Chapter 4

Business Cycle Synchronization and Financial Integration in the Asia-Pacific Region

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Abstract

This paper explores the factors that drive business cycle synchronization (BCS) in the Asia–Pacific region. Three main factors that figure prominently in the literature, viz., trade intensity, similarity of industrial structure and financial integration, are analyzed, with emphasis on the impact of financial integration on BCS. We employ a dynamic panel GMM approach in our estimation in order to control for biases associated with simultaneity and unobserved country-pair specific effects. Our results agree with theoretical predictions from the benchmark international business cycle models – greater financial integration leads to divergent BCS. This strongly suggests that controlling for biases associated with simultaneity and unobserved country-pair heterogeneity using a panel-based IV estimator is crucial in unraveling the contrasting evidence found in the empirical literature. Once such biases are accounted for, the prediction by theory regarding the inverse relationship between financial integration and BCS becomes apparent.

Keywords: financial integration, business cycle synchronization, Asia–Pacific countries, dynamic panel GMM, international business cycle models

JEL Classifications: C23, E32, E44, F36

Introduction

In the past decade, the buzzword amongst policymakers, observers and academics has been "globalization". Arising from the strong growth of trade, finance, transportation and communication, the impression has been that the world has increasingly become "borderless" or to operate in sync as a single global market. Fast forward to the past year or so and, against the backdrop of the global financial turmoil, a different theme has emerged – decoupling. The term emanates from the observation that despite weakening economic conditions in the US and a number of industrial countries, economic conditions in emerging-market countries have been surprisingly strong and resilient. This perception challenges conventional wisdom that when the US economy sneezes, the rest of the world economy, especially emerging-market countries, catches a cold.

Initially one could think that greater trade and financial linkages would lead to tighter business cycles instead of being divergent or asynchronous. A perfunctory inspection of the data would indeed suggest so (Lane and Milesi-Ferretti, 2004, 2007). However, theoretical predictions suggest that trade and financial linkages can influence business cycle synchronization (henceforth BCS) either way. For example, standard international business cycle theories predict that greater financial integration should lead to a lower degree of BCS, whereas models of financial contagion, such as Allen and Gale (2000) show how international financial integration can lead to financial panic and thus to synchronized business cycles. The question therefore is an empirical one, but the current empirical literature does not help as it fails to find a robust and systematic relationship. Pure cross-sectional studies, for instance, find a significantly positive relationship between financial integration and BCS, e.g. Imbs (2004, 2006). However, recent studies using panel data show a strong negative effect, e.g. Cerqueira and Martins (2009) and Kalemli-Ozcan et al. (2009), and thus are in conformity with the basic ideas of standard international business cycle theories.

This paper contributes to the empirical literature in a number of ways. First, we take the pragmatic approach of using quantity- and price-based indicators of financial integration to ensure robustness in our results. This is important as the lack of mandatory reporting of actual bilateral information on FDI and portfolio investment data, which limits many previous studies (including ours), makes this a sensible and logical strategy.[1] Second, in view of the data limitation just mentioned, we exploit the availability for a number of years of survey-based data on portfolio investment from the Coordinated Portfolio Investment Survey (CPIS) gathered by the IMF. To the best of our knowledge no study employs this dataset to ascertain the direction of the BCS–finance link in a panel context. Third, as we are using panel data, we adopt a panel-based GMM approach in our estimation. We believe that this technique gives us two important advantages compared to previous approaches, namely, controlling for biases associated with simultaneity, and unobserved country-pair specific effects. Finally, we examine the impact of trade and financial linkages on BCS across a broad cross-section of Asian and Pacific countries. The evidence is mostly for OECD or industrial countries, and little empirical work has been done in this part of the world. Previous empirical works by Park and Shin (2009) and Shin and Sohn (2006) find either a negligible or an insignificant effect of financial integration on BCS.[2]

Our results agree with the benchmark standard international business cycle models. This strongly suggests that controlling for biases associated with simultaneity, let alone unobserved country-pair heterogeneity, using a panel-based IV estimator is crucial in unraveling the contrasting evidence found in the empirical literature. In a pure cross-sectional instrumental-variable regression any unobserved time-invariant country-pair specific effects would be part of the error term, leading to biased estimates of the coefficients, and as such previous studies that examine the direction of the BCS–finance link using pure cross-sectional data are afflicted by this problem. Once biases associated with simultaneity and unobserved country-pair specific factors are accounted for, the theoretical prediction regarding the inverse relationship between financial linkages and BCS becomes apparent in the results.

The rest of the paper is structured as follows. The next section reviews the literature on the relationship between financial integration and BCS. Section 3 describes the data and

[1] Exceptions to this limitation are Kalemli-Ozcan et al. (2009), who use BIS proprietary data on OECD countries, and Garcia-Herrero and Ruiz (2008), who use a novel dataset on bilateral flows between Spain and a large number of countries taken from the Spanish Balance of Payments.
[2] An exception is by Lee and Azali (2010), who look at a smaller sub-section of countries than we examine in this paper and who emphasize OCA criteria.

the measures used for each of the variables. Section 4 describes the estimation strategy. Section 5 presents the empirical results. Section 6 offers some policy implications emanating from the results, and Section 7 concludes.

2. Literature Background

Financial integration has been argued to affect business cycle synchronization but as to whether it should lead to tighter or synchronized business cycles is unclear. Standard international macro theories predict that greater financial integration should lead to a lower degree of business cycle synchronization. As shown by Backus et al. (1992) and Baxter and Crucini (1995), in a two-country general equilibrium model with complete financial markets, a country hit by a positive productivity shock receives capital from the other country, resulting in a negative output correlation between the two. Similarly, Backus et al. (1994) document that complete markets result in negatively correlated GDP because an economy hit by a positive technology shock will attract capital flows away from the no-shock economy.

Heathcote and Perri (2001, 2003, 2004) also find results that are in line with the above mentioned studies through a model in which international financial market integration occurs endogenously in response to less-correlated shocks. They argue that increasing globalization in financial markets is the key for less international co-movement. A combination of less-correlated shocks coupled with the resulting deepening of international asset markets can account for the less-correlated international real business cycle. Likewise, international specialization theories along the lines of Obstfeld (1994) yield a similar prediction.

However, surprisingly, the empirical evidence is still mixed as to the relationship between financial integration and business cycle correlation. In a pure cross-sectional context, studies find a significantly positive relationship between financial integration and BCS. Using a system of simultaneous equations, Imbs (2004) finds evidence that economic regions with strong financial links are significantly more synchronized. Employing the same approach but using the 2001 Coordinated Portfolio Investment Survey (CPIS) gathered by the IMF, Imbs (2006) finds similar results. Kose et al. (2003) using a cross-country sample of 76 countries – 21 industrial and 55 developing –

find evidence that financially open developing economies have synchronized cycles with the core rich G7 countries over the period 1960–1999. Davis (2009) also finds a positive relationship between international credit and BCS over a cross-country sample of 58 countries from 1991 to 2004. Jansen and Stokman (2004) investigate the relationship between FDI and BCS over the period 1982–1991 using data from Canada, France, Germany, the Netherlands, the UK and the US, and find that the rapid expansion of FDI can be related to the phenomenon of more synchronized business cycles. Finally, Kose et al. (2008a) and Kose et al. (2008b) find the same results with a dynamic latent factor model.

A notable exception that documents a negative relationship between financial integration and BCS in a pure-cross section context is by Garcia-Herrero and Ruiz (2008). Using actual bilateral financial flows from Spain these authors estimate a system of simultaneous equations and find that greater financial integration leads to divergent BCS. Bordo and Heibling (2003) also find a long-run increase in cycle synchronization, but conclude that little of it can be ascribed to a proxy of financial integration using the removal of capital control.

Although one can reconcile these studies that find a positive relationship between financial integration and BCS based on models of financial contagion, such as Allen and Gale (2000), which show how international financial integration can lead to financial panic and thus to synchronized business cycles, these studies conflict with the benchmark international business cycle models discussed above.

Opposed to cross-country studies that consistently find a positive effect of financial integration on BCS, some recent studies using panel data show a robust negative effect, and thus are in conformity with the basic ideas of standard theories. For instance, using panel three-stage least squares estimation for 15 OECD countries from 1984 to 2003, Xing and Abbott (2007) find that economic regions with strong financial links are significantly less synchronized. Using GMM methods for 20 OECD countries from 1970 to 2002 Cerqueira and Martins (2009) also find a negative and significant effect of financial integration on BCS. Using data for eight Asian countries – ASEAN5 (Indonesia, Malaysia, the Philippines, Singapore and Thailand), plus three additional members of East Asia (China, Japan and Korea) – Lee and Azali (2009) also find a similar significantly negative effect. Finally, Kalemli-Ozcan et al. (2009), using a rich panel data structure on banks' international bilateral exposure over the past three decades across 20 developed countries, show that once country-pair and time-fixed effects are controlled for, a higher degree of cross-border financial integration leads to less synchronized, more divergent business cycles.

