Trade linkages and the globalisation of inflation in Asia and the Pacific[1]

Raphael A Auer[2] and Aaron Mehrotra[3]

This version: September 12, 2013

Abstract

Some observers argue that increased real integration has led to greater co-movement of prices internationally. We examine the evidence for cross-border price spillovers among economies participating in the pan-Asian cross-border production networks. Starting with country-level data, we find that both producer price and consumer price inflation rates move more closely together between those Asian economies that trade more with one another, ie that share a higher degree of trade intensity. Next, using a novel dataset based on the World Input-Output Database (WIOD), we examine the importance of the supply chain for cross-border price spillovers at the sectoral level. We document the increasing importance of imported intermediate inputs for economies in the Asia-Pacific region and examine the impact on domestic producer prices of changes in costs of imported intermediate inputs. Our results suggest that real integration through the supply chain matters for domestic price dynamics in the Asia-Pacific region.

Keywords: globalisation, inflation, Asian manufacturing supply chain, price spillovers

JEL classification: E31, F62, F14

1. Introduction

Some observers argue that increased real integration, ie increased international trade in goods and services, has led to greater co-movement of prices internationally. This could occur directly, through import prices, or more indirectly, due to the effect of increased international competition on domestic price mark-ups and overall wage and price setting dynamics. Most of the literature on the impact of globalisation on prices has focused on inflation in the advanced economies, treating Asian economies as a source of low-cost exports that could put downward pressure on inflation in advanced economies. Auer and

[1] The views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank for International Settlements or the Swiss National Bank. Emese Kuruc, Bat-el Berger and Pierre-Yves Deléamont provided excellent research assistance. We are grateful to Eli Remolona and James Yetman, and seminar participants at the Bank for International Settlements for helpful comments and suggestions.

[2] Swiss National Bank. Email: raphael.auer@snb.ch

[3] Bank for International Settlements, Representative Office for Asia and the Pacific. Email: aaron.mehrotra@bis.org

Fischer (2010) and Auer et al. (2013) find strong downward impacts from import competition in the emerging markets on producer prices in the United States and Europe, respectively.[4]

Greater real integration could also increase the sensitivity of inflation to cross-border shocks. The average economy is now substantially more integrated to world trade than was the case some decades ago. Worldwide, exports of goods as a share of global GDP have increased from 17% to 25% during 1980–2012. For twelve economies in the Asia-Pacific region, the increase has been even more prominent, with the share climbing from 15% to 26% during the same time period.[5] As a result, it seems fair to argue that the average economy is now more prone to international shocks *via* the trade channel than before. Such an impact is in addition to the effect of globalisation on the level of inflation due to import competition.

The increased real integration in Asia is reflected especially in the region's manufacturing supply chains. In closely integrated supply chains, any shock to domestic production costs or exchange rates could be easily passed through to economies in the supply chain, affecting intermediate prices in other economies, with potential implications for headline inflation as well.[6] While previous research has addressed various issues related to international production networks (see Baldwin and Lopez-Gonzalez, 2013, on the global pattern of supply-chain trade), the impact on inflation has not been analysed.[7] An exception is Auer and Sauré (2013) who analyse price spillovers specifically focusing on the supply chain and present a theoretical model capturing the channels through which spillovers occur. Our paper draws partly on their analysis.

In this study, we examine the evidence for cross-border price spillovers among economies participating in the pan-Asian supply chain. Instead of treating the Asian economies as sources of low-cost exports, we consider them importers themselves, and as such prone to cross-border shocks resulting from closer real integration. We study both aggregate (country-level, final goods) prices, and disaggregate (sector-and-country level, producer) prices. In the sectoral analysis, we draw on the novel World Input-Output Database (WIOD, 2012, and Timmer et al., 2013 a and b). This database is an extension of the national input-output tables, and was developed to analyse, *inter alia*, the effects of globalisation on trade patterns.

In our framework, as in Amiti et al (2012), the presence of imported intermediate goods implies that the exchange rate affects the domestic cost of production. Moreover, mark-ups are variable, so firms may not fully pass cost shocks through to prices. In the empirical analysis, we evaluate the extent to which domestic producer prices react to changes in costs of imported intermediate inputs, the latter possibly caused by exchange rate movements. A

[4] Lipińska and Millard (2012) show in the context of a theoretical model how productivity increases in the developing economies could lead to higher inflation in the advanced economies, depending on oil demand elasticities and the structure of labour markets. See also Holz and Mehrotra (2013) and the discussion in BIS (2009).

