An Empirical Investigation of Real Exchange Rate Responses to Foreign Currency Inflows: Revisiting the Dutch Disease phenomenon in South Asia

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Abstract

Inflows of foreign currencies into the developing economies, in particular, have been associated with the Dutch disease phenomenon whereby a surge in such inflows is believed to stimulate real appreciation of the real exchange rate. As a result, there could be deindustrialization impacts on the recipient economies following a growth in the non-tradable sector at the expense of the tradable sector's contraction. This paper empirically investigates the dynamics of real exchange rate responses to official development assistance, foreign direct investments and international remittances flowing into the four emerging South Asian economies Bangladesh, India, Pakistan, and Sri Lanka. The results from the extensive econometric analyses show that a 1% rise in the total volume of official development assistance and remittances received appreciates the real exchange rate by 0.18% and 0.23% respectively. In contrast, a 1% rise in FDI inflows was found to trigger a 0.19% depreciation of the real exchange rate. Furthermore, the Dumitrescu and Hurlin (2012) test results reveal unidirectional long run causalities running from official development assistances and FDI inflow to real exchange rate while certifying a bidirectional causal association between inward international remittances and the real exchange rate.

Keywords: Dutch disease; foreign exchange inflows; real exchange rate; foreign aid; remittance; foreign direct investment; causality

JEL Classification: C32, F21, F24, F30, F31, F35

1. Introduction

In the contemporary era of globalization, economies have become more exposed to engagement in multilateral trade (Borchert and Yotov 2017; Staníčková, Vahalík, and Fojtíková 2018; Murshed, Jannat and Amin 2018). Economies, traditionally characterized as closed economies, have gradually opened up to attain the welfares associated with international trade which could warrant the concerned economies to consume beyond their respective production possibility frontiers (Heckscher 1919; Ohlin 1933; Vanek 1968). Moreover, countries can make best use of their indigenous comparative advantages to structure their export and import decisions in order to maximize such welfares. However, the benefits from international trade engagements largely rely on the sound functioning of the appropriate theoretical trading-frameworks. Although the patterns of trade are expected to be governed by the World Trade Organization (WTO), the roles of the WTO seem to have been replaced via the individual decision-making of the trading partners. Regional trading pacts are the ones usually synthesizing the terms and conditions of trade. Thus, economists have often probed into the examination of factors that determine trade engagements between economies (Kale 2001; Kónya 2006; Jayasinghe and Sarkar 2008; Yang, Asche and Anderson 2019).

Amongst the numerous macroeconomic factors governing international trade flows, the critically important role of the Real Exchange Rate (RER)[1] with respect to influencing the magnitude of trade movements across the national boundaries is well documented in the existing literature (De Gregoria and Wolf 1994; Boyd, Caporale and Smith 2011). For instance, volatile RER can impose trade-deficit consequences which, to a large extent, can oin down the export-led growth strategies of the underdeveloped nations in particular (Murshed and Elahi 2019). Although the impact of RER on trade had exhibited ambiguity in country-specific studies, many empirical papers have commented in favor of a volatile RER being responsible for trade levels being below par (Koray and Lastrapes 1969; Bini-Smaghi 1991; Feenstra and Kendall 1991). Hence, it is pertinent to probe into the dynamics adhering to RER movements within the economy, particularly in the context of the developing economies that pursue export-led growth strategies. This is because a real appreciation of the exchange rate is expected to dampen exports while simultaneously amplifying the imports due to the relative price of the export baskets outweighing that of the import baskets (Bahmani-Oskooee and Goswami 2003; Bahmani-Oskooee 2005; SaangJoon 2008).

A plethora of existing studies in the relevant discourse have examined the determinants of RER movements within the economy. Amongst these, a sudden surge in the inflow of foreign currencies is often assumed to trigger unanticipated movements in the RER. In the case of such incoming foreign funds appreciating the RER, Corden and Neary (1982) asserted that it could lead to a trade-off between a nation's exports and imports particularly due to such RER appreciations causing deindustrialization, particularly causing the export sector to contract. This phenomenon of foreign currency inflow-induced RER appreciations, instigating havoc for a booming export sector in particular, is termed as the Dutch Disease (DD) problem[2]. The problem of DD, emerging first in the Netherlands following a discovery of indigenous Dutch natural gas reserve which shrunk the nation's booming export sector, is extremely crucial from the point of view of the emerging economies due to these economies being apprehensively reliant on external financing and foreign investments.

However, empirical evidence regarding the nature of the nexus between RER appreciation and the different modes of foreign currency inflows, attributing to the DD problem, does exhibit ambiguity with respect to the countries of origin[3]. Against this milieu, this paper aims at probing into the factors attributing to RER movements across the four South Asian emerging economies: Bangladesh, India, Pakistan and Sri Lanka. The motivation behind this empirical analysis is sourced from the fact that all these four nations are similar in terms of their classification as lower-middle-income economies which, to

[1] RER is calculated from the Nominal Exchange Rate, measured in terms of local currency units per US dollar, via multiplying it with the general price ratios of the two concerned countries. RER=NER(P*/P) where P* and P are the general price levels of the foreign and domestic countries respectively.

[2] For more information on Dutch Disease see Fielding and Gibson (1992); Nyoni (1998); Vos (1998); and Acosta, Lartey and Mandelman (2009); Amin and Murshed (2017, 2018a).

[3] The relevant studies are discussed in the literature review section in this paper.

some extent, portrays the reliance of these nations on external financing and therefore justifies their inclusion into the analyses. In addition, there are immense potentials among these countries with respect to engaging in cross-border power trade through greater trade openness (Murshed 2018a, 2019a) whereby the movements in the RERs could plausibly exhibit key functions as well. On the other hand, these nations differ in terms of their respective national exchange rate policy which would be ideal in implicitly commenting on the impacts of particular exchange rate regimes on the trends in RER movements following foreign currency influx into these economies. Furthermore, the fact that these countries also have also attracted for large volumes of foreign exchange inflows over the years, mainly in the form of foreign aids (Murshed 2019b), foreign direct investments (Sinha, Tirtosuharto and Sengupta 2019) and international remittances (Jawaid and Raza 2016), making them an ideal test bed for analyses. Although country-specific studies in this issue have been conducted for these nations, it is expected that examining the foreign currency inflow-public expenditure nexus within a panel framework would generate robust results to corroborate the existing findings. The following questions are specifically addressed in this paper:

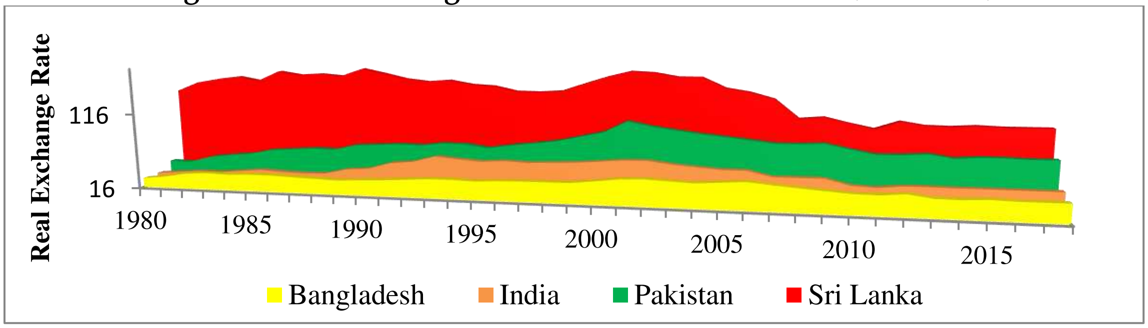
1. Is there any long-run association between foreign exchange inflows and RER movements across the selected South Asian economies?
2. Is there any heterogeneous impact on RER with respect to the types of foreign exchange inflows?
3. Does the nexus between the RER and foreign exchange inflows exhibit any causal association?

The remainder of this paper is structured as follows. Section 2 compares the contrasting exchange rate policies across the selected South Asian economies and also sheds light on the trends in foreign inflows within these economies. The literature review is outlined in section 3 while section 4 highlights the econometric models and the dataset considered in this paper. Section 5 discusses the methodology of the econometric analyses considered while the corresponding results are reported in section 6. Finally, section 7 provides the concluding remarks.

2. Stylized facts on Exchange Rate Policies and Foreign Currency Inflows across Bangladesh, India, Pakistan and Sri Lanka

The historical trends in the RER movements across the four South Asian economies are explicitly illustrated in Figure 1. The exchange rate policies faced by Bangladesh and India have undergone transitions from being fixed in the past to relatively more market-based in recent times. Although India floated its exchange rate regime completely, Bangladesh still pursues a managed floating system. Thus, these nations, to some extent, have succeeded in safeguarding the economies against impulsive RER movements. This scenario can particularly be attributed to the fact that the central banks of Pakistan and Sri Lanka, in contrast to those of Bangladesh and India, have failed to impose a tight grip over their respective RER. Thus, shielding these two nations against the macroeconomic adversities associated with RER volatilities was comparatively cumbersome.

