DO DIFFERENT TYPES OF CAPITAL INFLOWS HAVE DIFFERENTIAL IMPACT ON OUTPUT?

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DO DIFFERENT TYPES OF CAPITAL INFLOWS HAVE DIFFERENTIAL IMPACT ON OUTPUT?

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Abstract

Emerging market economies have experienced an unprecedented rise in cross-border capital flows and the existing literature provides us evidence of both expansionary and contractionary effects of inflows on domestic output. In this context, we make an attempt to answer the following questions: (1) Do capital inflows lead to expansionary or contractionary effect on emerging countries? (2) Do different types of capital inflows have different impacts? and (3) Do absorptive capacities influence the effect of capital flows on the host countries? To answer these questions, we carry out a comparative analysis for India and China using quarterly data for the period 1998Q1 to 2020Q1. The results reveal that total gross capital inflows as well as disaggregated capital inflows exhibit expansionary effect on domestic output in case of both India and China. We supplement the time series data with panel analysis for the top ten capital flows recipient EMEs over the period 1998-2019. We find that capital inflows at aggregate level and also at the disaggregate level except debt flows have an expansionary effect on output.

Keywords: Gross capital inflows, FDI, Emerging economies, Structural breaks

JEL: F21; F32; F43; C22; C23

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1. Introduction

This article tries to examine whether gross capital inflows lead to expansionary or contractionary effect on domestic output. Our study is motivated by the increased global financial integration in the past three decades which has seen an unprecedented rise in gross capital inflows in emerging market economies (EMEs). Many EMEs (borrower) have benefited from this transition as these capital-scarce economies achieved a higher economic growth trajectory by borrowing at low cost from abroad, to finance their investment needs. In comparison, the developed (lender) economies benefited with easy access to larger markets, cheap labor, higher returns, and diversified country-specific risks. Thus, cross-border capital flows benefit both the lender and the borrower country since free flow of global capital leads to more efficient allocation of the global saving to the most productive uses.

Obstfeld (1994, 2012), Blanchard et al. (2017) and Koepke (2018) argues that there are welfare gains from capital flows. On the production side, cross-border capital flows will enable movement of resources from low returns countries to high return countries, leading to a more productive allocation of global saving. For an EME, free access to international capital markets would mean availability of investment at low cost of borrowing, which reduces marginal cost of capital, accelerates investment and growth. Further, free capital flows would facilitate portfolio diversification by investors in developed as well as EMEs alike. On the consumption side, capital flows can result in a Pareto superior consumption path by allowing countries to smooth their consumption to an optimal level, i.e. capital flows enable countries to intertemporally trade its present consumption for future consumption through international capital lending and borrowing. Thus, capital flows can result in welfare gains for both the providers of capital and its recipient (Koepke, 2018). However, increased financial integration usually transfers risks too through its ill-effects on economic growth. Several studies have argued that capital flows may entail potential costs in terms of financial stability in the recipient country and risks associated with capital flows reversals, leading to a reduction in welfare[3] (Razin et al. 1999; Rodrik and Velasco, 1999; Ibarra 2011a,b; Saborowski, 2011; Ghosh et al. 2016; Obstfeld and Taylor, 2017). International capital

[3] Prasad et al. (2003), Edison et al. (2004) and Henry (2006) provides an excellent survey on the costs and benefits of capital flows.

flows expose a country to external shocks and can be a source of inherent risk for developing economies (Garg and Prabheesh, 2018, 2021).

Hence, capital inflows can lead to either expansionary or contractionary effect on the recipient economy's output. One strand of explanation can be found in the Mundell-Fleming model under which if an economy follows a flexible exchange rate regime then, for a given monetary policy rate, capital inflows will lead to an appreciation of the domestic currency, a contraction in net exports and domestic output. On the contrary, an increase in capital outflow will lead to a depreciation of the domestic currency, making the domestic goods more competitive. In this case, net exports increases and the effect on domestic output is expansionary in nature (Blanchard, 2017). Krugman (2014) also offers important insights into this debate why capital inflows can sometimes be expansionary. The study offers explanation by analyzing the effect of sudden stops in capital flows, i.e., capital outflows, on domestic country's output. Their study emphasized on two important factors that could affect the above relationship between capital flows and domestic output. First, whether the country has its own currency, unlike Greece and other Eurozone economies. Second, whether the currency regime is fixed exchange rate or floating exchange rate. He argues that for a floating exchange rate regime a sudden capital outflow will lead to depreciation of the domestic currency and at any given interest rate, net exports will increase. However, central bankers will try to reduce the inflationary pressures as a result of currency depreciation, i.e. leaning against the wind. Hence, a monetary tightening will increase the interest rates and reduce demand for domestic goods. Thus, if the increase in interest rates is more than required to offset the inflationary concerns, then there will be a reduction in domestic output. Symmetrically, capital inflows can be expansionary, only if the interest rate is decreased sufficiently.

Similarly, policymakers in the EMEs believes that capital inflows may lead to easy availability of credit, creating a credit boom and a rise in economic activity. Thus, capital inflows can be expansionary which can only be compensated through an increase in interest rates. However, the above channel leads to a policy dilemma wherein the use of policy rate increase to limit the rise in output is offset by the large capital inflows which, in turn, leads to even higher output. Blanchard et al. (2017) tries to resolve this dilemma by making a distinction between types of capital inflows. Their study argue that capital inflows may lead to a fall in the rate on non-bond flows, thereby declining interest rates and an increase in domestic demand. This would offset the effect of currency appreciation on export demand. Thus, capital inflows need not be expansionary and its macroeconomic effects on output depends on their nature of the flows (Henry, 2007). For EMEs policymakers, capital flows that are more volatile in nature such as short-term capital flows and portfolio equity flows have been a point of concern. The increase in capital inflows to EMEs is seen with a concomitant rise in unfolding of financial crises since the 1990s (East Asian crisis of 1997, Mexican crisis of 1994, Brazilian crisis of 1999, the Global Financial Crisis of 2008). These crises exhibited that volatile capital flows not only are a source of fragility to the financial sector but also significantly affect the economic activity of an economy. With regards to this, a distinction should be drawn between different types of capital flows for a careful empirical analysis.

