Domestic Investment Responses to Changes in the Real Exchange Rate: Asymmetries of Appreciation versus Depreciation

BAHMANI-OSKOOEE, Mohsen and HALICIOGLU, Ferda and Neumann, Rebecca

University of Wisconsin-Milwaukee, Istanbul Medeniyet University, Turkey, University of Wisconsin-Milwaukee

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Mohsen Bahmani-Oskooee[a]

Ferda Halicioglu[b]

Rebecca Neumann[c]

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Abstract

We examine how movements in the real exchange rate impact private domestic investment. Importantly, we consider whether investment responds differently to real depreciations versus real appreciations. Using a sample of six emerging markets over 1980 to 2014, we show that considering asymmetric responses provides an important contribution to this literature. Previous mixed results that assume symmetric responses may be better explained by considering such asymmetric effects.

JEL classification: F31

Keywords: Domestic investment; Real exchange rate, Asymmetry Effects, NARDL

[a] Department of Economics, PO Box 413, University of Wisconsin at Milwaukee, Milwaukee WI 53201-0413, USA, bahmani@uwm.edu.

[b] Department of Economics, Istanbul Medeniyet University, Istanbul, Turkey. ferda.halicioglu@medeniyet.edu.tr

[c] Department of Economics, PO Box 413, University of Wisconsin at Milwaukee, Milwaukee WI 53201-0413, USA, rneumann@uwm.edu.

I. Introduction

Changes in exchange rates have far-ranging effects on economies in our globally interdependent world. A preponderance of literature examines how exchange rate changes may impact such variables as output or capital flows.[1] One strand of this literature explores how exchange rate changes may impact domestic investment. From a theoretical perspective, there are countervailing impacts on investment. Currency depreciation is generally expected to increase domestic investment due to increased domestic and foreign demand as exports become relatively cheap, leading to a healthy economic environment, thus to an increase in domestic investment. At the same time, depreciation impacts the price of imported variable inputs as well as the price of imported investment, potentially decreasing domestic investment due to a reduced profit margin. Empirically, previous literature shows mixed responses depending on the sample of countries and the time period considered. We add to this literature by providing a potential answer for why this previous literature shows mixed results. That answer lies in separating out the responses of domestic investment to exchange rate appreciations versus depreciations. While previous literature has assumed symmetric effects, utilizing new empirical methodology allows us to instead explore asymmetric effects. To that end, we review the literature in Section II and introduce the models and methods in Section III. Empirical results supporting asymmetric effects of exchange rate changes on domestic investment are discussed in Section IV with a summary and conclusion in Section V. Finally data definition and sources are cited in an Appendix.

II. The Literature

Studies that have assessed the impact of exchange rate changes on domestic investment have either used panel data or time-series data at the aggregate or industry level. Pooled data or

[1] For example, see Bahmani-Oskooee and Miteza (2003) for a survey of the literature on whether devaluations are expansionary or contractionary.

panel data has shown generally that real depreciations are associated with reductions in domestic investment (Landon and Smith, 2009; Forbes, 2002; Campa and Goldberg, 1995 and 1999). Campa and Goldberg (1995) examine industry-level data for the US while Campa and Goldberg (1999) compare results for the US, Canada, Japan, and the UK. In both papers, they show that a depreciation of the US dollar leads to fall in domestic investment in the manufacturing sector. They find similar results for Japan and the UK but insignificant responses for Canada. Campa and Goldberg (1999) focus on export and imported-input orientation of producers to explain their results – for example in the US and Japan, domestic depreciation has a positive effect on domestic investment that is increasing in an industry's export share and decreasing in an industry's imported input share. Similar to Campa and Goldberg (1999), Harchaoui *et al*. (2005) find insignificant effects of exchange rate movements on the Canadian manufacturing sector overall during the period 1981-1997. Landon and Smith (2009) use an error-correction model to examine a panel of 17 OECD countries at both the aggregate and sectoral level. For their aggregate data, they show generally that real appreciation is associated with an increase in domestic investment overall, describing the result as follows: "the impact on investment of an exchange rate-induced decrease in the cost of imported capital and other inputs overwhelms the impact of any exchange rate-induced fall in the demand for domestic output" (p. 815). They find more persistent responses in the services sectors, which they attribute to these sectors not being large exporters and thus such sectors do not see a large exchange rate-induced contraction in demand for their output. They argue that this relationship implies that the increase in the cost of imported inputs "outweighs the impact of any exchange rate-induced rise in the demand for domestic output" (p. 824). At the industry level, however, they show that depreciations positively impact investment when exchange rate volatility is low.[2]

[2] See Kandilov and Leblebicioglu (2011) and Serven (2003) for related literature on the impact of exchange rate volatility or uncertainty on investment decisions.

At the firm level, Nucci and Pozzolo (2001) show that real depreciation positively impacts domestic investment via a revenue channel (reflecting export price competitiveness and import competition) and negatively impacts investment via a cost channel (coming from imported input prices). Berg, *et al.* (2015) also focus on firm level data from 66 developed and developing countries, introducing an additional channel via freeing up internal funds that can be put toward investment. In particular, they show that real depreciation is associated with labor-cost savings that allow internal financing of fixed capital, which translates into higher investment and growth in countries with less developed financial markets due to relaxing internal financing constraints. However, these results may mask the differential effects both within an individual country and in response to currency appreciations versus depreciations.

