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Comparison of the Marshall-Lerner condition in OECD and Asian countries: new evidence from pooled mean group estimation

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Abstract

This paper scrutinizes the well-known Marshall-Lerner (M-L) condition for a group of OECD countries (non-Asian) and Asian countries over the period from 2000 to 2017. We apply a pooled mean group (PMG) estimator to overcome the shortcomings of the other methods, such as the most famous in the recent literature—the autoregressive distributed lag (ARDL), which performs over DOLS, FMLS, and MLE and works despite having endogenous variables in the model. However, the ARDL provides significant but economically implausible coefficients due to country-specific omitted variables, as confirmed by recent empirical studies. In addition, the ARDL ignores the economic convergence assumption that is reasonably included in the PMG approach, assuming homogeneity of the long-run coefficients and heterogeneity of short-run dynamics. Homogeneity of long-run coefficients appears reasonable due to technology and knowledge mobility as a result of globalization. Furthermore, PMG estimator performs better than mean group (MG) and dynamic fixed-effect (DFE) approaches, as it provides more sensible results. We found that the M-L condition holds for Asian countries but not for OECD countries and that there exist signals of J curves for Asian countries in our results. This indicates that currency devaluation works for Asian countries but not OECD countries.

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1 Introduction

The recent trade war between the United States and China reverts attention to the effectiveness of currency devaluation to improve trade flows and competitiveness of domestic industries. The United States condemns China for devaluating its currency in favor of trade balance. However, this raises a crucial question that if currency devaluation works for China, why does the United States avoid manipulating its currency? One can extend such a concern to other developed counties and ask the same question. Therefore, this paper attempts to answer the crucial question of whether currency devaluation works for non-Asian developed countries.

Ever since the dollar devaluation in 1971 worsened U.S. trade balance, thus causing a higher trade deficit, economists have been concerned with the reason behind this phenomenon. Magee (1973) and Junz and Rhomberg (1973) explained that deterioration in trade flows occurs in the short run due to the lagged effect of the exchange rate devaluation. Magee (1973) explained the short-run deterioration and long-run improvement in trade flows (J curve) based on "currency contract" analysis, while Junz and Rhomberg (1973) focused on the obstacles that cause a delay in observing the immediate effect of currency devaluation, such as recognition lags, decision lags, delivery lags, replacement lags, and production lags.[1]They claim that most of a price change effect can be seen in the first five years and that the exchange rate behaves in the same fashion.[2] It is worth mentioning that in the early 1980s, the International Monetary Fund (IMF) proposed currency devaluation as a potential solution to improve developing countries' trade balance.[3]

The discussion surrounding exchange rate has been developed long ago, when Robinson (1937) explained the condition under which currency devaluation could improve trade balance (Bahmani-Oskooee *et al.,* 2013). The argument later formulated mathematically claims that to improve trade balance, the sum of the absolute value of the elasticity of exports demand and imports demand should be greater than one; otherwise, currency devaluation would deteriorate the balance of trade.[4] However, after the deterioration of the U.S. dollar in 1971, economists considered a time horizon to explain why the condition holds in the long run but not in the short run. They explained that in the short run, consumers do not change their behavior, and demand does not respond to a price change and exchange rate changes because the adjustment process takes time. In the long run, the adjustment process occurs, and the larger elasticities of exports and imports provide sufficient empirical evidence for theoretical conjecture.[5]

A broad range of empirical studies has been conducted to evaluate the effect of exchange rate on trade flows, and some of these studies primarily explored the existence of the M-L condition. The literature surrounding trade flows followed an evolutionary process, as did other fields that rely upon methodological advances. To study trade flows in a variety of countries, economists applied different approaches, such as Cochrane-Orcutt (Houthakker and Magee 1969, Warner and Kreinin 1983), GLS (Wilson and Takacs 1979), Almon lag (Bahmani- Oskooee 1986), 2SLS (Arize 1987), Engel-Granger (Andersen 1993, Bahmani-Oskooee 2002), Johansen (Bahmani-Oskooee1996, Bahman-Oskooee and Ebadi 2014), FMLS (Sinha 2001), DOLS (Reinhart 1995 and Prawoto

[1] A recent study reveals that these lags have shortened due to technological advances. For more information, see Bahmani-Oskoee and Ebadi (2016).

[2] They implicitly rejected the Orcutt's hypothesis (1958).

[3] Owen J.R. (2005).

[4] For mathematical proof, see Appendix B.

[5] Magee (1973) and Junz & Rhomberg (1973).

2007), and ARDL (Bahmani-Oskooee and Kara 2005, Liu *et al*. 2007, Bahmani-Oskooee and Ebadi 2016).

The cointegration approach became a turning point in empirical studies, as it uncovered the spurious relationship problem due to the inclusion of nonstationary variables in previous models.[6] While the Engel-Granger and Johansen approaches overcome some specification issues, they suffer from an endogeneity problem as a considerable shortcoming. They make an assumption to have exogenous explanatory variables in the model, and such an assumption is not reliable. The autoregressive distributed lag approach proposed by Pesaran and Shin (2001) has empirical power over previous methods such as DOLS, FMLS, and MLE[7] as it works even when having endogenous variables in the model.[8] Therefore, it has attracted many researchers, who use the approach to study trade flows.