A notable exception that documents a positive relationship between financial linkages and BCS using panel data is by Schiavo (2008). Using a system of simultaneous equations for 190 developed country pairs over the period 1991–2002, the study shows that financial integration is associated with less-divergent business cycles. Inklaar et al. (2007), using a sample of 21 OECD countries over the period 1970–2003, find that the financial integration measures suggested by Imbs (2004) are not robustly related to BCS.

3. Data and Measurement

In order to examine empirically the effect of financial linkages as well as the effect of two other channels – trade linkages and similarity in industrial structure – on BCS, we estimate the following equation:

$$\rho_{ij,t} = \alpha_0 + \alpha_1 \rho_{ij,t-1} + \beta_1 T_{ij,t} + \beta_2 S_{ij,t} + \beta_3 F_{ij,t} + \upsilon_{ij,t}\,, \qquad (1)$$

where i,j represents country pair i and j; $\rho_{ij,t}$ denotes the business cycle correlation or synchronization (BCS) between countries i and j; $\rho_{ij,t-1}$ is the lagged BCS between countries i and j; $T_{ij,t}$ is a measure of bilateral trade intensity; $S_{ij,t}$ is an index of the similarity of the industrial structure between countries i and j; and $F_{ij,t}$ is a measure of financial integration between countries i and j. In Equation 1 we assume that $\nu_{ij,t}$ contains the following two effects: (i) the unobserved time-invariant country-pair specific effect, $\eta_{ij,}$, and (ii) a stochastic error term $\varepsilon_{ij,t}$, varying across time and cross-section.

We adopt a panel approach over the period 1980–2007. To acquire a meaningful measure of business cycle correlation as well as to purge any possible cyclical effects on the variables, the 28 years have been split into subperiods of four years each.[3] Our strategy of investigation assumes that countries in the Asia–Pacific region integrate within the region as well as with two major economic blocs – the United States and the EU-14. The EU bloc includes Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain, Sweden and the UK. The Asian–Pacific bloc comprises China, Hong Kong, India, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan, Thailand, Australia and New Zealand.

The business cycle correlations refer to the Pearson correlation of cyclical component of annual real GDP expressed in US dollars between countries *i* and *j* over the relevant subperiods. The cyclical components of GDP are computed in two ways after taking the natural logarithms of real GDP: (i) using the Hodrick–Prescott (HP) filter, and (ii) the Baxter–King (BK) band-pass filter.[4]

Two different measures of bilateral trade intensity will be considered. The first is a standard measure used in many recent studies, for instance, Clark and van Wincoop (2001), Frankel and Rose (1997, 1998) and Imbs (2004) among others, and is calculated as:

$$T_{i,j}^{1} = \frac{1}{T}\sum_{t}\frac{X_{ij,t} + M_{ij,t}}{Y_{i,t} + Y_{j,t}} \tag{2}$$

where $X_{ij,t}$ denotes total merchandise exports from country *i* to *j* in year *t*, $M_{ij,t}$ denotes imports to *i* from *j*, and Y_i denotes nominal GDP in country *i*. An alternative measure, proposed by Clark and van Wincoop (2001) based on the model in Deardoff (1998), and employed by Imbs (2004, 2006) among others, can be constructed as:

$$T_{i,j}^{2} = \frac{1}{2}\frac{1}{T}\sum_{t}\frac{(X_{ij,t} + M_{ij,t})Y_{t}^{W}}{Y_{i,t} \times Y_{j,t}} \tag{3}$$

where Y_t^W is world GDP. T^2 differs from T^1 in that it is independent of country size and depends only on trade barriers. In the empirics, we take the natural logarithm of both

[3] 1980-1983, 1984-1987, 1988-1991, 1992-1995, 1996-1999, 2000-2003 and 2004-2007.

[4] We do not broach the literature on the appropriate measure of synchronizations of business cycles, noting only that most studies employ simple bilateral correlation coefficients in order to measure the strength and direction of the association between the cyclical components of the annual real GDP of two countries.

measures of bilateral trade intensity. We interpret that the *higher* the values of T^1 and T^2 the greater the trade intensity between countries *i* and *j.* In terms of Equation (1), the sign of β_1 can be positive or negative as theory predicts that closer bilateral trade could result in synchronized (positive) or asynchronous (negative) business cycles.

There are no standard measures of similarity in industrial structures (Imbs, 2004). We use two measures of similarity in industrial structures for comparability with existing research. The first, a measure akin to a Herfindahl index of concentration, is employed in Clark and van Wincoop (2001), Imbs (2004, 2006) and Krugman (1991), among others, and is measured as:

$$S_{i,j}^{1} = \frac{1}{T} \sum_{t} \sum_{n}^{N} \left| S_{ni} - S_{nj} \right|, \tag{4}$$

where S_{ni} and S_{nj} denote the GDP shares for industry *n* in countries *i* and *j,* respectively. If two countries had identical industrial structures, that is, the industry shares in GDP were the same for countries *i* and *j*, then the index would be 0. On the other hand, when two countries have completely disjointed or different industrial structures, the index reaches a maximum value of 2. Therefore, *lower* values of S^1 imply more similarity in industrial structure between countries *i* and *j.*

Our second measure of similarity of industrial structure, suggested by Shea (1996) and used by Imbs (2001, 2003) and Baxter and Kouparitsas (2005) is the correlation of industry shares:

$$S_{i,j}^{2} = \frac{1}{T} \sum_{t} \frac{\sum_{n}^{N} S_{ni} S_{nj}}{\sqrt{\sum_{n}^{N} S_{ni}^{2}} \sqrt{\sum_{n}^{N} S_{nj}^{2}}} \tag{5}$$

If $S_{ni} = S_{nj}$, that is, GDP shares of each industry are the same in countries *i* and *j*, S^2 is equal to 1. In other words, *lower* values of S^2 imply less similarity in industrial structure between countries *i* and *j.* We also take the natural logarithms of both measures of similarity of industrial structure in our empirics. Theory clearly predicts that similar production patterns between countries should lead to synchronized business cycles, and as such the sign of β_2 changes depending on the measure used for similarity of industrial structure in Equation (1). That is, S^1 implies $\beta_2 < 0$, while S^2 implies $\beta_2 > 0$.

Bilateral financial integration can be difficult to measure effectively. On this basis, our measures of financial integration encompass quantity- and price-based measures to ensure robustness. We first discuss our quantity-based measures of financial integration and then turn to our price-based measures. Our quantity-based measure of financial integration is constructed in three ways. The first measure uses the recently updated dataset of Lane and Milesi-Ferretti (2007) and, following Cerqueira and Martins (2009), is calculated as the average of the sum of stocks of assets and liabilities of foreign direct investment (FDI) and portfolio investment between countries *i* and *j* scaled by nominal GDP:[5]

$$F^1_{ij,t} = \frac{1}{T}\sum_t \left(\frac{A_{i,t} + L_{i,t}}{Y_{i,t}} + \frac{A_{j,t} + L_{j,t}}{Y_{j,t}} \right), \qquad (6)$$

where *Ai,t* and *Li,t* are total assets and liabilities of country *i*, at time *t*. We can interpret F^1 as a measure of the extent of openness between pairs of countries to global financial markets, and as such we think of *higher* values of F^1 when pairs of countries are more financially integrated with world financial markets. An alternative measure of financial openness, which builds on the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER) and is thus based on information on controls on financial flows, is the index-based measure recently put together by Chinn and Ito (2008). To construct a bilateral measure, the average of the sum of the indices between countries *i* and *j* is calculated as:

$$F^2_{ij,t} = \frac{1}{T}\sum_t (ITO_{i,t} + ITO_{j,t}), \qquad (7)$$

where $ITO_{i,t}$ and $ITO_{j,t}$ denote the Chinn–Ito indices for countries *i* and *j*, respectively. Likewise, F^2 takes higher values when pairs of countries have officially lower restrictions on financial flows.

[5] Also using the Lane and Milesi-Ferretti dataset, Imbs (2004) constructs his own alternative measure of bilateral financial integration by using the difference between the stocks of assets and liabilities (net foreign asset) between countries *i* and *j*.

