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Fiscal stance, foreign capital inflows and the behavior of current account in the Asian countries

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Abstract This paper investigates the causality relationship between budget and current account deficits for 14 Asian countries. The major findings based on bootstrap Granger causality tests in heterogenous mixed panels are: First, investments (and foreign capital inflows) have a notable impact on the current account prior to the Asian financial crisis of 1997–1998. Second, we detect direct causality between the budget and the current account imbalances before but not after the crisis. The data suggest that the current account targeting hypothesis, that is, the reverse causality, prevails after the crisis. Third, budget deficits have a significant influence on both investment and foreign capital inflows in the second period. Fourth, Asian countries are less susceptible to the influence of FDI inflows in the aftermath of the crisis. Finally, structural break turns out crucial in assessing the causal patterns of the twin deficits nexus in the Asian countries.

Keywords Global imbalances · Twin deficit hypothesis · FDI · Crisis · Asian

JEL Classification F41 · F32 · C10

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1 Introduction

The twin deficits hypothesis (TDH), which assumes that budget deficits have a direct effect on current account deficits, has been a topic of interest in the empirical literature (see, among many others, Salvatore 2006; Daly and Siddiki 2009; Kalou and Paleologou 2012; Makin and Narayan 2013; Trachanas and Katrakilidis 2013). The bulk of the literature, however, has focused mainly on large current account deficit countries with growing fiscal imbalances; see a summary by Xie and Chen (2014, Table 1). From a theoretical perspective, fiscal expansion leads to exchange rate appreciation, which could worsen current account deficits, and ultimately, this could lead to slower economic growth. Abell (1990, US), Bahmani-Oskooee (1995, US) and Salvatore (2006, G-7), just to name three, offer empirical support for the Keynesian view that fiscal deficits fuel current account deficits using data from the industrial countries. While the last two papers confirm a direct link between the fiscal and current account deficits, Abell (1990) reports that US budget deficits widen trade deficits indirectly via the US interest rate and exchange rate channels. Salvatore (2006) in his assessment argues that reducing the size of the budget deficit may be as effective as an exchange rate intervention for reducing the size of trade deficits. Salvatore (2006) also emphasizes the importance of foreign capital inflows: The entire current account deficit and part of the budget deficit in the advanced countries are financed by net capital inflows. On the contrary, Xie and Chen (2014), who apply new panel Granger causality proposed by Emirmahmutoglu and Kose (2011), document a two-way causal relationship between the two deficits for the Organization for Economic Cooperation and Development (OECD) countries.[1] Meanwhile, the paper by Papadogonas and Stournaras, (2006) provides support for the Ricardian equivalence hypothesis (REH) for EU-15 member states. Precisely, they conclude that improvement in government balance has a positive effect on the current account balance, but the effect is rather small. Changes in government balances are strongly associated with opposite changes in the private sector savings–investment gap (p. 601).

Several papers extend the empirical analysis to the Asian developing countries. These studies based on data that ended in the late 1990s point to either the bidirectional relationship between the two deficits (Baharumshah et al. 2006) or Summers (1988)'s current account targeting (CAT) as reported in Anoruo and Ramchander (1998). Accordingly, the CAT hypothesis asserts that there is unidirectional causality from the current to the budget deficit. The latter evidence is characterized by a link in the opposite direction of the one predicted by the TDH. Focusing on the Indian's twin deficits dilemma (1970–2000), Parikh and Rao (2006), basing their results on a single country setup, find that causality runs from budget to current account deficit,

[1] The budget and current account deficits tend to move together when capital mobility is high (especially in the international financial centers); see, e.g., the work of Xie and Chen (2014), Daly and Siddiki (2009), among others.

evidence which departs from the findings of Anoruo and Ramchander (1998).[2] Are these findings suggesting that the macroeconomic effects of fiscal policy changing?[3]

The primary aim of this paper is twofold. The first is to examine the effects of budget balance (BD) and investment on current account balance (CAD) in 14 Asian countries, including the fast-growing economies of the People's Republic of China and India and the advanced Asian countries (Japan, South Korea, Hong Kong, Taiwan and Singapore). Many of these countries have done well in terms of their average growth in the recent decades, with China and India now at the top of the ranking list. China has run large current account surpluses since the mid-1990s and has been seen as the most serious source of global imbalances. China's current account as a share of GDP touched a double-digit figure in 2004 but fall abruptly to around 5% in 2010 with the outbreak of 2007–2008 global financial crises (hereafter, GFC). Meanwhile, the hardest hit East Asian countries (Thailand, Indonesia, South Korea and Malaysia) have moved from a deficit to the current account surplus after the Asian financial crisis (AFC) of 1997–1998. Currently, there is a fear that the space for continued fiscal stimulus is already—or will soon be—exhausted in a majority of the Asian countries. The sharp swings in fiscal policy and large current account imbalances in these countries to cushion the blow of financial crises over the recent years offer an opportunity to open up an important debate on the existence (if any) of the twin deficit dilemma. Previously, Baharumshah et al. (2006) found that TDH holds only for Thailand, while the evidence points to reverse causality run for Indonesia. For the other ASEAN countries, bidirectional causality between the twin deficits is unveiled.

A question typically raised in these studies is: How susceptible is the current account to foreign capital inflows? Apart from assessing which of the two competing hypotheses—CAT or TDH—is most appropriate for the Asian countries, this paper also aims to examine the impact of the large surge of foreign capital inflows on the CAD. The waves of capital inflows in the past decades are particularly pronounced in these countries and thus shaped the course of current account developments. The results that emerged from earlier investigations agree that current account imbalances in emerging market economies (EMEs) in the era of high capital mobility are largely caused by foreign capital inflows. As already mentioned, different countries tend to differ with respect to their experience with foreign capital flows and they do not systematically gain from capital account liberalization; see Wong and Carranza (1999), Doraisami (2007), Yan (2007) and recently Yan and Yang (2012). Therefore, an understanding of the relationship between these variables will shed additional insights on the appropriate policy response (typically sterilization) when confronting with massive foreign capital inflows in a globalized economy.[4]

[2] Similarly, Kim and Kim (2006), who examine the twin deficits link for South Korea, find a unidirectional causal relation running from current account to budget deficits.

[3] The data used in our study are likely to lead to size distortion in the test results with asymptotic method simply because our sample size is too small.

[4] Developing countries with unsophisticated financial systems are prone to currency crisis when capital account is liberalized since foreign capital can worsen or improve national savings or domestic investment. This means that the current account is vulnerable to foreign capital inflows. The AFC is a point in case.

The findings of the extant literature with respect to the effect of fiscal deficits on current account are mixed. The present study contributes to the existing literature in the two ways. First, we apply up-to-date panel methods, including the panel Granger causality test advocates by Konya (2006) to analyze the relationship between CAD and BD, rather than limiting the empirical analysis to an individual country setup as usually done in the past. This technique has gained popularity in their application in recent years (see Chang et al. 2015 and Menyah et al. 2014), especially for individual country where the data span is not long enough. Furthermore, rather than relying on asymptotic critical values, we perform the Wald test for Granger causality with country-specific bootstrap critical values generated through simulations. This paper aims to complement existing studies by examining the causal relationships using the new empirical methodology. We apply a new methodology to deal with the major critics of panel analysis, namely the issue of heterogeneous mixed panels. Specifically, we apply the bootstrap panel causality test as outlined in Kónya (2006) to accommodate for both cross-sectional dependency and heterogeneity across the Asian countries. Most of the previous studies are based on single equation methods and the conclusion on whether the deficits are twined or not is based on asymptotic rather than simulated critical values. Causality tests based on asymptotic critical values tend to spuriously reject the null hypothesis. In this paper, we argue that it is more accurate to conduct inference using the bootstrap method. This method is shown to have less size distortion and to provide more precise test inferences than the asymptotic critical values if the sample size is small, as in many applications, including ours (Ko and Ogaki 2015).

The seemingly conflicting findings on the causal relations between the current account and budget balances lead us to an important question: Has the nexus between the two variables changed (or strengthen) over the past decades or are the differing results due to the use of different econometric methods? A recent strand of the literature has also emerged arguing that structural breaks in the data (say due to shifts in monetary regime) if ignored may bias the parameter estimates and hide the underlying relationship between the variables in a model. Recent studies (Hatemi-J and Shukur 2002; Daly and Siddiki 2009; Holmes 2011; Eldemerdash et al. 2014; Xie and Chen 2014; Nikiforos et al. 2015, among others) have taken up this issue. Their findings suggest that the divergence in the results could be the result of structural breaks and regime shifts that have not been appropriately accounted for in earlier studies.