Thus, the present study examines whether gross capital inflows lead to an expansionary effect or a contractionary effect on domestic output in presence of financial development. Further, we investigate if there is any aggregation bias in using total gross capital inflows as compared to disaggregated capital inflows. Our empirical approach towards testing the above hypotheses is as follows. First, we conduct two analyses; one with a time-series comparison of India and China and then a panel analysis of selected EMEs that have received substantial capital inflows since the GFC of 2008. Second, we specify three model specification wherein we consider both aggregated capital inflows and disaggregated capital inflows. All these models are tested with financial development as a control variable. Our empirical findings conclude as follows. First, we find the period of GFC as the period when there are structural breaks in capital inflows in India and China. Second, we find capital inflows lead to expansionary effect on domestic output in both India and China but there is no aggregation bias. Third, we find contrasting results with regards to aggregation bias in our panel analysis. While we find that capital inflows such as FDI, FPI, and equity inflows have an expansionary effect, the debt inflows lead to a contractionary impact on domestic output. Of note, our finding suggests aggregation bias since the equity and debt inflows exhibit opposite impact whereas total inflows have an expansionary impact on domestic output. Overall, our findings from panel estimation implies that distinction between types of inflows are necessary since the policy recommendations based on aggregate capital inflows may be ill-advised.

Hence, our study departs from the existing literature on several counts. First, most of the studies net capital inflows as a driver of domestic output growth. However, we consider gross capital inflows since net inflows may give a misleading picture of how much a country is borrowing capital from the rest of the world. Second, previous studies have used capital inflows either at the aggregated level or disaggregated level. We utilize both in order to provide insights on aggregation bias in using total capital inflows. Finally, while there is an abundant literature on the linkages between FDI and domestic output growth with the help of financial development indicators, the literature on the impact of portfolio inflows, equity inflows, and debt inflows in presence of financial development is scanty. We contribute to the present literature by filling these gaps. The rest of the paper is structured as follows: Section 2 outlines the empirical model and presents the data utilized. Section 3 discusses the empirical results. Section 4 concludes with policy implications.

2. Review of Literature

The empirical literature on capital inflows and its impact on output growth of the recipient country has been mixed at best. While some argue that foreign capital inflows contribute to economic growth in developing countries, others hold the view that not all capital flows are beneficial; some types of capital inflows can harm more, especially in the context of developing economies, and culminate in a financial crisis (Rodrik, 1998; Rodrik and Velasco, 1999; Eichengreen & Leblang, 2003; Baharumshah et al. 2015).

Most of the previous work is focused on identifying the linkages between FDI inflows and economic growth. One line of research supports the neoclassical theory that FDI inflows offer a dynamic vehicle to achieve higher growth. The reasons for this claim are made by many studies. A number of studies have argued that FDI enhances capital formation and thus results in employment growth, and promote manufacturing exports (Baharumshah & Thanoon, 2006; Balasubramanyam et al. 1996; Borensztein et al. 1998; Grossman & Helpman, 1991; Hansen & Rand, 2006). Especially in the context of Asian countries, where exports played a vital role in stimulating the growth, FDI has led to increase in manufacturing exports (Feder, 1992; Rodriguez & Rodrik, 1999). Other studies, including Baharumshah and Thanoon (2006) Markusen & Venables (1999), Wei (1995), Zhang (2001a), Sahoo (2005), Sahoo et al. (2011), among others, have discussed the importance of FDI in technology and spillover effects.

On the contrary, some studies found that FDI inflows are unstable and can lead to destabilizing effects on growth of a country. Some studies even find no significant relationships between the two. For instance, Rand and Trap (2002) found that there exists no significant relationship between FDI and output growth. They further argued that FDI inflows are very volatile and can have destabilizing effects leading to business cycle fluctuations. Similarly, Razin et al. (1999) highlighted the importance of signaling wherein a bad signal, in presence of a sound domestic credit market, may result in a welfare decline of the recipient country. Herzer et al. (2008) and Alvarado et al. (2017) looked at the relationship between FDI and economic growth in case of developing and Latin-American countries, respectively. They found that FDI does not have any short-run or long-run impact on growth. Similar results were found by Carbonell and Werner (2018) for Spain. In fact, the growth impact of FDI inflows is not related to level of per capita, financial market development, degree of openness, level of education.

The empirical literature on FDI-growth nexus can further be classified in terms of data and econometric methodologies[4]. A large number of studies are cross-country and panel studies. Some of the cross-country studies include Blomstrom et al. (1994), Balasubramanyam et al. (1996), Borensztein et al. (1998), Alfaro et al. (2004). However, cross-country studies are not reliable in terms of policy implications since they assume identical production technologies and institutions while these differ and likely to be heterogeneous in nature. Under these circumstances, results from regression may be unreliable and sensitive to the choice of countries selected (Erricson et al. 2001; Carkovic & Levine, 2005).

As a remedy to this issue, panel estimation takes care of the unobserved country-specific effects and endogeneity bias (Herzer et al. 2008). Hence, a number of panel studies were conducted in this regard. Carkovic and Levine (2005) implemented the GMM dynamic panel model for 68 countries over the period 1960-95. They controlled for endogeneity and omitted variables and found contrasting results to earlier work such as Blomstrom et al. (1994), Balasubramanyam et al. (1996), and Alfaro et al. (2004). They found that FDI does not positively impact economic growth. However, panel data estimators also suffer from the imposition of homogeneity on the coefficients of lagged dependent variables (Herzer et al. 2008). To ameliorate this issue, Nari-Reichert and Weinhold (2001) employed a mixed fixed and random coefficient approach and found that FDI

[4] Herzer et al. (2008) have provided an excellent survey on this topic.

has significant positive impacts on economic growth however the relationships are heterogeneous across the set of countries.

A number of studies also utilized panel cointegration techniques. For instance, Basu et al. (2003) tested a panel of 23 developing countries for the period 1978-96 and found a significant cointegrating relationship between FDI and growth. Likewise, Hansen and Rand (2006) tested for a sample of 31 developing countries over the period 1970-2000 and found similar results. Despite the improvement over cross-country studies, heterogeneity remains a serious concern in panel cointegration studies and may lead to false rejection of the null of no cointegration when within country relationships are not cointegrated (Guiterrez, 2003; Banerjee et al. 2004, 2005). Further, Strauss and Wohar (2004) argued that mixing of cointegrated and non-cointegrated relationships may lead to misleading results.

Several studies were conducted using time series techniques for investigating FDI and growth relationships. Zhang (2001b) looked at 11 developing countries over the period 1970-95. He found long-run causality from FDI to growth in case of five countries, and one country out of the other six countries without cointegration exhibit short-run causality from FDI to growth. Cudros et al. (2004) utilized quarterly series between 1980 and 2000, and found cointegrating relationship for two out of the three Latin American countries. Some of the other time series studies include Ramirez (2000) for Mexico, Xiaohui et al. (2002) for China, Fedderke and Romm (2006) for South Africa, to name a few. Nonetheless, even the time series studies should be careful in identifying the econometric methods for investigation. For instance, system-based cointegration procedure such as in Johansen (1995) may overreject the null of no cointegration and have poor small sample properties (Reinsel and Ahn, 1992; Cheung and Lai, 1993; Herzer et al. 2008). In that case, single equation cointegration techniques or other robust procedures need to be utilized to account for small-sample bias.