Finally, one of this paper's contributions is to use in a panel context the Coordinated Portfolio Investment Survey (CPIS) gathered by the IMF that provides direct observations on bilateral asset holdings from 2001 until 2007.[6] At the outset, however, one should also recognize that the data have their own limitations. Problems of under-reporting (Lane and Milesi-Ferretti, 2004), non-inclusion of foreign direct investment (FDI), and the absence of some countries in the collection in view of the non-mandatory nature of the reporting, are some of the concerns that plague these dataset (Imbs, 2006).[7] Nonetheless, given our awareness of the problems inherent in the dataset, which is borne out of the lack of data sources based on mandatory reporting of bilateral information on FDI and portfolio investment, our third quantity-based measure of financial integration is computed, as in Garcia-Herrero and Ruiz (2008):

$$F^3_{ij,t} = \frac{I_{ij,t} + I_{ji,t}}{Y_{i,t} + Y_{j,t}}, \quad (8)$$

where $I_{ij,t}$ represents financial flows from country i to country j at time t. Similar to F^1 and F^2, F^3 takes higher values when pairs of countries are more financially integrated.

In order to measure financial integration through a price-based indicator, we follow Schiavo (2008) in defining financial integration as the Euclidean distance between the spread among long- and short-term interest rates as well as the spread among long-term bank lending rates, respectively:

$$F^{LS}_{ij,t} = \sqrt{(loir_i - loir_j)^2 + (sir_i - sir_j)^2} \quad (9)$$

$$F^{LL}_{ij,t} = \sqrt{(loir_i - loir_j)^2 + (blir_i - blir_j)^2}, \quad (10)$$

where $loir_i$, sir_i, $blir_i$ are the long- and short-term and bank lending rates in country i. The underlying argument behind both measures is that it utilizes the notion that interest rate equalization can be expressed in the form of distance from the law of one price starting from long- and short-term interest rates (F^{LS}) or from long-term bank lending rates (F^{LL}). The appeal of both measures is that these rates span "orthogonal" markets

[6] Imbs (2006) uses the 2001 CPIS survey data, which effectively makes it a pure cross-sectional study.
[7] For instance, two of the 13 Asia–Pacific countries that are examined in this paper, namely, China and Taiwan, are both absent from the collection.

and therefore give a multi-faceted or complete picture of closer financial integration rather than focusing only on a single rate, which may produce a distorted picture (Schiavo, 2008). Being a measure of distance from the law of one price, *higher* values of F^{LS} and F^{LL} imply less financial integration between countries i and j.

As discussed in Section 2 above, standard international business cycles theories show that greater financial integration leads to divergent business cycles and as such the sign of β_3 changes, depending also on the measure used for financial integration in Equation (1). That is, F^1, F^2 and F^3 all imply $\beta_3 < 0$, while F^{LS} and F^{LL} both imply $\beta_3 > 0$.[8]

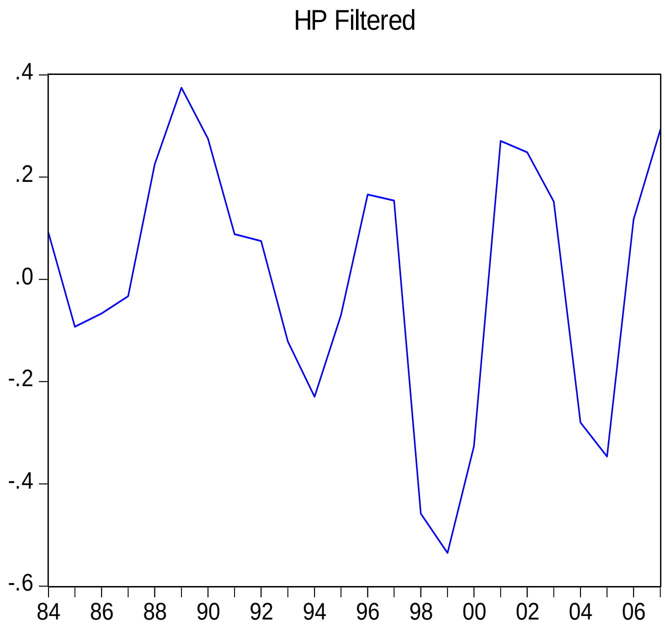
3.1. Some Stylized Trends

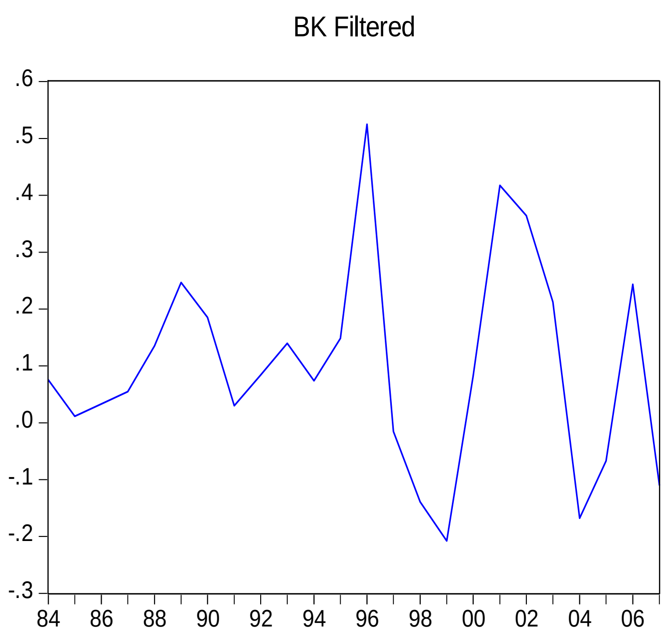
As a preliminary step, this section presents a graphical evolution of our measures of BCS as well as the measures of financial integration that we employ later in our analysis. We begin by examining the stylized trends in our BCS measures over the period considered in our analysis.

Figure 1 presents the business cycle synchronization between the Asia–Pacific economies and the EU-14. The figure has two graphs, one for each of detrending techniques used – the Hodrick–Prescott (HP) filter and the Baxter–King (BK) band-pass filter. Each graph contains time-series plots of the average BCS for all possible pairs of Asian and EU-14 countries at a point in time. The average BCS level varies over time between –0.5 to 0.5, but there are no obvious trends. To be sure, the Asian financial crisis of 1997–1998 made the business cycle asynchronous between these two regions as well as during the mid-2000s. However, by the end of the period of interest, there is a divergence in outcome on whether Asian business cycles became more (HP filter) or less (BK filter) correlated with the EU-14 countries.

[8] Obviously, in relation to cross-country empirical studies that find a positive relationship between financial integration and BCS, the sign of β_3 is the reverse of the above. That is, $\beta_3 > 0$ for F^1, F^2 and F^3, and $\beta_3 < 0$ for the case of F^{LS} and F^{LL}.

Figure 1. Asia–EU Business Cycle Correlation across Time

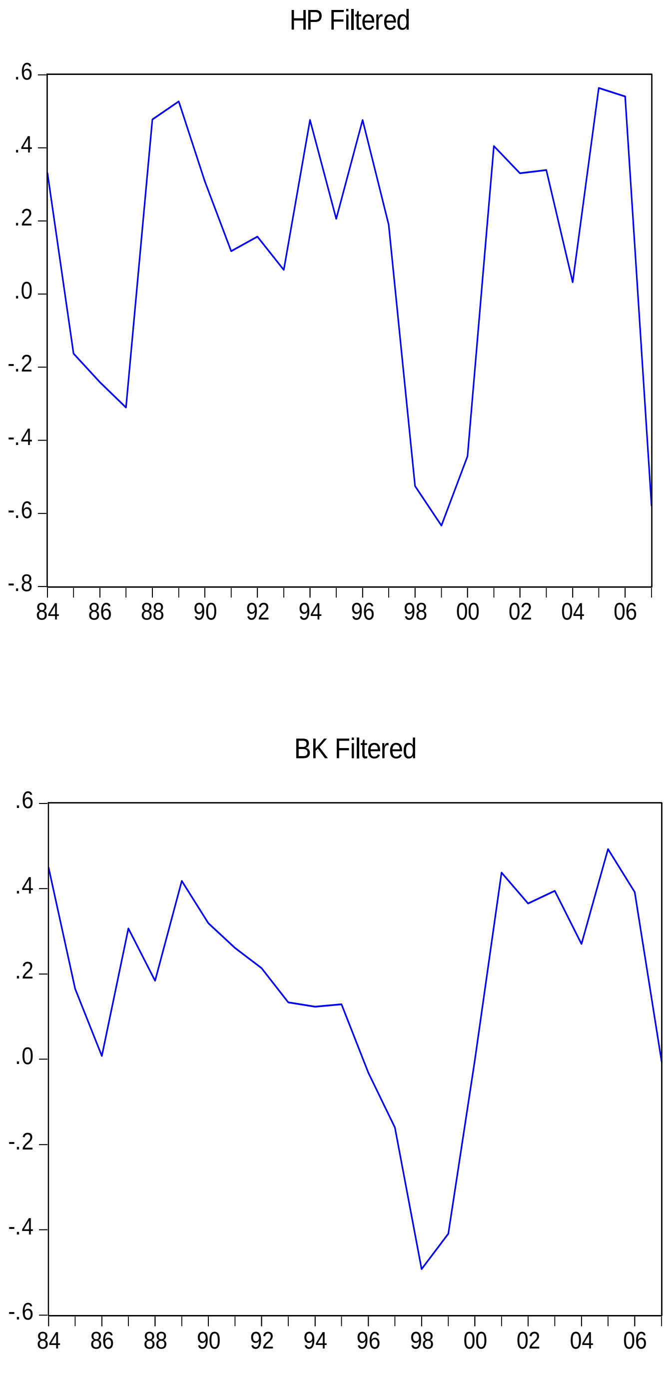




Note: The correlation is estimated with a four-year rolling window.