Surprisingly, panel studies did not attract researchers as strongly as the ARDL approach had. Although the panel data approach in time-series econometrics has some advantages, such as more data variation, less collinearity, and more degrees of freedom, it offers a feature that helps to consider economic convergence. Since globalization has spread worldwide, economists justified economic convergence in the long run due to capital and knowledge mobility.[9]However, in econometrics, including economic convergence encountered practical issues. Panel studies either assume different intercepts and slopes for all groups in the study (pooled ordinary least square) or the same intercepts and different slopes for all groups (fixed effect and random effect), ignoring the short-run and long-run dynamics. Therefore, it is impossible to scrutinize the economic convergence as we move from different short-run effects (heterogeneity of the short-run coefficients) across countries to similar long-run effects (homogeneity of the short-run coefficients) due to capital and knowledge mobility.

The pooled mean group (PMG) estimator, proposed by Pesaran *et al*. (1999), allows short-run coefficients to differ across groups but constrains long-run coefficients to be identical. It has empirical power over conventional fixed-effect estimators, as it is robust to the order of the lag and lag determination criterion, such as SBC and AIC. Although the ARDL approach for specific countries sometimes provides significant but economically implausible coefficients due to too much aggregation, sample-specific omitted variables or measurement errors correlated with the regressors, the PMG estimates of the long-run coefficients tend to be sensible.[10]

Therefore, in this paper, we apply the PMG estimator to discover the existence of the M-L condition in a group of OECD countries (non-Asian) and Asian countries, using the homogeneity assumption of the long-run coefficients for the price elasticity of exports and imports demand. Our hypothesis is that the M-L condition holds for Asian countries, as currency devaluation worked efficiently to improve their trade balance, but not for other OECD countries.

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[6] For a comprehensive literature review on empirical works around Marshall-Lerner condition, see Bahmani-Oskooee *et al*., (2013).
[7] Panopoulou and Pittis (2004).
[8] Pesaran and Shin (2001).
[9] Piketty, T. (2014)
[10] Pesaran *et al*. (1999).

The following section discusses the model and methodology. Afterward, Section 3 provides the empirical results, and Section 4 concludes the study.

2 Model and Methodology

Following Bahmani-Oskooee and Kara (2003) and Bahmani-Oskooee and Kara (2005), this paper applies a "well-known model"[11] to test the existence of the M-L condition for 22 OECD and 10 Asian countries[12] over the period from 2000 to 2017. We use the reduced-form equations[13]:

$$lnM = \alpha + \beta lnY^D + \gamma lnREER \qquad (1)$$

$$lnX = \alpha' + \beta' lnY^W + \gamma' lnREER \qquad (2)$$

In these models, X and M represent exports and imports demand, respectively. We include domestic income (Y^D) for imports demand and foreign income (Y^W) for exports demand to capture the effect of economic growth. We expect a positive sign for Y^W, as economic growth around the world would increase the demand for exports. We expect the same sign for domestic income, as economic growth in a specific country increases its demand for other countries' imports. In addition, one would expect a negative sign for the real effective exchange rate in the exports model because an increase in the real effective exchange rate (appreciation) hurts the export. However, the exchange rate appreciation rises imports demand as other countries' products become cheaper for domestic consumers. Therefore, we expect a positive sign for $REER$ in the imports demand model. Furthermore, we wish to estimate the ARDL (p, q_1, q_2) models:

$$\Delta lnM_{it} = \alpha_i + \lambda_{1i} lnY^D_{it} + \lambda_{2i} lnREER_{it} + \sum_{j=1}^{p-1} \varphi_{ij\Delta} lnM_{it-j} + \sum_{j=0}^{q_1-1} \beta_{ij} \Delta lnY^D{}_{it-j} + \sum_{j=0}^{q_2-1} \gamma_{ij} \Delta lnREER_{it-j} + \mu_{it} \quad (3)$$

$$\Delta lnX_{it} = \alpha'_i + \lambda'_{1i} lnY^W_{it} + \lambda'_{2i} lnREER_{it} + \sum_{j=1}^{p-1} \varphi'_{ij} \Delta lnX_{it-j} + \sum_{j=0}^{q_1-1} \beta'_{ij} \Delta lnY^W{}_{it-j} + \sum_{j=0}^{q_2-1} \gamma'_{ij} \Delta lnREER_{it-j} + \upsilon_{it} \quad (4)$$

To estimate the long-run coefficients and the group-specific error-correction coefficients, Pesaran *et al.* (1999) use the maximum likelihood (ML) estimation, considering homogeneity restrictions on the estimate of the long-run coefficients. For group-wide mean estimates of the short-run parameters and the error-correction coefficients, they use the averages across groups. They propose two different likelihood-based algorithms for the computation of PMG estimators: the "back-substitution" algorithm, which uses the first derivatives of the likelihood function in the

[11] Bahmani-Oskooee *et al.* (2013).