Considering time-series aggregate data, Bahmani-Oskooee and Hajilee (2010) is perhaps the most comprehensive study that examine the impact of real currency depreciation on domestic investment in each of the 50 countries in their sample. For 21 countries, they show a significant long-run impact with real depreciations associated with an increase in domestic investment in 10 countries and with a decrease in domestic investment in 11 countries. Their short-run results are also similarly mixed with real depreciations associated with an increase in investment in 16 countries and a decrease in investment in 23 countries.

A common feature of all studies mentioned above is that they have all assumed that exchange rate changes have symmetric effects on domestic investment, i.e., real appreciations are described as having opposite and equal effects as compared to real depreciations. Recent literature, however, has begun to consider whether exchange rate changes may have asymmetric effects on export and import prices. If true, we suspect exchange rate changes could have asymmetric effects on domestic investment as well. Elbejaoui (2013) examines the reaction of export and import prices to a change in the nominal exchange rate. Using quarterly data for four advanced economies over 1981-2011, Elbejaoui shows greater exchange rate pass through from exchange rate appreciations than from depreciations. By contrast, Delatte *et al.* (2012) find that nominal exchange rate depreciations are passed through to domestic prices more than exchange rate appreciations. They focus on the response of CPI prices in the G7 countries over the period 1970-2009. Others have shown that export shares may respond more strongly to real exchange rate appreciations than depreciations (Demian and di Mauro, 2015; Cheung and Sengupta, 2013) or that firm level output may respond differently to real appreciations than to real depreciations (Dhasmana, 2015). The trade balance is also shown to respond in an asymmetric manner to exchange rate changes when Bahmani-Oskooee and Fariditavana (2016) estimated the U.S. bilateral trade balance models with its six largest partners in assessing the J-curve effect.

Cheung and Sengupta (2013) utilize Indian firm-level data to examine how exporting firms respond to real exchange rate changes and real exchange rate volatility. They show that real currency appreciations reduce firms' export shares while real depreciations have little impact. Further, firms that export less face a larger negative impact on exports from exchange rate appreciation. Presumably, firms that export more are more productive and can choose to be more insensitive to exchange rate changes. Demian and di Mauro (2015) use sectoral level data for ten EU countries and show that exports react mostly to exchange rate appreciations rather than depreciations. They describe two rigidities as key to explaining these asymmetric responses: i.) prices are rigid downwards, and ii.) quantities are rigid upwards. Thus, real depreciation would allow an exporter to earn extra profits for each unit sold or to reduce prices and sell more to gain market share. Rigidities in quantities may mean that the firm cannot ramp up production quickly and thus cannot increase the amount exported. Real appreciation should make exports less competitive. A firm may wish to lower prices to maintain market share but cannot reduce prices too much before profits go negative. Hence, instead, the firm may reduce the quantity of exports. In their results regarding asymmetric effects, Demian and di Mauro (2015) show that a 10% real appreciation reduces the value of exports by 10%. Real depreciation has no noticeable impact, with either an insignificant coefficient or a coefficient of size zero. Others have shown that the response of exports to depreciation is insignificant or imprecisely measured (Rahman and Serletis, 2009; Grier and Smallwood, 2013).[3]

Dhasmana (2015) goes beyond exports to examine how the real exchange rate impacts firm level output. To get at potentially asymmetric effects, Dhasmana estimates the sample separately for real exchange rate appreciations versus depreciations. He also takes account of whether firms are in a high or low concentration industry using a Herfindahl index. He shows that firms in low concentration industries are significantly affected by real depreciations but not by real appreciations. Firms in low concentration industries also show evidence of both the export competitiveness channel and the import cost channel. Firm in high concentration industries are significantly affected by real appreciations but not by real depreciations and this effect comes through the export competitiveness channel. Thus, real appreciation makes exports less competitive and reduces firm output for those firms in high concentration industries.

The evidence of asymmetric responses to exchange rate changes on import and export prices as well as trade flows by the above studies points to the possibility of asymmetric responses to exchange rate changes of domestic investment, which we investigate in this paper. Therefore, we provide a first analysis that allows the effects of real appreciation to differ from those of real depreciation on domestic investment. In doing so we follow the approach developed by Shin *et al*. (2014) to allow for asymmetric effects within a standard ARDL (autoregressive distributed lag) approach of Pesaran *et al*. (2001). We present the model and these methods in the next section.

III. The Model and Methods

[3] While Rahman and Serletis (2009) focus on the impact of exchange rate uncertainty on exports, they also show that appreciation of the nominal exchange rate is associated with lower exports while depreciation is insignificant. Accounting for real exchange rate uncertainty, Grier and Smallwood (2013) show that real appreciations tend to reduce export growth. Real depreciations have asymmetric effects and may increase or decrease export growth. They argue these asymmetries come through increased real exchange rate uncertainty that arises from real depreciations.