Source: Authors' calculations.

Figure 2. Asia–US Business Cycle Correlation across Time



Note: The correlation is estimated with a four-year rolling window.

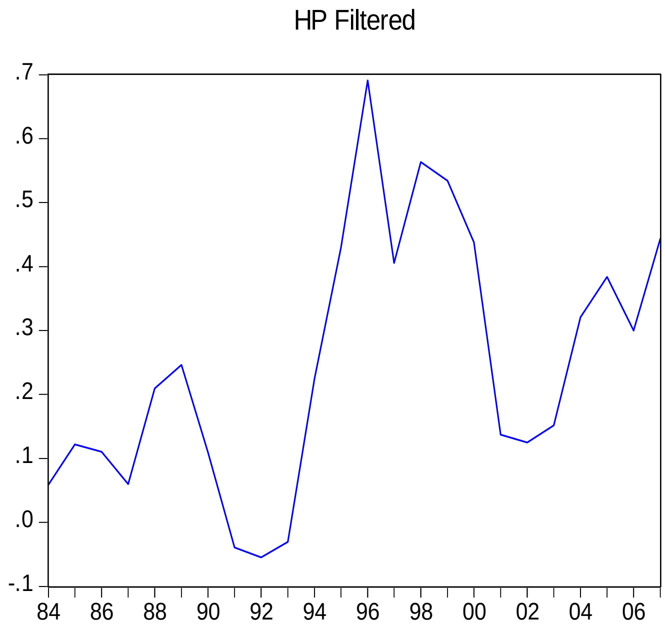
Source: Authors' calculations.

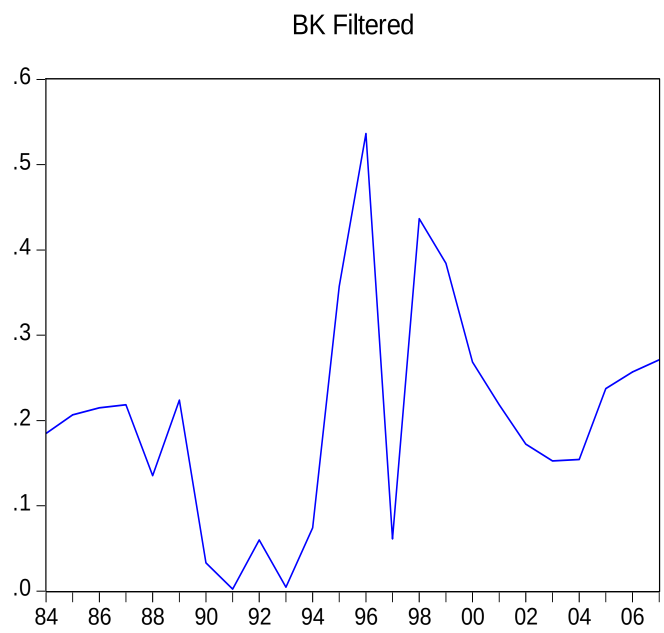
Figure 2 is an analogue that considers the BCS of the Asia–Pacific economies to the United States. In the aftermath of the steep decline in BCS levels around the time of the Asian financial crisis, BCS levels typically hovered around 0.4 or so in the 2000s. However, as portrayed in the figure, there is a tendency between business cycles across the Asia–Pacific countries and the United States in 2007 to be *less* correlated.

Finally, Figure 3 is an analogue that considers the intra-Asia–Pacific countries' BCS levels over time. The single most striking observation of Figure 3 is the marked contrast in average BCS levels in this figure as compared, in particular, with the average BCS levels presented in Figure 2. This suggests a shrinking relationship between the business cycles of the Asia–Pacific countries and with major countries outside of it, most especially the United States, whereas, at the intra-regional level, decoupling does not seem to be a phenomenon over time.

Obviously, the above analysis has its limitations, and thus should be taken as indicative only on whether business cycles in the Asia–Pacific economies are moving asynchronously or are "decoupled" over time with the two major economic blocs considered in the analysis. For one, the analysis is unconditional – no other factors have been taken into account as possibly affecting BCS. This will be the focus of our attention in the subsequent sections, but before that we take a brief look at the trends of our financial integration measures over time.

Figure 3. Asia–Asia Business Cycle Correlation across Time

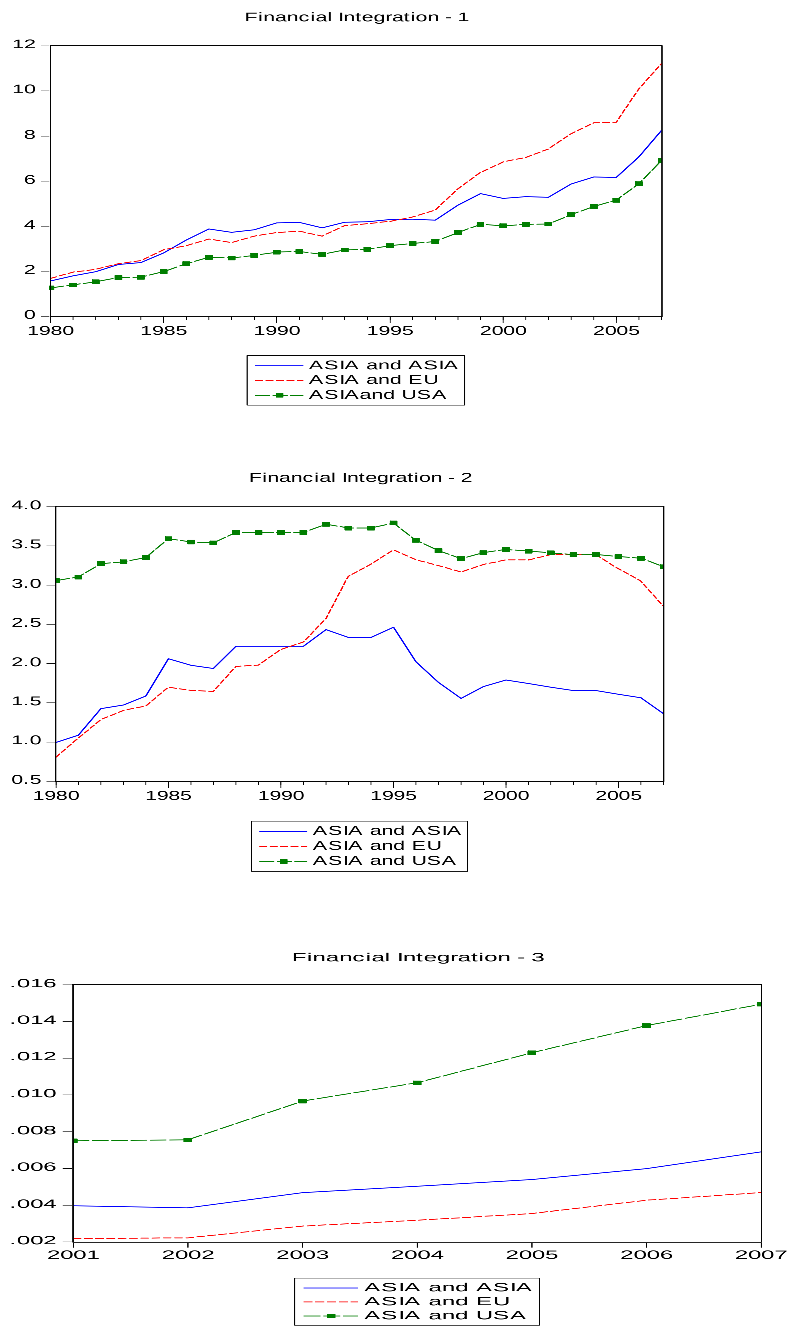




Note: The correlation is estimated with a four-year rolling window.