[12] See Appendix A for country name and data sources.

[13] Bahmani-Oskooee and Kara (2013) use the inverse of $REER$ for the imports demand function to obtain a negative coefficient revealing a theoretically reasonable effect of currency appreciation. However, we prefer to use the $REER$ by itself and expect a positive coefficient. Mathematically, there is no difference between $\gamma \ln\left(\frac{1}{REER}\right)$ and $\gamma lnREER$ as $ln1 = 0$ is zero. However, from an econometrical perspective, the signs of the coefficients would be different.

optimization procedure, and the "Newton-Raphson" algorithm, which applies first and second derivatives.[14]

The computational econometrics program was written using the GAUSS platform provided by Pesaran *et al*. (1999) and can be used to estimate models of interests. We follow the computation process assuming maximum lags of one. Primarily, we use the Akaike information criterion (AIC) to determine the optimal lags for the group-specific ARDL models used in the estimation process. To determine whether the models are robust to the choice of lag order, we use Bayesian information criterion (SBC) and maximum lags of two. In addition, we employ both algorithms to explore any possible changes in the coefficients. The GAUSS program provides an opportunity to consider trends together with the intercept in the model. Therefore, we include and exclude a trend to observe how our models respond to these kinds of inclusions and exclusions.

3 Empirical Results

According to country-specific estimates and diagnostic results based on the ARDL (p, q_1, q_2) specification for OECD countries' exports, we observe misspecification only in Spain. This confirms that the "well-known model"[15] performs well in this aspect as it does for normality. However, serial correlation in five countries (Australia, Belgium, Greece, Norway, and the US) provides evidence for country-specific omitted variables. However, 17 countries still perform well in terms of serial correlation. As Pesaran *et al*. (1999) mentioned in the paper, ARDL sometimes provides significant but economically implausible coefficients. The observation provided in Table 1 confirms this idea, and interestingly, the results of Bahmani-Oskooee *et al*. (2013) failed in the same manner.

The long-run coefficients for income and price elasticities are 1.45 and -0.42, respectively, which are significant (Table 2). Although MG estimates reveal the same elasticity for income, the price elasticity carries the wrong sign. However, the Hausman test confirms the homogeneity of the long-run coefficients, indicating that PMG performs better in this model.[16] We changed our lag selection criterion and our imposed maximum lags of two, but the price elasticity coefficient slightly changed. In addition, both algorithms provide similar results for price elasticity; however, "back substitution" alters the income elasticity from 1.45 to 1.1.

Surprisingly, the model is not stable without a trend, the reason for which requires further investigation.[17] We randomly dropped three countries and observed slight changes in the coefficients. The negative and significant error-correction coefficient in the exports model confirms cointegration among variables, and the speed of adjustment toward steady state is approximately 60%.

[14] Pesaran *et al*. (1999).

[15] Bahmani-Oskooee *et al*. (2013).

[16] The likelihood ratio statistic is distributed following chi-squared, with 42 degrees of freedom and a p-value of 0.00. Since the estimated joint Hausman test (h-t) p-value (0.27) is greater than 0.00, we cannot reject the null hypothesis of the joint homogeneity of the long-run coefficients. For more information, see Pesaran *et al*. (1999).

[17] Pesaran *et al*. (1999) did not include a trend in their provided applications of consumption function and energy demand. However, it seems that trade flows behave differently.

Table 1. Country-Specific Estimates and Diagnostics Results Based on ARDL Specification for OECD Exports[c]