While our goal is to study the different impacts of real exchange rate changes on private domestic investment by allowing for differential effects from exchange rate depreciations versus appreciations, we start with an initial specification that mirrors that in Bahmani-Oskooee and Hajilee (2010):

$$LnI_t = \alpha + \beta LnY_t + \gamma Lnr_t + \varphi LnREX_t + \varepsilon_t \tag{1}$$

The dependent variable, I, is private real domestic investment. The independent variables consist of real income, Y, the nominal interest rate, r, and the real exchange rate, REX (where an increase indicates real appreciation for the domestic economy). Since in a growing economy, firms become more optimistic about the future course of the economy and invest more, we expect an estimate of β to be positive. On the other hand, since an increase in the interest rate raises borrowing costs, we expect firms to invest less; hence a negative estimate of γ is expected. As discussed in the introductory section, real depreciation could have a positive or negative impact on domestic investment, hence an estimate of φ could be positive or negative, depending if more firms in a country are export oriented or import oriented. Clearly, a real depreciation that makes exports more attractive will boost exports and through multiplier effects will also boost real GDP and eventually domestic investment. On the other hand, a real depreciation that raises the cost of imported inputs will hurt profit margins and discourage investment by firms that rely heavily on imported inputs, though this effect could be tempered due to slow adjustment of wages to inflationary effects of depreciation. If wages do not keep up with inflationary effects of depreciation, income and profits are redistributed from workers to producers (Bahmani-Oskooee and Hajilee, 2010). Higher income and profits then provide incentives to domestic firms to invest more in domestic capital.

Estimation of equation (1) by any method yields only the long-run effects of exogenous variables on investment. In order to assess the short-run effects as well, equation (1) must be specified in an error-correction format. A one-step procedure that provides estimates of short-run and long-run effects is that of Pesaran *et al.*'s (2001) ARDL bounds testing approach, outlined by equation (2):

$$\Delta LnI_t = \alpha_0 + \sum_{i=1}^{n1} \alpha_{1i} \Delta LnY_{t-i} + \sum_{i=0}^{n2} \alpha_{2i} \Delta Lnr_{t-i} + \sum_{i=0}^{n3} \alpha_{3i} \Delta LnREX_{t-i} + \beta_0 LnI_{t-1} + \beta_1 LnY_{t-1} + \beta_2 Lnr_{t-1} + \beta_3 LnREX_{t-1} + \omega_t \qquad (2)$$

Specification (2) is an error-correction model and once it is estimated by OLS, short-run effects are inferred by the estimates of coefficients attached to first-differenced variables. The long-run effects are judged by the estimates of β_1 to β_3 normalized on β_0.[4] However, for the normalized estimates to be valid, we must establish cointegration among all four variables. Pesaran *et al.* (2001) propose applying the F test with critical values that account for the integrating properties of all variables. They demonstrate that their upper bound critical values could be used even if some variables are I(0) and some I(1), ruling out pre unit-root testing.[5]

As mentioned before, different variants of (1) or (2) have been estimated by previous studies by assuming that exchange rate changes have symmetric effects. However, as discussed, it is more likely for exchange rate changes to have asymmetric effects. Real appreciation (an increase in REX) makes imported inputs less expensive, and thus would be expected to increase the amount available for investment in the long run. However, real depreciation that makes exports more competitive could result in an increase in investment by export oriented firms. Thus, exchange rate appreciation could have the same or opposite effects on investment compared to depreciation. Even if the effects are in the same direction, the relative magnitude may differ; hence asymmetric effects could be seen in the size of the response of domestic investment.

The procedure that we follow to test for the asymmetric effects of exchange rate changes on domestic investment is that of Shin *et al.* (2014) who propose decomposing exchange rate

[4] For exact normalization procedure see Bahmani-Oskooee and Tanku (2008) and for another application see De Vita and Kyaw (2008).

[5] This relies on the assumption that macroeconomic variables are either *I*(0) or *I*(1) which is typically the case.

changes into two components, one representing only real appreciation and one representing only real depreciation. To that end we first form *ΔLnREX* (exchange rate changes) which includes positive changes denoted by $\Delta LnREX^{+}$ (appreciations) and negative changes denoted by $\Delta LnREX^{-}$ (depreciations). Our two new time-series variables are then constructed as:[6]

$$POS_t = \sum_{j=1}^{t} \Delta \ln REX_j^{+} = \sum_{j=1}^{t} \max\left(\Delta \ln REX_j, 0\right)$$

$$NEG_t = \sum_{j=1}^{t} \Delta \ln REX_j^{-} = \sum_{j=1}^{t} \min\left(\Delta \ln REX_j, 0\right) \qquad (3)$$

where POS_t, the partial sum of positive changes, represents currency appreciation and NEG_t,, the partial sum of negative changes, represents depreciation. We then shift back to error-correction model (2) and replace *LnREX* by POS and NEG variables. This results in a new error-correction model as follows:

$$\Delta LnI_t = \theta_0 + \sum_{i=1}^{n1} \theta_{1k} \Delta LnI_{t-i} + \sum_{i=0}^{n2} \theta_{2i} \Delta LnY_{t-i} + \sum_{i=0}^{n3} \theta_{3i} \Delta Lnr_{t-i} + \sum_{i=0}^{n4} \theta_{4i} \Delta POS_{t-i}$$
$$+ \sum_{i=0}^{n5} \theta_{5i} \Delta NEG_{t-i} + \lambda_0 LnI_{t-1} + \lambda_1 LnY_{t-1} + \lambda_2 Lnr_{t-1} + \lambda_3 POS_{t-1} + \lambda_4 NEG_{t-1} + \mu_t \qquad (4)$$

Shin *et al*. (2014) demonstrate that specification (4) can be estimated using the same approach of Pesaran *et al*. (2001). Since introducing the POS and NEG variables introduces nonlinearity into the model, specification (4) is labelled as a nonlinear ARDL model (NARDL) whereas specification (2) is referred to as a linear ARDL model.