Figure 1. Real Exchange Rate trends in South Asian (1980-2018)

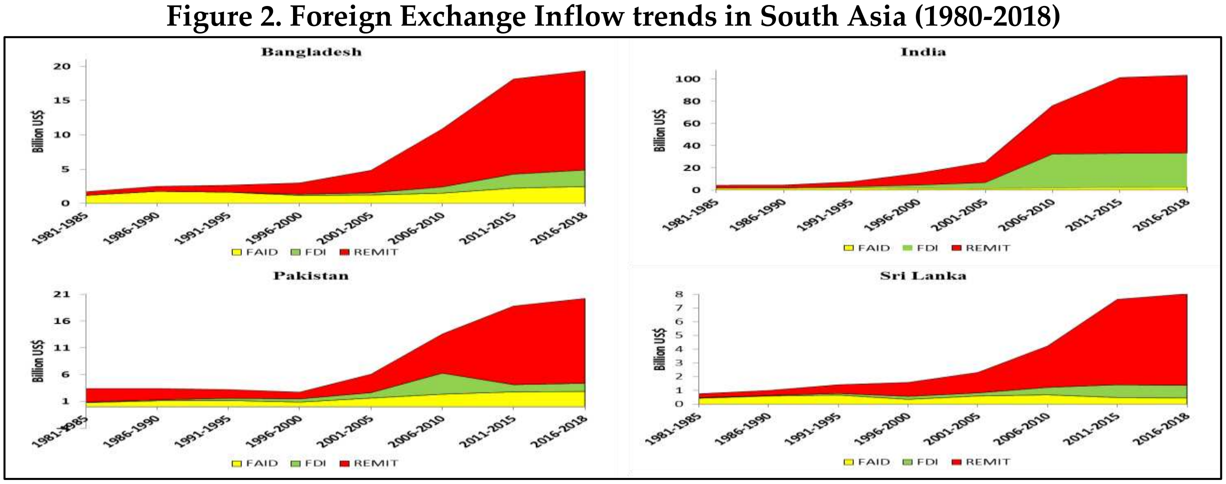


Source: Authors own calculation

Inflows of foreign currencies into developing economies have always been critically important in harnessing the respective development strategies pursued by these nations. Figure 2 provides a graphical analysis of the trends in foreign currency inflows in the form of Official Development Assistances (ODA), Foreign Direct Investments (FDI) and foreign remittances (REMIT) across Bangladesh, India, Pakistan and Sri Lanka. A common scenario from the illustration depicts that REMIT from the expatriates have been the leading source of foreign exchange receipts for all the four countries. India leads in terms of the aggregate REMIT influx, particularly due to it being the most populated country out of the four. As far as the growth in REMIT is concerned, between 2000 and 2014, REMIT in Bangladesh, India, Pakistan and Sri Lanka increased by almost 7.3, 5.4, 14.3 and 5.6 times respectively which seem to explicitly portray the tremendous growth in the volume of foreign currency influx into these countries.

ODA inflows across South Asia have grown over the years. Between 2000 and 2015, Pakistan and Bangladesh registered the highest ODA growth rates of 270% and 104% respectively. In contrast, ODA grew by merely 72% and 66% for India and Sri Lanka respectively. Simultaneously, inflows of FDI have also escalated all the three South Asian economies with Sri Lanka being the only exception. Net inflows of FDIs in Bangladesh, India and Pakistan increased by 2.08, 14.55 and 1.43 billion US dollars respectively over the period of 2006 to 2014. In contrast, Sri Lanka registered a negative FDI growth as perceived from the drop in Sri Lanka's net FDI inflows, within the aforementioned timespan, reduced by almost 0.30 billion US dollars. The dismal performance of Sri Lanka in attracting FDIs could be attributed to the nation's relentless civil war which came to an end only in 2009 (Kapferer 2011).

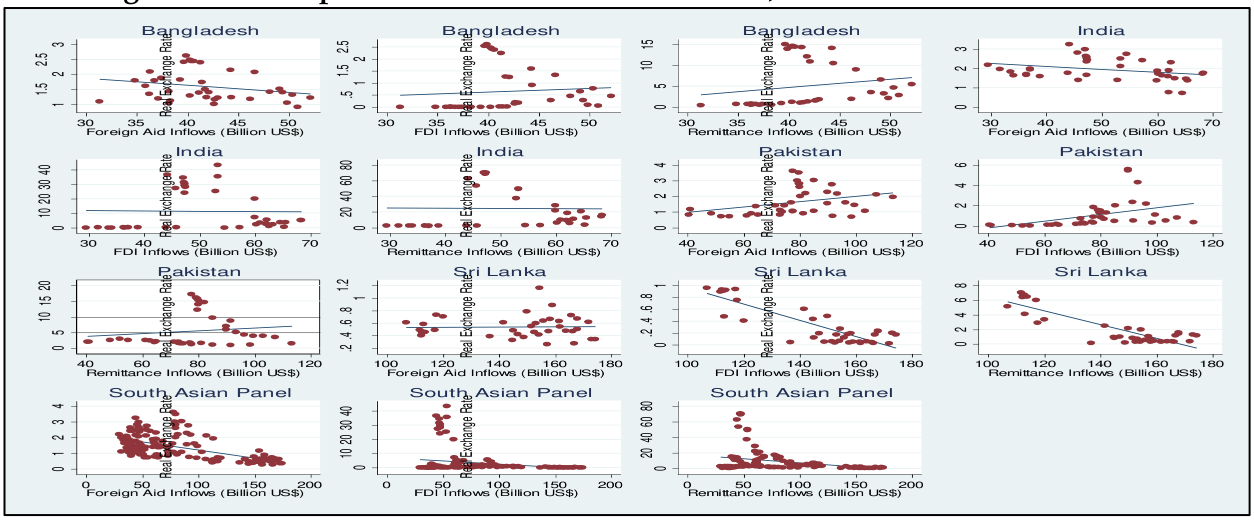
Figure 2. Foreign Exchange Inflow trends in South Asia (1980-2018)



Source: World Development Indicators (WDI 2018).

Figure 3 provides scatter plots to comprehend the nature of the correlation between RER and inflows of ODA, FDI and REMIT into Bangladesh, India, Pakistan and Sri Lanka. It is evident from the graphs that although the natures of correlation differ with respect to country-specific aspects, the bottom three plots tend to show that RER is negatively correlated to all the three sources of foreign currency influx in the context of the entire panel of the four countries. This provides a motivation to further examine the foreign inflow-induced DD phenomenon controlling for other key macroeconomic aggregates.

Figure 3. Scatter plots of RER and inflows of ODA, FDI and REMIT (1980-2018)



Source: World Development Indicators (WDI 2018)

3. Literature Review

This section is segmented into two subsections with the former highlighting the relevant theoretical frameworks while the latter reviewing the empirical findings.

3.1. Theoretical Framework

The negative impact of foreign currency inflow on the trade balance of the recipient economy, via the appreciation[4] of the recipient nation's RER, can be summed up by the DD phenomenon which first came to the limelight in 1977 through the contraction of the Dutch manufacturing sector following the discovery of the Groningen natural gas field in the Netherlands (Corden 1984; Gylfason 1984; Barder 2006). In the international economics literature, the DD problem is specifically referred to as the worsening of the trade balance, of a developing nation in particular, caused due to devaluation of the US dollars against the local currency of the developing economy courtesy the sudden rise in the foreign exchange reserves resulting from a surge in the foreign currency inflows (Younger 1992; Rajan and Subramanian 2011; Amin and Murshed 2018b). Thus, it is evident that influx of foreign exchange, particularly US dollars, tends to exert appreciative pressures on the RER of the recipient economy which makes the investigation of the RER response to multidimensional sources of incoming foreign exchanges pertinent in the context of the developing economies.

[4] In the context of the exchange rate being measured in terms of local currency units per US dollar, a RER appreciation is viewed as a decline in the value of the RER. In contrast, when the exchange rate is measured in terms of US dollar per local currency unit, the RER appreciation is viewed as an increase in the value of RER. Under both circumstances, a RER appreciation attributes to potential DD problems for the foreign exchange recipient economy.