Source: Authors' calculations.

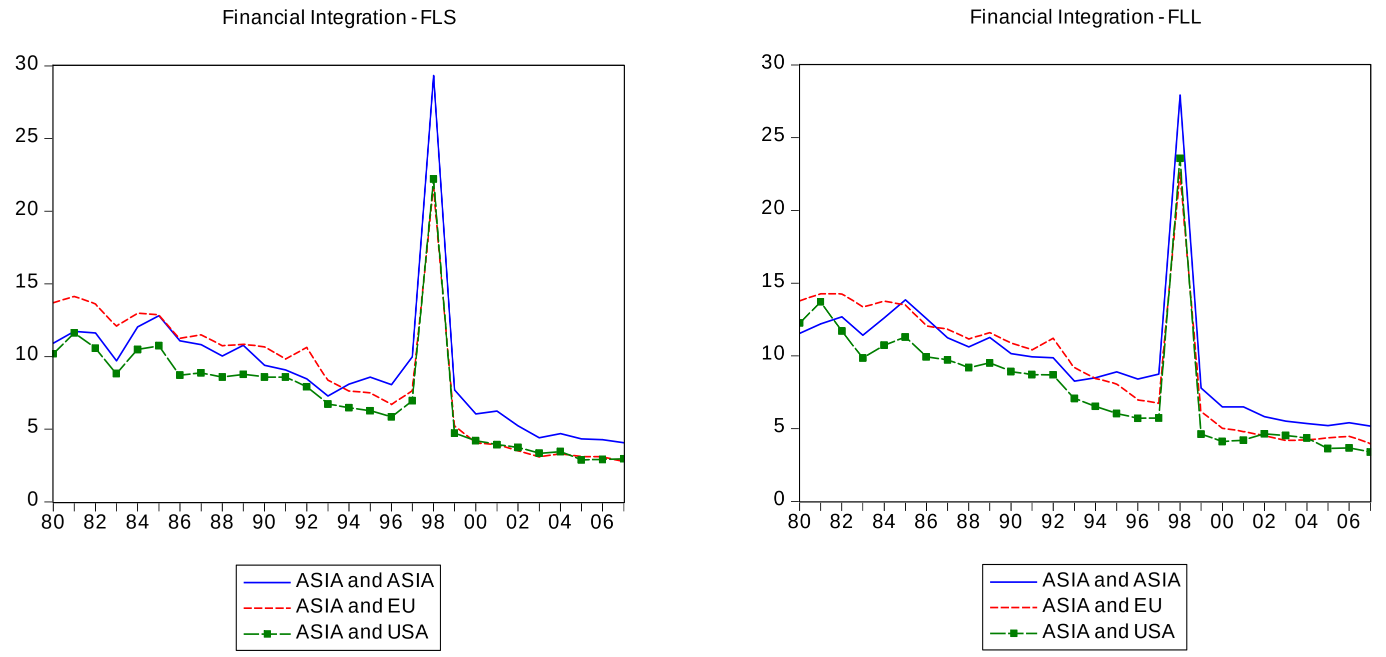
Figure 4. Financial Integration over Time (Quantity-Based)



Note: The correlation is estimated with a four-year rolling window.

Source: Authors' calculations.

Figure 5. Financial Integration over Time (Price-Based)



Note: The correlation is estimated with a four-year rolling window.

Source: Authors' calculations.

Figures 4 and 5 give a graphical depiction of the evolution over time of our quantity-based and price-based measures of bilateral financial integration, respectively. Figure 4 has three graphs, one for each of our bilateral financial integration measures (F^1, F^2 and F^3). Each graph contains time-series plots of the cross-country-pair average of the relevant financial integration measure.[9] Figure 4 clearly shows that, according to our first measure of financial integration, Asia's integration with world financial markets has increased considerably over the past three decades. Indeed, a more detailed inspection of this particular plot of F^1 indicates that the dramatic rise occurred around the middle of the 1990s. Whereas there is a clear and persistent upward trend in F^1, there is a reversal in Asia's openness to financial flows sometime in the mid-1990s according to our second bilateral measure (F^2). From this point until the end of the period considered, our second measure seems to fluctuate around an approximately constant mean, which intuitively suggests that Asia has not made much further progress in reducing formal restrictions on financial flows in recent years. With regards to our last quantity measure of bilateral financial integration (F^3), in consistency with Kim and Lee (2008), its evolution portrays Asia's increasing though limited intra-regional financial integration as it falls behind that of its integration with the US.

Finally, Figure 5 presents a graphical depiction of our price-based measures of financial integration. Clearly, with the exception of the period around the time of the 1997–1998 Asian crises, cross-country dispersion or "distance" from the law of one price is declining over time, but remains sizable, and is, likewise, suggestive of Asia's greater integration with the EU-14 and the US than within the region.

Supplementary results to the above discussions are provided by the unconditional correlations of the respective bilateral variables in Table 1.[10] While most of our quantity-based indicators of financial integration tend to show a very weak (almost zero) negative relationship with our two BCS measures, this stands in contrast to the unconditional correlation results between most of our price-based indicators and our

[9] Notice that the time-scale axis for the graph containing the F^3 measure is not in synch with the other two measures. Recall from our previous discussion that the F^3 measure is constructed using the CPIS data which are only available for 1997 and then from 2001 to 2007.

[10] The unconditional correlations for the smaller sample period of 2001–2007 when using the CPIS data in constructing the third quantity-based indicator of financial integration, i.e. F^3, is presented separately in Table A1 in the Appendix.

two BCS measures – a weak positive relationship. Furthermore, the low correlation (< 0.5) among our measures of the explanatory variables in Equation (1) suggests that multicollinearity is not an issue here.[11]

Table 1. Correlation Matrix of the Variables

	ρHP	ρBK	T1	T2	S1	S2	F1	F2	FLS	FLL
ρHP	1.0000									
ρBK	0.5875	1.0000								
T1	0.2151	0.1400	1.0000							
T2	0.1729	0.1443	0.8751	1.0000						
S1	0.0088	0.0149	0.0758	0.0919	1.0000					
S2	–0.0124	–0.0034	–0.0559	0.0025	–0.8465	1.0000				
F1	0.0353	–0.0247	0.3404	0.4223	0.0220	0.0176	1.0000			
F2	–0.0922	–0.0178	0.1568	0.1767	–0.2702	0.2273	0.5305	1.0000		
FLS	–0.0766	0.0055	–0.1953	–0.1139	0.1301	–0.0989	–0.3032	–0.2335	1.0000	
FLL	–0.0741	–0.0028	–0.2033	–0.1395	0.1444	–0.1252	–0.3157	–0.2038	0.8611	1.0000

Note: All variables are in logarithmic form with the exceptions of ρ^{HP}, ρ^{BK}, F^2, F^{LS} and F^{LL}.

Source: Authors' calculations.

4. Methodology

4.1. Dynamic Panel GMM Technique[12]

In a pure cross-sectional regression any unobserved time-invariant country-pair specific effects would be part of the error term, leading to biased estimate of the coefficients. Previous studies of the direction of the BCS–finance link using pure cross-sectional data are afflicted by this problem. A panel context, however, allows us to control for these unobserved time-invariant country-pair specific effects and, as a result, the problem of biased coefficient estimates are either reduced or eliminated. This is important as there is growing evidence in the literature that cultural biases and differences, for instance, have a substantial impact on a variety of financial flows – portfolio and direct investment (Ekinci et al., 2008; Guiso et al., 2009) as well as on foreign bank lending (Giannetti and Yafeh, 2008; Mian, 2006). In addition, the GMM estimator does not require any particular distributions of the error term. This veers away from the

[11] It should be noted at this point that the high correlation between T^1 and T^2, S^1 and S^2 and F^1 and F^2 in Table 2 does not pose any multicollinearity problems in the estimation of Equation (1) as two respective measures of, for instance, trade intensity, are entered separately as explanatory variables in Equation (1).

[12] The discussion that follows draws in part on Calderon and Chong (2001), Chong and Gradstein (2007) and Levine et al. (2000).

complicated inference introduced by using a Pearson correlation coefficient (bounded at –1 and 1) to measure BCS as the error term is unlikely going to be normally distributed.