A few issues related to equation (4) deserve mention. First, once equation (4) is estimated by OLS, short-run adjustment asymmetry is inferred if ΔPOS takes different lags than ΔNEG variables. This is judged only by observation. Second, short-run asymmetry size effects are inferred if for a given i, $\hat{\theta}_{4i}$ is different than $\hat{\theta}_{5i}$. Third, short-run impact asymmetry is established if $\sum \hat{\theta}_{4i} \neq \sum \hat{\theta}_{5i}$. To judge this we apply the Wald test, denoted by W_{Short} in the results. Fourth,

[6] For another application of the partial sum concept see Apergis and Miller (2006).

long-run asymmetry effects of exchange rate changes are established if $\hat{\lambda}_3 \neq \hat{\lambda}_4$. To this end, again, the Wald test is applied and is denoted by W_{Long} in the results section.[7] Fifth, if long-run estimates are to be meaningful, cointegration has to be established by applying the F test for joint significance of all lagged level variables. Although the nonlinear model (4) has one more variable than the linear model (2), Shin *et al.* (2014, p. 291) recommend treating POS and NEG as one variable and using the critical values for the F test when there are four exogenous variables and not five. Finally, in case the F statistic is insignificant, following Banerjee *et al.* (1998) they also recommend an alternative test. In this alternative test, estimates of $\hat{\lambda}_0 - \hat{\lambda}_1$ are used to construct the lagged error-correction term included in (4) as:

$$ECM_{t-1} = \hat{\lambda}_0 LnI_{t-1} + \hat{\lambda}_1 LnY_{t-1} + \hat{\lambda}_2 Lnr_{t-1} + \hat{\lambda}_3 POS_{t-1} + \hat{\lambda}_4 NEG_{t-1} \qquad (5)$$

After replacing the linear combination of lagged level variables in (4) by ECM_{t-1}, the new specification is estimated after imposing the same optimum lags. A significantly negative coefficient obtained for ECM_{t-1} supports cointegration. Note that the t-test that is used to judge the significance of this coefficient also has upper bound and lower bound critical values that Pesaran *et al.* (2001, p. 303) tabulate.[8] Both the linear and nonlinear models are estimated for several countries in the next section.

IV. The Results

In this section we estimate both the linear model (2) and the nonlinear model (4) for six emerging economies (Brazil, Hungary, Mexico, Malaysia, South Africa, and the Philippines) using quarterly data. Due to availability of data, the study period differs from one country to another, which is noted in the Appendix. The Appendix also provides definitions of the variables and

[7] Note that both Wald tests have χ^2 distributions with one degree of freedom.
[8] Note that for large samples, Pesaran *et al*.'s (2001) upper bound critical values are the same as Banerjee *et al.*'s (1998, p. 276) critical values. Thus, for small samples we need to use Banerjee et al.'s values.

sources of the data. In estimating each model, a maximum of eight lags are imposed on each first-differenced variable and Akaike's Information Criterion (AIC) is used to select the optimum lags or an optimum model. Results for each optimum model and for each country are reported in Tables 1-6. For ease of exposition, a significant coefficient or statistic at the 10% (5%) level is identified by * (**).

Tables 1-6 go about here

In what follows we review the results for the first country, Brazil, and then summarize the results for all countries. Note that in each table, Part I reports the results associated with the linear model and Part II reports the results associated with the nonlinear model. From the short-run coefficient estimates in Panel A of Part I (Table 1) we gather that each exogenous variable carries at least one significant coefficient, implying that all variables have short-run effects on domestic investment in Brazil. However, none of the effects last into the long run since none of the long-run normalized estimates are significant in Panel B. Thus, it is not surprising that cointegration is not confirmed by either the F test or the ECM_{t-1} test reported in Panel C. In Panel C we also report the Lagrange Multiplier (LM) and Ramsey's RESET statistics. The first statistic is used to test for autocorrelation and the second one for misspecification. Neither statistic is significant, supporting autocorrelation free residuals and correct specification of the optimum linear model. Stability of short run and long run estimates together are established by applying the well-known CUSUM and CUSUMSQ tests. Both show stable estimates, which are indicated by "S".[9] Finally, in Panel C we also report the size of adjusted R^2, which shows a good fit.

Turning to estimates of nonlinear model (4) reported in Part II of Table 1, we gather that again all variables carry at least one short-run significant coefficient, implying that all variables have short-run effects on domestic investment in Brazil. Clearly, appreciation and depreciation of

[9] Unstable coefficients are indicated by "UNS". For graphical presentation of CUSUM and CUSUMSQ tests see Bahmani-Oskooee and Fariditavana (2015).

the Brazilian real have short-run effects on domestic investment. The short-run effects seem to be asymmetric since the size of the estimates attached to ΔPOS variable are different from those attached to ΔNEG variable. Even the sum of short-run coefficient estimates of ΔPOS is significantly different than the sum of coefficients attached to ΔNEG variable as supported by a significant W_{Short} statistic of 2.85 reported in Panel F, supporting short-run impact asymmetry. From the long-run estimates in Panel E we gather that short-run effects of the interest rate and real depreciation last into the long run since these two variables carry significant coefficients. The long-run effects of exchange rate changes are asymmetric since the NEG variable carries a significant positive coefficient but the POS variable does not. Indeed, application of the Wald test supports our argument since the W_{Long} statistic of 3.18 is significant. However, these long-run effects seem to be spurious since cointegration is not supported by either the F test or by ECM_{t-1}. We can conclude that in Brazil real exchange rate changes do have short-run asymmetric effects on domestic investment but we cannot say anything about the long-run effects.