Country	φ_i^a	lnY^w	$lnREER$	X_{SC}^2	X_{FF}^2	X_{NO}^2	X_{HE}^2	$\overline{R^2}$
Australia	-1.00	-0.109	-0.262	4.04	0.66	0.5	0	0.73
	(NA)	(0.181)[b]	(0.062)					
Austria	-0.465	1.209	0.056	0.1	0	0.58	0	0.94
	(0.127)	(0.309)	(0.660)					
Belgium	-1.000	1.060	0.430	4.89	0.26	0.12	0	0.88
	(NA)	(1.060)	(0.137)					
Canada	-0.193	0.527	-0.244	0.07	1.45	2.53	0	0.95
	(0.311)	(2.097)	(0.328)					
Denmark	-0.623	1.371	0.200	2.83	0	0.63	0	0.82
	(0.139)	(0.302)	(0.337)					
Finland	-0.851	2.353	0.643	1.56	0.55	1.04	0	0.91
	(0.105)	(0.298)	(0.345)					
France	-1.000	0.917	-0.442	0.13	3.45	0.11	0	0.93
	(NA)	(0.113)	(0.104)					
Germany	-0.390	1.768	0.396	2.74	0.03	0.31	0	0.94
	(0.167)	(0.420)	(0.396)					
Greece	-0.687	-0.175	-0.371	8.35	0.87	0.24	0	0.62
	(0.249)	(0.700)	(0.478)					
Iceland	-0.369	6.498	-1.327	0.41	6.81	0.22	0	0.53
	(0.293)	(5.406)	(0.835)					
Ireland	-0.2000	3.951	-2.684	0.68	0.02	0.4	0	0.39
	(0.194)	(4.332)	(2.635)					
Italy	-1.000	1.535	-0.395	0.77	2.38	1.31	0	0.97
	(NA)	(0.095)	(0.085)					
Luxembourg	-0.701	2.105	5.042	1.72	1.78	2.41	0.01	0.59
	(0.190)	(0.816)	(1.146)					
Netherland	-0.331	1.490	0.551	0.01	0.49	0.6	0	0.82
	(0.146)	(0.603)	(0.534)					
New Zealand	-1.000	-0.140	0.163	1.08	2.25	0.69	0	0.39
	(NA)	(0.145)	(0.078)					
Norway	-0.575	0.587	-0.496	4.37	2.27	0.81	0	0.69
	(0.130)	(0.227)	(0.196)					
Portugal	-1.000	1.235	-0.227	0.95	0.69	0.65	0	0.94
	(NA)	(0.105)	(0.145)					
Spain	-1.000	0.777	-0.534	0.03	11.43	0.8	0	0.93
	(NA)	(0.108)	(0.082)					
Sweden	-0.219	2.063	-0.716	1.29	3.84	1.1	0	0.87
	(0.222)	(1.362)	(1.670)					
Switzerland	-0.569	1.598	1.641	2.76	0.43	2.76	0.03	0.39
	(0.215)	(1.385)	(1.306)					
UK	-1.000	-0.033	0.588	1.59	3.49	1.03	0	0.52
	(NA)	(0.540)	(0.273)					
US	-0.395	1.423	-0.894	4.17	0.49	0.62	0	0.95
	(0.101)	(0.373)	(0.225)					

Notes a: error-correction coefficient. , b: the standard errors. , c: AIC is the lag order selection criterion.

Table 2. Alternative Pooled Estimates of the Long-Run Income and Price Elasticities of OECD Exports Demand

AIC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.451	0.049	29.735	1.455	0.315	4.617	0	0.99
$lnREER$	-0.416	0.037	-11.363	0.049	0.298	0.164	2.47	0.12
φ	-0.631	0.067	-9.4	-0.662	0.065	-10.118		
						Joint Hausman test:	2.58	0.27

SBC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.411	0.049	28.782	1.252	0.205	6.106	0.64	0.42
$lnREER$	-0.426	0.038	-11.229	0.114	0.277	0.412	3.87	0.05
φ	-0.660	0.071	-9.344	-0.685	0.066	-10.381		
						Joint Hausman test:	5.15	0.08

Maximum Lags (2)

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.315	0.053	24.961	1.448	0.308	4.704	0.19	0.66
$lnREER$	-0.466	0.035	-13.184	0.15	0.373	0.402	2.75	0.1
φ	-0.705	0.068	-10.356	-0.79	0.073	-10.778		
						Joint Hausman test:	2.75	0.25

Three Countries Dropped Randomly

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.417	0.052	27.397	1.215	0.237	5.127	0.76	0.38
$lnREER$	-0.448	0.04	-11.101	0.093	0.337	0.276	2.62	0.11
φ	-0.616	0.076	-8.122	-0.662	0.072	-9.183		
						Joint Hausman test:	3.26	0.20

Back-Substitution Algorithm

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.115	0.049	22.681	1.455	0.315	4.617	1.19	0.27
$lnREER$	-0.386	0.037	-10.431	0.049	0.298	0.164	2.15	0.14
φ	-0.637	0.068	-9.399	-0.662	0.065	-10.118		
						Joint Hausman test:	4.14	0.13

Table 3. Country- Specific Estimates and Diagnostics Results Based on ARDL Specification for OECD Imports[c]

