83)*		-3.18(2.90)**
$\Delta lnY_{US,t-3}$	3.92(1.29)	2.50(0.81)	-2.89(1.40)	-2.84(1.35)	0.09(0.50)	-0.67(0.36)		-3.10(2.83)**
$\Delta lnY_{US,t-4}$	6.90(2.28)**	8.38(2.69)**			2.94(1.60)	3.72(2.12)**		-2.85(2.73)**
$\Delta lnY_{US,t-5}$	1.02(0.33)				3.52(1.97)**	2.96(1.66)		-3.79(3.79)**
$\Delta lnY_{US,t-6}$	-9.46(3.12)**					1.75(0.99)		-2.13(2.18)**
$\Delta lnY_{US,t-7}$						2.44(1.37)		
$\Delta lnY_{i,t}$	-0.74(3.04)**	-1.02(1.63)	-0.02(0.21)	-0.01(0.18)	-0.07(0.09)	-1.77(2.04)**		-0.01(0.02)
$\Delta lnY_{i,t-1}$		1.59(2.55)**	-0.11(1.28)	-0.11(1.29)	-2.13(3.10)**	-1.04(1.15)		1.35(2.31)**
$\Delta lnY_{i,t-2}$		0.17(0.25)	0.05(0.58)	0.04(0.53)	-1.95(2.93)**	-2.21(2.78)**		1.85(3.26)**
$\Delta lnY_{i,t-3}$		0.69(1.15)	-0.19(2.28)**	-0.19(2.27)**	0.58(0.89)	-0.59(0.76)		1.92(3.52)**
$\Delta lnY_{i,t-4}$		2.35(3.77)**	-0.09(1.15)	-0.09(1.16)	-1.52(2.25)**	1.56(1.89)*		1.72(3.24)**
$\Delta lnY_{i,t-5}$		-0.52(0.86)	-0.13(1.54)	-0.13(1.54)	-0.06(0.09)	-1.51(2.06)**		1.43(3.10)**
$\Delta lnY_{i,t-6}$		1.37(2.02)**			-0.25(0.43)	-0.81(1.37)		1.23(3.31)**
$\Delta lnY_{i,t-7}$					-2.33(4.00)**	-2.47(3.97)**		0.60(2.03)**
$\Delta lnREX_{i,t}$	0.55(5.58)**		0.45(3.03)**		0.97(4.83)**		0.23(3.16)**	
ΔPOS_t		1.55(6.52)**		1.17(1.82)*		0.82(1.28)		0.29(1.06)
ΔPOS_{t-1}						-1.24(1.49)		0.45(1.15)
ΔPOS_{t-2}						-2.06(2.29)**		0.29(0.75)
ΔPOS_{t-3}						-2.24(2.83)**		0.84(2.35)**
ΔPOS_{t-4}						-0.11(0.12)		0.58(1.67)
ΔPOS_{t-5}						-1.79(2.02)**		-0.35(0.91)
ΔPOS_{t-6}								0.65(1.99)*
ΔPOS_{t-7}								0.54(1.95)*
ΔNEG_t		0.76(0.84)		1.08(2.83)**		3.09(2.80)**		-0.95(1.60)
ΔNEG_{t-1}		-1.77(2.01)**				-2.84(2.45)**		0.59(0.91)
ΔNEG_{t-2}		-1.40(1.55)				1.77(1.62)		1.10(1.76)*
ΔNEG_{t-3}		0.77(0.87)				0.19(0.18)		0.24(0.38)
ΔNEG_{t-4}		-2.68(3.24)**				1.17(1.10)		0.55(0.76)
ΔNEG_{t-5}						3.94(3.67)**		0.96(1.25)
ΔNEG_{t-6}						1.73(1.64)		
ΔNEG_{t-7}						2.77(2.83)**		
Panel B: Long–Run Estimates								
lnY_{US}	3.02(5.34)**	2.56(4.32)**	0.28(1.00)	0.12(0.15)	-13.62(0.27)	3.49(4.51)**	2.66(5.02)**	2.05(8.66)**
lnY_i	-1.01(5.55)**	-1.09(9.30)**	0.15(2.17)**	0.16(1.77)*	6.67(0.28)	-2.07(4.29)**	-2.29(5.12)**	-1.13(3.62)**
$lnREX_i$	0.74(4.03)**		0.76(3.88)**		8.39(0.40)		0.15(1.05)	
POS		1.36(5.67)**		1.17(1.82)*		3.93(5.05)**		-0.35(1.42)
NEG		0.85(2.81)**		1.08(2.83)**		2.95(3.96)**		0.31(1.04)
Constant	-62.03(5.31)	-39.43(2.3)**	-13.03(1.86)*		132.94(0.22)	-36.8(2.09)**	-11.63(2.9)**	-27.0(4.04)**
Panel C: Diagnostic Statistics								
F	11.94**	12.32**	6.20**	4.90**	3.85	6.44**	5.74**	7.59**
ECM_{t-1}	-0.74(5.18)**	-1.14(6.52)**	-0.59(4.70)**	-0.59(4.59)**	-0.44(0.38)	-0.65(3.27)	-0.36(4.47)**	-0.99(5.82)**
LM	21.49**	13.67**	10.23**	8.94*	0.91	1.56	5.96	9.16*
RESET	3.04*	1.37	1.54	2.25	0.006	0.08	5.18**	0.02
AdjustedR2	0.42	0.49	0.29	0.29	0.45	0.52	0.26	0.27
CS (CS2)	S(S)	S(S)	S(S)	S(S)	S(S)	UNS(S)	S(UNS)	S(S)
WALD-S		2.86*		0.35		14.37**		0.05
WALD-L		4.35**		0.01		4.52**		8.33**