[5] These are the twelve BIS member economies in the Asia-Pacific region: Australia, China, Hong Kong SAR, India, Indonesia, Japan, Korea, Malaysia, New Zealand, the Philippines, Singapore and Thailand.

[6] In some cases, a supply chain may decrease the sensitivity of prices to exchange rate changes. If production is divided between a large number of firms located in different countries each adding a level amount of value to the finished product, then a depreciation in the local exchange rate implies, *ceteris paribus*, a nearly offsetting increase in both costs and revenues. In a sticky price context, a finely-divided supply chain may therefore be less sensitive to exchange rate changes.

[7] In a recent paper, Saito and Ruta (2013) find that higher economic growth is positively associated with participation in global supply chain networks.

crucial role is played by the importance of imported inputs as a fraction of a sector's total variable costs.

To preview our findings, we find that both headline inflation rates and producer prices move more closely together between those Asian economies that trade more with one another, ie that share a higher degree of trade intensity. Moreover, the impacts through higher costs of imported intermediate inputs on domestic producer prices are statistically and economically significant for economies participating in the supply chain. We show that the share of imported intermediate inputs in total costs is roughly 17% on average, for the seven Asia-Pacific economies for which data are available in the WIOD database.[8] When prices of intermediate imports change by 1%, domestic producer prices change cumulatively by close to 0.3% over the following two years.[9] These results suggest that real integration through the supply chain matters for domestic price dynamics in the Asia-Pacific region.

Our paper is not the first to investigate the importance of real integration – or international factors more broadly – for inflation dynamics. Using a factor model and data for 22 OECD economies, Ciccarelli and Mojon (2010) show that a global factor accounts for nearly 70% of the variance of national inflation rates. Moreover, they identify an error correction mechanism, whereby national inflation rates converge to global inflation. Mumtaz and Surico (2012) also suggest, using a factor model, that an international factor tracks both the level and persistence of national inflation rates well.

On the importance of the trade channel, Monacelli and Sala (2009) find a significant relationship at the sectoral level between the importance of the common international factor in driving prices and trade openness. Trade links are also relevant in the literature on global output gaps, whereby (trade-weighted) measures of external output gaps have been found to significantly contribute to domestic inflation equations (eg Borio and Filardo, 2007; earlier work includes eg Tootell, 1998).[10] Nevertheless, previous evidence on the importance of the trade channel for Asian inflation dynamics is sparse. The lack of previous research is particularly evident in the case of sectoral data – but it is only at this level where the supply chain links can be effectively analysed.

This paper is structured as follows. The next section presents some stylised facts about comovement of inflation in the Asia-Pacific region and the increasing importance of intraregional trade. In Section 3, we describe our dataset and the empirical framework, and present the results of the analysis of sectoral price spillovers along the supply chain. Section 4 concludes the paper.

2. Some stylised facts

In this section, we present some stylised facts about the co-movement of inflation in the Asia-Pacific region in the past three decades, together with the increased importance of intraregional trade. Our focus in the country-level analysis is on the twelve BIS member economies in the Asia-Pacific region: Australia, China, Hong Kong SAR, India, Indonesia, Japan, Korea, Malaysia, New Zealand, the Philippines, Singapore and Thailand. We use

[8] Australia, China, Chinese Taipei, India, Indonesia, Japan and Korea.
[9] Computed at the median imported-input intensity in our sample.
[10] Milani (2009) confirms the reduced-form evidence on the importance of global output gaps in the context of a structural model.