Glancing through Tables 1-6 we summarize our findings by noting first that the real exchange rate (ΔLnREX) has short-run effects on domestic investment in Brazil and the Philippines. However, when appreciations are separated from depreciations, at least one of the two variables (ΔPOS or ΔNEG) carries at least one significant coefficient in all countries except South Africa. This discovery by itself supports nonlinear adjustment of the real exchange rate and use of the nonlinear model. Second, asymmetric adjustment is observed in four countries (Brazil, Malaysia, Mexico, and the Philippines) since the order of lags on ΔPOS are different than the order of lags on ΔNEG variable. Third, significant short-run asymmetric effects are also observed in all countries since either the sign or the size of the short-run coefficients are different. However, significant short-run impact asymmetry is established only in the cases of Brazil and Malaysia since only in these countries is the $Wald_{Short}$ statistic significant.

Turning to the long-run effects, from the linear model we gather that the real exchange rate (ΔLnREX) has no significant long-run effects on domestic investment in any of the countries in the sample. However, when appreciations are separated from depreciations, either the POS variable or the NEG variable carries a significant long-run coefficient in all countries except the Philippines and South Africa. This further supports the importance of the nonlinear model. Indeed, since either the sign or size of these coefficients are different, exchange rate changes have significant long-run asymmetric effects on investment in these four countries given that the W_{Long} statistic is significant in all four cases. These long-run significant asymmetry effects are meaningful due to the fact that cointegration is established either by the F test or by ECM_{t-1} in Hungary, Malaysia, and Mexico. The long-run effects in these three countries are not uniform. In Hungary and Mexico, the POS variable carries a significant negative coefficient, implying that real exchange rate appreciation is associated with a decline in domestic investment. As exports become less competitive in these two countries, export oriented firm are hurt more than those firms that rely on imported inputs. On the other hand, in Malaysia the opposite is true. The NEG variable carries a significant positive coefficient, implying that as the Malaysian ringgit faces real depreciation, domestic investment declines. Apparently, real depreciation that raises the cost of imported inputs hurts import-oriented firms more than it helps export-oriented firms, resulting in a net negative impact on investment. These findings are masked by the linear model and can be attributed to nonlinear adjustment of the real exchange rate and nonlinear modelling.[10]

V. Conclusions

Since the advent of generally floating exchange rates in 1973, the effect of exchange rate changes on macroeconomic variable such as the trade balance, imports, exports, wages, demand

[10] All other statistics are similar to those of Brazil described above, indicating autocorrelation free residuals, correctly specified optimum models, stable coefficients, and good fits.

for money, etc. have received a great deal of attention; domestic investment is no exception. Currency depreciation is said to boost exports and through multiplier effects, domestic output, leading to increased investment by firms to take advantage of higher domestic and foreign demand. On the other hand, due to increases in the cost of imported inputs, firms that rely heavily on imported inputs could experience a decline in their profits leading such firms to invest less. The ultimate impact on domestic investment is ambiguous and country specific. Indeed, previous research has supported both outcomes.

In assessing the impact of exchange rate changes on domestic investment, previous research has assumed that the effects are symmetric. However, due to differential responses of trade flows and import as well as export prices to currency depreciation as compared to currency appreciation, exchange rate changes could have asymmetric effects on domestic investment. In this paper we examine this issue in six emerging markets (Brazil, Hungary, Malaysia, Mexico, the Philippines, and South Africa) using quarterly data over the period 1980-2014. We include a set of standard variables that are typically thought to explain the level of domestic investment in a country. In particular, in addition to the real exchange rate we include the level of economic activity measured by domestic income and financing cost measured by the interest rate. Our focus is on the role that the real exchange rate plays in explaining investment. In particular, we add to this literature by separating out the effects for an appreciation of the real exchange rate versus a depreciation of the real exchange rate, expecting that there may be asymmetric impacts on domestic investment. This allows us to examine whether previous results are masked by assuming symmetric effects.

Policymakers are keen to promote domestic private investment since it has strong implications for economic growth. Competitive devaluations of a currency have been proposed as a way to boost exports and thus economic growth. We show here that the potential impacts may differ depending on whether a country's real exchange rate is appreciating or depreciating.

Generally, we find short-run asymmetric effects of exchange rate changes on domestic investment in almost all countries. Significant long-run asymmetric effects are established using the nonlinear cointegration model in three countries: Hungary, Malaysia, and Mexico. We find that in Hungary and Mexico, real appreciation has significant negative effects on domestic investment but real depreciation does not. In Malaysia, however, the opposite is true. Real depreciation is found to hurt domestic investment in Malaysia but real appreciation is found to have no long-run effect. Such asymmetric findings are attributed to nonlinear adjustment of the real exchange rate and nonlinear modelling. None of these asymmetric effects are apparent when relying upon the standard linear model. Indeed, using the linear model indicates no long-run effect from the real exchange rate on domestic investment in each of these countries. Thus, the nonlinear model reveals important information regarding changes in the real exchange rate and domestic investment.

Appendix

Data Definition and Sources

Quarterly data over the period 1980Q1-2014Q4 are used for Mexico, the Philippines, and South Africa. Due to missing observations, the period was restricted to 1995Q1-2014Q3 for Brazil, 1995QI-2014QIV for Hungary, and 1991Q1-2014Q4 for Malaysia.

All data come from the International Financial Statistics of the IMF.

Variables

I: Gross capital formation in real terms. Nominal figures are deflated by GDP deflator.

Y: Real GDP. Nominal figures are deflated by GDP deflator.

r: Domestic interest rate defined as 3 months deposit rate.