In order to estimate our model (Equation 1) consistently and efficiently, we use a generalized method of moments (GMM) estimator for dynamic panel data models that was introduced by Holtz-Eakin et al. (1990) and Arellano and Bond (1991), and further developed in a series of papers including Arellano and Bover (1995) and Blundell and Bond (1998). This estimator encompasses a regression equation in both differences and levels, each one with its specific set of instrumental variables. Consider the following regression equation for BCS:

$$y_{ij,t} = \alpha y_{ij,t-1} + \beta X_{ij,t} + \eta_{ij} + \varepsilon_{ij,t} \,, \tag{11}$$

where y is the business cycle correlation measure, X represents the set of explanatory variables apart from the lagged business cycle correlation measure, η is an unobserved, time-invariant country-pair specific effect, ε is the error term, and the subscripts i, j and t represent country pairs and time period, respectively.

We eliminate country-pair specific effects (η_{ij}) by taking first differences of Equation (11):

$$y_{ij,t} - y_{ij,t-1} = \alpha(y_{ij,t-1} - y_{ij,t-2}) + \beta(X_{ij,t} - X_{ij,t-1}) + (\varepsilon_{ij,t} - \varepsilon_{ij,t-1}) \tag{12}$$

The use of own suitable lagged levels of $y_{ij,t}$ as instruments is required to deal with the problem that by differencing the lagged dependent variable ($y_{ij,t-1} - y_{ij,t-2}$) is correlated with the error term, $\varepsilon_{ij,t} - \varepsilon_{ij,t-1}$. The same strategy is applied to form instruments for other explanatory variables that are allowed to be endogenous in the sense that they can be affected by current and past realizations of business cycle correlations. This feature enables us to avoid simultaneity bias due to the endogeneity of our financial linkages variables. Strictly speaking, under the assumption that (i) the explanatory variables, X, are weakly exogenous (no correlation with future realizations of the error term), and (ii) the error term, ε, is not serially correlated, the dynamic panel GMM estimator exploits the following moment conditions:

$$E[y_{ij,t-s} \cdot (\varepsilon_{ij,t} - \varepsilon_{ij,t-1})] = 0 \qquad \text{for} \qquad s \geq 2;\ t = 3, \ldots, T \tag{13}$$

$$E[X_{ij,t-s} \cdot (\varepsilon_{ij,t} - \varepsilon_{ij,t-1})] = 0 \qquad \text{for} \qquad s \geq 2;\ t = 3, \ldots, T. \tag{14}$$

The resulting GMM estimator based on these conditions is known as the *difference*-GMM estimator. There is, however, an issue with the *difference*-GMM estimator. If lagged dependent variables and explanatory variables are persistent over time, the lagged levels likely represent weak instruments for the first-differenced variables. This causes finite sample bias and low accuracy, which leads to the need to complement the regression in first differences with a regression in levels. The instruments for the regression in first differences are the same as above. The instruments for the regression in levels, in turn, are the lagged *differences* of the same corresponding variables, under the assumption that although there may be correlation between the levels of the right-hand side variables and the country-pair specific effect in Equation (11), none exists between the differences of these variables and the country-pair specific effect.

The additional moment conditions for the regression in levels are:

$$E[y_{ij,t-s} - y_{ij,t-s-1}) \cdot (\eta_{ij} + \varepsilon_{ij,t})] = 0 \ \text{for } s = 1 \tag{15}$$

$$E[X_{ij,t-s} - X_{ij,t-s-1}) \cdot (\eta_{ij} + \varepsilon_{ij,t})] = 0 \ \text{for } s = 1. \tag{16}$$

The consistency of the GMM estimator depends on whether lagged values of the explanatory variables are valid instruments in the regression. To address this issue, we consider two specification tests: the first is the Hansen test of over-identifying restrictions, which tests the overall validity of the instruments. Failure to reject the null hypothesis supports the model. The second test examines the hypothesis that the error term is not serially correlated. We test whether the differenced error term, that is, the residual of the regression in differences, is second-order serially correlated.[13] If the test fails to reject the null hypothesis of absence second-order serial correlation, we conclude that the original error term is serially uncorrelated and use the corresponding moment conditions.

[13] Second-order serial correlation of the differenced residual indicates that the original error term is serially correlated and follows a moving-average process at least of order one.

5. Estimation Results

Table 2 presents the dynamic panel GMM estimation results using business cycle correlation or synchronization as the dependent variable. The table has six columns. The first column refers to the explanatory variables used in the estimation of a particular specification of the dynamic panel GMM. The second to fourth columns contain the coefficient estimates of the explanatory variables, that is, measures of trade, similarity of industrial structure, and financial, respectively, along with the p-values in square brackets; estimates that are significantly different from zero at the 0.10, 0.05 and 0.01 are marked by one, two and three asterisks, respectively. The last two columns contain the specification tests results. The table also contains two panels. The upper panel contains the results using the Hodrick–Prescott filter; the lower panel contains the estimates using the Baxter–King band-pass filter. Each panel has four sets of results, and each set is delineated based on the measure of financial integration (*F*) used as the explanatory variable in the GMM estimation. In turn, each set contains four rows of results that pertain to all possible permutations of *F* with the other explanatory variables *T* (bilateral trade) and similarity of industrial structure (*S*).

We begin by examining the specification test results. The Hansen test of overidentifying restrictions in all cases does not reject the null hypothesis that the instruments are not correlated with the error process. In addition, tests of serial correlation fail to reject the null that the error term, expressed in first differences, is not second-order serially correlated in all cases. This supports using lags of the explanatory variables as instruments.[14]

[14] In addition, apart from using "internal" instruments , i.e. lags of the explanatory variables, we also included the pair-wise sum of the recently updated exchange rate regime classification of Reinhart et al.(2010), the pair-wise sum of GDP per capita, and the absolute value of differences in GDP per capita as "external" instruments in the GMM estimations.

Table 2. Dynamic Panel GMM estimations of Business Cycle Synchronizations (Crisis Years Included)

Variables	Trade ($T_{ij,t}$)	Specialization ($S_{ij,t}$)	Finance ($F_{ij,t}$)	Hansen test (p–value)	AB test for AR(2) (p-value)
Hodrick–Prescott Detrended					
T^1, S^1, F^1	1.153 [0.00]***	–0.448 [0.04]**	–0.446 [0.01]**	0.38	0.29
T^2, S^1, F^1	0.448 [0.01]**	0.633 [0.01]**	–0.443 [0.00]***	0.78	0.26
T^1, S^2, F^1	0.654 [0.00]***	1.27 [0.03]**	–0.415 [0.00]***	0.61	0.08
T^2, S^2, F^1	0.534 [0.01]**	–1.85 [0.03]**	–0.494 [0.00]***	0.87	0.17
T^1, S^1, F^2	0.448 [0.00]***	–0.439 [0.00]***	–0.109 [0.00]***	0.14	0.66
T^2, S^1, F^2	1.516 [0.05]**	2.361 [0.12]	–0.902 [0.03]**	0.78	0.84
T^1, S^2, F^2	0.609 [0.00]***	2.780 [0.00]***	–0.139 [0.00]***	0.81	0.76
T^2, S^2, F^2	0.349 [0.02]**	2.68 [0.00]***	–0.458 [0.00]***	0.35	0.40
T^1, S^1, F^{LS}	1.397 [0.01]**	–0.832 [0.02]**	0.121 [0.05]*	0.89	0.26
T^2, S^1, F^{LS}	0.660 [0.01]**	–0.101 [0.42]	0.089 [0.04]**	0.16	0.33
T^1, S^2, F^{LS}	1.070 [0.00]**	3.088 [0.04]**	0.078 [0.05]*	0.41	0.64
T^2, S^2, F^{LS}	1.195 [0.02]**	6.098 [0.07]*	0.060 [0.03]**	0.29	0.62
T^1, S^1, F^{LL}	1.547 [0.04]**	–1.031 [0.02]**	0.149 [0.07]*	0.72	0.42
T^2, S^1, F^{LL}	0.478 [0.04]**	–0.299 [0.02]**	0.128 [0.00]***	0.27	0.12
T^1, S^2, F^{LL}	1.163 [0.01]**	3.469 [0.04]**	0.088 [0.07]*	0.37	0.67
T^2, S^2, F^{LL}	0.892 [0.02]**	–2.192 [0.27]	0.128 [0.02]**	0.14	0.71
Baxter–King Detrended					
T^1, S^1, F^1	0.307 [0.01]**	–0.311 [0.00]***	–0.302 [0.00]***	0.92	0.63
T^2, S^1, F^1	0.228 [0.00]***	–0.287 [0.00]***	–0.208 [0.00]***	0.75	0.51
T^1, S^2, F^1	0.282 [0.01]**	1.571 [0.00]***	–0.280 [0.00]***	0.86	0.56
T^2, S^2, F^1	0.226 [0.00]***	1.491 [0.00]***	–0.195 [0.00]***	0.71	0.51
T^1, S^1, F^2	0.003 [0.98]	–0.223 [0.30]	–0.081 [0.01]**	0.87	0.25
T^2, S^1, F^2	0.037 [0.02]**	–0.189 [0.28]	–0.052 [0.00]***	0.53	0.12
T^1, S^2, F^2	0.071 [0.51]	0.620 [0.21]	–0.094 [0.01]**	0.96	0.31
T^2, S^2, F^2	0.040 [0.02]**	0.453 [0.25]	–0.058 [0.00]***	0.48	0.08
T^1, S^1, F^{LS}	0.523 [0.09]*	–0.038 [0.72]	0.063 [0.00]***	0.29	0.11
T^2, S^1, F^{LS}	0.208 [0.02]**	–0.644 [0.00]***	0.063 [0.00]***	0.20	0.97
T^1, S^2, F^{LS}	0.836 [0.02]**	0.304 [0.72]	0.058 [0.01]**	0.56	0.11
T^2, S^2, F^{LS}	0.208 [0.01]**	1.846 [0.00]***	0.022 [0.01]**	0.20	0.20
T^1, S^1, F^{LL}	0.222 [0.22]	–0.055 [0.53]	0.046 [0.02]**	0.20	0.72
T^2, S^1, F^{LL}	0.214 [0.01]**	–0.691 [0.00]***	0.068 [0.00]***	0.29	0.59
T^1, S^2, F^{LL}	0.496 [0.08]*	0.040 [0.95]	0.051 [0.02]**	0.30	0.23
T^2, S^2, F^{LL}	0.194 [0.01]**	2.028 [0.00]***	0.022 [0.01]**	0.19	0.10