While there have been numerous studies on the effect of FDI inflows on growth, the empirical literature on impact of non-FDI flows on growth is scanty. Investigating the effects of the different components of gross capital inflows is important since the empirical results may be more detailed and specific policy recommendation can be made. Although few, studies that have focused on non-FDI capital flows, such as portfolio equity and debt inflows, the evidence is mixed at best (Reisen and Soto, 2001; Soto, 2003; Laureti and Postiglione, 2005; Baharumshah and Thanoon, 2006; Klein and Olivei, 2008).

Reisen and Soto (2001) investigated the impact of private capital inflows on economic growth for 44 countries. They found that FDI and portfolio equity flows impacts positively whereas long-term and short-term bank lending negatively impacts on per capita income growth. Along the same lines, Soto (2003) explored the effects of various types of private capital inflows on economic growth in case of 72 countries. The study found some evidence that FDI, portfolio debt investment and bank lending leads to a positive impact while portfolio equity inflows has a negative effect on national income. Laureti and Postiglione (2005) examined the effect of different types of capital inflows on economic growth in 11 Mediterranean countries. They found that the Mediterranean countries weren't able to exploit the wave of capital flows, and thus did not lead to an improvement in the growth process. Baharumshah and Thanoon (2006) looked at 8 East Asian countries and explored the effect of different capital flows on the growth process. They found that FDI and domestic savings contribute to growth however FDI inflows are more productive than domestic investment, indicating positive spillover effects from FDI. The study also finds negative effect of short-term inflows on growth. Klein and Olivei (2008) utilized a sample of 21 OECD and 74 non-OECD countries to investigate the effect of libealized capital account on financial depth and economic growth. Their overall findings indicate that a freer capital account can have a significant positive impact on financial depth which, in turn, exerts a positive effect on economic growth. However, the benefits from KA liberalization are profoundly significant in OECD countries as compared to developing countries. One reason could be the effect of policy change and ability to reap the benefits of more liberalized capital accounts in the presence of adequate institutions and macroeconomic policies.

In the last two decades, studies have also identified the importance of absorptive capacity of the recipient country in utilizing the transfer of new technology and spillover efficiency through FDI flows (Reisen and Sotto, 2001; Balasubramanyam et al. 1996; Borensztein et al. 1998; Kohpaiboon, 2003; Hansen and Rand, 2006). For instance, Balasubramanyam et al. (1996) find that increased trade openness is crucial to attracting larger volumes of FDI, which acts as a vehicle of technology transfer, and acquiring the potential economic growth. Similarly, Borensztein et al.

(1998) in their research suggest that FDI inflows can lead to positive impact on economic growth only if a threshold level of human capital is achieved in the host country.

In another strand, the earlier works in McKinnon (1973) and Shaw (1973) postulate that an increased level of financial development, resulting from capital account liberalization, will lead to higher economic growth. However, until 1990s, studies did not focus on the role of finance in impacting the relationship between FDI inflows and economic growth. Earlier studies by Greenwood and Jovanovic (1990), King and Levine (1993a,b) and Pagano (1993) demonstrated that a higher level of financial development reduces frictions related to informational levels and improve resource allocation efficiency. Several studies suggest that a higher level of financial depth in the recipient country allows it to utilize FDI flows more efficiently (Hermes and Lensink 2003; Omran and Bolbol 2003; Alfaro et al. 2004; Durham 2004; Klein and Olivei, 2008; Ang 2009a, 2009b). Ang (2009a, 2009b) points out the different ways in which a higher level of financial development can enable FDI to achieve higher economic growth. First, FDI inflows will lead to credit expansion which in turn allows firms to invest in human capital and adopt better technology. Second, a sound financial system in the recipient country helps in absorbing the spillovers associated with FDI. In this way, better financial development of local markets improves the ability of the host country to reap the benefits from FDI.

By the above discussion, few points can be made about the existing empirical literature. First, distinction between different types of capital flows is important. Second, local absorptive capacities are an important factor for the recipient country in order to reap the benefits from capital inflows. Third, most of the studies have focused on developed countries. Fourth, there is a lack of time series study that is more appropriate when countries have different level of financial development and macroeconomic fundamentals.

3. **Empirical strategy**

Following Blanchard et al. (2017) and Koepke (2018) arguments in favor of gross inflows, we use aggregated and disaggregated gross inflows. Specification we estimate three model specifications:

Model I: $$PCI = f(TGCI) \tag{1}$$

Model II: $$PCI = f(GFDII, GFPII) \tag{2}$$

Model III: $$PCI = f(GEI, GDI) \tag{3}$$

Where PCI is per-capita income used as a proxy for domestic output, TGCI is total gross capital inflows, GFDII is gross foreign direct investment inflows, GFPII is gross foreign portfolio investment inflows, GEI is gross equity inflows, and GDI is gross debt inflows.

The model I in equation (1) takes per capita income as a function of total gross capital inflows. Since the results about the impact of TGCI on PCI could be affected with aggregation bias either due to the adding-up or cancelling-out effect, we decompose the total inflows in model II and model III. First, we decompose the TGCI into GFDII and GFPII to investigate whether FDI inflows or FPI inflows, which are broad indicators of long-term and short-term flows, lead to expansionary or contractionary impact on output. Then, we decompose the TGCI into equity inflows and debt inflows to identify if various types of capital flows have different impact on domestic output. Finally, we incorporate the role of financial development in improving absorptive capacity of the borrower country. To this end, we consider two measures of financial development, namely broad money to GDP and domestic credit to private sector as a ratio to GDP. These control variables are interacted with capital inflows in all the three model specifications.

With regards to analysis of the three model specifications, we consider both panel and time-series estimations. First, using quarterly data from 1998Q1 to 2020Q1, we conduct a comparative time-series analysis of India and China since they are the two largest EMEs and recipient of huge capital inflows since the GFC of 2008. Our chosen sample period is characterized by a period of liberalized capital flows. The sample covers the main global events such as the Bretton Woods II period where the global imbalances started widening in 1998 and peaked in 2006, the unfolding of GFC of 2008, and a decade of observation post-GFC crisis. For panel analysis, we utilize annual data from 1998-2019 and we select ten EMEs that have occurred consistently among the top recipient EMEs. These emerging countries are Brazil, Chile, China, India, Indonesia, Mexico, Poland, Russia, South Africa, and Thailand. The data for gross domestic output (GDP), gross total capital inflows (TCI), gross foreign direct investment inflows (FDI), gross foreign portfolio investment inflows (FPI), gross debt inflows (DEBT), and gross equity inflows (EQUITY) is extracted from IMF database. All data are converted into real terms and expressed in natural logarithms.