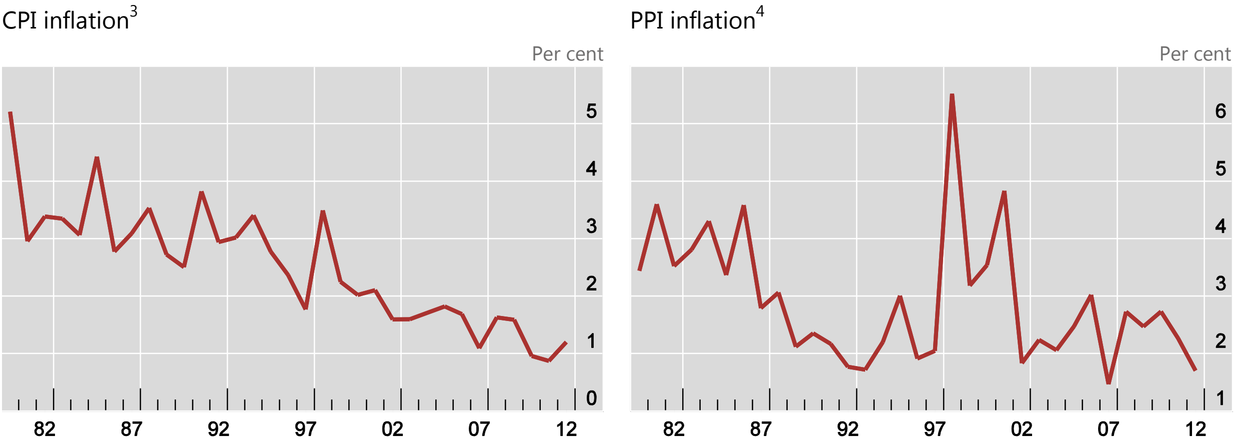
inflation rates at quarterly frequency given that monthly data are not available for Australia and New Zealand.

Graph 1 shows the standard deviation of economy-level inflation rates across the region, for both consumer and producer prices during 1980 to 2012. The further apart the inflation rates between the economies, the higher is the value shown in the graph. To remove outliers, we exclude the economies with both the lowest and highest inflation rates for each year from the construction of the standard deviation. The graph shows that the past three decades have seen a notable decline in divergence of the levels of inflation rates in the Asia-Pacific region, especially in terms of consumer prices (left-hand panel). This is consistent with the increased focus of macroeconomic policy on inflation control. Producer prices (right-hand panel) are typically more volatile, and are likely to be characterised by greater pass-through of shocks, due to cross-border supply chains, for example. The impact of the Asian crisis is notable in our sample in this regard.

Co-movement of prices in Asia[1]

Standard deviation of inflation rates across the economies[2]

Graph 1



[1] Australia, China, Hong Kong SAR, India, Indonesia, Japan, Korea, Malaysia, New Zealand, the Philippines, Singapore, Thailand. [2] Standard deviation is computed across the different economies, using average annual inflation rate for each economy for each year; the lowest and highest inflation rates are excluded for each year. [3] Wholesale prices for India. [4] Where data are available.

Sources: CEIC; Datastream; national data; authors' calculations.

A more formal test of inflation convergence could be provided by a test for panel cointegration, effectively testing for the presence of a long-run relationship among integrated variables with both a time-series dimension, T, and a cross-sectional dimension, N.[11] Our test consists of examining whether national inflation rates (both CPI and PPI) are cointegrated with a measure of Asian regional inflation over different time periods. The latter is defined as the simple average of the inflation rates in our sample of Asia-Pacific economies, excluding the home economy in question.

In the panel cointegration test suggested by Westerlund (2007), the null hypothesis of no cointegration is tested by examining whether the error correction term in a panel error correction model can be set to zero. The test is general enough to accommodate economy-specific short-run dynamics, unit-specific trend and slope parameters, as well as cross-

[11] Holmes (2002) uses this approach to investigate inflation convergence within the European Union.

sectional dependence.[12] A finding of cointegration, by definition, implies that the series share a common stochastic trend – but need not move strictly in unison over time. This could be taken as an indication of weak convergence, and is arguably not dissimilar to analysing the importance of a common global factor in driving domestic inflation dynamics (eg Ciccarelli and Mojon, 2010).

In order to check whether the co-movement between inflation rates has changed over time we divide the sample into two different periods: 1980Q1-1991Q4 and 1992Q1-2012Q4.[13] We allow for the existence of time trends. The *p*-values of the test statistics are computed by a bootstrapping procedure – this in effect deals with cross-sectional dependence between the different economies. The results from the tests are shown in Table 1.

Common stochastic trends in inflation rates in Asia-Pacific

Table 1

	CPI				PPI			
	Q1 1980–Q4 1991		Q1 1992–Q4 2012		Q1 1980–Q4 1991		Q1 1992–Q4 2012	
Statistics	Z-value	*p*-value	Z-value	*p*-value	Z-value	*p*-value	Z-value	*p*-value
G_T	–3.101	0.09*	–6.588	0.00***	–0.956	0.62	–7.003	0.00***
G_α	–1.969	0.33	–8.164	0.00***	–2.080	0.86	–7.709	0.00***
P_T	–7.774	0.02**	–6.384	0.03**	–2.140	0.24	–4.207	0.15
P_α	–6.233	0.10	–11.349	0.01***	–5.184	0.37	–9.207	0.06*

Note: The null hypothesis of no cointegration is tested by inferring whether the error correction term in the error correction model between a domestic inflation rate and Asian regional inflation (excluding the domestic economy in question) can be set to zero. The test statistics are defined in Westerlund (2007). Robust *p*-values are computed by a bootstrapping procedure with 200 replications. The lag structure of the error correction model is chosen by the Akaike information criterion.