Corden and Neary (1982) explained the DD problem in terms of the 'spending effect'[5] and 'resource movement effect'[6] in response to a surge in the volume of foreign currency inflows. The spending effect is derived from the fact that ODA and FDI are key means for catalyzing development within the underdeveloped economies. In addition, REMIT also play a critically important role in relaxing the budget constraints of the remittance-receiving household. Thus, foreign currency inflows can justifiably be linked to economic development causing betterment of the purchasing power capacities of the people. As a result, demand for both tradable and non-tradable goods can be expected to rise, while simultaneously driving their respective prices. However, under such circumstances, the price ratio of tradable to non-tradable goods is likely to decline. This is because the prices of the non-tradable goods are locally determined, unlike the case for tradable goods due to their prices being determined in the international markets. Thus, the rise in demand for both tradable and non-tradable goods ultimately pushes up the price of the non-tradable goods only while the tradable goods' prices exhibit relatively more rigidity. Hence, it can be said that the spending effect, resulting in the relative price ratio of tradable to non-tradable goods to decline, tend to appreciate the RER and possibly trigger the DD phenomenon. On the other hand, the resource movement effect is hypothesized from the understanding that a surge in the foreign exchange inflows, causing the relative price ratio to fall, would reallocate resources from the tradable to the non-tradable sector, simply because of the returns to investment in the non-tradable sector being relatively more than that in the tradable sector. Thus, the tradable sector is likely to shrink and attribute to worsening of the trade balance.

3.2. Empirical Literature

The impact of foreign inflows on the RER movements leading to the DD problem is well documented in the existing literature. However, the results do show ambiguity with respect to country and region-specific empirical exercises. The subsequent sections shed light on the different sources of foreign exchange inflow and their effects on the RER of the recipient economies.

3.2.1. Literature on Foreign Aid inflows and Dutch Disease

In a study by Nyoni (1998), the aid influx-DD nexus in the context of Tanzania was explored using annual time series data between 1969 and 1993. The econometric model in this paper expressed equilibrium real exchange rate, measured as a ratio of tradable to non-tradable goods, was expressed as a function of ODA, government expenditure and economic openness. The long-run estimates from the error-correction modeling suggested that aid inflows to Tanzania do not attribute to the DD problem and rather cause the RER to depreciate. In contrast, the short run estimated coefficient attached to aid inflow is found to be statistically significant in both the lags of ODA inflows.

Vos (1998) analyzed the DD phenomenon in the context of Pakistan. The author considered the impacts of development aid received for international donors on the RER of Pakistan. The empirical examination was executed using multisectoral computable general equilibrium framework suited to account for the imperfections existing in the financial and goods markets in Pakistan which also incorporated the structural difference in the savings and investment behavior of the stakeholders into the analysis. The results indicated towards the validity of the DD problem originating from ODA inflows to Pakistan. For robustness check, similar empirical exercises were conducted for Philippines, Mexico and Thailand. The results indicated that foreign capital influx to Mexico and Thailand promote expansion of the respective traded goods sector and therefore such inflow of foreign currencies do not exert appreciative pressures

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[5] For more information on the spending effect see Corden (1984) and Benjamin, Devarajan and Weiner (1989).

[6] For more information on the resource movement effect see Lartey (2008) and Acosta, Lartey and Mandelman (2009).

on the RER of these two nations. In contrast, inflow of foreign currency to Philippines was associated with the DD problem.

Focusing on the possible appreciation of RER following the influx of ODA in Bangladesh, Amin and Murshed (2018b) employed Johansen cointegration analysis and Granger causality technique to explore the aid induced DD effect on the Bangladesh economy. The results, although validating a long run association between the concerned variables, failed to establish a causal linkage between ODA inflows and RER movements in Bangladesh over the period of 1980 to 2014. Based on the findings, the authors voiced in favor of Bangladesh's transition from a fixed to a managed-float type of exchange rate policy has enabled the nation to shield the economy from the DD effects of foreign currency influx, particularly in the form of ODA.

In a study by 10 Pacific island states, Fielding (2007) concluded that the impact of ODA on RER movements depicts ambiguity across the countries considered in the investigations. Using annual data from 1970 to 2003, simple conditional vector autoregressive methodology was tapped to draw relevant conclusions on the DD phenomenon in these states. The estimates showed that aid inflows into the Cook Islands and Tuvalu depreciated the respective RER while influx of aid in Tonga did appreciate Tonga's RER initially before depreciating the RER in one period lag of aid inflow. The author also asserted that the small countries not having a currency of their own, thus being unable to adjust the exchange rate movements, are comparatively more susceptible to DD impacts of aid inflows.

As far as panel studies are concerned, Adenauer and Vagassky (1998) found statistical evidence of DD effects of Official Development Assistance (ODA) inflows into the four African nations Burkina Faso, Cote d'Ivoire, Senegal and Togo. A log-log function was considered in this paper in which the Real Effective Exchange Rate (REER) of each of the countries were pooled and regressed as a function of the pooled figures of real GDP, one lag differenced growth rates, terms of trade, and ODA at both its current level and one period lag. The generalized least squares estimates from the regression analyses showed that inflow of ODA appreciated the REER across the four African economies. However, the rate of appreciation was found to diminish with time as the estimated coefficient attached to the one period lag of ODA inflows was lower than the estimated coefficient attached to the current level of ODA influx. This implied that the lagged effect tends to undermine the contemporaneous effect of foreign currency inflows on the REER of the African nations.

3.2.2. Literature on Foreign Direct Investment inflows and Dutch Disease

Financial openness is believed to play a critically important role in escalating the investment profiles of the developing economies in particular. Thus, a financially open economy is expected to attract FDI necessary for boosting investments to harness their development strategies to a large extent. However, obnoxiously high levels of net FDI inflows can be expected to trigger the DD problem which, although is not anonymously validated for all countries, can be detrimental to the development prospects of the recipient economies. In a study by Lartey (2011), the impacts of FDI inflows to 109 developing and transition nations, between 1990 and 2003, on the REER were explored using the Generalized Method of Moments (GMM) panel data estimator. The study considered REER in terms of the price ratio of domestic to international commodities whereby a rise in the REER is expected to dampen exports and ultimately trigger the DD problem. The regression results showed that FDI inflows tend to reduce the REER in the panel of countries considered in this paper. However, the combined impact of financial account openness and FDI inflows on the REER was found to be positive which indicated that financial openness leads to the appreciation of the REER and it also offsets the impacts of FDI inflows on the REER to a large extent.

Thus, the author referred to the overall impact of FDI inflows on REER movements can be expected to vary across countries with dissimilar degrees of financial openness.

Using the behavioral equilibrium real exchange rate methodology, Jaffri and Ahmed (2010) analyzed the impacts of FDI inflows on Pakistan's real exchange rate. This method basically used the Johansen (1988) test for cointegration on a regression model in which the REER was regressed using FDI inflows and other real fundamentals of REER. Monthly data stemming from 1993:M7 to 2009:M3 was compiled for the analysis. Both the trace and the maximum Eigen value tests confirmed the presence of one cointegrating equation in the model. The statistical estimates found FDI inflows into Pakistan to appreciate its REER. Moreover, the fact that almost 70% of the total FDIs flowing into the country within the time period of the study were aimed at the non-tradable sector, the appreciation of the REER can be associated with the validity of the DD phenomenon in Pakistan.

Linking capital flows with RER movements in Asian economies, Jongwanich and Kohpaiboon (2013) explored the nexus between RER and both FDI inflows and outflows in People's Republic of China, India, Indonesia, the Republic of Korea, Malaysia, Philippines, Singapore, Chinese Taipei and Thailand. The study encompassed a dynamic panel data regression approach using biannual data from 2000 to 2009. The dynamic panel data model considered the RER as the dependent variable while FDI, REMITTANCES, government spending, portfolio investment, terms of trade, trade openness, other international financial investments and productivity differentials across the FDI making and receiving economies. The relevant results from the GMM panel data regression analyses specified that both inflows and outflows of FDIs are found to be associated with the appreciation of the RER which can be linked to the DD problem in the context of the aforementioned emerging Asian economies.

3.2.3. Literature on Remittances and Dutch Disease

The rapid growth of inward international REMITTANCES across the globe has attracted the attention of policymakers, particularly linking inflows of such foreign earnings to the DD problem. In an empirical work by Ameudo-Dorantes and Pozo (2004), the impacts of workers' remittance on the RER of 13 Latin American and Caribbean countries. The country-specific data of these 13 nations from 1979 to 1998 were pooled to conduct the econometric analyses. This study tapped the instrumental variable with fixed effects panel data regression method to draw statistically certified conclusions on the RER-remittance nexus. The econometric model expressed RER, measured in terms of US dollars per unit of local currencies of the 13 countries, against workers' remittance, ODA, technological advancement, government expenditures, terms of trade and world interest rate. The results pointed out that the marginal impact of a 1 US dollar increase in workers' remittance leads to a 0.22% appreciation of the RER, on average, ceteris paribus. Thus, inward remittance can be linked to loss of export competitiveness in these nations leading to the DD problem.

Lartey, Mandelman and Acosta (2012) attempted to link REMIT inflows to DD effects on the 109 REMIT-receiving developing economies. The study compiled annual data of all these nations from 1990 to 2003 and employed the GMM panel data estimation technique to handle the problem of endogeneity in the dataset. The study not only analyzed the impact of remittance inflows on the direction of change in the REER but it also isolated the impacts of remittance with respect to the sectoral output composition. The results indicated that inflow of remittance into the developing nations appreciated the REER which could be explained by the spending effect hypothesis put forward by Corden and Neary (1982). Moreover, inward remittance was also found to be associated with the expansion of the non-tradable sector relative to the growth of the tradable sector which further reflected towards the existence of the DD problem via validating the resource movement hypothesis of Corden and Neary (1982).