REX=Real effective exchange rate. A decline reflects a real depreciation of domestic currency.

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Table 1: Estimates of Both the Linear and Nonlinear Models for Brazil.

Part I: Full-Information Estimate of the Linear ARDL Equation (2)

Panel A: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.26 (1.75)*	-0.74 (4.93)**	-0.36 (2.34)**	-0.03 (0.23)	-0.20 (1.84)*		
Δln Y	1.72 (6.25)**	1.92 (5.72)**	1.12 (2.78)**	2.10 (5.43)**	0.08 (0.19)	0.68 (2.36)**	0.22 (2.86)**	0.27 (1.17)
Δln r	-0.03 (1.71)*	0.01 (0.93)	-0.05 (0.33)	-0.09 (0.13)	0.09 (2.23)**	-0.04 (1.33)	0.03 (1.03)	-0.06 (1.84)*
Δln REX	-0.001 (0.01)	0.18 (2.75)**	0.004 (0.06)	-0.03 (0.50)	0.19 (3.05)**			

Panel B: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	Ln REX
	-24.90 (0.65)	3.56 (0.86)	0.78 (0.51)	0.77 (0.19)

Panel C: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	
	0.64	-0.04 (0.60)	2.58	0.04	S	S	0.89	

Part II: Full-Information Estimate of Nonlinear ARDL Equation (4)

Panel D: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	0.08 (0.52)	-0.54 (3.68)**	-0.36 (2.51)**	-0.03 (0.24)	-0.19 (1.89)*		
Δln Y	1.74 (6.31)**	1.36 (3.79)**	0.93 (2.37)**	2.09 (5.69)**	0.38 (0.92)	0.87 (3.15)**	0.35 (1.91)*	0.25 (1.67) *
Δln r	-0.06 (1.72)*	0.06 (1.34)	0.02 (0.59)	-0.07 (1.64)*	0.11 (2.64)**	-0.04 (1.12)	0.08 (2.34)**	-0.03 (1.09)
ΔPOS	0.18 (2.03)**	-0.09 (0.85)	0.03 (0.30)	0.29 (2.90)**				
ΔNEG	0.03 (4.78)**	-0.006 (0.08)	-0.03 (0.48)	-0.02 (0.36)	-0.13 (2.04) **	0.11 (1.63)		

Panel E: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	POS	NEG
	-5.11 (0.61)	1.59 (1.58)	-0.32 (1.83)*	-0.07 (0.28)	0.29 (3.03)**

Panel F: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	W_{Short}	W_{Long}
	2.51	-0.28 (2.16)	1.6	0.0027	S	S	0.90	2.85*	3.18*

Notes:

a. Numbers inside the parentheses are absolute value of t-ratios. *, ** indicate statistical significance at the 10% and 5% levels respectively.

b. The upper bound critical value of the F-test for cointegration when there are four exogenous variables is 3.52 (4.01) at the 10% (5%) level of statistical significance. These come from Pesaran *et al*. (2001, Table CI, Case III, p. 300).

c. The critical value for significance of ECM_{t-1} is -3.66 (-3.99) at the 10% (5%) level when k =4. The comparable figures when k = 5 in the nonlinear model are -3.86 and -4.19 respectively. These come from Pesaran *et al*. (2001, Table CII, Case III, p. 303).

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%) significance level. Both Wald statistics also have χ^2 distribution with one degree of freedom.

Table 2: Estimates of Both the Linear and Nonlinear Models for Hungary

Part I: Full-Information Estimate of the Linear ARDL Equation (2)

Panel A: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.42 (0.76)	-0.26 (1.25)	-0.33 (4.02)**	0.52 (1.32)	0.22 (2.26)**		
Δln Y	1.39 (3.89)**	0.66 (2.17)**	0.79 (2.63)**	0.97 (3.07)**				
Δln r	0.04 (1.61)							
Δln REX	-0.05 (0.31)							

Panel B: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	Ln REX
	-3.93 (1.95)*	1.45 (3.44)**	0. 21 (1.68)*	-0.25 (0.31)

Panel C: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	
	1.53	-0.22 (2.62)	1.08	0.65	S	UNS	0.98	

Part II: Full-Information Estimates of the Nonlinear ARDL Equation (4)

Panel D: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.24 (1.98)**	-0.16 (1.31)	-0.25 (2.23)**	0.46 (4.06)**	0.16 (1.43)		
Δln Y	0.86 (3.93)**							
Δln r	0.001 (0.04)							
ΔPOS	-0.26 (2.06) **							
ΔNEG	-0.10 (0.71)							

Panel E: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	POS	NEG
	-3.88 (2.74)**	1.70 (4.96)**	0.002 (0.04)	-0.52 (1.91)*	-0.20 (0.68)

Panel F: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	W_{Short}	W_{Long}
	3.29	-0.51 (4.57)**	1.33	1.52	S	S	0.98	0.64	6.14**

Notes:

a. Numbers inside the parentheses are absolute value of t-ratios. *, ** indicate statistical significance at the 10% and 5% levels respectively.

b. The upper bound critical value of the F-test for cointegration when there are four exogenous variables is 3.52 (4.01) at the 10% (5%) level of statistical significance. These come from Pesaran *et al*. (2001, Table CI, Case III, p. 300).

c. The critical value for significance of ECM_{t-1} is -3.66 (-3.99) at the 10% (5%) level when k =4. The comparable figures when k = 5 in the nonlinear model are -3.86 and -4.19 respectively. These come from Pesaran *et al*. (2001, Table CII, Case III, p. 303).