We examine a break date in the relation to the linkage between the current accounts and its determinants. The two factors that can potentially affect the current account dynamics include exchange rate flexibility and large accumulation of foreign exchange reserves that follows after the outbreak of the AFC.[5] To account for the episodes and the changing causal relationship as well, we create two subsamples, one prior to the crisis up to 1997 and another for the period after the AFC. Failure to account for structural break not only biases the estimated parameters of the model, but could also lead to incorrect inference concerning the direction of a causal nexus between budget

[5] The AFC apparently played a key role in promoting surpluses (or lower deficits), both directly by restraining domestic spending (particularly investment) and indirectly by encouraging the authorities to take measures to keep their exchange rates competitive.

and current account imbalances. As it is shown later, our findings clearly illustrate this point.

Since the late 1990s, many of the Asian countries have experienced episodes of large budget deficits and growing debts, mainly as the result of huge and persistent fiscal stimulus in response to the economic slowdown due to the negative macroeconomic shocks associated with the aftermath AFC and GFC.[6] The growth in public debt in the second half of 1990s has been rapid in Asia (China, India, Indonesia, South Korea, Malaysia and the Philippines) from 40% in 1996 to about 65% of aggregate gross domestic product (GDP) in 2002, significantly larger than those of countries in Latin America. Growing fiscal deficits can be problematic if they turn out to be long lasting in nature since they have the potential to affect other macroeconomic variables, especially the current account. Persistent budget deficits can lead to current account deterioration, and the impact is expected to be much more severe for the smaller economies with high debt-to-GDP ratio (Corsetti and Müller 2006).

The remainder of the paper is organized as follows. In the next section, we briefly review the literature on the twin deficit dilemma. Section 3 introduces the analytical framework and sets out the methodology used in the empirical analysis. In Sect. 4, we describe the data and details of the empirical results. Finally, Sect. 5 summarizes the major findings.

2 Brief review of the relevant literature

The above discussion focused mainly on the TDH. Apart from the TDH that is based on the Keynesian absorption theory, there are other competing views on the causal relationship between budget deficit and CAD. In fact, the bulk of the studies has reported a positive link, no external impact of a higher fiscal deficit (Ricardian equivalence proposition)[7] and even a negative relationship (twin divergence referred by Kim and Roubini 2008) between the two deficits; see also Enders and Lee (1990) and Alse and Bahmani-Oskooee (1992). The (small) negative influence of budget deficit on the CAD (for the USA) is also reported by other related study (Corsetti and Müller 2006).[8] Finally, Anoruo and Ramchander (1998), Khalid and Guan (1999), Marinheiro (2008), and more recently Kalou and Paleologou (2012) and Nikiforos et al. (2015) found support for the CAT hypothesis. The last two articles focus on Greece, where current account and fiscal deficits coexist for a long period. According to this hypothesis, the causal ordering runs from external to internal deficits and could be justified for a country with deterioration in their external position. The fiscal deficit

[6] Conventional wisdom holds that fiscal expansion during a downturn can stimulate growth and return the economy to full employment, provided that debts are not so high as to preclude new borrowing. This according to many scholars have led to rapid and robust recovery (also dubbed as V-recovery).

[7] According to this proposition, movement between taxes and budget deficits has no effect on the real interest rate or the CAD. In other words, there is no relationship between the two deficits.

[8] The literature has also found the link between the two deficits is weak or even nonexistence; see, for example, Bussiére et al. (2010) and Algieri (2013). Evidence in Algieri (2013) based on traditional Granger test and the approach suggests by Toda–Yamamoto is in conformity with the Ricardo equivalence theorem for Portugal, Ireland, Italy, Greece and Spain (labeled as GIIPS countries).

responds to this deterioration and adjusts—either through automatic stabilizers or active policy decisions of the government—in order to stabilize the economy (Darrat 1988). It should be mentioned that this type of stabilization requires sufficient foreign capital inflows and relatively low-interest rate. Nikiforos et al. (2015), who find the case to be strong in Greece, explain along these lines of argument. A continuous accumulation of public and external debts associated with debt crises will eventually lead to current account deterioration. Governments may offset the deflationary effect of the current account by increasing its expenditure to assist the domestic industry; see Kalou and Paleologou (2012). This could be another explanation for the observed reverse causality between the twin deficits.

Trachanas and Katrakilidis (2013), who apply asymmetric cointegration methods, find strong evidence supporting the TDH for five European economies that are under financial market pressure and insolvency (Portugal, Ireland, Italy, Greece and Spain). The finding shows that fiscal deficit deceases have a greater impact on the current account deficit as opposed to that of budget increases. Meanwhile, Nikiforos et al. (2015), using the latest available data for Greece, demonstrate that trade balance and primary deficits are independent prior to 1995, while the trade deficit has been the primary driving factor for the primary deficit in the post-1995 period. The change in the direction of causality patterns is closely connected with the European monetary unification and deregulation of capital movements in 1993 and 1994. It appears that the evidence on the budget deficit-current account deficit linkage in Euroland from the accumulated studies is neither robust nor stable over time.

Candelon and Lieb (2013) apply a regime switching framework to identify recession and expansion periods. In their analysis, they confirm that in times of recession active spending policies have a stronger impact and should be preferred to deficit-finances tax cuts in a recession. In the short term, deficits usually stimulate economic activities and policy actions to reduce the deficit dampen growth. Over the longer term, if the deficits are not corrected, the stimulative effects tend to wear off.

Various studies have considered structural breaks to assess the co-movements between the two deficits (see among others, Bagnai 2006; Grier and Ye 2009; Daly and Siddiki 2009; Rafiq 2010; Kalou and Paleologou 2012) and have drawn our attention to the issue. Kalou and Paleologou (2012) provide further empirical evidence in favor of CAT, but only after taking into account shifts in the data generating process for Greece, an over-indebted country (with high possibility of default) and facing serious twin deficit dilemma (Trachanas and Katrakilidis 2013). Similarly, Marinheiro (2008) taking Egypt (another highly indebted country) as a case study report a twin deficit dilemma.[9] Rafiq (2010) also recognized the importance of structural breaks in the relationship and observed that the overall impact of fiscal deficits in the OECD countries has fallen in magnitude over the past 2 decades after controlling for the movements in their exchange rates. Rafiq's 2010 findings suggest that future budget deficit reduction alone may be unable to correct the global imbalances—the root of the recent financial crisis.

[9] Many countries around the globe, including those under review, have accumulated a large debt in the aftermath of the GFC (Baldacci et al. 2010).

Yuan and Chen (2015) find no evidence in favor of either the twin deficits or the twin divergence hypotheses (i.e., expansionary fiscal policy improves the current account position) in five major EMEs of Brazil, Russia, India, China and South Africa (the so-called BRICS). With regard to the role of monetary and fiscal policies in managing the external imbalances, they document a significant impact of a monetary shock on real economic growth. Their findings show that fiscal balances are extremely sluggish in affecting the external balances, suggesting no evidence supporting neither twin deficits nor twin divergence hypotheses. When looking at the BRICS-US bilateral trade balances, their analysis reveals that trade surpluses in BRICS countries tend to weaken their currencies. Low financial openness, nominal rigidity in the exchange rate and lack of price flexibility are among the potential factors explaining this finding. Another important insight offered by this paper is that pushing the currencies in the EMEs vis-a-vis the US dollar can hardly improve the US trade deficits. In other words, depreciating the currencies in BRICS countries will have a disappointing (small) effect on the US trade imbalances. Finally, Eldemerdash et al. (2014) find a positive relationship between fiscal and external balances with causality running from budget deficits to trade deficits in oil-producing countries. They find support for the Ricardian view in which lower public savings are met by equal increases in private savings for non-oil-producing countries.

In short, there is considerable disagreement about the causal ordering between twin deficits, especially in the last few decades due to the fact that the current account has been increasing determined by short-term factors. From a policy perspective, it is crucial to determine whether the BD can influence the CAD or vice versa in a predictable pattern. If it is known that the TDH holds, then the external imbalances cannot be resolved unless appropriate policies to address BD are first put into place. The characterization of the hypothesis has crucial implication in terms of best practice to resolve problems associated with budget and current account deficits in order to avoid future policy failures. Specifically, Yan and Yang (2012) examine the causal relationship between the two variables for the period from 1987 to 2006, a year prior to the GFC.[10] In particular, the studies show that current account movements in the EMEs—including some under review in this current study—are largely caused by foreign capital inflows. The opposite result is observed for the panel of advanced industrial countries (particularly the USA). The different behaviors are attributed to two major factors: exchange rate system and financial institutions. The advanced economies with highly flexible exchange rate system and more sophisticated financial system are less likely to suffer a capital account crisis that the emerging economies experience in the late 1990s due to the sudden cut off from international capital markets (dubbed sudden stop); see Caballero et al. (2005). As financial markets became globalized, we expect EMEs to receive more of the net foreign capital inflows. The puzzle with financial

[10] Two major episodes of foreign capital inflows to Asia over the past 2 decades affect the current account in significant manner. The first began in early 1990s and ended in abruptly with the AFC. The second episode began in early 2000 and ended with the GFC. There were significant variations in both the magnitude and types of capital inflows experienced by these countries; see Yan and Yang (2012). Unlike the advanced countries, they find that current account in the EMEs (India, Indonesia, South Korea, Malaysia, Pakistan, the Philippines, Sri Lanka and Thailand) is susceptible to foreign capital inflows.

globalization is that capital does not flow from rich to poor countries but instead tends to flow from capital-scarce poor to capital-abundant rich countries (the so-called Lucas paradox). The recent trend shows the EMEs (China and others) are exporting capital to advanced economies (including USA).