In a recent study by Amin and Murshed (2018b), the authors evaluated the causal impacts of incoming foreign remittances on Bangladesh's RER movements over the period 1980 to 2014. The results from the cointegration analysis revealed long-run association between the two macroeconomic variables. The Granger causality estimates showed that in the long-run there is unidirectional causality stemming from the nation's RER to remittance inflows which implied that RER appreciations can possibly facilitate greater inflows of foreign currencies into the economy in the form of remittances received from the expatriates.

4. Empirical Model and Data

The empirical model relevant to this paper expressed the RER as a linear function of foreign currency inflows and other key control variables attributing to RER movements across Bangladesh, India, Pakistan and Sri Lanka. The functional form of empirical model can be expressed as:

$$RER = f(ODA, FDI, REMIT, GOV, TOT, OPEN) \quad (1)$$

where RER, ODA, FDI, FDI, GOV, TOT and OPEN refer to the real exchange rate, inflow of development assistances, foreign direct investments received, international remittances receipts, government expenditure, terms of trade and trade openness respectively. The RER is calculated multiplying the nominal exchange rate by the ratio of the international and domestic general price levels, proxied by the consumer price indices of the United States and the four South Asian economies respectively.[7]

The three main sources of foreign exchange receipts considered in this paper are ODA, FDI and REMIT only[8]. ODA comprises of the net official development assistance received.[9] As per the theoretical frameworks building to the DD phenomenon, a rise in the inflow of ODA is expected to appreciate the RER whereby a negative correlation between RER and FDI can be expected (Adenauer and Vagassky 1998; Barder 2006; Addison and Baliamoune-Lutz 2017). Moreover, studies have also linked ODA inflows to stimulation of RER misalignments within the aid-recipient economies (Terra and Valladares 2010; Elbadawi, Kaltani and Soto 2012). FDI includes the net foreign direct investments flowing into the economy. Much like the case of the ODA inflow, the DD phenomenon is believed to be triggered by a rise in FDI influx whereby a negative association between FDI inflow and RER can be anticipated (Kosteletou and Liargovas 2000; Jongwanich and Kohpaiboon 2013; Athukorala and Rajapatiran 2003). International REMIT accounts for the foreign currencies remitted by the expatriate workers and excludes the corresponding outflows. A negative relationship between REMIT flowing into the home economy of the emigrant and the RER of the recipient economy can also be expected as per the DD hypothesis (Bourdet and Falck 2006; Chowdhury and Rabbi 2014).

The regression model controls for GOV, TOT and OPEN. GOV comprises of all current public expenditures for purchases of goods and services and also includes most expenditure. The nexus between GOV and RER movements can be explained by linking public expenditure to the 'resource movement

[7] RER = NER*(P*/P) where NER is nominal exchange rate measured in terms of local currency units per US $ and P* and P are the consumer price indices of the United States and the South Asian economies respectively.

[8] Following the unavailability of portfolio investment and other investment data this paper limits the examination of foreign currency inflows to ODA, FDI and REMIT only.

[9] Due to unavailability of disaggregated data, this paper considers only net official development assistance received to account for foreign aid inflows.

effect' whereby relative expansion of the non-traded sector takes place due to the fact that the public sector mainly purchases non-tradable goods and services, thus increasing the demand and price of non-tradables. Such disproportionate and biased government expenditure in the non-traded sector is likely to appreciate the RER whereby a negative correlation between these variables is ought to be anticipated (Penati 1987; Pegg 2010; Di Giorgio, Nisticò and Traficante 2018; Miyamoto, Nguyen and Sheremirov 2019).

TOT takes into account the net barter terms of trade index which is calculated using the ratio of export to import value indices of the respective South Asian economies. In line with the existing literature summarizing the TOT-RER nexus, a rise in the TOT index is expected to induce 'spending effect' resulting in appreciation of the RER (Cashin, Céspede and Sahay 2004; Ricci, Lee and Milesi-Ferretti 2008). In contrast, some studies have also referred TOT volatility to exert depreciative pressures on the RER of developing countries (Aizenman and Riera-Crichton 2008). Finally, OPEN is measured in terms of the sum of total trade as a percentage of GDP of the respective economies. Similar to TOT index enhancements, an increment in the OPEN is expected to stimulate RER appreciation, primarily via inducing the 'spending effect' while indirectly attributing to the 'resource movement effect' also' to drive up the overall demand for both tradable and non-tradable commodities (Torvik 2001; Bleaney 2008). Thus, a negative association between OPEN and RER can be expected.
The empirical model (1) is log-transformed whereby the relevant variables are expressed in terms of their respective natural logarithms in order to estimate the elasticities. Hence, the functional form of the empirical model can be shown by:

$$\begin{aligned} \boldsymbol{lnRER_{it}} = {} & \boldsymbol{\beta_0} + \boldsymbol{\beta_1 lnODA_{it}} + \boldsymbol{\beta_2 lnFDI_{it}} + \boldsymbol{\beta_3 lnREMIT_{it}} + \boldsymbol{\beta_4 lnGOV_{it}} + \boldsymbol{\beta_5 lnTOT_{it}} \\ & + \boldsymbol{\beta_6 OPEN_{it}} + \boldsymbol{\varepsilon_{it}} \end{aligned} \tag{2}$$

where the subscripts i and t denote the individual cross-sections and the time period, respectively; ε represents the error-term. Annual data stemming from 1981 to 2018 for all the aforementioned variables, in the context of Bangladesh, India, Pakistan and Sri Lanka, have been sourced from the World Development Indicators (WDI 2018) database of the World Bank.

5. Methodology

Prior to the section of the appropriate regression and causality techniques to analyze the data, a set of pre-estimation tests are applied in this paper.

5.1. Cross-sectional Dependency analysis

At first, the dataset is tested for possible cross-sectional dependencies using the appropriate cross-sectional dependency tests proposed by Pesaran, Friedman and Frees[10] and the Breusch-Pagan Lagrange Multiplier test for cross-sectional correlation (Baltagi, Feng and Kao 2012).[11] Cross-section dependence can originate from observed and/or unobserved spatial or spillover impacts common factors between cross-sections/panels whereby (Baltagi and Pesaran 2007). Testing dependency across cross-sections is critically important in the sense that in the case such dependence across the panels, the first generation panel data unit root tests do not perform properly since most of these unit root testing techniques assume

[10] For more information on cross-sectional dependencies in panel data see Pesaran (2015).
[11] It is to be noted that due to the number of cross-sections in the dataset being smaller than the number of time period (N<T), the cross-sectional dependency analyses proposed by Pesaran, Friedman and Frees are more appropriate compared to the Breusch-Pagan LM test .

cross-sectional independence (Basak and Das 2018). Thus, such dependency analysis helps to understand the convergence process of the panel data variables efficiently. The results from the cross-sectional dependency analyses, reported in table 1 in the appendix, provide statistical support of cross-sectional independence in the dataset used in this paper and therefore validate the use of the first generation panel unit root tests.

5.2. Panel group wise Heteroscedasticity Test

Presence of heteroscedasticity in the panel dataset violates the Ordinary Least Squares (OLS) assumptions of homoscedastic variance whereby the OLS estimator is no longer Best Linear Unbiased Estimator (BLUE). Thus, it is pertinent to check for possible heteroscedasticity in the empirical model. Table 2 in the appendix reports the results from the panel group wise heteroscedasticity test. This test basically comprises of three panel data heteroscedasticity identification analyses namely Lagrange multiplier, likelihood ratio and Wald tests (Judge et al. 1982; Greene 1993). The corresponding results from these three aforementioned tests indicate the problem of heteroscedasticity in the empirical model considered in this paper.

5.3. First Generation Panel Unit Root Tests

Following the statistical certification of cross-sectional independence across the panel data set considered in this paper, the first generation panel unit root analyzing tools are employed.