Country	φ_i^a	lnY^D	$lnREER$	X_{SC}^2	X_{FF}^2	X_{NO}^2	X_{HE}^2	$\overline{R^2}$
Australia	-0.102	33.421	3.792	4.36	6.05	0.65	0	0.55
	(0.224)[b]	(80.8)	(7.132)					
Austria	-0.753	0.576	0.778	0.04	0.93	2.29	0	0.92
	(0.173)	(0.181)	(0.321)					
Belgium	-0.256	0.115	-0.453	4.16	2.14	0.52	0	0.8
	(0.208)	(0.868)	(0.321)					
Canada	-1.000	0.889	0.528	0.86	5.28	0.67	0	0.89
	(NA)	(0.113)	(0.059)					
Denmark	-0.408	2.254	-0.013	0.23	4.11	1.84	0	0.75
	(0.147)	(0.784)	(0.910)					
Finland	-0.658	0.999	0.545	3.57	5.27	2.83	0	0.9
	(0.270)	(0.219)	(0.563)					
France	-0.253	0.714	1.850	5.44	0.15	1.62	0	0.95
	(0.132)	(0.380)	(1.089)					
Germany	-0.550	0.401	0.712	0.15	2.25	1.42	0	0.91
	(0.201)	(0.299)	(0.314)					
Greece	-1.000	2.051	0.130	2.57	0.2	1.36	0	0.75
	(NA)	(0.230)	(0.308)					
Iceland	-1.000	0.752	1.845	0.04	1.76	1.86	0	0.94
	(NA)	(0.133)	(0.104)					
Ireland	-0.411	1.094	-0.432	2.82	1.33	0.9	0	0.2
	(0.247)	(0.438)	(0.813)					
Italy	-0.608	0.829	0.706	0.08	8.14	0.31	0	0.95
	(0.203)	(0.219)	(0.241)					
Luxembourg	-0.729	0.900	1.964	7.23	0.84	1.41	0	0.61
	(0.213)	(0.320)	(0.984)					
Netherland	-0.299	-0.571	0.933	0.44	12.05	1.12	0	0.6
	(0.207)	(1.665)	(1.051)					
New Zealand	-1.000	1.378	0.417	0.23	5.81	1.09	0	0.8
	(NA)	(0.301)	(0.145)					
Norway	-0.565	2.172	1.419	0.07	5.76	0.93	0	0.14
	(0.230)	(1.162)	(0.671)					
Portugal	-0.208	2.591	6.327	2.52	2.4	0.3	0	0.78
	(0.134)	(1.611)	(4.492)					
Spain	-0.709	1.214	0.536	2.3	1.78	1.3	0	0.98
	(0.178)	(0.070)	(0.156)					
Sweden	-0.574	0.976	0.448	1.68	2.88	1.03	0	0.88
	(0.234)	(0.220)	(0.372)					
Switzerland	-0.721	2.302	2.902	11.4	0.08	0.55	0	0.77
	(0.144)	(0.702)	(0.886)					
UK	-1.000	1.182	0.227	0.02	2.94	4.35	9	0.58
	(NA)	(0.452)	(0.194)					
US	-0.286	0.880	0.547	0.03	0.61	1.19	0	0.82
	(0.224)	(0.938)	(0.659)					

Notes a: error-correction coefficient. , b: the standard errors. , c: AIC is the lag order selection criterion.

Table 4. Alternative Pooled Estimates of the Long-Run Income and Price Elasticities of OECD Imports Demand

AIC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	1.024	0.04	25.604	2.596	1.477	1.758	1.14	0.29
$lnREER$	0.615	0.035	17.371	1.169	0.329	3.553	2.86	0.09
φ	-0.615	0.058	-10.654	-0.595	0.062	-9.597		
						Joint Hausman test:	3.01	0.22

SBC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	1.008	0.037	27.273	2.59	1.477	1.754	1.15	0.28
$lnREER$	0.601	0.034	17.547	1.16	0.333	3.482	2.84	0.09
φ	-0.673	0.064	-10.567	-0.642	0.069	-9.314		
						Joint Hausman test:	3.01	0.22

Maximum Lags (2)

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	1.144	0.025	46.152	0.479	0.745	0.644	0.8	0.37
$lnREER$	0.598	0.033	18.348	1.069	0.298	3.594	2.54	0.11
φ	-0.700	0.210	-3.334	-1.175	0.4015	-2.928		
						Joint Hausman test:	3.89	0.14

Three Countries Dropped Randomly

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	0.986	0.042	23.556	2.842	1.707	1.665	1.18	0.28
$lnREER$	0.613	0.036	17.248	1.281	0.364	3.521	3.41	0.06
φ	-0.622	0.057	-10.925	-0.601	0.065	-9.202		
						Joint Hausman test:	3.51	0.17

Back-Substitution Algorithm

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	0.982	0.041	23.836	2.596	1.477	1.758	1.2	0.27
$lnREER$	0.595	0.035	17.196	1.169	0.329	3.553	3.08	0.08
φ	-0.618	0.057	-10.84	-0.595	0.062	-9.597		
						Joint Hausman test:	3.23	0.20

Table 5. Country-Specific Estimates and Diagnostics Results Based on ARDL Specification for Asian Exports[c]