Notes: Numbers in square brackets are p-values. Significance levels: *10%, **5%, ***1%.

The GMM estimations include lags of the dependent variable as well as time-dummies.

The estimates of the impact of bilateral trade intensity on BCS are strongly positive (more trade between two countries induces higher BCS). Of the 32 coefficients (= 2 detrending technique × 2 measures of bilateral trade linkages × 2 measures of similarity of industrial structure × 4 measures of financial integration,[15] two are significant at the 10% level, 21 are significantly so at the 5% level, and six at the 1% level. Of these statistically significant coefficients, these coefficients range from 0.04 (using T^2 as one of the regressors in the GMM estimation) to 1.547 (using T^1 as one of the regressors), and this means that an increase by our relevant measure of bilateral trade intensity by one standard deviation implies that BCS starting from its mean would increase from 0.025 (= 1.547*0.016) to 0.254 (= 0.04*6.35). On average, this represents moving the correlation of output by about 22% of one standard deviation, quite an economically significant effect.

Concerning similarity of industrial structure, most estimated coefficients are significant, although once we control for endogeneity some of the estimated coefficients have incorrect signs (we should expect that $S^1 < 0$, $S^2 > 0$; more similar countries have higher BCS). Of the 32 coefficients, 28 have correct signs; of these 28 coefficients, one is significant at the 10% level, seven are significant at the 5% level, and 11 at the 1% level.

Finally, we found in our estimation results that the higher is the financial linkages between countries (higher F^1 and F^2, lower F^{LS} and F^{LL}) the less synchronization of business cycles. This finding is strikingly consistent across all estimated coefficients, all of which are statistically significant at conventional significance levels. These estimated coefficients range from –0.902 to –0.052 (both using F^2 as one of the regressors), and from 0.022 to 0.149 (both using F^{LL} as regressor).[16] This means that an increase by one standard deviation by our relevant measure of financial linkages implies that BCS starting from its mean would be reduced from 0.786 to 0.045 (using F^2 as one of the regressors), while BCS increases from 0.123 to 0.836 (using F^{LL} as regressor). In turn, this represents moving the correlation of output from 20% to 60% of one standard deviation, again quite an economically significant effect.

[15] Note that there are five measures of financial integration used in this paper. The results using the fifth measure are presented in a separate table.
[16] We recognize that we are dealing with quantity- and price-based indicators of financial integration, and as such we separate the economic interpretation of the size coefficients based on this distinction.

Table 3 is an analogue to Table 2, but excludes the crisis years of 1997–1998 from the GMM estimation. Although some of the estimated coefficients lost their statistical significance, the general features and story of the results are still strikingly similar to its Table 2 analogue. Table 4 is also an analogue to Table 2 that considers our final indicator of financial integration (which is also our third quantity-based indicator of financial integration) using the 2001 to 2007 CPIS data.[17] Though two as well as four of the eight estimated coefficients of the indicators for financial and bilateral trade linkages, respectively, are statistically indistinguishable from zero, this additional sensitivity test does not undermine the initial findings: both specification tests are fulfilled; a slightly weaker although consistent result indicates that more trade between two countries induces higher BCS; there are some indications that more similar countries have higher BCS; and; finally, increased financial linkages between countries leads to divergent BCS.

[17] In order to create the measure of BCS, i.e. the correlation between cyclical output in countries i and j, we follow Rose (2009) by using 20 quarterly observations (five years) preceding through time τ.

Table 3. Dynamic Panel GMM estimations of Business Cycle Synchronizations (Crisis Years Excluded)

Variables	Trade ($T_{ij,t}$)	Specialization ($S_{ij,t}$)	Finance ($F_{ij,t}$)	Hansen test (p-value)	AB test for AR(2) (p-value)
			Hodrick–Prescott Detrended		
T^1, S^1, F^1	1.092 [0.00]***	–0.048 [0.76]	–0.435 [0.01]**	0.39	0.42
T^2, S^1, F^1	0.494 [0.01]**	0.636 [0.01]**	–0.573[0.00]***	0.74	0.36
T^1, S^2, F^1	0.638 [0.00]***	1.082 [0.03]**	–0.468 [0.00]***	0.42	0.80
T^2, S^2, F^1	0.623 [0.01]**	–2.078 [0.02]**	–0.662 [0.00]***	0.91	0.19
T^1, S^1, F^2	0.430 [0.00]***	–0.119 [0.19]	–0.114 [0.00]***	0.17	0.82
T^2, S^1, F^2	0.954 [0.02]**	0.802 [0.39]	–0.202 [0.51]	0.35	0.14
T^1, S^2, F^2	0.408 [0.00]***	0.865 [0.10]	–0.109 [0.00]***	0.66	0.86
T^2, S^2, F^2	0.300 [0.06]**	3.069 [0.00]***	–0.448 [0.00]***	0.22	0.98
T^1, S^1, F^{LS}	1.150 [0.01]**	–0.061 [0.73]	0.101 [0.05]**	0.88	0.88
T^2, S^1, F^{LS}	1.398 [0.03]**	–2.877 [0.00]***	0.011 [0.89]	0.10	0.40
T^1, S^2, F^{LS}	0.699 [0.00]***	0.296 [0.66]	0.063 [0.04]**	0.38	0.67
T^2, S^2, F^{LS}	0.150 [0.86]	17.583 [0.00]***	0.142 [0.10]	0.15	0.92
T^1, S^1, F^{LL}	1.218 [0.03]**	–0.133 [0.45]	0.115 [0.06]*	0.66	0.67
T^2, S^1, F^{LL}	1.011 [0.03]**	–1.179 [0.10]	0.011 [0.89]	0.25	0.77
T^1, S^2, F^{LL}	0.701 [0.01]**	0.640 [0.38]	0.066 [0.04]**	0.39	0.84
T^2, S^2, F^{LL}	0.883 [0.09]*	9.417 [0.02]***	0.106 [0.24]	0.36	0.72
			Baxter–King Detrended		
T^1, S^1, F^1	0.172 [0.01]**	–0.101 [0.06]*	–0.271 [0.00]***	0.43	0.85
T^2, S^1, F^1	0.190 [0.00]***	–0.103 [0.04]**	–0.222 [0.00]***	0.32	0.93
T^1, S^2, F^1	0.202 [0.01]**	0.535 [0.09]*	–0.278 [0.00]***	0.42	0.78
T^2, S^2, F^1	0.194 [0.00]***	0.591 [0.05]*	–0.216 [0.00]***	0.28	0.96
T^1, S^1, F^2	0.021 [0.85]	–0.291 [0.08]*	–0.082 [0.00]***	0.58	0.22
T^2, S^1, F^2	0.042 [0.01]**	–0.620 [0.01]**	–0.154 [0.00]***	0.96	0.13
T^1, S^2, F^2	0.086 [0.40]	0.967 [0.04]**	–0.123 [0.00]***	0.97	0.89
T^2, S^2, F^2	0.046 [0.01]**	1.391 [0.01]**	–0.176 [0.00]***	0.94	0.33
T^1, S^1, F^{LS}	0.524 [0.09]*	–0.157 [0.17]	0.084 [0.01]**	0.60	0.85
T^2, S^1, F^{LS}	0.003 [0.99]	–0.108 [0.05]*	0.048 [0.00]***	0.39	0.29
T^1, S^2, F^{LS}	0.975 [0.02]**	1.224 [0.17]	0.095 [0.00]***	0.34	0.83
T^2, S^2, F^{LS}	0.025 [0.86]	0.580 [0.12]	0.063 [0.02]**	0.48	0.59
T^1, S^1, F^{LL}	0.160 [0.37]	–0.118 [0.17]	0.055 [0.03]**	0.13	0.46
T^2, S^1, F^{LL}	0.083 [0.67]	–0.108 [0.08]*	0.048 [0.00]***	0.19	0.14
T^1, S^2, F^{LL}	0.591 [0.10]	0.636 [0.37]	0.061 [0.04]**	0.13	0.94
T^2, S^2, F^{LL}	0.006 [0.96]	0.703 [0.07]*	0.068 [0.01]**	0.52	0.93

Notes Numbers in square brackets are p-values. Significance levels: *10%, **5%, ***1%.