5.3.1. Levin, Lin and Chu (LLC) Test

The LLC test (Levin, Lin and Chu 2002) panel unit root test hinges on the assumption that unit root is a homogeneous process. The term 'homogeneous' denotes that the test is estimated assuming a common Autoregressive (AR) structure for all the cross-sectional units in the form of countries considered in the panel. Let us consider the Augmented Dickey-Fuller (ADF) regression model (3) below to get a clear understanding of the LLC test:

$$\Delta \boldsymbol{y_{it}} = \boldsymbol{\alpha_i y_{i,t-1}} + \sum_{\boldsymbol{L=1}}^{\boldsymbol{\rho_i}} \boldsymbol{\theta_{iL}} \,\Delta \boldsymbol{y_{i,t-L}} + \boldsymbol{\delta_{mi} d_{mt}} + \boldsymbol{\epsilon_{it}} \tag{3}$$

where Δyit = yi,t-1 , αi = -(1-ϱi), dmt is the vector of deterministic variables, δ_miis the corresponding vector of coefficients for model m and εit is a white noise error term for i = 1, …, N cross-sections and t = 1, …, T time periods. The homogeneous unit root assumption implies that αi = α for all i. The LLC test null hypothesis is that each individual series of the panel cross-sections contain a unit root (H0: α = 0 for all i). The null is tested against the alternative hypothesis that the individual series does not contain a unit root (H1: $\alpha \neq 0$ for all i). The probability value of the estimated t-statistic for each of the series provides the result of stationarity with the rule of thumb being if the probability value, with respect to a particular series across all cross-sections, is below 10% level of significance, then the null hypothesis can be rejected implying the series to be stationary. Due to the limitations of the LLC test in the form of being heavily dependent on the assumption of homogeneous unit root across all the cross-sections and being more restrictive in the sense that is assumes all cross-sections to have or not have a unit root which needs ϱ to be homogeneous across all i, the other panel unit root tests are conducted as well.

5.3.2. The Im, Pesaran and Shin (IPS) Test

Unlike the LLC test for panel unit root which assumes a homogeneous unit root process, the IPS test (Im, Pesaran and Shin 2003) allows for a heterogeneous value of αi. The IPS suggests a unit root testing method based on averaging individual unit root test statistics. The basic equation for IPS is as follows:

$$\Delta y_{i,t} = \propto_i + \rho_i y_{i,t} + \sum_{j=1}^{\beta} \varphi_{ij} \Delta y_{i,t-j} + \epsilon_{i,t} \quad (4)$$

where yi,t represents each of the variables under consideration in the model, αi is the individual fixed effect, and β is selected to make the residuals uncorrelated over time. The null hypothesis is that each individual series of the panel cross-sections contain a unit root (H0: $\alpha = 0$ for all i) which is tested against the alternative hypothesis is that for each individual series at least one of the cross-section does not contain a unit root (H1: $\alpha 1 < 0$, for i = 1, 2, …, N1; H1: $\alpha 1 = 0$, for i = N1 + 1, N1 +2, …, N). The probability value of the estimated w-statistic for each of the series provides the result of stationarity with the rule of thumb being if the probability value, with respect to a particular series across all cross-sections, is below 10% level of significance, then the null hypothesis can be rejected implying the series to be stationary.

5.3.3. The Breitung Test

The Breitung (2000) test is referred to be second generation panel unit root test that studies the local power of the LLC and IPS test statistics and finds them to be very sensitive to the inclusion of the individual-specific trends. This is because the LLC and IPS tests employ a bias correction. The Breitung test statistic avoids the bias adjustment and has been found to have the capability that is greater than the LLC test, where the capability is the probability of rejecting a false null hypothesis. The Breitung test statistic is obtained going through similar steps to the LLC, till obtaining the residuals, where LLC uses Δyi, t-L and dmt both, the vector deterministic variables, but the Breitung test uses only the Δyi, t-L excluding the dmt. Similarly to the LLC test, the Breitung test assumes that all the panels in the paper have a common AR parameter. The null hypothesis is that each of the series is non-stationary (Ho: $\alpha = 0$ for all i) which is tested against an alternative hypothesis is that each of the series is stationary (H1: $\alpha \neq 0$ for all i). The probability value of the estimated t-statistic for each of the series provides the result of stationarity with the rule of thumb being if the probability value, with respect to a particular series across all cross-sections, is below 10% level of significance, then the null hypothesis can be rejected implying the series to be stationary.

5.3.4. Maddala and Wu Test

The Maddala and Wu (1999) panel unit root test, a first generation non-stationarity test, is actually a Fisher-type test combining the probability values from unit root tests for each cross-section in the panel. In similarity to the IPS test, the heterogeneity of the unit root process is considered in this test. This can be shown using the following equation:

$$P = -2 \sum_{i=1}^{N} ln\, p_i \quad (5)$$

where p_iis the probability value from any individual unit root test for any cross-section and P is distributed as Chi-square with 2N degrees of freedom where N is the total number of cross-sections considered in the panel. The probability values are obtained from the estimated Augmented Dickey-Fuller (ADF)-Fisher and the Phillips-Perron (PP)-Fisher Chi-square test statistics. The null hypothesis is that each individual series of the panel cross-sections contain a unit root (H0: pi = 1 for all i) which is tested against the alternative hypothesis is that for each individual series at least one of the cross-section does not contain a unit root (H1: pi < 1). The probability values of the estimated ADF-Fisher Chi-square and PP-Fisher Chi-square statistics for each of the series provide the result of stationarity with the rule of thumb being if the probability value, with respect to a particular series across all cross-sections, is below 10% level of significance, then the null hypothesis can be rejected implying the series to be stationary.[12]

5.3.5. Hadri Test

Unlike the aforementioned panel unit root tests, the Hadri (2000) test is based on the null hypothesis of stationarity. The test is an extension of the stationarity test developed by Kwiatkowski et al. (1992) in the context of time series study. The test is a first generation panel unit root test and considers a residual-based Lagrange multiplier test for the null hypothesis that the individual series are stationary around a deterministic level or around a deterministic trend (H0: σit2=0, for all i), tested against an alternative hypothesis of the presence of a unit root in the panel data of each series (H1: σit2>0, for all i). The probability values from the estimated Hadri z-statistic and the estimated Heteroskedasticity consistent Hadri z-statistic are considered to draw conclusions on the stationarity of all the series considered. If the probability values are more than 10%, meaning that the null hypothesis cannot be rejected at the conventional 10% level of significance, it implies the presence of stationarity in the panel data. The panel unit root tests were followed by the fixed effects panel estimation techniques to estimate the elasticities of the explanatory variables with respect to the dependent variable CA balance.

5.4. Panel Cointegration analyses

Analysis of cointegration among the concerned variables in the econometric model is pertinent to unearth the possible long-run associations between them. Statistical support to the existing of cointegrating equations implies that the variables are associated and move together in the long run. This paper considers two types of panel data cointegration approaches:

5.4.1. Pedroni residual-based Cointegration Test

The Pedroni (2004) residual-based test of cointegration employs the Engle-Granger two-step cointegration tests (Engle and Granger 1987) that examine the residuals of a spurious regression performed using variables that are found to be stationary at the first differences, I(1). It uses seven test statistics that are tested for the null hypothesis of no cointegration against the alternative hypothesis of cointegration for panels in which the estimated slope coefficients are permitted to vary across individual cross-sections of the panels. Thus, these statistics allow for the heterogeneous fixed effects and deterministic trends and also for heterogeneous short-run dynamics. In the context of a panel of N countries, M number of regressors (Xm) across T time period, the Pedroni test considers the following regression model:

$$y_{it} = \alpha_i + \gamma_{it} + \sum_{m=1}^{M} \beta X_{m,it} + \epsilon_{it}, \qquad for\ t = 1, \dots, T; i = 1, \dots, N; m = 1, \dots, M \tag{6}$$

where the variables yit and Xm,it are assumed to be integrated of the same order I(d), for each cross-sectional unit i in the panel. The parameters αi, γ_iand βm,i account for heterogeneous fixed effects, deterministic trends and heterogeneous slope coefficients respectively. Eit are estimated residuals indicating deviations from the long-run relationship. In order to carry out the cointegration test, Pedroni conducts unit root tests on the residuals as follows:

$$\boldsymbol{\epsilon_{it} = \gamma_i \varepsilon_{i,t-1} + \omega_{it}} \tag{7}$$

[12] Maddala and Wu (1999) find that for high values of T and N the Maddala and Wu-Fisher-test is chosen over the IPS test as size distortions are smaller at comparable power. For smaller values of T and N, however, IPS and LLC seem to be preferable over Maddala and Wu-Fisher-tests.

The tests are classified into two categories. The first set of tests is the panel cointegration based on the within-dimension approach which contains eight panel statistics (v-statistic, ϱ-statistic, ADF-statistic, PP-statistic and the weighted statistics of these four panel statistics) that pool the AR coefficients across different cross-sections for the unit root tests on the estimated residuals. Accordingly, these panel statistics are tested for the null hypothesis of no cointegration (H0: γi=1 for all i) against the alternative hypothesis of cointegration in the panel (H1: γi= γ<1 for all i), which assumes a homogeneous γ across all cross-sections. If the null hypothesis is rejected in these panel statistics case then the variables are said to be cointegrated for all the cross-sections in the panel (Ramirez, 2006).[13]

The second set of tests is the group cointegration tests based on a between-dimension approach that includes three group panel statistics (ϱ-statistic, ADF-statistic and PP-statistic). These statistics simply average the individually estimated coefficients for each cross-section, i. For the between-dimension approach, the null hypothesis of no cointegration (H0: γi=1 for all i) is tested against the alternative hypothesis (H1: γi= γ<1 for all i), which allows for heterogeneity in the AR coefficients. If the null hypothesis is rejected in these group panel statistics case then the variables are said to be cointegrated for at least one cross-section in the panel (Ramirez, 2006).