Notes: See notes at the end.

Table 1c. Full-Information Estimates of Both Linear ARDL (L-ARDL) and Nonlinear ARDL (NL-ARDL) Models

	i=Malaysia		i=Philippines		i=South Africa	
	L-ARDL#	NL-ARDL#	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL
Panel A: Short–Run Estimates						
$\Delta lnY_{US,t}$	-1.12(0.63)	-0.94(0.48)	-1.35(0.75)	-2.09(1.14)	1.39(2.33)**	0.86(1.59)
$\Delta lnY_{US,t-1}$	0.84(0.46)	0.27(0.14)	2.59(1.39)	2.74(1.50)		
$\Delta lnY_{US,t-2}$	-1.25(0.68)	-3.74(1.99)*	-0.53(0.28)	-1.76(0.93)		
$\Delta lnY_{US,t-3}$	4.87(2.65)**	5.64(3.07)**	2.56(1.36)	3.35(1.79)*		
$\Delta lnY_{US,t-4}$			-0.09(0.05)			
$\Delta lnY_{US,t-5}$			3.64(2.06)**			
$\Delta lnY_{i,t}$	0.19(1.22)	0.21(1.07)	0.11(2.02)**	0.09(1.03)	0.52(0.47)	0.16(0.14)
$\Delta lnY_{i,t-1}$	0.02(0.13)	0.33(1.80)*			2.67(2.55)**	2.49(2.39)**
$\Delta lnY_{i,t-2}$	-0.31(1.97)*	-0.19(1.22)			1.95(1.77)*	0.64(0.58)
$\Delta lnY_{i,t-3}$	0.17(1.09)	0.27(1.81)*			2.87(0.67)	2.14(2.01)**
$\Delta lnY_{i,t-4}$	-0.34(2.31)**	-0.32(2.23)**				
$\Delta lnY_{i,t-5}$	0.32(2.35)**	0.43(3.34)**				
$\Delta lnY_{i,t-6}$		-0.11(0.80)				
$\Delta lnY_{i,t-7}$		0.30(2.17)**				
$\Delta lnREX_{i,t}$	0.27(0.81)		0.02(0.30)		0.16(0.66)	
$\Delta lnREX_{i,t-1}$	0.15(0.51)				0.44(1.82)*	
$\Delta lnREX_{i,t-2}$	-0.22(0.74)				-0.02(0.09)	
$\Delta lnREX_{i,t-3}$	1.01(3.44)**				0.17(0.66)	
$\Delta lnREX_{i,t-4}$	-0.23(3.67)**				-0.34(1.34)	
$\Delta lnREX_{i,t-5}$					-0.44(1.69)*	
$\Delta lnREX_{i,t-6}$					-0.16(0.68)	
$\Delta lnREX_{i,t-7}$					-0.72(3.03)**	
ΔPOS_{t}		0.33(0.36)		2.81(2.31)**		0.27(0.29)
ΔPOS_{t-1}		0.01(0.01)		-0.69(0.54)		-3.36(3.49)**
ΔPOS_{t-2}		-1.30(1.19)		-0.55(0.47)		-0.13(0.15)
ΔPOS_{t-3}		4.99(4.57)**		3.21(3.06)**		-2.33(2.54)**
ΔPOS_{t-4}		-1.61(1.08)		-0.05(0.04)		
ΔPOS_{t-5}		5.25(3.87)**		1.14(1.05)		
ΔPOS_{t-6}		-1.94(1.59)		0.57(0.54)		
ΔPOS_{t-7}		2.79(2.34)**		2.33(2.23)**		
ΔNEG_{t}		2.34(1.12)		-1.20(0.93)		0.87(2.02)**
ΔNEG_{t-1}		4.77(2.53)**		3.14(2.44)**		
ΔNEG_{t-2}		2.71(1.41)				
ΔNEG_{t-3}		4.25(2.22)**				
ΔNEG_{t-4}		1.85(1.04)				
ΔNEG_{t-5}		-1.44(0.77)				
ΔNEG_{t-6}		4.17(2.14)**				
Panel B: Long–Run Estimates						
lnY_{US}	2.90(1.89)*	4.26(5.20)**	-1.56(3.99)**	0.29(0.39)	4.62(2.47)**	2.26(1.77)*
lnY_{i}	-0.42(1.27)	-0.32(1.22)	0.18(2.16)**	0.14(1.15)	-2.22(3.07)**	-2.35(1.84)*
$lnREX_{i}$	1.25(2.82)**		0.04(0.32)		-0.13(0.43)	
POS		0.05(0.08)		-0.59(0.93)		3.61(2.96)**
NEG		2.89(3.64)**		0.62(2.09)**		2.29(1.95)*
Constant	-77.09(2.05)**	-117.75(5.97)**	41.90(4.16)**	-11.48(0.53)	-77.19(1.84)*	-3.43(0.09)
Panel C: Diagnostic Statistics						
F	5.09**	8.25**	3.86	4.41*	3.52	6.67**
ECMt-1	-0.35(3.34)	-0.67(5.27)**	-0.59(3.56)*	-0.69(3.78)*	-0.30(3.16)	-0.38(4.09)**
LM	6.23	0.92	12.44**	8.22*	5.09	4.65
RESET	0.33	1.53	2.94*	0.18	1.10	0.02
AdjustedR2	0.35	0.33	0.47	0.49	0.43	0.41
CS (CS2)	S(UNS)	S(S)	S(UNS)	S(UNS)	S(S)	S(S)
WALD – S		0.03		3.32*		1.53
WALD – L		2.83*		2.16		0.50