3.1 Time-series analysis

Unit root test

Before testing the relationships between different types of gross inflows and domestic output, we examine the stationarity properties of the macroeconomic variables. We apply standard unit root tests such as ADF, PP and KPSS. However, these tests are reliable if a time series has undergone structural change. Since our sample span covers various events such as GFC in 2008 or the taper tantrum in 2013, the time series are likely to contain structural breaks. Thus, we apply the Narayan and Popp (NP) (2010) endogenous structural breaks unit root test, which differs from other popular break tests such as Zivot and Andrews (1992), Lumsdaine and Papell (1997), and Lee and Strazicich (2003). The NP test is more accurate in detecting the break dates and invariant to the break magnitude (Narayan and Popp, 2013). As with standard two break tests, the NP test also allows for testing of two models, Model 1 and Model 2. Model 1 is the 'crash' model and assumes two in the intercept whereas Model 2 assumes two breaks both in the intercept and trend:

$$\begin{aligned} y_t^{M1} &= \rho y_{t-1} + \alpha_1 + \beta^* t + \theta_1 D(T_B')_{1,t} + \theta_2 D(T_B')_{2,t} + \delta_1 DU_{1,t-1}' \\ &\quad + \delta_2 DU_{2,t-1}' + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_t \end{aligned} \tag{4}$$

$$\begin{aligned} y_t^{M2} &= \rho y_{t-1} + \alpha^* + \beta^* t + \kappa_1 D(T_B')_{1,t} + \kappa_2 D(T_B')_{2,t} + \delta_1^* DU_{1,t-1}' \\ &\quad + \delta_2^* DU_{2,t-1}' + \gamma_1^* DT_{1,t-1}' + \gamma_2^* DT_{2,t-1}' \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_t \end{aligned} \tag{5}$$

with $DU_{1,t}' = 1(t > T_{B,i}'),$ $\qquad DT_{1,t}' = 1(t > T_{B,i}')\,(t - T_{B,i}') \quad \forall i = 1,2$

The null hypothesis of $\rho = 1$ against the alternative hypothesis of $\rho < 1$ is tested using the t -statistics of $\hat{\rho}$, denoted $t_{\hat{\rho}}$, in equations (4) and (5) (Narayan and Popp, 2010).

Cointegration test

Next, we test the presence of long-run equilibrium relationship between *PCI* and different types of gross inflows. We implement the ARDL bound testing approach developed by Pesaran and Shin (1999) and Pesaran et al. (2001). The advantage of ARDL approach over other cointegration approach is that it can be applied if there is a mix of I(0) and I(1) variables. The approach usually involves first establishing a cointegrating relationship and then estimating the long-run and short-run parameters using ECM. The ECM form of Eq. (1), (2) and (3) is given as:

$$\triangle \ln PCI_t = \alpha_0 + \theta_1 \ln PCI_{t-1} + \theta_2 \ln TGCI_{t-1} + \sum_{i=1}^{p} \beta_i \triangle \ln PCI_{t-i} + \sum_{j=1}^{p} \beta_j \triangle \ln TGCI_{t-j} + \varepsilon_t \quad (6)$$

$$\triangle \ln PCI_t = \alpha_0 + \theta_1 \ln PCI_{t-1} + \theta_2 \ln FDII_{t-1} + \theta_3 \ln FPII_{t-1} + \sum_{i=1}^{p} \beta_i \triangle \ln PCI_{t-i} + \sum_{j=1}^{p} \beta_j \triangle \ln FDII_{t-j} + \sum_{m=1}^{p} \beta_j \triangle \ln FPII_{t-m} + \varepsilon_t \quad (7)$$

$$\triangle \ln PCI_t = \alpha_0 + \theta_1 \ln PCI_{t-1} + \theta_2 \ln GEI_{t-1} + \theta_3 \ln GDI_{t-1} + \sum_{i=1}^{p} \beta_i \triangle \ln PCI_{t-i} + \sum_{j=1}^{p} \beta_j \triangle \ln GEI_{t-j} + \sum_{m=1}^{p} \beta_j \triangle \ln GDI_{t-m} + \varepsilon_t \quad (8)$$

In Eq. (6), (7) and (8), the parameter θ represents the long-run relationship between *PCI* and inflows while β represents the short-run dynamics. First, we conduct an *F*-test to identify the presence of cointegration between *PCI* and inflows through a joint significance test of the lagged coefficients. Thus, for eq. (6) we test $H_0: \theta_1 = \theta_2 = 0$ and for eq. (7) and (8), we test $H_0: \theta_1 = \theta_2 = \theta_3 = 0$. If the null hypothesis of no cointegration is rejected through the bounds test, then it indicates the presence of at least one cointegrating relationship. In the next step, the long-run coefficients are estimated by choosing an appropriate lag length (*m*). Since the ARDL model assumes that there is no serial correlation among residuals, an appropriate lag length (*m*) is selected to ensure this. Finally, we estimate an ECM model to estimate the short-run parameters and the speed of adjustment.

Finally, we test long-run and short-run causality separately. While other widely used causality tests such as VAR Granger-causality test and MWALD Granger-causality test does not distinguish between long-run and short-run causality, ARDL procedure allows for this innovation (Fatai et al. 2001; Narayan and Smith, 2005). The long-run causality is tested through the *t*-statistics on the coefficients of the lagged ECM term while the *F*-statistics on the lagged independent regressors of the ECM indicates short-run causality.

3.2 Panel analysis

Unit root tests

Before testing for panel unit root, we conduct Cross-Sectional (CD) Pesaran (2015) test that accounts for the presence of cross-sectional dependence in panel date. Usually, panel estimation assumes that disturbances are cross-sectionally independent. However, there may be cross-sectional dependence might occur due to various reasons and are more relevant for a group of countries. Then, we conduct several unit root tests such as Im et al. (2003)(IPS) test, Breitung (2000) test, and second generation IPS test (CIPS) developed by Pesaran (2007) wherein the first two tests are conventional panel unit root tests while the CIPS test does not assume cross-sectional independence of the contemporaneous correlations.

Cointegration test

In the second stage, we apply the dynamic panel ARDL cointegration to identify the long-run relationship between the variables. The panel ARDL is considered superior as it can be used whether the variables exhibit I (0), I (1) or the mixture of both. Further, the panel ARDL test also has the capacity to deal with endogeneity issues in econometric models. Other dynamic panel models such as GMM is not implemented here because our data set has more time periods (T) than cross sections (N). Further, the long-run equilibrium in the PMG-panel ARDL is confined to be homogenous across countries, but allowing for heterogeneity in short-run dynamics. Pesaran et al (1999) also suggested that the PMG estimator is reliable, robust and strong to lag orders and outliers.