Country	φ_i^a	lnY^w	$lnREER$	X_{SC}^2	X_{FF}^2	X_{NO}^2	X_{HE}^2	$\overline{R^2}$
China	-0.104	-3.414	-3.156	0.84	0.06	0.34	0	0.92
	(0.062)[b]	(4.723)	(1.5)					
India	0.014	-5.423	-25.8	1.89	0.53	0.64	0	0.51
	(0.14)	(81.9)	(261)					
Indonesia	-0.499	-1.230	0.3	0.56	0.23	1.18	0	0.65
	(0.18)	(0.9)	(0.312)					
Japan	-0.143	-2.125	-2.014	6.37	1.96	0.55	0	0.93
	(0.133)	(6.294)	(2.505)					
Korea	0.209	5.960	-0.642	8.07	4.02	2.22	0	0.71
	(0.216)	(3.175)	(0.347)					
Malaysia	-0.351	0.865	-2.093	1.34	4.86	0.83	0	0.77
	(0.143)	(1.271)	(0.848)					
Pakistan	-0.253	0.309	-0.862	2.07	3.18	5.59	0	0.55
	(0.148)	(2.149)	(1.267)					
Philippines	-1.337	1.665	-0.644	0.23	0.14	1.5	0	0.72
	(0.204)	(0.379)	(0.200)					
Singapore	-0.171	-3.989	-4.058	0	3.8	3.51	0	0.73
	(0.123)	(5.596)	(2.826)					
Thailand	-0.011	27.123	-20.6	0.03	0.46	0.74	0	0.95
	(0.06)	(129.8)	(106.7)					

Table 6. Country-Specific Estimates and Diagnostics Results Based on ARDL Specification for Asian Imports[c]

Country	φ_i^a	lnY^w	$lnREER$	X_{SC}^2	X_{FF}^2	X_{NO}^2	X_{HE}^2	$\overline{R^2}$
China	-0.674	2.009	-0.414	0.7	0.09	1.36	0.01	0.43
	(0.244)	(0.666)	(0.359)					
India	0.139	27.744	-2.785	0.95	0.13	0.09	0	0.11
	(0.269)	(44.606)	(4.287)					
Indonesia	-0.400	4.602	1.466	4.72	2.21	0.03	0	0.42
	(0.233)	(2.867)	(1.127)					
Malaysia	-1.000	3.512	-0.085	0.2	2.11	0.46	0	0.53
	(NA)	(0.776)	(0.392)					
Pakistan	-1.000	4.267	1.132	1.81	2.2	2.06	0	0.82
	(NA)	(0.414)	(0.181)					
Philippines	-0.433	8.887	1.360	0.28	0.04	0.97	0	0.58
	(0.221)	(3.378)	(2.089)					
Thailand	-1.000	0.678	0.879	0.73	0.44	1.03	0	0.3
	(NA)	(0.690)	(0.275)					

Notes a: error-correction coefficient. , b: the standard errors. , c: AIC is the lag order selection criterion.

Table 7. Alternative Pooled Estimates of the Long-Run Income and Price Elasticities of Asian Exports Demand

AIC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.444	0.317	4.549	1.974	2.979	0.663	0.03	0.86
$lnREER$	-1.043	0.172	-6.063	-5.956	2.927	-2.035	2.83	0.09
φ	-0.25	0.121	-2.061	-0.265	0.134	-1.968		
						Joint Hausman test:	3.03	0.22

SBC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.444	0.317	4.549	1.974	2.979	0.663	0.03	0.86
$lnREER$	-1.043	0.172	-6.063	-5.956	2.927	-2.035	2.83	0.09
φ	-0.25	0.121	-2.061	-0.265	0.134	-1.968		
						Joint Hausman test:	3.03	0.22

Maximum Lags (2)

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	2.641	0.3	8.807	3.481	2.8	1.243	0.09	0.76
$lnREER$	-1.16	0.142	-8.197	-3.621	2	-1.811	1.52	0.22
φ	-0.233	0.106	-2.188	-0.229	0.13	-1.756		
						Joint Hausman test:	2.99	0.22

Three Countries (Japan, South Korea, and Singapore) Dropped

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.659	0.319	5.203	2.842	4.156	0.684	0.08	0.78
$lnREER$	-0.966	0.174	-5.549	-7.55	4.099	-1.842	2.58	0.11
φ	-0.285	0.179	-1.591	-0.363	0.177	-2.055		
						Joint Hausman test:	2.69	0.26

Back-Substitution Algorithm

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^w	1.444	0.317	4.548	1.974	2.979	0.663	0.03	0.86
$lnREER$	-1.043	0.172	-6.063	-5.956	2.927	-2.035	2.83	0.09
φ	-0.25	0.121	-2.061	-0.265	0.134	-1.968		
						Joint Hausman test:	3.03	0.22

Table 8. Alternative Pooled Estimates of the Long-Run Income and Price Elasticities of Asian Imports Demand

AIC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	4.505	0.298	15.113	7.386	3.529	2.093	0.67	0.41
$lnREER$	1.067	1.133	7.996	0.222	0.57	0.389	2.32	0.13
φ	-0.597	0.154	-3.866	-0.624	0.162	-3.862		
						Joint Hausman test:	2.81	0.24

SBC Lag Selection Criterion

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	4.536	0.284	15.953	7.485	3.522	2.125	0.71	0.4
$lnREER$	1.061	0.130	8.168	0.221	0.573	0,385	2.27	0.13
φ	-0.712	0.148	-4.817	-0.674	0.170	-3.977		
						Joint Hausman test:	2.58	0.28

Maximum Lags (2)