Sources: authors' calculations.

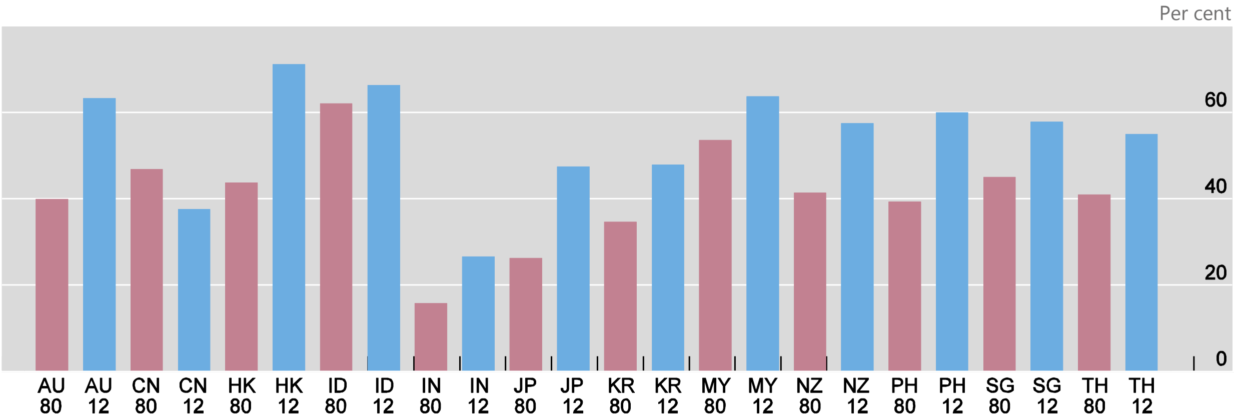
Most of the test statistics (using robust *p*-values) indicate that the null hypothesis of no cointegration cannot be rejected at the 5% level in the earlier sample period for either CPI or PPI inflation rates. In contrast, the results are indicative of cointegration between the national and regional inflation rates in the latter part of the sample. The finding of cointegration in the most recent sample holds most strongly for consumer prices, but evidence of a common stochastic trend can be found for producer prices as well. Overall, these results suggest that the inflation rates between the individual economies and the region as a whole share a common stochastic trend in the latter sample.

[12] The reader is referred to Westerlund (2007) for details. The tests have good small-sample properties with small size distortions and high power compared to other, residual-based, panel cointegration tests.

[13] When considering producer prices during the 1980Q1-1991Q4 subsample we omit China, Hong Kong SAR and the Philippines due to lack of data. In contrast, these economies are included in the more recent sample (1992Q1-2012Q4).

At the same time as the co-movement of inflation has increased, trade between these same economies has accounted for an increasing share of their total international trade. Graph 2 shows that for all economies except China, the share of trade with other Asian economies was larger in 2012 than in 1980. For many of them (all except China, India, Japan and Korea), the trade with our sample of Asian economies accounts for over 50% of total international trade. Part of the increase reflects the deepening of the cross-border supply chains in Asia, which we investigate formally in Section 3.

Increasing role of intraregional trade in Asia[1] Graph 2



AU = Australia; CN = China; HK = Hong Kong SAR; ID = Indonesia; IN = India; JP = Japan; KR = Korea; MY = Malaysia; NZ = New Zealand; PH = Philippines; SG = Singapore; TH = Thailand.

[1] Share of trade with Asia as a share of total international trade. Trade is the sum of exports and imports. The Asian aggregate includes Australia, China, Hong Kong SAR, India, Indonesia, Japan, Korea, Malaysia, New Zealand, the Philippines, Singapore, Thailand. For Singapore in 1980, trade with Indonesia is missing. China's exports to India in 1980 are proxied by their corresponding value in 1981.

Source: IMF, *Direction of Trade Statistics*; authors' calculations.