The main model of the panel ARDL can be represented as follows:

$$y_{it} = \alpha_i + \sum_{k=1}^{p} \beta_0 y_{i,t-k} + \sum_{k=0}^{q} \beta_1 x_{i,t-k} + \sum_{k=0}^{q} \beta_2 z_{i,t-k} + u_{it} \qquad (1)$$

Equation 1 represents the long-run panel ARDL model where subscript *i* and *t* represents country and time respectively, y is a dependent variable, x is s set of independent variable and z being the set of control variables.

$$\Delta y_{it} = \alpha_i + \varphi_i\left(y_{i,t-k} - \theta_1 x_{i,t-k} + \theta_2 z_{i,t-k}\right) + \sum_{k=1}^{p-1} \gamma_{il} \Delta y_{i,t-k} + \sum_{k=0}^{q-1} \gamma'_{il} \Delta x_{i,t-k} + \sum_{k=0}^{q-1} \gamma''_{il} \Delta y_{i,t-k} + u_{it} \qquad (2)$$

Equation 2 represents the short-run dynamics where γ, γ' and γ'' represents the short-run coefficients of the lagged dependent variable, independent variable and control variable respectively. The φ_i shows the speed of adjustment. It is important to note that we find cross-sectional dependence among disturbances and therefore the dynamic panel ARDL model may not be results. As a remedy, we apply the augmented panel ARDL model that accounts for the common correlated effect. Specifically, we focus on Common Correlated Effect Pooled Mean Group (CCEPMG) method (Pesaran 2006).

In the final stage, we conduct dynamic panel causality tests to identify the presence as well as direction of causality. For this, we apply panel causality tests developed by Dumitrescu and Hurlin (DH) (2012) wherein they extended the model of Granger (1969) to detect the causality in panel data. The DH causality test can be applied when *T* is greater than *N* and thus a good fit for our empirical analysis. The panel data causality model can be represented as:

$$y_{i,t} = \alpha_i + \sum_{k=1}^{K} \gamma_{ik} y_{i,t-k} + \sum_{k=1}^{K} \beta_{ik} x_{i,t-k} + \varepsilon_{i,t} \qquad (2)$$

With, $i = 1, \ldots., N \; and \; t = 1, \ldots., T$

where $x_{i,t}$ and $y_{i,t}$ are the observations for i in period t. k depicts the lag length, γ_{ik} represents the autoregressive parameter while β_{ik} represents the regression coefficients that vary with the group but is time invariant. The DH causality test generates fixed coefficient model, normally distributed and allows for heterogeneity.

4. Empirical Results and discussion

5.1 Time series analysis: A comparison of India and China

First, we test the stationarity of the variables by applying ADF and PP test and Table 1 reports the results. We find that the dependent variable, PCI, is non-stationary at levels while different capital flows variables are a combination of I(1) and I(0) processes. Specifically, unit root test results suggest that *PCI* is integrated of order one, I(1), in case of both India and China. Among capital inflows, *GFDII* is non-stationary in case of both India and China whereas *GEI* is non-stationary in case of China. For all other capital inflows, we confirm stationarity at levels. Thus, the results of unit root tests are consistent with the precondition of applying the ARDL model wherein none of the variable should be I(2) and the dependent variable is I(1).

Table 1. Unit root test results.

The table shows the unit root test of the variables based on ADF and PP tests. The null and the alternative hypotheses for ADF and PP tests are series is non-stationary (contains unit root) and series is stationary (no-unit root), respectively. The sample period used is from 1998Q1-2020Q1 where *pci*, *total*, *fdi*, *fpi*, *equity*, and *debt* represents per-capita GDP, gross total inflows, gross FDI inflows, gross FPI inflows, gross Debt inflows, and gross Equity inflows. *, ** and *** represents significance at the 1%, 5%, and 10% level, respectively.

Countries/Variables	ADF		PP		
	Level	First difference	Level	First difference	
India					
PCI	-1.782	-5.010*	-1.676	-9.690*	I(1)
TGCI	-7.946*	NA	-7.940*	NA	I(0)
GFDII	-1.994	-9.817*	-2.203	-15.549*	
GFPII	-6.805*	NA	-6.800*	NA	I(0)
GEI	-7.206*	NA	-7.297*	NA	I(0)
GDI	-5.764*	NA	-5.854*	NA	I(0)
China					
PCI	-2.157	-2.997**	-2.378	-2.688***	I(1)
TGCI	-6.927*	NA	-7.039*	NA	I(0)
GFDII	-2.005	-4.826*	-3.714*	NA	I(0)
GFPII	-6.486*	NA	-6.462*	NA	I(0)
GEI	-2.061	-9.676*	-3.364**	NA	I(0)
GDI	-6.457*	NA	-6.505*	NA	I(0)

Source: Author's calculations

Since our sample includes the period of peaking of global imbalances, GFC of 2008, etc., *PCI* and capital inflows may have gone significant structural change in their data generating process. To account for the possible structural breaks, we conduct the NP (2010) test to detect the break dates in the time series. The results for both Model 1 and Model 2 are reported in Table 2. The results show that the null of unit root with breaks could not be rejected for *PCI* in case of both China and India. Similar to standard unit root test results, we find that capital inflows are a mix of I(0) and I(1) variables. Of note, while we detect unit root for all time series, the identified structural breaks are also statistically significant, indicating the importance of break dummies even if the unit root is present. Majority of the detected break dates are clustered around 2006–2009, indicating that there has been a structural shift in the capital flows since the peaking of global imbalances and the onset of the GFC.

Table 2. Narayan and Popp (2010) endogenous structural break test results

Narayan and Popp (2010) test methodology is used, with critical values for M1 and M2 given respectively as: at the 10% level (*), -4.958 and -5.576; at the 5% level (**), -4.316 and -4.937 (p.1429). [a] represents variables in which the order of integration is inconclusive. In that case, we consider the graphical analysis and choose the model that best fits the structural break properties.

	M1 Two breaks in intercept				M2 Two breaks in intercept and trend				Order of integration
	Lag	t-stat	TB1	TB2	Lag	t-stat	TB1	TB2	
India									
PCI	4	-1.981	2008Q3	2009Q3	4	-4.94**	2008Q3	2011Q3	Inconclusive[a]
TGCI	2	-1.643	2008Q3	2009Q3	0	-40.83*	2003Q1	2008Q3	Inconclusive[a]
GFDII	0	-5.01**	2006Q3	2012Q3	2	-2.709	2006Q3	2008Q3	Inconclusive[a]
GFPII	3	-5.835	2007Q4	2013Q1	3	-6.644	2007Q4	2013Q1	I(0)
GEI	0	-8.509	2003Q3	2008Q3	5	-7.173	2006Q1	2008Q3	I(0)
GDI	0	-8.014	2008Q3	2009Q3	3	-4.990	2004Q1	2008Q3	I(0)
China									
PCI	4	-1.547	2006Q4	2008Q4	4	-2.028	2008Q4	2015Q2	I(1)
TGCI	0	-6.582*	2014Q4	2015Q2	4	-5.222*	2008Q3	2014Q4	I(0)
GFDII	4	-2.746	2003Q2	2006Q3	5	-5.588*	2006Q3	2015Q2	Inconclusive[a]
GFPII	0	-6.625*	2010Q1	2015Q2	0	-6.451*	2010Q1	2015Q2	I(0)
GEI	3	-1.340	2006Q3	2015Q2	5	-4.497	2006Q3	2015Q2	I(1)
GDI	0	-5.902*	2011Q3	2012Q2	4	-4.361	2011Q3	2012Q2	Inconclusive[a]