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	4.561	0.211	21.582	7.855	3.458	2.272	0.91	0.34
$lnREER$	1.138	0.118	9.613	0.315	0.687	0.458	1.48	0.22
φ	-0.769	0.195	-2.952	-0.938	0.222	-4.223		
						Joint Hausman test:	3.89	0.14

Three Countries (Japan, South Korea, and Singapore) Dropped

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	4.57	0.303	15.086	6.698	2.563	2.613	0.7	0.4
$lnREER$	0.951	0.142	6.714	-0.136	0.653	-0.208	2.91	0.09
φ	-0.419	0.139	-3.021	-0.538	0.13	-4.152		
						Joint Hausman test:	3.17	0.21

Back-Substitution Algorithm

	PMG			MG			Hausman Test	
	Coef.	St.	t-ratio	Coef.	St.	t-ratio	h-test	p-val
lnY^D	4.318	0.299	13.828	7.386	3.529	2.093	0.85	0.36
$lnREER$	0.939	0.133	7.039	0.222	0.570	0.389	1.67	0.20
φ	-0.603	0.152	-3.969	-0.624	0.162	-3.862		
						Joint Hausman test:	2.44	0.29

Although Bahmani-Oskooee (2013) believes that the "well-known model" appears to be misspecified, it appears that the problem arises not because of the model but due to the methodological shortcoming of the ARDL approach, which does not consider economic convergence in the model. The PMG estimates of the exports demand provide significant and economically plausible coefficients.

Table 3 presents the country-specific estimates and diagnostic results for imports demand. As can be seen, we have evidence of serial correlation due to country-specific omitted variables appearing to be the main problem resulting in economically implausible but significant coefficients. Misspecification can be seen in a few countries, and normality holds for all groups. The PMG estimator provides significant long-run income and price elasticities of 1.02 and 0.62, respectively, and performs over the MG estimator, according to the Hausman (Table 4).

We changed the lag selection criterion, assumed maximum lags, and computation algorithms, but the coefficients slightly responded to the changes. In addition, we randomly dropped some countries to see whether the model would collapse, but we did not see any evidence to support that, as the coefficients did not change considerably (Table 4). The negative and significant error-correction coefficient in the exports model confirms cointegration among variables, and the speed of adjustment toward steady state is approximately 60%.

To satisfy the M-L condition, one must sum the price elasticity of exports demand and imports demand to be greater than one. However, the sum by itself is not reliable proof of M-L existence. As Bahmani-Oskooee (2013) suggested, it is better to determine whether the sum is significantly greater than one. Using the observed coefficients and the standard errors, the t statistic (0.515) does not exceed the critical value (2.01), so we cannot reject the null hypothesis of the sum of the price elasticities being equal to one, in favor of the alternative hypothesis, which claims that the sum is greater than one.[18] This suggests that for OECD countries, currency devaluation does not improve their trade balance.

Table 5 provides the country-specific estimates and diagnostic results for Asian countries' exports. As with the exports of OECD countries, the results reveal some evidence of serial correlation and misspecification in the model, and normality is rejected for only one country. The PMG estimator again performs over the MG estimator according to the Hausman test and provides significant income and price elasticities of 1.44 and -1.04, respectively (Table 7). It is worth mentioning that when we impose maximum lags of two, the short-run coefficients carry positive signs but are not significant. This might serve as evidence for the J curve in Asian countries but not OECD countries. The negative and significant error-correction coefficient in the exports model confirms cointegration among variables, and the speed of adjustment toward steady state is approximately 25%.

We changed the lag selection criterion, assumed maximum lags, and algorithms, but the coefficients did not respond considerably, save for income elasticity with maximum lags of two. We dropped three Asian developed countries (Japan, South Korea, and Singapore) to observe how the exports demand responded. However, the coefficients did not appear to be sensitive to that change, and the model remained stable (Table 7).

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[18] For more information, see Bahman-Oskooee *et al*. (2013).

Finally, Table 6 presents the country-specific estimates and diagnostic results for Asian countries' imports. Interestingly, this model presents no sign of serial correlation or misspecification, except in Korea; however, we still observe economically implausible and significant coefficients. It is worth mentioning that this model appeared to be sensitive when we used raw data rather than cross-section demeaned data. This reveals the common factor effect in imports demand among Asian countries.

Again, the PMG estimator performs better than the MG estimator according to the Hausman test and provides significant income and price elasticity of 4.57 and 0.95, respectively (Table 8). Furthermore, changing the lag selection criterion, maximum lags, and computation algorithm did not change the coefficients considerably, nor did dropping developed Asian countries (Japan, South Korea, and Singapore). However, when we dropped Asian developed countries, the model became sensitive to the maximum lags of two. The negative and significant error-correction coefficient in the exports model confirms cointegration among variables, and the speed of adjustment toward steady state is approximately 40%.

The observed t statistic (5.11) confirms that the sum of price elasticities of exports and imports demand exceeds the critical value of 2.01. This indicates that for Asian countries, the M-L condition exists, and currency devaluation improves their trade balance.