In sum, an important policy implication of the Yan and Yang's findings concerns the capital account liberalization, and according to them, "EMEs are susceptible to the whimsical movement of foreign capital inflows, and therefore such countries should be cautious when liberalizing the capital account" (p. 27). With specific reference to the Asian EMEs, they confirm that after the 1997–1998 AFC foreign capital inflows have changed the nature of their effects on the current account. The extent, as well as the drivers of the current account, may have changed after the AFC.[11]

3 A brief analytical framework and statistical strategy

The public sector may meet its financing needs through domestic and/or international financial markets. Considering the significant role of private investment in the intertemporal approach to the balance of payment, the relationship between current account balance (CAD), budget balance (BD) and investment (INV) (all expressed as percentage of GDP) can be presented as:

$$\mathrm{CAD}_{it} = \beta_1 + \beta_2 \mathrm{BD}_{it} + \beta_3 \mathrm{INV}_{it} + \varepsilon_{it} \tag{1}$$

The model given is a starting point for our analysis and draws from the past literature. It provides a basic analytical framework for investigating the link between CAD, BD and INV and on the quantitative effect of fiscal on current account imbalances.[12] Briefly, the model derives from the national accounting framework to show that current account and fiscal balance are directly interrelated (or twinned). The model has been applied by Marinheiro (2008) for Egypt, Chen (2010) for USA, Kalou and Paleologou (2012) for Greece and Litsios and Pilbeam (2017) for Greece, Portugal and Spain, among others. Following these as well as the other studies on twin deficits, we have included investment in our model.[13] We expect the coefficients for BD and INV in the equation to be positive and negative, respectively. Large and highly sophisticated capital markets make it possible for countries to finance public and private domestic needs via global capital markets. It should be noted that in the absence of capital mobility, the two deficits cannot be "twins" and parameters are close to zero as detailed by Bagnai (2006) and Xie and Chen (2014). The addition of the INV ratio in the model is to capture the effect of future productivity gains due to the investment that could improve CAD

[11] Duncan (2015) has addressed these issues. The author finds that productivity has become an important driver of the US current account. The issue of fiscal policy as a stabilization tool is discussed in Candelon and Lieb (2013). Briefly, they suggest that active spending policies have a stronger impact in times of recession.

[12] For the details of the model, see, e.g., Fidrmuc (2003).

[13] Parikh and Rao (2006) include real interest rate in the relationship. We consider interest rate and exchange rates in the preliminary analysis. Results are close to the one without the two variables, and hence, we choose not to report them here.

through larger exports. It is reasonable to expect a high correspondence between CAD and INV in developing countries and less so in more advanced countries. Finally, the appropriate method to estimate the model depends on the time-series property of the individual series.

To investigate the causal relationship(s) between variables of interest, we employ the procedure described by Konya (2006). To briefly demonstrate the application of the Granger causality test, consider the following system:

$$\begin{aligned}
\mathrm{CAD}_{1,t} &= \alpha_{1,1} + \sum_{l=1}^{ml\mathrm{CAD}_1} \beta_{1,1,l}\mathrm{CAD}_{1,t-l} + \sum_{l=1}^{ml\mathrm{BD}_1} \lambda_{1,1,l}\mathrm{BD}_{1,t-l} + \sum_{l=1}^{ml\mathrm{IINV}_1} \eta_{1,1,l}\mathrm{INV}_{1,t-l} + \varepsilon_{1,1,t} \\
\mathrm{CAD}_{2,t} &= \alpha_{1,2} + \sum_{l=1}^{ml\mathrm{CAD}_1} \beta_{1,2,l}\mathrm{CAD}_{2,t-l} + \sum_{l=1}^{ml\mathrm{BD}_1} \lambda_{1,2,l}\mathrm{BD}_{2,t-l} + \sum_{l=1}^{ml\mathrm{IINV}_1} \eta_{1,2,l}\mathrm{INV}_{2,t-l} + \varepsilon_{1,2,t} \\
&\vdots \\
\mathrm{CAD}_{N,t} &= \alpha_{1,N} + \sum_{l=1}^{ml\mathrm{CAD}_1} \beta_{1,N,l}\mathrm{CAD}_{N,t-l} + \sum_{l=1}^{ml\mathrm{BD}_1} \lambda_{1,N,l}\mathrm{BD}_{N,t-l} + \sum_{l=1}^{ml\mathrm{IINV}_1} \eta_{1,N,l}\mathrm{INV}_{N,t-l} + \varepsilon_{1,N,t}
\end{aligned} \tag{2}$$

and

$$\begin{aligned}
\mathrm{BD}_{1,t} &= \alpha_{2,1} + \sum_{l=1}^{ml\mathrm{CAD}_2} \beta_{2,1,l}\mathrm{CAD}_{1,t-l} + \sum_{l=1}^{ml\mathrm{BD}_2} \lambda_{2,1,l}\mathrm{BD}_{1,t-l} \times \sum_{l=1}^{ml\mathrm{IINV}_2} \eta_{2,1,l}\mathrm{INV}_{1,t-l} + \varepsilon_{2,1,t} \\
\mathrm{BD}_{2,t} &= \alpha_{2,2} + \sum_{l=1}^{ml\mathrm{CAD}_2} \beta_{2,2,l}\mathrm{CAD}_{2,t-l} + \sum_{l=1}^{ml\mathrm{BD}_2} \lambda_{2,2,l}\mathrm{BD}_{2,t-l} \times \sum_{l=1}^{ml\mathrm{IINV}_2} \eta_{2,2,l}\mathrm{INV}_{2,t-l} + \varepsilon_{2,2,t} \\
&\vdots \\
\mathrm{BD}_{N,t} &= \alpha_{2,N} + \sum_{l=1}^{ml\mathrm{CAD}_2} \beta_{2,N,l}\mathrm{CAD}_{N,t-l} + \sum_{l=1}^{mlBD_2} \lambda_{2,N,l}\mathrm{BD}_{N,t-l} \times \sum_{l=1}^{ml\mathrm{IINV}_2} \eta_{2,N,l}\mathrm{INV}_{N,t-l} + \varepsilon_{2,N,t}
\end{aligned} \tag{3}$$

where $CAD_{i,t-l}(l = 1, \ldots, ml\mathrm{CAD}_i)$, $\mathrm{BD}_{i,t-l}(l = 1, \ldots, ml\mathrm{BD}_i)$ and $\mathrm{INV}_{i,t-l}(l = 1, \ldots, ml\mathrm{INV}_i)$ denote the lagged values of the current account, budget and investment, respectively, for country $i = 1,\ldots,N$, and time period $t = 1,\ldots,T$. l refers to the lag and the residuals are supposed to be white noise that may be correlated for a given country, but not across countries. In the presence of contemporaneous correlation across countries, seemingly unrelated regression (SUR) estimator utilized.

The Konya's (2006) panel causality analysis has been applied by many studies (e.g., Menyah et al. 2014; Puente-Ajovín and Sanso-Navarro 2015; Zhang et al. 2016) to account for both cross-sectional dependence and heterogeneity across countries in the sample. These studies conclude that the choice of statistical tools in analyzing the causality among variables may play an important role in determining the causal direction (unidirectional, bidirectional or nonexistence) between the variables. With respect to these systems, in country i there is one-way Granger causality running

from BD to CAD if in Eq. (2) not all $\lambda_{1,i}$'s are zero, but in Eq. (3) all $\beta_{1,i}$'s are zero—TDH. There is one-way Granger causality from CAD to BD (called CAT) if in Eq. (2) all $\lambda_{1,i}$'s are zero, but in Eq. (3) not all $\beta_{1,i}$'s are zero. Equations (2) and (3) have different predetermined variables within SUR systems, and the bootstrap critical value unbiased to the time-series properties of the variables. This approach allows for different predetermined, i.e., lagged exogenous and endogenous variables, with different maximal lags for variables, but do not allow them to vary across countries. The lag length selection depends on Akaike information criterion (AIC) and Schwartz criterion (SC). The bootstrap distribution of the test statistic derived from 10,000 replications. For the details procedure to generate and the usage of the bootstrap samples, see Konya (2006, pp. 985–986).