Source: Author's calculations

As a next step, we examine if gross capital inflows lead to an expansionary or contractionary impact on domestic output, and whether the capital inflows at disaggregated level exhibit different linkages. Thus, we test for the presence of a long-run equilibrium relationship between the *PCI* and capital inflows at the aggregated and disaggregated level using the ARDL approach. The *F*-statistic and results of the bounds test are reported in Table 3. The results imply that the null of no cointegration is strongly rejected for all three model specifications. Thus, there is presence of a long-run relationship between the *PCI* and capital inflows in all three specifications.

Table 3. Test for cointegrating relationship.

Notes: *k*=1 for all ARDL bound tests. Critical values are from Narayan (2005). ** and * represents significance at the 5% and 1% level, respectively.

F-statistic	India	China	Critical value 95% level		Critical value 99% level	
F_{PCI} (*PCI* | *TGCI*)	5.754**	0.417	3.74	4.30	5.15	5.91
F_{PCI} (*PCI* | *GFDII*, *GFPII*)	6.212*	10.432*	3.23	4.05	4.35	5.39
F_{PCI} (*PCI* | *GEI*, *GDI*)	17.746*	8.341*	3.23	4.05	4.35	5.39

After establishing presence of cointegration between the *PCI* and capital inflows, we estimate the long-run coefficients of the three model specifications. The estimated long-run parameters of the ARDL model are presented in Table 4. The long-run coefficient estimates all show the expected signs at either 1%, 5% or 10% level of statistical significance.

Table 4. Estimated long-run coefficients from the ARDL model.

Numbers in the parenthesis are the std errors. *, ** and *** represents significance level at 1%, 5% and 10%, respectively. All capital inflow variables are interacted with a control variable in the respective three models.

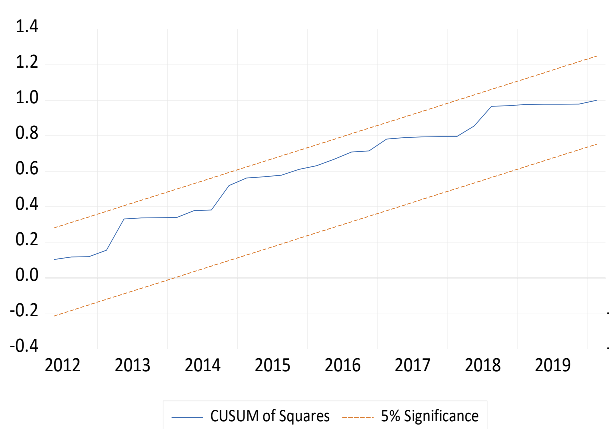
	India			China		
	Model I	Model II	Model III	Model I	Model II	Model III
C	4.368*	2.846*	4.196*	7.390	-9.249	-1.708
	(0.280)	(0.559)	(0.187)	(6.958)	(6.544)	(1.010)
TGCI	0.195*			-		
	(0.029)					
GFDII		0.397*			1.579**	
		(0.062)			(0.661)	
GFPII		0.024**			0.041***	
		(0.010)			(0.022)	
GEI			0.209*			0.831*
			(0.020)			(0.108)
GDI			0.011*			0.027**
			(0.002)			(0.012)
$J\text{-}B_{Norm}$	1.185	0.031	0.830	-	1.606	1.821
	(0.552)	(0.984)	(0.660)		(0.447)	(0.402)
F_{Auto}	0.968	0.013	1.883	-	1.342	1.615
	(0.386)	(0.986)	(0.162)		(0.270)	(0.218)
F_{ARCH}	0.000	0.365	0.201	-	0.081	0.183
	(0.993)	(0.547)	(0.654)		(0.777)	(0.669)

Many interesting results emerge from the time series estimation. First, the estimated coefficient of the gross capital inflows, *TGCI*, is statistically significant for India and greater than zero which implies that gross capital inflows results in expansionary impact on domestic output in presence of financial development. However, for China we do not find any significant impact of gross capital inflows on domestic output. Then, we disaggregate TGCI to further interrogate if there is any aggregation bias due to the cancelling-out or adding-up effect with the use of aggregate gross capital inflows. For this, we first decompose the *TGCI* into *GFDII* and *GFPII* and check if the direct investment, a metric of long-term capital inflows, or portfolio investment, a metric of short-term capital inflows, exhibit similar linkages or not. The results are reported in column 3 and column 6 of Table 4. We find that both the *GFDII* and *GFPII* leads to an expansionary effect on domestic output for both countries. While there is a positive impact of both *GFDII* and *GFPII* inflows, the coefficient on *GFDII* inflows is greater than that of *GFPII* inflows, indicating that *GFDII* inflows have a stronger impact on domestic output in presence of financial development in the borrower country. The results are in line with the arguments made in earlier studies that absorptive capacities such as development financial markets and provision of more credit facilitates FDI to create higher domestic output growth (Ang 2009b; Alfaro et al. 2004; Durham 2004). Then, we again decompose the *TGCI* into *GEI* and *GDI* in our third specification to check if we find similar results or not. The results are reported in column 4 and column 7, and we find that for both countries, equity inflows and debt inflows exhibit positive impact on domestic output. It is interesting to note that in case of China, we find that *TGCI* do not have any significant impact on domestic output but at the disaggregated level, gross inflows have a significant positive impact on domestic output. These contrasting findings with disaggregated capital flows indicates a possibility of aggregation bias because of a cancelling-out or adding-up effect.