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%) significance level. Both Wald statistics also have χ^2 distribution with one degree of freedom.

Table 3: Estimates of Both the Linear and Nonlinear Models for Malaysia

Part I Full-Information Estimate of Linear ARDL Equation (2)

Panel A: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.32 (2.70) **	-0.25 (2.48)**					
Δln Y	1.11 (3.21)**	-0.12 (0.38)	1.93 (5.80)**	0.37 (1.24)	1.03 (2.83)**			
Δln r	-0.04 (0.42)	-0.29 (2.83)**						
Δln REX	-0.30 (1.14)	-0.49 (1.58)						

Panel B: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	Ln REX
	-4.15 (0.54)	1.19 (1.32)	1.80 (0.99)	-28.67 (0.69)

Panel C: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	
	5.77**	-0.04 (0.71)	9.78	0.01	UNS	S	0.61	

Part II Full-Information of Estimates of Nonlinear ARDL Equation (4)

Panel D: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-							
Δln Y	1.14 (3.46)**	-0.89 (3.12)**	1.22 (4.18)**	-0.06 (0.24)	0.59 (1.78)*			
Δln r	0.04 (0.43)	-0.24 (2.48)**						
ΔPOS	-1.42 (2.30)**	1.09 (1.62)						
ΔNEG	0.71 (4.51)**							

Panel E: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	POS	NEG
	-5.62 (1.65)*	1.92 (3.36)**	-0.08 (0.53)	1.09 (1.17)	2.57 (5.48)**

Panel F: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	W_{Short}	W_{Long}
	4.83**	-0.27 (3.75)	2.97	5.84	UNS	S	0.64	2.74*	2.92*

Notes:

a. Numbers inside the parentheses are absolute value of t-ratios. *, ** indicate statistical significance at the 10% and 5% levels respectively.

b. The upper bound critical value of the F-test for cointegration when there are four exogenous variables is 3.52 (4.01) at the 10% (5%) level of statistical significance. These come from Pesaran *et al*. (2001, Table CI, Case III, p. 300).

c. The critical value for significance of ECM_{t-1} is -3.66 (-3.99) at the 10% (5%) level when k =4. The comparable figures when k = 5 in the nonlinear model are -3.86 and -4.19 respectively. These come from Pesaran *et al*. (2001, Table CII, Case III, p. 303).

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%) significance level. Both Wald statistics also have χ^2 distribution with one degree of freedom.

Table 4: Estimates of Both the Linear and Nonlinear Models for Mexico

Part I Full-Information Estimate of Linear ARDL Equation (2)

Panel A: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.04 (0.45)	-0.22 (2.73)**	-0.20 (2.50)**	0.25 (4.03)**			
Δln Y	1.55 (6.22) **	0.28 (0.98)	0.62 (2.19)**	1.00 (3.62)**				
Δln r	-0.06 (0.53)							
Δln REX	-0.008 (1.09)							

Panel B: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	Ln REX
	-2.07 (1.27)	1.09 (3.35)**	-0. 03 (0.55)	0.41 (0.98)

Panel C: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	
	2.26	-0.20 (2.93)	3.82	4.76	S	UNS	0.64	

Part II Full-Information of Estimates of Nonlinear ARDL Equation (4)

Panel D: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	0.03 (0.38)	-0.17 (1.98) **	-0.15 (1.81)*	0.31 (4.79)**			
Δln Y	1.54 (5.88)**	-0.16 (0.51)	0.31 (1.05)	0.92 (3.15)**				
Δln r	-0.01 (0.55)	-0.08 (2.55)**						
ΔPOS	-0.33 (1.30)	-0.39 (1.87)*	0.04 (0.23)	-0.20 (1.09)	0.47 (2.51)**	-0.41 (2.22)**		
ΔNEG	-0.22 (1.88)*							

Panel E: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	POS	NEG
	-7.53 (3.49)**	2.43 (4.86)**	0.008 (0.20)	-0.51 (2.45)**	-0.11 (0.91)

Panel F: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	W_{Short}	W_{Long}
	3.53*	-0.30 (3.69)	1.26	2.13	UNS	S	0.68	0.25	8.71**

Notes:

a. Numbers inside the parentheses are absolute value of t-ratios. *, ** indicate statistical significance at the 10% and 5% levels respectively.

b. The upper bound critical value of the F-test for cointegration when there are four exogenous variables is 3.52 (4.01) at the 10% (5%) level of statistical significance. These come from Pesaran *et al*. (2001, Table CI, Case III, p. 300).

c. The critical value for significance of ECM_{t-1} is -3.66 (-3.99) at the 10% (5%) level when k =4. The comparable figures when k = 5 in the nonlinear model are -3.86 and -4.19 respectively. These come from Pesaran *et al*. (2001, Table CII, Case III, p. 303).

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%) significance level. Both Wald statistics also have χ^2 distribution with one degree of freedom.