There are several advantages of Konya's (2006) methodology. First, the methodology does not require testing the variables for unit root and cointegration, in this case, the variables being used in their levels. It is widely acknowledged that standard unit-root and cointegration tests suffer from lower power and different tests often lead to contradictory results. The testing procedure circumvents some pretesting biases that researchers may encounter with the standard unit-root and cointegration tests usually encountered in empirical work. Hence, Toda and Yamamoto (1995) as well as Konya (2006) approaches minimize the risk associated with identifying the orders of integration of the series, or the presence of cointegration that is often discussed in the empirical literature. Secondly, the tool considers the existence of contemporaneous correlations across countries and offers additional panel information. The standard panel approaches evaluate the causality for the whole panel, but are unable to reveal which cross-sectional unit is potentially driving the results. Konya approach addresses the issue by giving different critical values for each country.

We also use the procedure recently described by Emirmahmutoglu and Kose (2011). This new panel causality method combines the lag- augmented vector autoregressive (LA-VAR) approach of the Toda and Yamamoto (1995) in heterogeneous mixed panels. It is based on meta-analysis, and it is useful to investigate how the causality test performs in mixed panels involving cross-section units, which have different time-series characteristics. A level VAR with $k_i + d\max_i$ lags in heterogeneous mixed panels can present as:

$$z_{i,t} = \mu_i + A_{i1} z_{i,t-1} + \cdots + A_{ik} z_{i,t-k_i} + \sum_{l=k_i+1}^{k_i+d\max_i} A_{il} z_{i,t-l} + u_{i,t} \qquad (4)$$

where μ_i is a p-dimensional vector of fixed effects, $A_{i1}, \ldots, A_{ik}$ are fixed $p \times p$ matrices of coefficients to be estimated that are allowed to vary across cross-section units and $u_{i,t}$ is a column vector of p error terms for $i = 1, 2, \ldots, N, t = 1, 2, \ldots, T$. The lag structure (k_i) is selected based on SIC, and it allowed to differ across cross-section units. A joint restriction test, which has an asymptotic Chi-square distribution on the coefficient of VAR system, used to determine whether, e.g., CAD does not Granger-cause BD. Emirmahmutoglu and Kose (2011) employ a Fisher's (1932) approach to conduct the Granger non-causality hypothesis in heterogeneous mixed

panels. Fisher (1932) combined several significant levels of identical, but independent tests. The Fisher test statistic (λ) is given as the following:

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

where p_i is the p value corresponding to the Wald statistic of the $i-$th individual cross-section units. The test statistic has a Chi-square (χ^2) distribution with $2N$ degree of freedom under the cross-section independence assumption. The test is valid for $T \rightarrow \infty$ with N fixed. In what follows, the bootstrap empirical distribution of the Fisher test statistics is derived from 10,000 replications to avoid serious size distortion for a small value of T. The bootstrap critical value is specified by selecting the appropriate percentiles of the sampling distributions. The Emirmahmutoglu–Kose test is very powerful even if N and T are small under both cross-section independence and the cross-section dependency. However, the approach has serious size distortion when T is small under the cross-section dependency assumption. The limit distribution of the Fisher test statistic is not valid in the presence of cross-section dependence. Therefore, the bootstrap method is utilized here to generate the empirical distributions of Fisher test. Details of the bootstrap procedure are given in Emirmahmutoglu and Kose (2011) and Xie and Chen (2014).

4 Empirical results

Our empirical analysis employs data for the period 1980–2013 (34 annual frequency) in 14 Asian countries, including China (CHN), Indonesia (IDN), the Philippines (PHL), Thailand (THA), Malaysia (MYS), Pakistan (PAK), Sri Lanka (LKA), Nepal (NPL), India (IND) and five advanced economies—Japan (JPN), Hong Kong SAR (HKG), South Korea (KOR), Singapore (SGP) and Taiwan (TWN). We make use of annual data because of unavailability of higher frequency in most of the countries being studied. Unlike most of the literature that omits the crisis of late 2000, we extend the sampling period to capture the historical events after the AFC. The series of current account balance/GDP (CAD/GDP) (surplus=+, deficit=−) and investment/GDP (INV/GDP) are drawn from World Development Indicators (WDI), and budget balance/GDP (BD/GDP) (surplus=+, deficit=−) was collected from the Asian Development Bank.[14] All series are constructed as GDP ratio, and Table 1 displays the summary statistics for the three series—CAD, BD and INV. Based on the median as the cutoff point, we notice that half of the sample of countries incurred deficits (India, Indonesia, Sri Lanka, Nepal, Pakistan, the Philippines and Thailand). Countries with surpluses in their current account include Malaysia, China, plus the five advanced economies (Japan, Hong Kong SAR, South Korea, Singapore and Taiwan).[15]

[14] The data for Taiwan are sourced from *DataStream*.

[15] China has a trade surplus, mainly with the USA and the European Union. However, it has a relatively large trade deficit with some Asian countries and regions the main origin of its FDI.

We begin by investigating the integration properties of BD, CAD and INV (recall that these are relative to GDP). From a modeling perspective, the order of integration will indicate if the classical regression approach (all variables are $I(0)$) or the cointegration analysis should be considered instead (if they are $I(1)$). Given the short span of the time series, panel unit-root tests have been used as a way to increase test power by exploiting the information from the cross-sectional units. For example, Pesaran (2004) develops a panel unit-root test that takes account of cross-sectional dependence. In our empirical analysis, we pretested our variables by testing the null hypothesis of no cross-sectional dependence against the alternative of cross-sectional dependence as suggested by Pesaran (2004).[16] As expected, the test statistics for all the series are significant at the 1% significance, suggesting a high degree of cross-sectional dependence among panel members due to their trade dependence or common random shocks. The existence of cross-sectional dependency indicates that a shock to the current account in a country is likely to affect the other countries in the panel—one of the points stressed by Xie and Chen (2014). Following Pesaran (2007), we applied the cross-sectionally augmented Im et al. (2003) test that allows for cross-sectional dependence (columns labeled CIPS) to enhance the power of the tests and report the empirical results in Table 1. The CIPS statistics for the series are statistically significant at the 10% significance level or better, suggesting all the series in our model can be characterized as $I(0)$ variable.[17] In contrast, the standard univariate and panel unit-root tests reveal just the opposite—all the variables are all nonstationary in their level.[18] Unlike the standard panel regressions, the CIPS test uses a continuous time-varying intercept term to accommodate the changing macro fundamentals in the countries under investigation.[19]

The above findings also validate the notion that both the external and internal balances in the Asian countries are on a sustainable path, at least for the sample period that ended in 2013. Thus, contrary to what some others (Khalid and Guan 1999; Baharumshah et al. 2006) have claimed, the current account and control variables follow a nonstationary process. It should be noted that the earlier studies are based on traditional unit-root tests which can have low power to detect a stationary series. When the variables are stationary, a spurious cointegrating relationship may exist between CAD and its determinants using cointegration techniques (Grier and Ye

[16] Pesaran (2004) develops a modification of the Bruesch–Pagan LM test (CD_{lm}) and employs a simple average of all pair-wise correlations of individual regression residuals from a panel. This test can be applied to any value of N and T and is correctly centered. In addition, this test is shown to be robust for multiple breaks and/or error variance as long as the unconditional means of the variables in the panel remain constant over time.

[17] Given that these series are expressed as fractions of GDP, they do not intrinsically seem likely to be unit-root processes. Moreover, it widely acknowledges in the literature that conventional unit-root tests have very low power when the sample size is small or structural breaks appear in the data.

[18] We also pretest the FDI/GDP series, and the series is also found stationary at level.

[19] We note here that recent papers (e.g., Lee et al. 2016) have argued that breaks can be smooth and ignoring them might biased estimates. Lee et al. (2016) suggest a test that accommodates cross-sectional dependency among the variables and smooth structural change in the deterministic components. The sizes (powers) of the test are good when $T \geq 50$ ($T \geq 100$). Unfortunately, our sample size is too small to conduct this test. We are grateful to a referee for pointing out the misspecification risk associated with the other tests to us.