The GMM estimations include lags of the dependent variable as well as time-dummies.

Table 4. Dynamic Panel GMM estimations of Business Cycle Synchronizations (Using CPIS Data)

Variables	Trade ($T_{ij,t}$)	Specialization ($S_{ij,t}$)	Finance ($F_{ij,t}$)	Hansen test (p-value)	AB test for AR(2) (p-value)
		Hodrick–Prescott Detrended			
T^1, S^1, F^3	0.051 [0.04]**	–0.042 [0.17]	–0.050 [0.09]*	0.53	0.31
T^2, S^1, F^3	0.089 [0.05]*	–0.058 [0.29]	–0.058 [0.27]	0.77	0.28
T^1, S^2, F^3	0.010 [0.81]	0.552 [0.09]*	–0.054 [0.05]**	0.86	0.32
T^2, S^2, F^3	0.062 [0.09]*	0.128 [0.73]	–0.019 [0.66]	0.30	0.23
		Baxter–King Detrended			
T^1, S^1, F^3	1.141 [0.03]**	–0.427 [0.01]***	–0.348 [0.03]**	0.23	0.20
T^2, S^1, F^3	1.066 [0.52]	–7.641 [0.02]**	–0.339 [0.09]*	0.41	0.36
T^1, S^2, F^3	1.722 [0.30]	–1.089 [0.94]	–0.581 [0.05]*	0.24	0.83
T^2, S^2, F^3	0.063 [0.93]	–5.743 [0.62]	–0.458 [0.04]**	0.16	0.34

Notes: Numbers in square brackets are p-values. Significance levels: *10%, **5%, ***1%.

The GMM estimations include lags of the dependent variable as well as time-dummies.

Source: Authors' calculations.

6. Policy Implications

The implications of the above results for Asian policy making are important and far reaching. Given that Asia's integration within itself and with countries outside the region has advanced steadily in recent years, the finding of an inverse relationship between bilateral financial linkages and BCS might indicate that financial integration makes it easier for these countries to transfer financial assets, which should help Asian countries diversify themselves against country-specific risks, and thus enable their decoupling (Flood et al, 2009; Garcia-Herrero and Ruiz, 2008). This is consistent with the basic argument underlying the theory of international risk sharing, as further financial integration offers better opportunities for risk diversification.

In view of the benefits of increasing financial integration by affording better risk sharing opportunities, the importance in actuality of developing and strengthening capital markets through which agents can diversify their portfolio via cross-border ownership of assets is essential and cannot be over-emphasized. Moving forward, reforms aimed at improving the capability of agents to adjust the size of their asset portfolio (savings rate), e.g. pension-fund schemes, in response to shocks as well as to being more receptive of cross-border M&As can further enhance risk-sharing opportunities between countries (Kalemli-Ozcan et al., 2004).

On the other hand, it is alleged that the downside of greater financial integration is that it can pose risks to financial stability. This claim takes on ever-increasing traction and prominence in discussions, especially in light of the recent painful experience with the global financial crisis, which tends to demonstrate the role that financial linkages play in the transmission of shocks between economies. However, despite concerns in policy circles, it appears that the jury is still out on whether greater financial integration increases the likelihood of crises. For instance, in a recent study Fecht et al. (2008) show that the availability of better risk-sharing mechanisms tend to offset the risk of spillover or transmission of shocks, and thus financial integration leads to an improvement in welfare as specialization benefits are magnified and realized.[18] Furthermore, a few studies argue that greater integration poses no risk to financial stability on its own, but when too-rapid a liberalization of financial markets *interact*, for instance, with certain distortions in the economy such as weak and lax supervisory regulations as well as problems of credibility and enforcements of contracts, these distortions get magnified and financial instability problems arise (Fecht et al., 2008).[19]

Nonetheless, as financial integration deepens and international capital flows become more intense the enduring and considerable challenge for policy makers is to be able to counter balance the benefits and alleged risk of greater financial integration. This trade-off will vary across economies as it will be a function of the level of financial market development of individual countries in the region. Thus, the appropriate policy approach to enable countries to reap the benefits of the opportunities that come with further financial integration, e.g. international risk sharing, and at the same time contain the alleged inherent risk of greater financial integration, will likely differ across countries in the region (BIS, 2009).

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[18] See also Bonfiglioli and Mendicino (2004) and Glick et al. (2006). For contrasting evidence, see for instance, Demirguc-Kunt and Detragiache (2001) and Hartmann et al. (2005).

[19] See, for instance, Ishii and Habermeier (2002).

7. Conclusion

This paper examines empirically the effect of three factors, viz., financial and trade linkages as well as similarity of industrial structure on BCS in the Asia–Pacific region. We adopt a panel approach and employ dynamic panel GMM estimation in order to unpack endogeneity problems as well as to control for country-pair specific unobservable time-invariant factors. Our modeling strategy assumes that the Asia–Pacific countries integrate not just from within the region, but also with major countries outside the region. As also documented in the paper, this strategy conforms with the stylized fact that Asia's integration within itself and countries outside the region have increased considerably over the past three decades. Furthermore, as opposed to previous studies, our measures of financial integration encompass both quantity- and price-based indicators to ensure robustness in our results. The main results are as follows: some indications that more similar countries have higher BCS; consistent support to the evidence that shows that more trade between two countries induces higher BCS as well as that increased financial linkages between countries leads to divergent BCS.

While the first two findings are in line with Imbs (2004, 2006), the third is contrary to his and other studies that use pure-cross sectional data to examine the direction of the finance–BCS link. Our third finding conforms with the basic ideas of benchmark international business cycle models. What perhaps drives the conflicting evidence found by pure cross-sectional studies from that of the theoretical benchmarks is that, apart from being able to control for simultaneity, one also needs to account for the role of country-pair specific unobservable time-invariant factors. Similar to our approach that used panel IV techniques, recent studies such as Cerqueira and Martins (2009) and Kalemli-Ozcan et al. (2009) have controlled for such influences. These works are now considered to be fundamental and basic in any study that tries to ascertain the link between BCS and finance. Once biases associated with simultaneity and unobserved country-pair specific factors are accounted for, the prediction by theory regarding the inverse relationship between financial linkages and BCS become apparent in the results.

In terms of future work, it would be useful to disentangle the possible differing effects that different types of financial integration have on synchronicity of business cycles. This would allow a close examination of whether the rapid expansion of equity and FDI positions may make the economy more susceptible to external economic influences, and thus lead to more synchronized business cycles. It would also be desirable to explore the regional dimension of the linkage between financial integration and BCS, and to ascertain whether our results “hold out” across various country configurations in the region. Finally, it would also be interesting to make use of an event-study approach to capture explicitly the period of the financial turmoil associated with the Asian financial crisis of 1997–1998 as well as with the global financial crisis of 2008–2009, and, accordingly, to pin down the extent of the transmission of financial crisis to business cycles, while controlling for the significant factors that drive BCS.

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Table A1. Correlation Matrix of Variables Using CPIS Data

	ρ^{HP}	ρ^{BK}	T^1	T^2	S^1	S^2	F^3
ρ^{HP}	1.0000						
ρ^{BK}	0.5791	1.0000					
T^1	0.2003	0.2168	1.0000				
T^2	0.1860	0.2275	0.8967	1.0000			
S^1	–0.0302	–0.0039	0.1332	0.2298	1.0000		
S^2	0.0726	0.1187	–0.0680	–0.1098	–0.8072	1.0000	
F^3	0.1956	0.1867	0.4128	0.2937	–0.1105	0.1435	1.0000

Note: All variables are in logarithmic form with the exceptions of ρ^{HP} and ρ^{BK}.

Source: Authors' calculations.