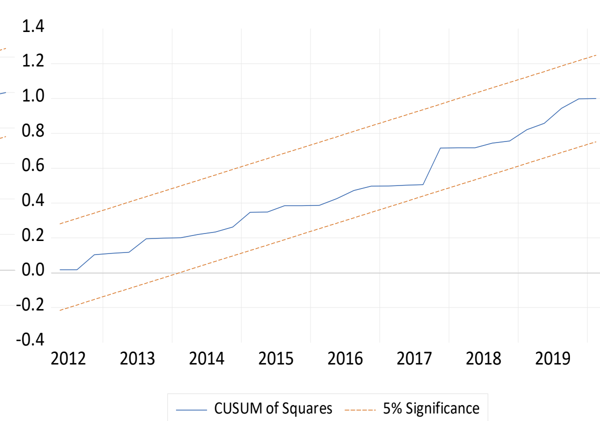
As a metric of soundness of our results, we conducted a battery of diagnostic checks, including tests of normality, autocorrelation, and heteroscedasticity in the residual term. From diagnostic test results, we find that the residuals are normally distributed and there is no autocorrelation or heteroscedasticity among residuals. The plots of the cumulative sum of squares (CUSUMQ) based on the recursive residuals is presented in Figure 1, and the results show that the coefficients estimates are stable across the sample span. Hence, we can conclude that the model is well-behaved.

Figure 1. Plots of cumulative sum of squares (CUSUMQ) of recursive residuals. Figure 1(a)-1(c) are plots of CUSUMQ for specification1-3 for India, respectively. Figure 1(d) and 1(e) are plots of CUSUMQ for specification 2 and specification 3 for China.

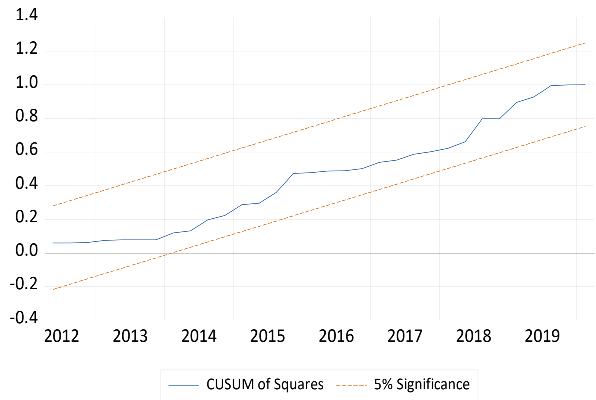
(a)



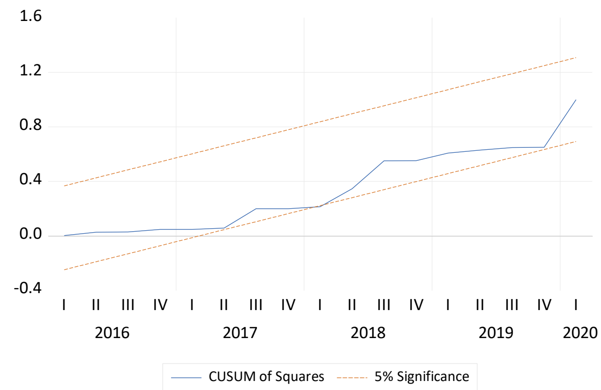
(b)



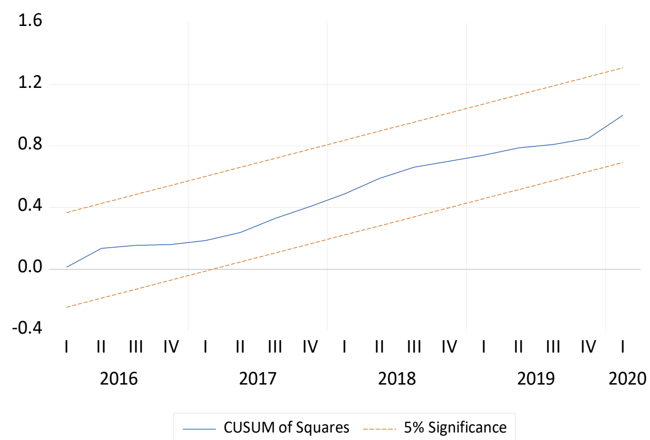
(c)



(d)



(e)



Next, we estimate the long-run causality and short-run causality using the ARDL error correction mechanism and the results are reported in Table 5. Specifically, we applied the Wald test with the assumption that the lagged independent regressors of the short-run equation are equal to zero. Thus, if the value of F-statistics is able to reject the null hypothesis of no causality, we confirm the significance of the short-run causal relations. For the long-run causality, the t-statistics on the lagged ECM terms indicates the long-run causality. The results of short-run and long-run causality are presented in Table 5. The results from the causality test implies that, in the short-run, TGCI Granger-causes PCI in case of India but not for China. Likewise, the long-run causality test indicates significant causal relations from TGCI to PCI in case of India. Thus, the results are in line with the ARDL long-run elasticities for Model I. Then, for Model II, we find that short-run causality from GFDII to PCI in case of China but not India however there exists significant long-run causal relations in case of both countries. For model 3, we find significant causality running from GEI to PCI and GDI to PCI for both countries. Hence, the short-run and long-run causality test results are quite consistent with the ARDL cointegration results.

Table 5. Granger-causality tests

The table shows the short-run and long-run Granger causality results between $\Delta tgci$ and Δpci for India and China from ARDL short-run dynamics. *, ** and *** represents rejection of the null hypothesis of Granger non-causality at the 1%, 5% and 10% significance level, respectively.

Country	Null Hypothesis	Short-run causality *F-statistic* (prob.)	Long-run causality *t-statistic* (prob.)
India	$\Delta tgci$ does not Granger-cause Δpci	4.313*	-0.335*
	$\Delta gfdii$ does not Granger-cause Δpci	0.489	-0.304*
	$\Delta gfpii$ does not Granger-cause Δpci	9.482*	
	Δgei does not Granger-cause Δpci	4.662*	-0.404*
	Δgdi does not Granger-cause Δpci	1.791***	
China	$\Delta tgci$ does not Granger-cause Δpci	1.266	NA
	$\Delta gfdii$ does not Granger-cause Δpci	1.997***	-0.055*
	Δgpii does not Granger-cause Δpci	1.982***	
	Δgei does not Granger-cause Δpci	2.134**	-0.345*
	Δgdi does not Granger-cause Δpci	0.821	

5.2 Panel analysis

Next, we consider a panel analysis of the selected ten EMEs that have received huge influx of capital inflows since the GFC of 2008. The selected countries are Brazil, Chile, China, India, Indonesia, Mexico, Poland, Russia, South Africa, and Thailand. The panel data estimation assumes that the disturbances are cross-sectionally independent however there might be instances where they are cross-sectionally dependent. Thus, we first test for the cross-section dependence (Pesaran, 2006). Table 6 presents the results for cross-sectional dependence test that does not assume cross-sectional independence of the disturbance terms. The results indicate that the null hypothesis of cross-section independence is strongly rejected at the 1% level, confirming the presence of cross-country dependence in the data. The dependence may be due to similar political and economic conditions; therefore, it is necessary to employ cross-section dependence IPS (CIPS) unit root test (Pesaran, 2006) which deals with cross-section dependence. However, we conduct both the first-generation and second-generation unit root tests. The first-generation unit root test which we have conducted here are Breitung (2000) and Im et al. (IPS) (1997) tests and assumes independence of cross-sections whereas the second-generation unit root test (CIPS) relaxes this assumption and considers cross-section dependence.