Table 1 Descriptive statistics and panel unit-root test

		CAD/GDP		BD/GDP		INV/GDP	
		Mean	Median	Mean	Median	Mean	Median
IDN		−0.005	−0.016	−0.011	−0.011	0.256	0.255
PHL		−0.005	−0.004	−0.019	−0.020	0.215	0.207
THA		−0.009	−0.014	−0.042	−0.041	0.295	0.277
MYS		0.033	0.045	−0.041	−0.036	0.289	0.263
PAK		−0.021	−0.024	−0.057	−0.058	0.163	0.169
LKA		−0.062	−0.061	−0.087	−0.082	0.250	0.248
NPL		−0.018	−0.022	−0.039	−0.046	0.199	0.199
IND		−0.014	−0.014	−0.074	−0.072	0.247	0.237
CHN		0.022	0.018	−0.019	−0.019	0.352	0.341
HKG		0.047	0.056	0.010	0.008	0.254	0.254
JPN		0.025	0.027	−0.043	−0.042	0.255	0.274
KOR		0.014	0.016	0.007	0.010	0.304	0.305
SGP		0.113	0.135	0.064	0.064	0.327	0.322
TWN		0.071	0.068	−0.043	−0.041	0.237	0.240
Panel unit-root test							
PCD		19.39 [0.000]		6.17 [0.000]		6.15 [0.000]	
	corr(p)	0.349		0.111		0.111	
CIPS	φ	−2.770 [0.003]		−1.933 [0.027]		−2.503 [0.006]	
	τ	−0.807 [0.210]		−0.387 [0.349]		0.099 [0.539]	

The Pesaran's (2004) PCD test is based on the residual cross-correlation of the augmented Dickey–Fuller (p) regressions, where p denotes the lag. The test follows a standard normal distribution under the null hypothesis of cross-sectional independence. Reject H_0 when the test statistic, PCD $= \sqrt{2T/N(N-1)}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij} > N(0,1)$. φ and τ refer to specification without trend and with the trend, respectively. The values in [ ] refer to p value

2009).[20] The fact that the indicators are characterized by stationary processes reflects the flexibility—countries which pursue sustainable fiscal policies also display a flexibility of the external account (Fidrmuc 2003, p. 149). The mean reverting behavior of CAD and BD provides little evidence of insolvency of the internal and external balances, at least for the sample period examined here. As explained earlier, policymakers may consider appropriate monetary and fiscal policies to counter large (temporary) imbalances in fiscal and current accounts for fear of default or to avoid a hard landing.[21] It should be mentioned that the non-lasting effects of fiscal shocks will have a

[20] Precisely, Grier and Ye (2009) noted that the statistical robust relationship between budget and trade deficits is difficult to untangle largely because of ignoring structural break in the series.

[21] The evidence suggests the current account of most Asian countries is on a sustainable path and is consistent with the long-run intertemporal budget constraint. This is also in line with the findings of Lau et al. (2006) and Kim et al. (2009). It should be mentioned that domestic banks hold high shares of government debt in a number of Asian economies, including the China (77%), Malaysia (41%), Indonesia (39%) and Japan (38%). These countries face the greatest risk of a "doom loop" cycle developing in the

strong negative effect on investment, thereby muting the effect fiscal expansion on the trade balance.

Next, we apply the SUR estimator to estimate the parameters of the model to highlight individual country experiences rather than Asian-wide trends. An important issue that has been ignored up to now is the impact of the AFC and the sharp depreciation of the currencies as well as regime changes corresponding to the episode on the underlying relationship. To probe further into this important issue, the Quandt–Andrews structural breakpoint test is applied to check whether there is a structural change in the subset of the parameters of the model. The test statistics along with the approximate asymptotic p values based on the Quandt–Andrews procedure as outlined in Hansen (1997) are presented in Table 2. As indicated in the table, the break is mainly located in the mid- and late 1990s (Indonesia, Malaysia, Thailand, Sri Lanka, Nepal, China, South Korea and Singapore) and early 2000s (the Philippines, Pakistan, Hong Kong, Japan and Taiwan).[22] We also observed from the data that a significant number of the Asian EMEs generally been running current account surpluses after having suffered currency crises.

Two points in Table 2 are worth highlighting here. First, the Asian crisis marked the beginning of a sequence of financial crisis among the EMEs. It also led to major structural changes in the exchange rate regimes and foreign exchange rate reserves in the EMEs. Second, the different timing of the break dates for the group of Asian countries largely to different economic policy responses to the economic turmoil that started in the late 1990s. As already mentioned, failure to account for the sterilization actions and other policy measures in this period may bias the empirical findings. Notice that all the statistics (Sup/MaxF, ExpF, and AveF) reject the null hypothesis of no structural breaks within the 15% trimmed data at the 10% significance level or better. This has prompted us to include an additional variable in our model—crisis dummy ($D = 1$ after the break date and zero otherwise).

Empirical results from running the regression model are presented in Table 3.[23] The size for the coefficient of the BD is around 0.76 in pre-crisis (period I), and it increases to 0.89 in the post-crisis periods. Meanwhile, the impact of investment on CAD has changed significantly. Notice that the size of the investment coefficient dropped sharply from 0.27 to 0.09 in the second sample period, suggesting the impact of investment on CAD has declined drastically after the crisis. Focusing on the individual countries (Panel B), there is a positive, statistically significant effect of BD on CAD (10% or better) in nine countries: Malaysia, Pakistan, Sri Lanka, Nepal, China, Japan, South Korea, Singapore and Taiwan. For these countries, a 1% increase in BD is associated with an about 0.36% (Nepal) to 1.05% (China) increase in CAD. Notice that for Thailand the coefficient of the BD variable is negative (-0.64), implying that a 1% increase in BD is associated with a declined in CAD by 0.64% of GDP. The

future and therefore need to strengthen frameworks for fiscal sustainability and diversify the holders of government debts.

[22] For India, we find a break in 1991 that is associated with the balance of payment crisis that surfaced in early 1990s, but it was found to be insignificant.

[23] The CD_{lm} (test statistic = 580.89; p value = 0.0000) is significant at 1%, suggesting that contemporaneous correlation between equation errors exists in our panel model.

Table 2 Parameter and structural instability test

	Break date	Quandt–Andrews unknown breakpoint tests statistics		
		Sup/MaxF	ExpF	AveF
IDN	1998	14.192[b] (0.044)	5.581[b] (0.016)	8.878[a] (0.006)
PHL	2004	12.980[c] (0.071)	4.877[b] (0.029)	7.540[b] (0.018)
THA	1998	28.160[a] (0.000)	11.006[a] (0.000)	5.851[b] (0.020)
MYS	1998	48.436[a] (0.000)	21.131[a] (0.000)	24.010[a] (0.000)
PAK	2000	17.280[b] (0.012)	6.044[b] (0.011)	7.165[b] (0.023)
LKA	1997	24.173[a] (0.000)	10.025[a] (0.000)	13.743[a] (0.000)
NPL	1997	23.654[a] (0.001)	9.516[a] (0.001)	8.663[b] (0.008)
IND	1991	5.510 (0.481)	1.620 (0.272)	2.924 (0.188)
CHN	1997	12.091 (0.100)	4.352[b] (0.047)	7.265[b] (0.022)
HKG	2002	8.535 (0.168)	2.832[c] (0.074)	3.305 (0.139)
JPN	2001	30.161[a] (0.000)	13.054[a] (0.000)	18.752[a] (0.000)
KOR	1997	32.835[a] (0.000)	13.554[a] (0.000)	12.906[a] (0.000)
SGP	1998	26.239[a] (0.000)	11.083[a] (0.000)	15.530[a] (0.000)
TWN	2002	16.114[b] (0.020)	6.456[a] (0.008)	7.961[b] (0.013)

The null hypothesis for structural stability test is no structural breaks within the possible dates (trimmed data). The Sup/MaxF is Sup or Maximum statistic, ExpF is Exp statistic and AveF is Ave statistic (See Andrews 1993; Andrews and Ploberger 1994). The values in () refer to *p* value obtained with Hansen's (1997) method. The break date is selected based on maximum statistic with 15% trimming
[a], [b] and [c] indicate statistically significant at the 1, 5 and 10% significance level, respectively

negative correlation between budget deficits and current account imbalance, however, is inconsistent with the TDH. For Hong Kong SAR, the estimates reveal that BD does not even enter in the regression, therefore rejecting the existence of the twin dilemma. It confirms the findings report in previous studies using different empirical analyses; see Bussiére et al. (2010) for the data from G-7 countries.