Table 1d. Full-Information Estimates of Both Linear ARDL (L-ARDL) and Nonlinear ARDL (NL-ARDL) Models

	i=Singapore		i=Thailand	
	L-ARDL	NL-ARDL	L-ARDL	NL-ARDL
Panel A: Short–Run Estimates				
ΔlnY US,t	0.45(0.58)	0.38(0.48)	-3.59(1.52)	-3.13(1.39)
ΔlnY US,t-1				
ΔlnY i,t	0.54(1.25)	0.38(0.86)	-0.33(0.78)	-0.03(0.07)
ΔlnY i,t-1			-0.68(1.48)	-0.26(0.64)
ΔlnY i,t-2			-0.62(1.49)	-0.31(0.79)
ΔlnY i,t-3			-0.23(0.59)	0.26(0.66)
ΔlnY i,t-4			-1.07(2.82)**	-0.79(2.04)**
ΔlnY i,t-5			-1.09(2.94)**	-0.95(2.57)**
ΔlnY i,t-6				
ΔlnY i,t-7				
ΔlnREXi,t	-0.38(0.70)		-0.27(0.80)	
ΔlnREXi,t-1	1.38(3.06)**		-0.90(2.64)**	
ΔlnREXi,t-2			0.28(0.82)	
ΔlnREXi,t-3			-0.66(1.84)*	
ΔlnREXi,t-4				
ΔPOSt		-0.95(1.31)		-0.98(0.89)
ΔPOSt-1				
ΔPOSt-2				
ΔNEGt		-1.13(0.54)		0.71(0.43)
ΔNEGt-1		5.26(2.74)**		-4.52(3.07)**
ΔNEGt-2				0.86(0.56)
ΔNEGt-3				-2.03(1.42)
ΔNEGt-4				-1.95(1.32)
ΔNEGt-5				-2.43(1.65)
ΔNEGt-6				
ΔNEGt-7				
Panel B: Long–Run Estimates				
ln Y US	0.85(0.59)	0.69(0.49)	0.74(1.08)	0.41(0.53)
ln Yi	-1.19(1.87)*	-1.09(1.62)	-0.21(0.52)	-0.02(0.04)
ln REXi	-0.68(1.40)		0.74(2.75)**	
POS		-1.71(1.39)		1.87(3.16)**
NEG		-1.49(1.23)		1.82(2.62)**
Constant	3.52(0.12)	5.97(0.20)	-18.21(1.92)*	-11.26(0.63)
Panel C: Diagnostic Statistics				
F	8.44**	6.91**	17.01**	12.82**
ECMt-1	-0.53(5.58)**	-0.55(5.70)**	-0.92(8.08)**	-0.92(7.61)**
LM	6.18	6.75	11.61**	13.78**
RESET	0.03	0.70	0.51	0.003
AdjustedR2	0.15	0.12	0.38	0.42
CS (CS2)	S(UNS)	S(UNS)	S(S)	S(S)
WALD – S		1.39		2.24
WALD – L		0.11		0.03