5.4.2. The Johansen Fisher Panel Cointegration Test

The Johansen Fisher panel cointegration test proposed by Maddala and Wu (1999) is a panel version of the individual Johansen (1998) cointegration test. The Johansen (1998) procedure is known to provide a unified framework for estimation and testing of cointegration relations in the context of VAR error correction models. It basically tells us whether or not the variables are associated in the long run. This paper estimates an Unrestricted Vector of Autocorrelation of the following form for this purpose:

$$\Delta x_t = \alpha + \theta_1 \Delta x_{t-1} + \theta_2 \Delta x_{t-2} + \theta_3 \Delta x_{t-3} + \cdots\cdots + \theta_{k-1} \Delta x_{t-k+1} + \theta_k \Delta x_{t-k} + u_t \tag{8}$$

where Δ is the difference operator; xt is a (n x 1) vector of non-stationary variables (in levels); and Ut is the (n x 1) vector of random errors. The matrix θk contains the information on the long-run relationship between variables, for instance, if the rank of θk = 0, the variables are not cointegrated. On the other hand if rank (usually denoted by r) is equal to 1, there exists one cointegrating vector and finally if 1 < r < n, there are multiple cointegrating vectors. Johansen (1998) derive two tests for cointegration, namely the trace test and the maximum Eigenvalue test. The trace statistic test evaluates the null hypothesis that there are at most r cointegrating vectors whereas the maximal Eigenvalue test, evaluates the null hypothesis that there are exactly r cointegrating vectors in xt. According to cointegration analysis, when two variables are cointegrated then there exists at least one direction of causality.

Johansen (1998) suggests a method for both determining how many cointegrating vectors there are and also estimating all the distinct relationships. Thus, it can be viewed as a multivariate generalization of the Dickey-Fuller test. Similarly, the Johansen Fisher panel cointegration test is founded on the same principles underpinning the Fisher ADF panel unit root test and aggregates the probability values of individual Johansen maximum eigenvalue and trace statistics. If pi is the probability value of from an individual cointegration test for cross-section i, under the null hypothesis of no cointegration then the test statistic for the panel is given by:

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[13] As a rule of thumb, if the majority of the eleven test statistics considered in the Pedroni test can be used to reject the null at 10% level of significance then it is said that the variables considered in the panel study are cointegrated in the long run and vice-versa.

$$-2\sum_{i=1}^{N} ln\, p_i \sim \times X^2_{2N} \tag{9}$$

5.5. Feasible Generalized Least Squares (FGLS) Regressions

The FGLS panel data estimator developed by Parks (1967) was chosen for regression purposes in this paper following the statistical certification of heteroscedasticity problems in the model whereby estimated coefficients, without accounting for the heteroscedasticity issue within the panel data set, lacks efficiency and acceptability. Apart from dealing with the problem of heteroscedasticity, the FGLS estimator can also generate efficient coefficients with robust standard errors (Reed and Webb 2011) in the presence of autocorrelation and cross-sectional dependency within the panel data set. For robustness check of the estimated coefficients, the random effects panel regression tools with both maximum likelihood and generalized least squares estimators (RE-MLE and RE-GLS) are applied. The regression analyses are followed by the short and long run causal analyses.

5.6. Panel Causality Analyses

From the point of view of policy implications, it is pertinent to investigate the causal dynamics of the concerned variables which would provide more sense to the findings from the coefficient estimates from the regression analyses.

5.6.1. Panel Vector-Error Correction Model approach to Short Run Causality

Panel Vector Error-Correction Model (VECM) is applicable only in the context of cointegration within the panel data set. It is a basically restricted vector autoregressive model structured to employ stationary as well as non-stationary series that are known to be cointegrated. It is restricted in the sense that the VECM has cointegrating relations built into the specification so that it restricts the long-run behavior of the endogenous variables to converge to their cointegrating relationships while allowing for short-run adjustment dynamics. The cointegration term is known as the Error Correction Term (ECT) which provides the pace at which any deviation from the long-run equilibrium in the previous lag is corrected in the next lag through a series of partial short-run adjustments. This is referred to as the Error Correction Mechanism (ECM).

Engle and Granger (1987) showed that a VECM is an appropriate method to model the long-run as well as short-run dynamics among the cointegrated variables. However, in the context of multivariate regression analysis, the VECM approach is preferred to provide only the short-run causality among the variables. Causality inferences in the multi-variate framework are made by estimating the parameters of the following VECM equations:

$$\Delta Y = \alpha + \sum_{i=1}^{m} \beta i \Delta Y_{t-i} + \sum_{j=1}^{n} \gamma_j \Delta X_{t-j} + \sum_{k=1}^{0} \delta \Delta M^s + \sum_{l=1}^{p} \zeta \Delta N + \theta Z_{t-1} + \varepsilon_t \tag{10}$$

$$\Delta X = a + \sum_{i=1}^{m} b_i \Delta Y + \sum_{j=1}^{n} c_j \Delta X_{t-j} + \sum_{k=1}^{0} d \Delta M^s + \sum_{l=1}^{p} e \Delta N + f Z_{t-1} + \xi_t \tag{11}$$

z t-1 is the error-correction term which is the lagged residual series of the cointegrating vector. The error-correction term measures the deviations of the series from the long run equilibrium relation. For example, a null hypothesis of X not Granger-causing Y is rejected if the set of estimated coefficients on the lagged values of X is jointly significant. Furthermore, in those instances where X appears in the cointegrating relationship, the hypothesis is also supported if the coefficient of the lagged error-correction term is significant. Changes in an independent variable may be interpreted as representing the short run causal impact while the error-correction term provides the adjustment of Y and X toward their respective long-run equilibrium. Thus, the VECM representation allows us to differentiate between the short- and long-run dynamic relationships. The Chi-Square test statistic is used to determine the short run causalities between pairs of variables in the model. In the context of a panel of N countries, three regressors (X, Y and Z) across T time period, the panel VECM model can be given by:

$$\begin{bmatrix}\Delta X_{it}\\ \Delta Y_{it}\\ \Delta Z_{it}\end{bmatrix} = \begin{bmatrix}\omega_{1i}\\ \omega_{2i}\\ \omega_{3i}\end{bmatrix} + \sum_{k=1}^{q}\begin{bmatrix}\alpha_{11ik} & \alpha_{12ik} & \alpha_{13ik}\\ \alpha_{21ik} & \alpha_{22ik} & \alpha_{23ik}\\ \alpha_{31ik} & \alpha_{32ik} & \alpha_{33ik}\end{bmatrix}\begin{bmatrix}\Delta X_{it-k}\\ \Delta Y_{it-k}\\ \Delta Z_{it-k}\end{bmatrix} + \begin{bmatrix}\gamma_{1i}\\ \gamma_{2i}\\ \gamma_{3i}\end{bmatrix} ECT_{it-1} + \begin{bmatrix}\mu_{1it}\\ \mu_{2it}\\ \mu_{3it}\end{bmatrix} \quad (12)$$

where Δ denotes first difference transformation of the variables. Although the VECM approach can also be relied on to generate estimates of long-run causal associations between the variables considered in the empirical model, it does not provide pairwise causal estimates. Thus, this study limits the VECM approach to causality investigations in the short run only.

5.6.2. Dumitrescu and Hurlin Granger Non-Causality Test

In contrast to the simple Granger (1969) Granger long-run causality analysis, this paper considers the Dumitrescu and Hurlin Granger Non-Causality Test (DHGNCT) proposed by Dumitrescu and Hurlin (2012). The DHGNCT was chosen for its ability to provide statistical evidence of long-run pairwise causal relationships between variables taking the possibility of heterogenous panels existing in the panel data set. The striking difference behind the Granger (1969) causality test and the DHGNCT is that the former assumes the panels to be homogenous whereby the null hypothesis of a particular variable X Granger causing Y is valid for all cross-sections, while the latter presumes the null hypothesis for only a particular sub panel and not the entire panel data set. Moreover, the DHGNCT also takes into account the possible cross-sectional dependencies within the data. Moreover, Dumitrescu and Hurlin (2012) also mentioned in their paper that this long run causality test is efficient in the case of examinations incorporating small time periods and cross-sections which coincides with the case in this paper.

6. Results and Discussions

Table 3 reports the panel unit root test results. According to the estimates, it can be seen that all the variables included in the dataset are non-stationary at their respective levels but they do become stationary at their first differences. Thus, it can be concluded that all these variables are mean reverting whereby the possibility of regression analyses, being conducted with this data, being spurious is nullified. Moreover, since all the variables are stationary at the same order it provides empirical support to perform the Pedroni residual-based and Johansen-Fisher tests for cointegration in the panel data.

Table 3. Panel Unit Roots Test Results.