Turning to the investment variable, we notice that it has a statistically negative effect on CAD for 12 out of 14 countries. Clearly, this supports the idea that the collapse of investment after AFC has contributed to the current account surplus. In most cases, the magnitude of the investment variable is significantly larger than BD and in some cases (e.g., Malaysia, Hong Kong SAR and Taiwan) close to unity. This finding echoes Ito's (2009) view: The current account surplus in the Asian countries was largely due to the investment drought. However, there is one counter example and that is China that contributes to the growing global current surplus starting after AFC due to over-saving. Precautionary saving behavior attributed to a relatively low level of financial development can potentially explain the empirical fact. The greater macroeconomic uncertainty (especially after the AFC), the greater will be the incentive for precautionary saving, ceteris paribus, and hence the larger the current account surplus. Finally, we find that the BD coefficient falls between the 0 to 1 range and it fits well with the Makin's (2004) medium-run open macroeconomic model. Fiscal expansion has no effect on trade balance (current account) if the economy is not open

Table 3 SUR estimation for 14 Asian countries

	Model: $CAD_{it} = \beta_1 + \beta_2 BD_{it} + \beta_3 INV_{it} + \varepsilon_{it}$					
	Panel A: without break			Panel B: with break		
	$\hat{\beta}_{1i}$	$\hat{\beta}_{2i}$	$\hat{\beta}_{3i}$	$\hat{\beta}_{1i}$	$\hat{\beta}_{2i}$	$\hat{\beta}_{3i}$
IDN	0.082[a]	− 0.107	− 0.347[a]	0.080[a]	− 0.030	− 0.411[a]
PHL	0.146[a]	− 0.067	− 0.710[a]	0.075[a]	− 0.381[b]	− 0.475[a]
THA	0.178[a]	− 0.628[a]	− 0.723[a]	0.143[a]	− 0.640[a]	− 0.648[a]
MYS	0.414[a]	0.658[a]	− 1.224[a]	0.325[a]	0.537[a]	− 1.005[a]
PAK	0.130[a]	0.575[a]	− 0.730[a]	0.135[a]	0.681[a]	− 0.691[a]
LKA	0.161[a]	0.734[a]	− 0.633[a]	0.135[a]	0.702[a]	− 0.596[a]
NPL	− 0.007	1.168[a]	0.177	− 0.026	0.363[b]	− 0.057
IND	0.001	− 0.232[b]	− 0.133[a]	0.016	− 0.188[b]	− 0.234[a]
CHN	− 0.011	0.921[a]	0.144[a]	0.028	1.054[a]	− 0.016
HKG	0.328[a]	0.107	− 1.110[a]	0.364[a]	0.108	− 1.236[a]
JPN	0.025[b]	0.062	0.009	0.004	0.113[b]	− 0.072[b]
KOR	0.128[a]	0.833[a]	− 0.393[a]	0.177[a]	0.676[a]	− 0.586[a]
SGP	0.351[a]	0.528[a]	− 0.831[a]	0.276[a]	0.674[a]	− 0.723[a]
TWN	0.340[a]	0.314[b]	− 1.084[a]	0.436[a]	0.546[a]	− 1.394[a]
Panel	0.101[a]	0.825[a]	− 0.246[a]	0.071[a]	0.795[a]	− 0.207[a]
Period I	0.085[a]	0.758[a]	− 0.272[a]			
Period II	0.086[a]	0.889[a]	− 0.088[a]			
	Diagnostic tests			Diagnostic tests		
R^2	0.8246			0.9674		
Aj$-R^2$	0.8238			0.9669		
D-W	1.1544			1.7756		

Iterate coefficient after one-step weighting matrix and the coefficient covariance method is based on cross-section SUR (PCSE). D–W refers to Durbin–Watson *d* statistic. The dependent variable is CAD (see Eq. 1). Break date from Table 2

[a], [b] and [c] indicate statistically significant at 1, 5 and 10% significance level, respectively

to trade and/or fiscal shocks are not too persistent. In other words, the crowding-out effect of fiscal shocks on private investment is much stronger when the two conditions are at work (a pointed out by Corsetti and Müller 2006). Finally, we also replace investment with FDI net inflows and find that the major results tend to replicate the main results reported in Table 3.

Before we present the results of the Toda and Yamamoto (1995) causality test based on LA-VAR, two points are worth mentioning. First, we split the sample into two sub-periods, given the large swings in the variables under investigation; see Quandt–Andrews test (Table 2).[24] Second, we apply the Hausman test to determine whether a

[24] The literature on TDH in Asian countries is based on the pre-crisis period, which correspond to period I of our sample. Given the small size of our sample, we restrict to only single break test.

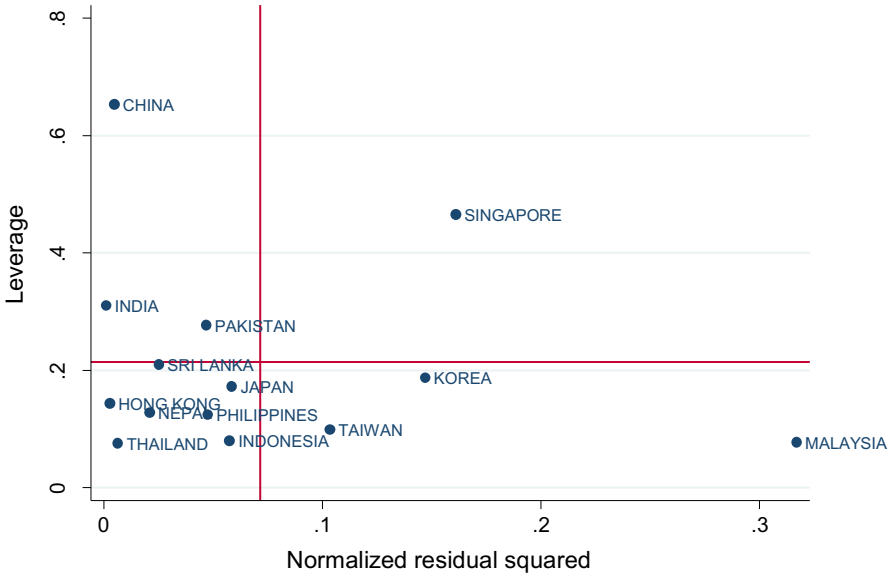

Fig. 1 Scatter plot of leverage and residual squared (post-crisis). Cook's Distance (or Cook's D) shows Singapore is an outlier

set of homogeneous coefficients applies across all the countries in the panel. Results in Table 3 show that the null hypothesis of cross-sectional homogeneity is not rejected for the full sample period [test statistic = 2.92; p-value = 0.2326] and period I [test statistic = 3.30; p value = 0.1918]. However, the test rejects null for period II [test statistic = 5.72; p value = 0.0574], meaning that the data reject the restriction of a common set of coefficients among the panel members. There are severe limitations associated with the test as discussed in the literature, and for completeness, we also run the bias-adjusted cross-sectional dependency and homogeneity tests due to Pesaran et al. (2008) and Pesaran and Yamagata (2008) to the same data set. The bias-adjusted test is a modified version of the LM test by using the exact mean and variance of the LM statistics. The null hypothesis of no cross-sectional dependency [test statistic = 10.146; p value = 0.000] is overwhelming rejected by the test statistic. This means that any shock from one country is transmitted to another in the panel.[25] Further analysis based on Cook's distance (or Cook's D) reveals that Singapore is an outlier (see Fig. 1). In what follows, we exclude Singapore from the panel for robust checks. The unreported results do not change in any significant way.

[25] We also conduct the slope homogeneity test for the full sample period, but the outcome from the delta test [test statistic = −0.441; p value = 0.671] and the bias-adjusted version test statistic [test statistic = −0.462; p value = 0.678] of Pesaran and Yamagata (2008) are statistically insignificant even at the 10% level. Many authors (e.g., Emirmahmutoglu and Kose 2011) have argue that countries differ in their institutional, political and economic structure. Besides that, these countries also differ in their reactions to external shocks. For these reasons, we apply the Konya's bootstrap panel causality test to determine the relationship between two deficits.

Table 4 Granger causality test with bootstrapping

H_0	BD↛CAD			INV↛CAD			CAD↛BD			INV↛BD			CAD↛INV			BD↛INV		
	τ			τ			τ			τ			τ			τ		
Panel A: Period I (pre-crisis)																		
CHN	−	0.004	(0.947)	−	1.112	(0.292)	+	3.919	(0.048)	+	0.041	(0.839)	+	5.012	(0.025)	+	1.399	(0.237)
HKG	−	0.900	(0.343)	−	8.478	(0.004)	+	11.820	(0.001)	−	6.220	(0.013)	−	6.355	(0.012)	+	28.837	(0.000)
INDIA	+	6.689	(0.010)	−	3.904	(0.048)	+	0.174	(0.677)	+	1.029	(0.310)	+	4.841	(0.028)	+	0.001	(0.971)
INDO	+	16.704	(0.000)	+	0.953	(0.329)	−	1.070	(0.301)	+	10.534	(0.001)	+	38.145	(0.000)	+	6.298	(0.012)
JAPAN	+	27.355	(0.000)	−	30.250	(0.000)	−	49.238	(0.000)	−	56.919	(0.000)	−	2.209	(0.137)	+	15.685	(0.000)
KOR	+	24.699	(0.000)	−	3.632	(0.057)	+	0.730	(0.393)	+	0.020	(0.886)	+	11.718	(0.001)	−	0.008	(0.927)
MYS	+	4.687	(0.030)	−	10.843	(0.000)	+	0.273	(0.602)	+	0.108	(0.743)	+	23.333	(0.000)	−	1.901	(0.168)
NEPAL	+	8.008	(0.005)	−	28.170	(0.000)	+	0.715	(0.398)	+	0.057	(0.811)	+	0.349	(0.555)	+	0.867	(0.352)
PAK	−	3.073	(0.080)	−	2.198	(0.138)	−	4.130	(0.042)	+	19.972	(0.000)	+	0.441	(0.507)	−	12.919	(0.000)
PHL	+	24.649	(0.000)	−	3.280	(0.070)	−	1.047	(0.306)	+	0.001	(0.994)	−	5.578	(0.018)	−	1.134	(0.287)
LKA	−	0.256	(0.613)	+	0.978	(0.323)	+	2.108	(0.147)	−	0.075	(0.784)	−	43.765	(0.000)	+	35.622	(0.000)
THAI	+	3.538	(0.060)	−	0.182	(0.670)	−	0.898	(0.343)	−	287.440	(0.000)	+	26.166	(0.000)	−	0.035	(0.849)
TWN	+	22.682	(0.000)	−	21.635	(0.000)	+	2.182	(0.140)	+	2.485	(0.115)	+	2.424	(0.120)	−	8.379	(0.004)
SGP	+	6.061	(0.014)	−	1.256	(0.262)	+	0.596	(0.440)	−	1.370	(0.242)	−	3.369	(0.066)	+	5.213	(0.022)