Table 6. Cross-sectional Dependence Test

Null: Cross-sectional Independence		
Test	Statistics	Prob
Breusch-Pagan LM	408.605	0.000
Pesaran Scaled LM	38.327	0.000
Pesaran CD	13.225	0.000

Table 7 presents the results of of Breitung (2000) and IPS (1997) panel unit root test. The results suggest that the *PCI* and *TGCI* are integrated of order one, I(1). However, all the other capital inflow variables, *GFDII*, *GFPII*, *GEI* and *GDI* are stationary at levels, I(0). We find the similar results for the second-generation unit test results except for *TGCI* which is stationary at levels. Since, we find that the variables are a mix of I(1) and I(0) hence we employ dynamic panel ARDL approach since it can be easily conducted when the independent regressors are either integrated of order one or zero and are robust.

Table 7. Panel unit root tests

*, **, *** represents significance level at 1%, 5% and 10% respectively.

Variables	Breitung (2000)		IPS (1997)		CIPS (2007)	
	Level	First Difference	Level	First Difference	Level	First Difference
PCI	2.9891	-4.0984*	0.2765	-4.6690*	-2.035	-3.261*
TGCI	1.4747	-7.7701*	-1.2480	-7.1661*	-3.374*	NA
GFDII	-2.4534*	NA	-2.9796*	NA	-3.644*	NA
GFPII	-3.3402*	NA	-10.2917*	NA	-4.808*	NA
GEI	-1.9282**	NA	-2.9235*	NA	-2.576**	NA
GDI	-4.8324*	NA	-2.3170**	NA	-3.190*	NA

Next, we report the long-run coefficients from panel ARDL in Table 8. One usefulness of panel ARDL over other approaches is that it can be utilized to identify long-run relationships even without cointegration. Column 2 and column 5 reports the Model I results for two proxies for financial development as control variables. The control variable 1 is financial development proxied by broad money to GDP and control variable 1 is financial development proxied by domestic credit to private sector to GDP. The results suggest that *TGCI* has an expansionary impact on *PCI* in large EMEs. Model II results are reported in column 3 and column 6 wherein total inflows are decomposed into *GFDII* and *GFPII*. With disaggregated capital inflows, we find similar results that both direct investment and portfolio investment inflows leads to an expansionary impact on domestic output. However, the coefficient on *GFDII* is greater than *GFPII* which implies that *FDI* inflows lead to a larger expansionary impact with the help of adequate financial development in the borrower country. With regards to decomposition of total inflows into equity and debt inflows, we find equity inflows, *GEI*, has an expansionary impact while debt inflows, *GDI*, has a contractionary impact on domestic output. Thus, contrasting signs with disaggregated inflows in Model III as compared to Model I in which we use total inflows suggests possibility of aggregation bias. Hence, our results present an interesting finding that type of capital inflows are important in identifying their effect on domestic output and for appropriate policy recommendations.

Table 8. Panel ARDL Estimation Results

Numbers in the parenthesis are the std errors. *, ** and *** represents significance level at 1%, 5% and 10%, respectively. Control variable 1 and 2 is broad money to GDP and domestic credit to private sector to GDP, respectively. All control variables are interacted with capital flows in the respective three models.

Long-run	Control Variable 1			Control Variable 2		
Variables	Model I	Model II	Model III	Model I	Model II	Model III
C	2.839*	2.879**	1.727***	4.658*	3.031*	0.334
	(0.549)	(1.439)	(0.946)	(1.149)	(0.618)	(0.213)
TGCI	0.032*			0.039*		
	(0.006)			(0.007)		
GFDII		0.326*			0.164*	
		(0.016)			(0.043)	
GFPII		0.013*			0.004**	
		(0.001)			(0.001)	
GEI			0.266*			0.083*
			(0.007)			(0.102)
GDI			-0.068*			-0.401*
			(0.002)			(0.065)
ECT(-1)	-0.324*	-0.937**	-0.318***	-0.522*	-0.482*	-0.220**
	(0.066)	(0.460)	(0.172)	(0.120)	(0.075)	(0.092)

Finally, we identify causal relations among different types of gross capital inflows and domestic output by applying Dumitrescu and Hurlin (2012) panel causality tests. Table 9 reports the panel causality test results. We find the evidence that there exists a bidirectional causality between *TGCI* and *PCI*, and *GFDII* and *PCI*. Further, we find unidirectional causality running from *GFPII* and *GEI* to *PCI*. However, we do not find causality running from *GDI* to *PCI*. Overall, capital inflows at the aggregate and disaggregated level causes domestic output, except for debt inflows.

Table 9. Dumitrescu Hurlin Panel Causality Tests
The values in the parenthesis are the p-values. *, ** and *** represents significance level at 1%, 5% and 10%, respectively.

Null Hypothesis	W-statistic
$FDII \Rightarrow PCI$	4.473**
	(0.011)
$PCI \Rightarrow FDII$	4.820*
	(0.003)
$TGCI \Rightarrow PCI$	6.308*
	(0.000)
$PCI \Rightarrow TGCI$	6.904*
	(0.000)
$FPII \Rightarrow PCI$	6.164*
	(0.000)
$PCI \Rightarrow FPII$	3.215
	(0.483)
$EI \Rightarrow PCI$	5.808*
	(0.000)
$PCI \Rightarrow EI$	3.215
	(0.286)
$DI \Rightarrow PCI$	3.235
	(0.276)
$PCI \Rightarrow DI$	5.127*
	(0.000)

5. Conclusion

With the majority of the previous studies focusing on the impact of net capital inflows on domestic output, literature on gross capital flows as well as the impact of disaggregated gross capital inflows is scanty. We fill this gap in our study. First, we conduct a comparative analysis of India and China using time-series data. We find that gross capital inflows are expansionary in nature both in case of India and China, with FDI inflows having the largest impact in presence of adequate financial developments in the borrower country. Our time-series results do not confirm aggregation bias since the results with both aggregate and disaggregated capital inflows are consistent.

Second, we select ten large EMEs that have received substantial gross capital inflows since the GFC of 2008, and conduct a panel analysis. Overall, we find that gross capital inflows lead to an expansionary effect, except for gross debt inflows which exhibit a contractionary effect. Thus, our results show possibility of aggregation bias due to the adding-up or cancelling-out effect when total gross capital inflows are used. Our findings also imply that distinction between different types of capital inflows are necessary since the policy recommendations based on aggregate capital inflows may be ill-advised. For EMEs, the policymakers' impetus should be on encouraging FDI and FPI rather than debt inflows, and improving absorptive capacities of local financial markets in borrower countries for effective utilization of capital inflows.

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