Notes: a. Numbers inside the parentheses next to coefficient estimates are absolute value of t-ratios. *, ** indicate significance at the 10% and 5% levels, respectively.

b. The upper bound critical value of the F-test for cointegration when there are three exogenous variables is 3.88 (4.51) at the 10% (5%) level of significance. These come from Narayan (2005). Pesaran et al. (2001) critical values are for large samples.

c. The critical value for significance of ECM_{t-1} is -3.47 (-3.82) at the 10% (5%) level when k=3. The comparable figures when k=4 are -3.67 and -4.03, respectively. These come from Banerjee et al. (1998).

d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with 4 degrees of freedom. The critical value is 7.77 (9.48) at the 10% (5%) level.

e. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) level.

f. Symbol , # , show that dummy is significant. (To account for the Asian financial crisis of 1997)

g. Both Wald tests are distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) level.

These estimates are meaningful due to the fact that cointegration in all six models is supported either by the F test or by ECM_{t-1} test.[10] However, when we consider estimates of the nonlinear model, the POS or NEG variable carry meaningful, significant, and positive coefficient in 10 models that belong to Argentina, Chile, India, Indonesia, Israel, South Korea, Malaysia, the Philippines, South Africa, and Thailand. This supports the new definition of the J-curve due to Bahmani-Oskooee and Fariditavana (2016) and increase support is clearly due to nonlinear adjustment of the real bilateral exchange rate in each case. Clearly, the results are partner specific. For example, in the linear model with India, the exchange rate has no long-run effects on the U.S. trade balance with India. If we were to rely upon the linear model, the process would have stopped right here and we would have concluded that the real bilateral rupee-dollar rate plays no role. However, estimates from the nonlinear model reveals that dollar appreciation could hurt the U.S. trade balance with India, since the POS variable carries a positive and significant coefficient. Or consider the case of South Korea. Again no significant long-run effects of exchange rate changes are discovered from the linear model. However, the nonlinear model reveals that dollar appreciation will hurt the U.S. trade balance with Korea and dollar depreciation will improve it. Again, this must be due to nonlinear adjustment of the real dollar-won rate. All in all, whereas the linear model supports the J-curve in six models, the nonlinear model provides support in 10 cases.

As for the asymmetric effects of exchange rate changes, short-run adjustment asymmetry is supported in all models except in the U.S.-Israel model. Only in this case the ΔPOS and the ΔNEG variables take the same lag order. Furthermore, short-run effects seem to be asymmetric almost in all models since the sign or size of the coefficient estimate attached to ΔPOS at any given lag is different than the estimate attached to ΔNEG at the same lag. However, short-run cumulative or impact asymmetry is evidenced only in the models of the U.S. with Argentina, Ecuador, Indonesia, South Korea, and the Philippines, since the Wald statistic reported as Wald-S in Panel C is significant. Finally, significant long-run asymmetry is established in the U.S. models with Ecuador, Indonesia, South Korea, Mexico, and Malaysia since in these models the Wald test reported as Wald-L is significant.