	Level, I(0)							
	Levin, Lin & Chu	**Im, Pesaran & Shin**	**Breitung**	**Maddala and Wu**		**Hadri**		**Decision on Stationarity**
Variables	**t-stat**	**W-stat.**	**t-stat.**	**Fisher Chi-Square Stat.**	**Fisher Chi-Square Stat.**	**Hadri Z-stat**	**Heter. Consistent Z-Stat.**	
lnRERt	1.354 (0.912)	2.786 (0.997)	2.879 (0.998)	1.243 (0.996)	1.644 (0.990)	4.482 (0.000)	4.421 (0.000)	Non-Stationary
lnODAt	-1.603 (0.605)	-1.193 (0.614)	-1.239 (0.108)	1.544 (0.725)	1.381 (0.719)	3.647 (0.000)	2.914 (0.002)	Non-Stationary

lnFDIt	2.995 (0.999)	2.973 (0.999)	3.463 (1.000)	2.811 (0.946)	26.743 (0.001)	6.130 (0.000)	4.623 (0.000)	Non-Stationary
lnREMITt	1.820 (0.996)	1.912 (0.972)	2.042 (0.979)	4.668 (0.792)	4.718 (0.787)	5.286 (0.000)	3.869 (0.000)	Non-Stationary
OPENt	-0.844 (0.199)	-0.786 (0.216)	-0.203 (0.420)	11.141 (0.194)	11.918 (0.155)	4.367 (0.000)	3.759 (0.000)	Non-Stationary
lnGOVt	2.422 (0.992)	2.304 (0.989)	2.518 (0.994)	9.248 (0.322)	4.794 (0.799)	4.692 (0.000)	4.229 (0.000)	Non-Stationary
lnTOTt	6.520 (1.000)	8.211 (1.000)	6.817 (1.000)	0.045 (1.000)	0.049 (1.000)	6.425 (0.000)	6.279 (0.000)	Non-Stationary
				1st difference, I(1).				
Variables	**Levin, Lin & Chu**	**Im, Presaran & Shin**	**Breitung**	**Maddala and Wu**		**Hadri**		**Decision on Stationarity**
	t-stat	**W-stat.**	**t-stat.**	**Fisher Chi-Square Stat.**	**Fisher Chi-Square Stat.**	**Hadri Z-stat**	**Heter. Consistent Z-Stat.**	
lnRERt	-7.740 (0.000)	-6.811 (0.000)	-5.778 (0.000)	51.244 (0.000)	50.970 (0.000)	-0.008 (0.503)	0.466 (0.321)	Stationary
lnODAt	-9.767 (0.000)	12.338 (0.000)	-4.758 (0.000)	107.60 (0.000)	232.10 (0.000)	2.370 (0.009)	6.032 (0.000)	Stationary
lnFDIt	-4.129 (0.000)	-6.766 (0.000)	-3.114 (0.001)	51.912 (0.000)	302.228 (0.000)	2.542 (0.000)	2.338 (0.010)	Stationary
lnREMITt	-9.386 (0.000)	-9.471 (0.000)	-6.215 (0.000)	74.283 (0.000)	77.651 (0.000)	-0.453 (0.675)	1.440 (0.075)	Stationary
OPENt	-9.642 (0.000)	-9.725 (0.000)	-4.247 (0.000)	78.856 (0.000)	163.12 (0.000)	3.500 (0.000)	3.103 (0.001)	Stationary
lnGOVt	-0.802 (0.210)	-4.993 (0.000)	2.585 (0.995)	49.000 (0.000)	212.13 (0.000)	-0.886 (0.812)	1.764 (0.039)	Stationary
lnTOTt	-3.568 (0.000)	-3.152 (0.001)	-1.097 (0.136)	25.156 (0.002)	36.286 (0.000)	1.211 (0.113)	3.483 (0.000)	Stationary

Notes: Considering trend and intercepts; Automatic maximum lag and lag length selections based on Schwarz Information Criteria (SIC); The probability values are provided within the parentheses.

The results from the Pedroni residual-based and the Johansen-Fisher cointegration analyses are presented in tables 4 and 5 respectively. The results from both these cointegration analyses provide statistical evidence of long-run cointegration between the variables considered in the econometric models in this paper. This implies that all these variables move together in the long run. Hence, it can be stated that RER and all the three sources of foreign currency influx are associated in the long run. The results are similar to the findings by Amin and Murshed (2018b) in the context of Bangladesh.

Table 4 Pedroni Residual-Based Panel Cointegration Test

Within Dimension	**Statistic**	**Weighted Statistic**
Panel v-statistic	-1.230 (0.891)	-1.296 (0.903)
Panel ρ-statistic	1.308 (0.905)	1.539 (0.938)
Panel PP-statistic	-0.202 (0.420)	0.284 (0.612)
Panel ADF-statistic	-0.251 (0.401)	0.097 (0.539)
Between Dimension	**Statistic**	

Group ϱ-statistic	2.308 (0.990)
Group PP-statistic	0.607 (0.728)
Group ADF-statistic	0.328 (0.628)

Note: Null Hypothesis: No cointegration. Trend Assumption: No deterministic trend. Authentic lag length selection based on Schwarz Information Criteria (SIC); The probability values are provided within the parentheses.

Table 5: Johansen Fisher Panel Cointegration Test.

Unrestricted Cointegration Rank Test (Trace and Maximum Eigenvalue)				
Hypothesized No. of CE(s)	**Fisher Statistic (from trace test)**	**Probability**	**Fisher Statistic (from max-Eigen test)**	**Probability**
None	158.2	0.0000	87.02	0.0000
At most 1	96.00	0.0000	36.50	0.0000
At most 2	64.69	0.0000	22.89	0.0035
At most 3	46.80	0.0000	20.88	0.0075
At most 4	31.14	0.0001	18.40	0.0184
At most 5	20.19	0.0096	15.67	0.0473
At most 6	17.60	0.0244	17.60	0.0244

Note: Trend assumption: Intercept (no trend) in CE and VAR. Lags interval (in first differences)

The unit root and cointegration analyses are followed by the regression analyses to estimate the long run coefficients of the empirical model (2) explained in section 4. The regression results, reported in table 6, suggest that inflow of ODA and REMIT into the four South Asian economies tend to appreciate the RER as depicted from the negative and statistically significant coefficients attached to lnODAt and lnREMITt. The corresponding FGLS estimates, showing the expected signs, reveal that a 1% rise in the influx of ODA and REMIT decreases the RER on average by 0.18% and 0.23% respectively, ceteris paribus. A plausible explanation behind REMIT inflows exerting a relatively greater appreciative pressure on the RER can be provided by the fact that the volumes of the international REMIT flowing into the four South Asian economies overwhelmingly outweigh the corresponding ODA inflows. Moreover, the fact that ODA received at times have repayment obligations attached, unlike the REMIT which tends to generate a larger spending effect resulting in a comparatively greater appreciation of the RER.

In contrast, the results also imply that inward FDIs result in depreciation of the RER as perceived from the positive and statistically significant coefficient attached to lnFDIt. The result from the FGLS estimator shows that a 1% rise in the FDI receipts triggers a corresponding rise in the RER on average by 0.19%, ceteris paribus. This imposes a key policy implication in the sense that attracting large amounts of FDI can be a solution to the potential RER appreciation problems, following receipts of ODA and REMIT, across the four South Asian economies since the RER depreciation in the context of the FDI inflows can neutralize the RER appreciative pressures sourced from the other two sources of foreign exchange. Moreover, the FDIs should be aimed at the development of the tradable sector in order to negate the hypothesized resource movement effect and therefore avoid the deindustrialization effect discussed in the DD relevant literature. FDI receipts in the context of the tradable sector can be ideal in avoiding the foreign inflow-induced DD problems by effectively expanding the tradable sector with the South Asian economies considered in this paper.

Other results also highlight the importance of public investments in countering the RER appreciation caused by the influx of ODA and REMIT. The positively statistically significant estimated coefficient attached to lnGOVt predicts that a 1% rise in investments made by the government, on average, can attribute to 0.19% depreciation in the RER ceteris paribus. This particular finding, although suggesting an expansion of the non-tradable sector due to public investments usually fostering development of the non-tradable sector, is critically important from the perspective of the RER management in the sense that public investments are likely to develop the infrastructures within the South Asian economies which, in turn, would facilitate the attraction of FDI to neutralize the RER appreciations stemming from the incoming ODA and international REMIT. Furthermore, the FGLS estimates also denote in favor of open relatively more economies being less likely to encounter RER appreciations which is evident from the positive and statistically significant estimated coefficient attached to OPENt. This result tends to justify the export-led growth strategies pursued by a majority of the four South Asian economies considered in this paper.

The results found in this paper are robust across different panel data regression techniques. As seen from table 6, the estimates from the RE-MLE and RE-GLS estimation techniques depict similar signs and statistical significances as the FGLS estimates. However, from the perspective to further policy implications, it is also important to analyze the causal impacts of the foreign exchange inflows on the RER. Thus, the short and long run causality tests are performed.