The only existing study[19] in the literature before this paper that applies a different model (the ratio of exports over imports as a dependent variable) but the same methodology, using bilateral trade data between the US and her trade partners, found the coefficient of the real effective exchange rate to be 1.24. Although Bahmani-Oskooee *et al*. (2013) believe that this model "does not test the true M-L condition because no separate elasticities are obtained," this paper provides a mathematical proof that to satisfy the M-L condition in the model with a trade ratio of exports over imports, the relative price elasticity should be greater than one.[20] Therefore, that paper demonstrates that devaluation works for the United States, which contrasts with our findings. However, the paper has its own shortcomings. When we use bilateral trade data, there is no opportunity to compare two different groups of countries as we did for OECD and Asian countries, because the reasonable convergence assumption would be controversial.

For instance, if we wish to use bilateral trade data for China to investigate whether devaluation improves its trade balance, we must consider her trade partners from primarily OECD countries. Therefore, it is not consistent to apply PMG using bilateral trade data. The same occurs for studying other Asian countries with their trade partners.

We believe that our findings, based on the consistent comparison using PMG, are more reliable than using bilateral trade data, as we did not exclude one of the main U.S. trade partners to remain consistent with respect to the economic convergence assumption. That exclusion represents the main shortcoming of the mentioned paper.

[19] Goswami, G. G. and Junayed S.H. (2006).

[20] For mathematical proof, see Appendix B.

4 Conclusion

The recent trade war between the US and China reverts attention to the effectiveness of currency devaluation as a potential solution to improve trade balance. To observe the effect of currency devaluation, we must satisfy the M-L condition, claiming that sum of the absolute values of the price elasticities of exports and imports should be greater than one. To investigate the condition, we apply the PMG estimator considering economic convergence in the long-run, assuming homogeneity of the long-run coefficients. We use reduced form exports and imports demand to estimate the elasticities of income and relative prices.

We found that the sum of the absolute values of the price elasticities of exports and imports demand is significantly greater than one for Asian countries but not (non-Asian) OECD countries. The results confirm that currency devaluation works for Asian countries but not OECD countries. The income elasticities of Asian exports and imports demand demonstrate that the economic growth of Asian countries has a crucial rule in international trade, as the income elasticity of demand for import is approximately 4.6, which provides potential opportunities for other nations. The recent events of economic slowdown due to the trade war between the US and China confirms this statement.

Our findings appear sensible compared with other studies because the models are robust to different changes. Interestingly, the models collapse without including trend in the models. In addition, for Asian countries' import, we observe a common factor effect, as the model becomes stable using cross-section demeaned data rather than raw data.

This is the second study in the literature to use the PMG estimator. The first study applies the PMG estimator using bilateral trade data. However, the paper ignores one of the most important trade partners of the US (China) to hold the economic convergence and homogeneity of the long-run coefficients. Therefore, we believe our findings are more reliable, as we did not encounter that problem using aggregated data. For the first time in the literature, we provided a solid comparison to investigate the M-L condition using the "well-known model" in the literature. Although recent literature reveals that the models we use in this paper are misspecified, our findings confirm the model works well and appears to be stable and that the problem is of different estimation methods resulting in economically implausible coefficients.

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Appendix A

All data are annual over the period 2000-20017. We collected all data from the World Bank and the International Financial Statistics of the IMF. The variables and countries are as follows:

Variables:

M: For each country, M is index of the volume of imports.

X: For each country, X is index of the volume of exports.

Y^D: Measure of domestic income proxied by the index of industrial production in each country.

Y^W: World Real Income measured by the index of industrial production in industrial countries.

REER: Real effective exchange rate (a decline reflects depreciation of domestic Currency).

Countries:

OECD: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States.

Asian Countries: China, India, Indonesia, Japan, South Korea, Malaysia, Pakistan, Philippines, Singapore, Thailand

Appendix B

If we formulate trade balance (TB) as follows:

$$TB_x = \frac{X}{Qe} = \frac{X}{Q}\frac{1}{e}$$

Where X, $M, and\ e$ are exports, imports, and exchange rate respectively, assuming $Y = \frac{X}{M}$, normalizing relative prices to one, and differentiating with respect to e, we have

$$\frac{\partial N_y}{\partial e} = \frac{\partial Y}{\partial e}\frac{1}{e} - Y\frac{1}{e^2}$$

Then factoring $Y\frac{1}{e^2}$ from the equation and rearranging the equation, we obtain elasticities:

$$\frac{\partial N_y}{\partial e} = Y\frac{1}{e^2}\left(\frac{\partial Y}{\partial e}\frac{e}{Y} - 1\right)$$

$$\frac{\partial N_y}{\partial e}\frac{e}{Y} = \frac{1}{e}\left(\frac{\partial Y}{\partial e}\frac{e}{Y} - 1\right)$$

Hence, M-L condition requires $\frac{\partial Y}{\partial e}\frac{e}{Y} > 1 \Rightarrow \frac{\partial N_y}{\partial e}\frac{e}{Y} > 0$, which means $\xi_{Ye} > 1$