A few additional diagnostic statistics are also reported in panel C for each model. To test for autocorrelation among the residuals of each optimum model, we report the Lagrange Multiplier test as LM. As can be seen, it is insignificant in 50% of the models,

[10] ECM_{t-1} test is an alternative test under which we use normalized long-run estimates and Equation (1) to generate the error term, called ECM. We then go back to Equation (2) and replace the linear combination of lagged level variables by ECM_{t-1} and estimate the new specification after imposing the same optimum lags. A significantly negative coefficient obtained for ECM_{t-1} will be an alternative way of supporting cointegration. This test which was originally introduced by Banerjee et al (1998) within Engle-Granger (1987) framework is known as the t-test for cointegration and has new critical values tabulated by Banerjee et al. (1998) as well as by Pesaran et al. (2001).

supporting autocorrelation free residuals. To test for misspecification, Ramsey's RESET test is also reported. This statistic is insignificant in most models, implying that a majority of optimum models are correctly specified. The stability of short-run and long-run coefficient estimates is tested by applying the CUSUM (reported as CS) and CUSUMSQ (reported as CS2) to the residuals of each optimum model. Indicating stable estimates by "S" and unstable ones by "UNS", clearly most estimates are stable.

4. SUMMARY AND CONCLUSION

Most of the early studies that tried to test the short-run and the long-run effects of exchange rate changes on the trade balance or the J-curve effect, used aggregate trade flows between one country and the rest of the world. Rose and Yellen (1989) not only criticized these studies for suffering from aggregation bias, but also alerted us to a new definition of the J-curve effect that had its basis in error-correction and cointegration methods, i.e., short-run deterioration combined with long-run improvement. To demonstrate the two points, they used bilateral trade data between the U.S. and each of her six trading partners and Engle-Granger cointegration method but found no support for their definition of the J-curve effect.

Recently, Bahmani-Oskooee and Fariditavana (2016) criticized Rose and Yellen (1989) for assuming the effects of exchange rate changes to be symmetric. They argued that since traders' reactions to currency appreciations could be different than their reactions to depreciations, exchange rate changes could have asymmetric effects on the trade balance. Once they separated appreciations from depreciations and applied Shin et al.'s (2014) recent method of asymmetry cointegration to the same bilateral trade balance models, they not only found evidence of asymmetric effects, but also introduced a new definition of the J-curve effect that was based only on currency appreciation or currency depreciation.

In this paper, we follow the above studies and consider the bilateral trade balances of the U.S. with 13 partners from the developing world. The list of partners are: Argentina, Chile, Ecuador, Indonesia, Israel, India, Korea, Malaysia, Mexico, the Philippines, Singapore, South Africa, and Thailand. These countries, all together, engage in close to 30% of the U.S. trade and therefore they are as important as many partners from the developed world. After applying Pesaran et al.'s (2001) linear ARDL approach to error-correction modeling and cointegration and Shin et al.'s (2014) nonlinear ARDL approach to asymmetry cointegration, our findings could be best summarized by saying that from the linear model we found support for Rose and Yellen's definition of the J-curve in six partners with the U.S. However, when we shifted to the nonlinear model and asymmetry cointegration analysis, we were able to support Bahmani-Oskooee and Fariditavana's definition of the J-curve in the U.S. trade with 10 partners. Like other multi-country studies, our findings are partner specific and have important policy implications. The asymmetry analysis revealed that while dollar appreciation could hurt

U.S. trade balance, dollar depreciation may have no long-run effects. In some other cases, the opposite was true.

APPENDIX

A1. Variable Definition and Data Sources

Quarterly data over the period 1993I-2015IV are used to carry out the empirical analysis. They come from the following sources:

a. Direction of Trade Statistics by the IMF.
b. International Financial Statistics (IFS)

Due to the unavailability of data on some variables, the period was restricted to 1994I-2015III for Argentina, 1996I-2013IV for Chile, 1993I-2006III for Ecuador, 2005I- 2015IV for India, and 1993I-2015II for Singapore.

A.1.1. Variables:

TB_i=US trade balance with partner i, defined as US imports from partner i over her exports to partner i. The data come from source a.

Y_{US}=Measure of the U.S. income. It is proxied by index of real GDP. The data come from source b.

Y_i=Trading partner i's income. This is also proxied by the index of real GDP in country i and the data come from source b.

REX_i = The real bilateral exchange rate of the US dollar against the currency of partner i. It is defined as $REX_i = (P_{US}.NEX_i/P_i)$ where NEX_i is the nominal exchange rate defined as number of units of partner i's currency per US dollar, P_{US} is the price level in US. (measured by CPI) and P_i is the price level in country i (also measured by CPI with the exception of Argentina and Chile). Thus, a decline in REX reflects a real depreciation of the US dollar. All nominal exchange rates and price levels data come from source b.

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Mailing Address: The Center for Research on International Economics and Department of Economics, University of Wisconsin-Milwaukee, Milwaukee, WI 53201, USA. Email: bahmani@uwm.edu

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