Table 6. The Panel Data Regression analysis results

Dependent Variable: lnRERt			
Estimator	**FGLS**	**RE-MLE**	**RE-GLS**
Regressors			
lnODAt	-0.179*	-0.057***	-0.225**
	(0.072)	(0.024)	(0.081)
lnFDIt	0.185*	0.085*	0.130*
	(0.026)	(0.014)	(0.030)
lnREMITt	-0.232*	-0.165*	-0.207*
	(0.025)	(0.025)	(0.048)
lnGOVt	0.018*	0.034***	0.018*
	(0.004)	(0.017)	(0.005)
lnTOTt	-0.003	-0.012	-0.025
	(0.022)	(0.022)	(0.025)
lnOPENt	0.448*	0.115**	0.395*
	(0.090)	(0.057)	(0.107)
Intercept	2.739*	3.410*	3.536*
	(0.663)	(0.778)	(0.804)
Country Fixed Effects	Yes	Yes	Yes
Adj. R2			0.754
Wald chi2	344.95*		184.08*
Log Likelihood		53.745	
LR chi2		62.40*	
No. of observations	152	152	152

*Note: *, ** & *** denote statistical significance at 1%, 5% & 10% levels, respectively; The robust standard error are reported within the parentheses.*

The short-run causal estimates from the VECM analysis is reported in table 7. It is clear from the statistical evidence found that there is no short-run causal association between RER and the three sources of foreign exchanges flowing into the four South Asian nations. This implies that the movements in the

RER are not influenced by the incoming ODA, FDI and international REMIT into these countries which tend to nullify the possibility of these economies experiencing the DD phenomenon in the short run. It is also seen that RER causally influences TOT movements across the selected South Asian nations without the feedback. This contradicts the findings by Murshed (2018b) in the context of South Asian and Southeast Asian economies. The error-correction term is found to be negative and statistically significant as well and the predicted value conveys that any distortion from the equilibrium in the current period is corrected at a rate of 6.1% in the next period. Since there is lack of causality running from either direction in the context of RER and the foreign currency inflows in the short run, and also due to the statistical evidence of heteroscedasticity problem found to exist in the data set, the DHGNCT is tapped to generate the possible long-run causal linkages.

Table 7. The VECM short-run causality test results

Sources of Causation								
Dependent Variable	**Short run**							**Long Run**
	lnRERt	**lnODAt**	**lnFDIt**	**lnREMITt**	**lnGOVt**	**lnTOTt**	**lnOPENt**	**ECT**
lnRERt	-	0.433	0.064	0.018	1.393	0.388	1.774	-0.061
		(0.805)	(0.968)	(0.991)	(0.498)	(0.824)	(0.412)	(0.034)
lnODAt	1.472	-	1.096	8.555	3.402	5.555	6.661	-0.118
	(0.479)		(0.580)	(0.014)	(0.813)	(0.062)	(0.036)	(0.300)
lnFDIt	1.290	1.814	-	16.754	15.683	8.901	1.289	0.016
	(0.525)	(0.404)		(0.000)	(0.000)	(0.012)	(0.525)	(0.601)
lnREMITt	1.265	0.384	19.561	-	59.442	14.629	4.529	-0.152
	(0.531)	(0.825)	(0.001)		(0.000)	(0.001)	(0.104)	(0.053)
lnGOVt	0.635	1.275	10.941	20.305	-	12.663	2.536	-0.107
	(0.728)	(0.529)	(0.004)	(0.000)		(0.002)	(0.282)	(0.057)
lnTOTt	6.060	3.661	3.512	1.294	19.098	-	5.919	-0.406
	(0.048)	(0.160)	(0.173)	(0.523)	(0.000)		(0.052)	(0.001)
lnOPENt	0.104	2.099	1.287	1.071	6.710	2.313	-	-0.155
	(0.949)	(0.350)	(0.525)	(0.586)	(0.035)	(0.314)		(0.008)

Note: The probability values are provided within the parentheses; The optimal lags are selected using the Schwarz Information Criterion (SIC).

The results from the DHGNCT are illustrated in table 8. The estimated results advocate in favor of RER movements being triggered by ODA and FDI influx into the South Asian economies which is confirmed by the statistical evidence of unidirectional causalities, without the feedback effect, running from ODA and FDI to RER. Thus, these findings can be used to certify the regression findings reported in table 4 in conclude that inflows of ODA and FDI into the four selected South Asian economies do trigger appreciation and depreciation, respectively, of the RER. Therefore, as far as the policy implications are concerned, it is pertinent for these nations to slowly ease out their reliance on ODA inflows and rather attract greater amounts of FDIs, primarily aimed at the tradable sectors. A plausible explanation behind FDI inflows not leading to DD problem can be interpreted from the correlative association between these variables found in the aforementioned regression analysis, whereby inward FDI were found to rather depreciate the RER. This finding is of significant importance in the context of China's 'One Belt and Road Initiative' under which the nation is expected to handsomely invest in development projects across South Asia. Hence, in line with the findings mentioned earlier, it can be asserted that such foreign investments by China would not exert appreciative pressures on the RERs of the South Asian economies.

Another key result from the DHGNCT shows that there is bidirectional causality between RER and REMIT which implies that the international REMIT not only appreciates the RER in the long run, but this appreciation also spawns greater volumes of REMIT for these South Asian nations. The recent growth in global inflows of REMIT into the developing economies, in particular, have often been questioned the efficacy of such forms foreign exchange receipts in the attainment of macroeconomic and social development of the recipient economies. Although the regression results advocated in favor of international REMIT stimulating appreciative pressures on the RER, a possible solution to this problem could be through channeling of these foreign exchanges for investment into the tradable sector which, to a large extent, would negate the spending and the resource movement effects linked to foreign REMIT inflows within the concerned economies.

Apart from the three sources of foreign exchange inflows, the RER movements, in the long run, are also found to be influenced by government expenditures in these countries. The relevant long-run causal estimates validate the presence of a unidirectional causal chain stemming from government expenditure to RER movements which further strengthens the regression corresponding findings in order to conclude in favor of such public investments initiating depreciation of the RER within the South Asian nations. Thus, it is advisable to the concerned governments to scale up their respective public expenditure budgets, precisely developing the indigenous infrastructures to facilitate the expansion of both bilateral and multilateral trade volumes across the respective economies.

Table 8. The Dumitrescu & Hurlin (2012) Granger Non Causality Test results

Dependent Variable	**Independent Variables**						
	lnRERt	**lnODAt**	**lnFDIt**	**lnREMITt**	**lnGOVt**	**lnTOTt**	**lnOPENt**
lnRERt	-	3.750	3.623	3.488	2.016	0.540	1.470
		(0.000)	(0.000)	(0.001)	(0.043)	(0.589)	(0.142)
lnODAt	0.514	-	-0.466	0.612	5.172	3.015	3.571
	(0.607)		(0.642)	(0.541)	(0.000)	(0.003)	(0.000)
lnFDIt	-0.031	-0.088	-	8.451	5.172	3.015	3.571
	(0.975)	(0.930)		(0.000)	(0.000)	(0.003)	(0.000)
lnREMITt	6.697	4.129	3.914	-	4.116	0.271	5.554
	(0.000)	(0.000)	(0.001)		(0.000)	(0.311)	(0.000)
lnGOVt	1.893	1.361	1.504	3.500	-	2.952	6.175
	(0.304)	(0.174)	(0.133)	(0.001)		(0.003)	(5.298)
lnTOTt	0.356	1.291	0.178	0.349	0.406	-	1.119
	(0.721)	(0.211)	(0.112)	(0.244)	(0.149)		(0.154)
lnOPENt	3.786*	-0.302	3.293	0.601	6.837	0.319	-
	(0.002)	(0.763)	(0.001)	(0.549)	(0.000)	(0.123)	

Note: Since the number of cross sections is smaller than the time period the Z-bar statistics are reported; The probability values are provided within the parentheses; The optimal lags are selected using the Schwarz Information Criterion (SIC).

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Appendix

Table A1. Results from Cross-sectional Dependency analysis.

Tests	Statistic
Pesaran's Test of Cross-sectional independence	0.207 (0.834)
Friedman's Test of Cross-sectional independence	1.129 (0.786)
Frees' Test of Cross-sectional independence	0.028 (0.129)
Breusch-Pagan Lagrange Multiplier [b]	1.161 (0.106)

Note: a denotes the null hypothesis of cross-sections being independent; b denotes the null hypothesis of no cross-sectional dependence in residuals; The probability values are reported within the parentheses.

Table A2. Results from the Panel Group wise Heteroscedasticity tests.

Tests	Statistic
Lagrange Multiplier	50.538 (0.000)
Likelihood Ratio	58.719 (0.000)
Wald	265.038 (0.000)

Note: Null Hypothesis: Homoscedasticity; Alternate Hypothesis: Group wise heteroscedasticity; The probability values are reported within the parentheses.