Table 4 continued

H_0	BD↛CAD			INV↛CAD			CAD↛BD			INV↛BD			CAD↛INV			BD↛INV		
	τ			τ			τ			τ			τ			τ		
Panel B: Period II (post-crisis)																		
CHN	+	22.034	(0.000)	−	5.333	(0.021)	+	8.075	(0.004)	+	9.165	(0.002)	−	0.085	(0.770)	+	0.262	(0.608)
HKG	−	1.336	(0.248)	−	22.565	(0.000)	+	5.865	(0.015)	−	2.862	(0.091)	−	11.284	(0.001)	+	7.025	(0.008)
INDIA	−	1.386	(0.239)	−	32.415	(0.000)	+	2.950	(0.086)	−	0.206	(0.650)	+	15.452	(0.000)	−	0.047	(0.829)
INDO	+	1.475	(0.225)	−	8.099	(0.004)	+	5.459	(0.019)	−	4.323	(0.038)	−	9.567	(0.002)	+	0.608	(0.434)
JAPAN	−	0.400	(0.527)	−	4.249	(0.039)	+	4.419	(0.036)	+	1.998	(0.157)	+	101.218	(0.000)	−	35.862	(0.000)
KOR	−	2.051	(0.152)	−	0.006	(0.938)	+	3.156	(0.076)	−	28.746	(0.000)	−	96.389	(0.000)	−	7.665	(0.006)
MYS	−	1.677	(0.195)	−	0.036	(0.850)	+	7.635	(0.006)	−	1.792	(0.181)	−	11.326	(0.001)	−	0.873	(0.350)
NEPAL	−	3.842	(0.050)	−	17.336	(0.000)	+	2.060	(0.151)	−	0.035	(0.852)	+	12.268	(0.000)	+	25.434	(0.000)
PAK	−	0.406	(0.524)	−	1.527	(0.217)	+	31.555	(0.000)	+	15.986	(0.000)	−	1.040	(0.308)	+	17.535	(0.000)
PHL	−	0.003	(0.956)	−	72.785	(0.000)	+	18.693	(0.000)	−	2.962	(0.085)	−	33.467	(0.000)	−	0.252	(0.616)
LKA	−	60.633	(0.000)	−	56.365	(0.000)	+	7.815	(0.005)	+	0.152	(0.696)	+	42.550	(0.000)	+	34.606	(0.000)
THAI	+	0.509	(0.475)	−	0.950	(0.330)	+	28.703	(0.000)	+	75.951	(0.000)	−	4.523	(0.033)	−	1.787	(0.181)
TWN	+	0.004	(0.948)	−	9.127	(0.003)	+	5.515	(0.019)	+	9.709	(0.002)	−	0.558	(0.455)	+	6.448	(0.011)
SGP	−	107.2	(0.000)	−	64.628	(0.000)	+	37.515	(0.000)	−	0.194	(0.660)	+	12.543	(0.000)	+	45.049	(0.000)

The null hypothesis BD↛CAD represents budget balance does not Granger-cause current account balance. Figure τ refers to the sign for test value. The Granger causality test is based on SUR systems advocated by Konya (2006). The bootstrap distribution of each Wald test statistic with country-specific bootstrap critical values is derived from 10,000 replications. Period I covers from 1980 to 1997. Refer to table above. The sample for period II covers from 1998 to 2013.

Table 5 Emirmahutoglu and Kose's (2011) LA-VAR Granger causality test

Lag orders are selected by minimizing the Schwarz Bayesian criteria. The bootstrap distribution of the Fisher test statistics is derived from 10,000 replications [a], [b] and [c] indicate statistically significant at the 1, 5 and 10% significance level, respectively

	λ statistic	Bootstrapped critical value		
		1%	5%	10%
Period I: Pre-crisis				
BD↛CAD	68.710[c]	92.207	70.312	61.949
FDI↛CAD	67.311[b]	84.987	65.762	57.806
CAD↛BD	47.011	93.319	70.964	62.152
Period II: Post-crisis				
BD↛CAD	36.521	102.820	74.281	63.664
FDI↛CAD	32.598	108.445	74.234	63.470
CAD↛BD	74.961[c]	107.696	77.055	66.016

Turning to the specific results, first, we observe that there is no clear nexus between BD and CAD for the full sample (not reported). The MWALD test statistics for BD↛CAD and CAD↛BD are both statistically insignificant in most of the sample countries. The TDH is confirmed for Taiwan, whereas the CAT (reverse causality) is revealed by the Philippines, Thailand, Sri Lanka and Nepal. For these South Asian countries, Ravinthirakumaran et al. (2016) show the direction of causality is mixed. When the sample is split into pre-crisis (period I) and post-crisis (period II) periods, however, we untangle a direct relationship between the two variables for the first period (panel A, Table 4): BD↛CAD. In other words, it supports the notion that the two deficits are twined, except for three countries (China, Hong Kong SAR and Sri Lanka). In the case of China and Hong Kong, the results confirm a one-way causality running from CAD to BD. In contrast, our results decisively reject the notion the two deficits are a twin in 13 (out 14 countries) after the crisis (panel B, Table 4), but instead we find the two deficits are connected (positive) through the CAT hypothesis. The sole exception is Nepal, even with the updated data. Second, the investment (negatively) Granger caused CAD for eight countries at the indicated significance levels before AFC increased to 10 countries in the second sample period. With few exceptions, we detect a one-way Granger causality between investment and current account.

For robustness check, some studies (e.g., Algieri 2013) have considered alternative causality test to the same data set. The purpose is to determine whether the results remain the same, regardless of the causality technique used in the analysis. Next, we apply the Emirmahmutoglu and Kose's (2011) causality test and the main results are reported in Table 5. Analysis based on the panel causality test confirmed that the causality in the period after AFC runs from CAD to BD. This finding is consistent with the idea put forth by Ito (2009) and Corsetti and Müller (2006): The current account surplus in most of the emerging Asian countries is due to drop in investment, largely to the reaction of excessive investment activities prior to the eruption of the AFC in the late 1990s. Given that the saving rates have not changed much after AFC, they all started to run large external surpluses. In the case of Japan, investment falls steadily due to the bubble years. It is worth noting that Japan's experience in the 1990s is often cited as evidence that changes in private saving can offset changes in fiscal policy, leaving Japan's CAD largely unaffected (mentioned in Bartolini and Lahiri 2006).

Hatemi-J and Shukur (2002) also find that the TDH is supported for the period 1975–1989 for the USA through Keynesian hypothesis, but the twin deficits are positively linked via CAT in the later period (1990–1998). The changing pattern in the causal relationship between budget and current account deficits is also emphasized in Nikiforos et al. (2015). In particular, they find that the shift is from the nonexistent relationship (the Ricardian hypothesis) to the CAT type of the regime for Greece after 1995.

Focusing on China, the current account surplus, we observe that the relationship changed from CAT to two-way causality between budget and current account balances. The significance of this finding is that it suggests fiscal policy is not an instrument to help to unwind the large external imbalances. This is because the budget deficit is not a fully controlled (exogenous) variable as far as dealing with the imbalance problem with the USA. For India, the inter-linkages between the two deficits also found to be time dependent; TDH is observed in the earlier sample period, while strong support for CAT is found in more recent years. This indicates that the fiscal stimulus packages are introduced by the Indian government to lessen the economic difficulties to mitigate the consequences of large trade imbalances in that period. Finally, for robust checks perform by adding two control variables (interest rate and exchange rate) in the VAR system, and our main results (unreported) with regard to the causal ordering remain broadly unaffected.