Table 5: Estimates of Both the Linear and Nonlinear Models for the Philippines

Part I: Full-Information Estimate of Linear ARDL Equation (2)

Panel A: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.60 (6.02)**	-0.18 (1.66)*	-0.17 (1.80)*	0.38 (4.21)**	0.23 (2.91)**		
Δln Y	1.42 (4.54) **	1.41 (4.09)**	1.27 (3.74)**	1.18 (3.64)**				
Δln r	0.03 (0.84)	-0.07 (1.31)	0.03 (0.64)	-0.009 (0.18)	-0.04 (0.82)	-0.12 (2.64)**		
Δln REX	0.05 (0.27)	-0.37 (1.34)	-0.39 (1.66)**	-0.55 (2.97)**				

Panel B: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	Ln REX
	-3.69 (2.44)**	1.43 (3.43)**	0.33 (1.50)	2.27 (1.00)

Panel C: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	$Adj.R^2$	
	3.19	-0.16 (2.49)	6.66	11.17	S	UNS	0.64	

Part II: Full-Information of Estimates of Nonlinear ARDL Equation (4)

Panel D: Short-run Coefficient Estimates

	Lag Order							
	0	1	2	3	4	5	6	7
Δln I	-	-0.57 (5.92)**	-0.12 (1.15)	-0.14 (1.46)	0.37 (4.19)**	0.25 (3.22)**		
Δln Y	1.37 (4.15)**	1.32 (3.89)**	1.26 (3.68)**	1.14 (3.43)**				
Δln r	0.05 (1.15)	-0.06 (1.28)	-0.002 (0.04)	0.01 (0.27)	-0.02 (0.59)	-0.14 (3.22) **		
ΔPOS	0.12 (0.34)	-0.17 (0.51)	-0.37 (1.05)	-0.72 (2.01)**	1.43 (3.99)**			
ΔNEG	0.005 (0.08)							

Panel E: Long-run Coefficient Estimates

	Constant	Ln Y	Ln r	POS	NEG
	-3.45 (1.73)*	1.21 (1.66)*	0.37 (1.62)	0.24 (0.45)	0.03 (0.08)

Panel F: Diagnostic Statistics

	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	$Adj.R^2$	W_{Short}	W_{Long}
	2.61	-0.17 (2.51)	5.69	9.76	UNS	S	0.69	0.10	0.22

Notes:

a. Numbers inside the parentheses are absolute value of t-ratios. *, ** indicate statistical significance at the 10% and 5% levels respectively.

b. The upper bound critical value of the F-test for cointegration when there are four exogenous variables is 3.52 (4.01) at the 10% (5%) level of statistical significance. These come from Pesaran *et al.* (2001, Table CI, Case III, p. 300).

c. The critical value for significance of ECM_{t-1} is -3.66 (-3.99) at the 10% (5%) level when k =4. The comparable figures when k = 5 in the nonlinear model are -3.86 and -4.19 respectively. These come from Pesaran *et al.* (2001, Table CII, Case III, p. 303).

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%) significance level. Both Wald statistics also have χ^2 distribution with one degree of freedom.

Table 6: Estimates of Both the Linear and Nonlinear Models for South Africa									
Part I: Full-Information Estimate of Linear ARDL Equation (2)									
Panel A: Short-run Coefficient Estimates									
	Lag Order								
	0	1	2	3	4	5	6	7	
Δln I	-								
Δln Y	1.34 (4.31)**	-0.25 (0.77)	0.07 (0.21)	1.01 (3.02)**	0.018 (0.05)	-0.46 (1.40)	0.85 (2.85)**		
Δln r	0.07 (2.77)**	0.08 (3.28)**	-0.04 (1.65)*	0.05 (2.06)**					
Δln REX	-0.01 (0.61)								
Panel B: Long-run Coefficient Estimates									
	Constant	Ln Y	Ln r	Ln REX					
	-5.19 (1.02)	1.26 (3.42)**	-0. 16 (0.77)	0.22 (0.66)					
Panel C: Diagnostic Statistics									
	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2		
	3.84*	-0.04 (3.06)	8.46	1.95	S	S	0.54		
Part II Full-Information of Estimates of Nonlinear ARDL Equation (4)									
Panel D: Short-run Coefficient Estimates									
	Lag Order								
	0	1	2	3	4	5	6	7	
Δln I	-								
Δln Y	1.31 (4.09)**	-0.27 (0.80)	0.05 (3.32)**	1.01 (0.17)	-0.003 (2.90)**	-0.49 (0.01)	0.86 (1,48)		
Δln r	0.07 (2.72)**	0.08 (3.14)**	-0.04 (1.63)						
ΔPOS	-0.0004 (0.01)								
ΔNEG	0.01 (0.83)								
Panel E: Long-run Coefficient Estimates									
	Constant	Ln Y	Ln r	POS	NEG				
	-10.63 (1.05)	1.96 (2.14)**	-0.13 (0.75)	0.007 (0.01)	0.27 (1.01)				
Panel F: Diagnostic Statistics									
	F	ECM_{t-1}	LM	RESET	CUSM	$CUSM^2$	Adj.R^2	W_{Short}	W_{Long}
	2.42	-0.05 (2.18)	8.44	2.42	S	S	0.53	0.33	0.05

Notes:

a. Numbers inside the parentheses are absolute value of t-ratios. *, ** indicate statistical significance at the 10% and 5% levels respectively.

b. The upper bound critical value of the F-test for cointegration when there are four exogenous variables is 3.52 (4.01) at the 10% (5%) level of statistical significance. These come from Pesaran *et al*. (2001, Table CI, Case III, p. 300).

c. The critical value for significance of ECM_{t-1} is -3.66 (-3.99) at the 10% (5%) level when k =4. The comparable figures when k = 5 in the nonlinear model are -3.86 and -4.19 respectively. These come from Pesaran *et al*. (2001, Table CII, Case III, p. 303).

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%) significance level. Both Wald statistics also have χ^2 distribution with one degree of freedom.