4.1 Do foreign capital inflows affect current account imbalances?

We now shift the focus to the role of foreign capital on current account imbalances. The Asian countries under review relied heavily on foreign savings to finance their current account deficits prior to the 1997–1998 AFC. Unlike the first, the second waves of foreign capital flow to the EMEs after the crisis has resulted in a reversal of current account toward surplus (International Monetary Fund 2007). This section specifically looks at the role of capital inflows on current account imbalances in order to address some of the key hypotheses discussed in the literature. To this end, we replaced the investment variable in Eq. (1) with FDI inflows and conduct the causality test with FDI net inflows as a control variable in CAD–BD nexus.[26] There is no clear causal relationship running from BD and CAD in two subsample periods; instead, we observed a reverse causality in period II. Next, we replicate the work of Doraisami (2007) and Yan and Yang (2012) and find unlike these earlier studies, the results are a function of the different time-series method.[27] Given the evidence of the cross-sectional dependence (particularly in period II), the LA-VAR bootstrap Granger causality test advocated by Emirmahmutoglu and Kose (2011) is employed here. As mentioned earlier, the test is an extension of the Toda and Yamamoto (1995) via meta-analysis to test for Granger

[26] The net foreign direct investment is sourced from WDI, while data for Hong Kong SAR and Taiwan are drawn from UNCTAD.

[27] The EMEs consider by Yan and Yang (2012) include South Africa, Argentina, Brazil, Chile, Colombia, Egypt, Greece, India, Indonesia, Israel, Jordan, South Korea, Mexico, Morocco, Pakistan, Peru, the Philippines, Sri Lanka, Taiwan, Thailand and Turkey.

causality in heterogeneous mixed panels.[28] The bootstrap distribution of critical values is derived from 10,000 replications. The results are reported in Table 5. The major findings remain essentially unchanged. One should note, however, this is a bivariate analysis. Our findings are fairly robust: The new test has not reversed the underlying relationship between CAD and BD in sub-period I—TDH (either direct or indirectly), and sub-period II—CAT hypothesis. Again, the reverse causality that prevails after the crisis has been detected earlier is confirmed by the Emirmahmutoglu and Kose's 2011 panel causality bootstrap test statistics. Briefly, the bootstrap approach is similar in approach to Konya (2006). With regard to the power of the test, it tends to be more powerful even when N and T are small. Moreover, Emirmahmutoglu and Kose have shown that results drawn from bootstrap critical values are broadly different from the asymptotic Chi-square critical value.

The other interesting findings offered when the bootstrap distribution of the Fisher test statistics are applied are: first, a causal relationship that runs from FDI to CAD prior to the AFC. The evidence found in others such as Doraisami (2007) and Yan and Yang (2012) is supportive of this finding. Doraisami (2007) relied on bivariate, while Yan and Yang (2012) later consider the multivariate model. Yan and Yang (2012) found foreign capital inflows "push" the CAD toward an imbalance, and the capital flows have a positive causal relationship on CAD after the crisis. The surge in international flows to developing countries coincided with the widening in the CAD. Widening CAD is one of the less desirable macroeconomic effects of large capital inflows to the debtor countries (Calvo et al. 1996). Doraisami (2007) forcefully argues that countries should exercise caution and not abruptly remove their restrictions on capital mobility, especially without sophisticated and sound the financial system. Second, the Fisher test (λ) statistic (not reported) confirmed BD Granger-causes both FDI and investment in sub-period II (significant at the 10% level or better). For the EMEs where the financial systems are not highly sophisticated, perfect capital mobility is susceptible to exchange rate overshooting, and according to Bacchetta and Wincoop (2000), this could lead to high volatility of capital inflows. Our results seem to endorse the notion that Asian countries more susceptible to the effects of FDI inflows before than after the AFC. It also points to an important fact: Ignoring break might underestimate the role of some important drivers of current the current account. Another striking feature of our results is that we find no direct causal relationship between FDI and CAD in sub-period II. The results do not support the view that current account imbalances are due to foreign capital inflows in the later sample period.

As shown in Table 5, capital flows have changed the CAD dynamics through their effect on BD after the AFC. Specifically, large budget imbalances affect the capital flows and domestic investment which in turn worsened the external balances.[29] Policymakers should caution the macroeconomic effect of imbalances in the budget on FDI

[28] This is a bootstrap method to generate the empirical distributions of Fisher test. Simulation analysis showed the test under both the cross-section independency and the cross-section dependency are very powerful even if N and T are small.

[29] The current account surplus in China and the other Asian countries are known to be driven by other factors (e.g., distortions in factor markets and prices, a weak financial sector and investment-oriented government policies). The issue, however, is beyond the scope of this study.

and investment after the AFC. Massive modifications in discretionary fiscal policies to adjust external imbalances following the crisis may have contributed to this finding. Our results raise some concerns with regard to the fiscal stimulus packages launched after the GFC to jumpstart consumer demand and business investment. Eventually, a prolonged fiscal deficit will adversely affect either crowding in or crowding-out investment and in turn the current account imbalances (trade balances); see Abell (1990) on these issues. Clearly, macroeconomic policy should be aimed at fiscal austerity to reduce the BD to avoid a hard landing.

Yan and Yang (2012), for instance, empirically verify that it is mostly true that foreign capital inflows Granger-cause current account imbalances in the EMEs. However, no significant causal relationship is detected running from FDI to CAD (or the reverse) for sub-period II. According to Wong and Carranza (1999), the current account is susceptible to the influence of foreign capital inflows when capital account is liberalized and the domestic financial system is unsophisticated; see also Yan and Yang (2012). It also supports the notion that foreign capital, prior to the crisis, could not provide a viable source of financing domestic capital under liberalize capital account with pegged exchange rate regime since it is likely to be less susceptible (prone) to currency crisis compared to other exchange regimes.

5 Summary and conclusion

After examining the interaction between the current account deficit, budget deficit and investment for a sample of Asian countries separately and as a group, we find that the impact of investment on current account deficit is significant and comparatively larger than that of the budget deficit, both before and after the AFC. This point has been stressed in Chinn and Prasad (2003) and Papadogonas and Stournaras, (2006), among others, who have argued that only a small fraction of a budget deficit is reflected in the current account balances. The finding base on two alternative panel causality tests suggests that shocks originating from public activities may not be sufficient to correct the current account in the surplus countries and correct the global imbalances in the foreseeable future. The bootstrap test performs better than asymptotic test in many practical applications.

Turning to the twin deficit dilemma, the bootstrap panel Granger causality tests generate a number of notable findings worth mentioning. First, we find strong support for TDH via the Keynesian view before but not after the AFC. The evidence after the crisis largely points to the opposite direction, where we observed a strong (direct) evidence of causality running from a current account deficit to budget deficit in most of the countries. This is no surprise as the Asian countries in our sample have embraced greater caution for the financial globalization in the aftermath of the currency crisis. Hence, our findings support the conventional wisdom that prudent budget policies are essential when the external deficits are excessively large (i.e., prior to the AFC) as echoed in recent studies. In other words, there have been significant changes in the dynamic behavior of the current account from the period before the AFC (deficits years) to the post-AFC period (surpluses years). The crisis has led to not only increase

in government spending but also fall in tax revenues due to a decline in business activities and tend to support the causality from current account to budget imbalances.

Another striking feature of our results is that we find that there is a close link between a current account and investments running from investment (or FDI) to current account before the currency crisis. This means that the external (push) factors cause foreign capital inflows and push the current account toward imbalances. The empirical findings provide no indication to suggest that FDI causes current account imbalance (or vice versa) in the aftermath of the crisis. The relationship between the two variables has disappeared (decoupled) when the currency crisis erupted in the late 1990s and this feature needs to be recognized by policymakers in the Asian countries. This is because they raise questions about the extent to which reduction in the budget deficit can (if any) correct the current account imbalances. Our results are robust with respect to the two methods (Konya's and Emirmahutoglu and Kose's panel causality tests) that apply in the empirical analysis. While the results based on the earlier sample period tally with that of Yan and Yang (2012) and others, the findings from the latter sample are in sharp contrast to their conclusion. Yan and Yang (2012) find a current account reversal that follows ever since the AFC has changed the nature of its causal relationship between the two variables from negative toward the positive relation. The nonexistence of a causal relationship between FDI and current account in the later sample period balance is an indicative of macroeconomic measures undertaken after the crisis of the late 1990s much less vulnerable to the financial crises. Kim et al. (2007) claimed that sterilization measures undertake in the wake of the AFC have led to a significant reduction in capital mobility. Experience from the global financial crises suggests that the majority of these countries fared better in terms of the output decline during the crisis. Dominguez et al. (2012) have shown that countries that have a higher reserve accumulation are associated with higher post-crisis GDP growth.

Finally, one fruitful way to approach the issue is to apply nonlinear models using longer data series and include debt ratio. The impact of fiscal policy on the current account might vary with changing size public debt. The impact may be large following a large change in fiscal deficits. Clearly, much interesting research remains to be done in this line of investigation.

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