Is there any evidence that a greater interconnectedness through trade led to the observed co-movement of inflation rates at the economy-wide level? To answer this question, we estimate a model that explicitly links the correlation between headline inflation rates of different economies with their bilateral trade intensities – the latter measuring how closely two economies are connected through international trade. Our approach is similar to Artis and Okubo (2012), who examine the impact of trade on business cycle synchronisation using a gravity-type model.[14]

The estimated equation is of the form:

$$Corr\_\pi_{ijt} = \alpha + \beta_1(Trade_{ijt}) + \beta_2(Corr\_mon\_pol_{ijt}) + \beta_3\left(Corr\_y^{gap}{}_{ijt}\right) + \varepsilon_{ijt}, \qquad (1)$$

where $Corr\_\pi_{ij}$ is the correlation of inflation rates between economies *i* and *j*. We examine both headline (consumer price) and producer price inflation. $Trade_{ij}$ denotes the trade intensity between the two economies, measured by export intensity. Following Artis and Okubo (2012), this is defined as:

[14] In the specification by Artis and Okubo (2012), the correlation between business cycles is regressed on trade intensity, the product of GDPs of the two economies and a dummy variable indicating a common currency.

$$Trade_{ij} \equiv \frac{X_{ij}/X_i}{M_j/(M^W - M_i)} \ , \qquad (2)$$

where X_i denotes total exports from country *i* and M_i is total imports into country *i*. X_{ij} is exports from country *i* to country *j*. M^W denotes the total world imports. A higher value for this index indicates that the two economies are more closely linked by international trade. $Corr\_mon\_pol_{\,ij}$ denotes the correlation between broad money growth between economies *i* and *j*, used as a proxy for the commonality of the monetary policy stance, which is relevant due to the monetary nature of inflation in the long run.[15] Indeed, Mumtaz and Surico (2012) find that co-movements in money supply across countries are positively correlated with co-movements in inflation rates internationally. As a robustness test, we use the correlation of real interest rates between the two economies as a measure of the commonality of the monetary policy stance. Finally, $Corr\_y^{gap}{}_{ij}$ denotes the correlation between the output gaps of economies *i* and *j*, capturing the co-movement of inflation that is due to business cycle co-movement.[16]

The time dimension of the panel is comprised of three observations: 1980–1989, 1990–1999 and 2000–2012. The exact starting date depends on data availability for each economy-pair (see Appendix for details). Over these three time periods, the average CPI correlation in our sample is 0.32 and the average PPI correlation is 0.39; the standard deviations of these same correlations are 0.32 and 0.37, respectively.[17] The right-hand side variables are normalised by dividing them by their standard deviation. Hong Kong SAR and Singapore are omitted, due to the different structure of trade flows in these economies, in particular the importance of re-exports (see eg Feenstra & Hanson, 2004).

Table 2 shows that there is a positive and statistically significant correlation between the co-movement of inflation rates (both CPI and PPI) and trade intensity, for our sample of Asian economies. Regarding the magnitude of the estimated coefficients, specifications (1) to (3) suggest that a one standard deviation increase in trade intensity between two economies is associated with a 5 to 6 percentage point increase in the correlation between the CPI inflation rates of the same economies. Similarly, a rise in trade intensity of the same size leads to roughly 6 percentage point increase in the correlation between the PPI inflation rates of the same economies (specifications 4, 5 and 6). The findings remain robust to the inclusion of broad money growth and output gaps in the estimation, suggesting that more highly correlated inflation rates are not driven solely by commonality of monetary policy or

[15] Ciccarelli and Mojon (2010), in the working paper version (2008) of their article, emphasise that commonalities in monetary policy could *reduce* the importance of the global component of inflation at business cycle frequencies. This could arise if central banks around the world follow the same reaction function and offset those inflation movements that are due to global forces, for example.

[16] For the estimation of (1), $Corr\_\pi_{ijt}$ is expressed as the correlation of inflation rates (CPI or PPI; quarterly year-on-year) during the time period considered. The trade data are the averages of annual totals for the same periods. $Corr\_mon\_pol_{ij}$ is the correlation between broad money growth (quarterly year-on-year) in the two economies. M3 is used for Australia and Thailand; M2 for all others. The measures for the output gaps in $Corr\_y^{gap}{}_{ij}$ are obtained from data on annual real GDP, using the Hodrick-Prescott filter with a conventional smoothing parameter of 100.

[17] Omitting Hong Kong SAR and Singapore, as in the estimation reported in Table 2.

common business cycles. They are also robust to the use of real interest rates as a measure of the commonality of the policy stance.[18]

Inflation co-movement and trade intensity

Table 2

Model	(1)	(2)	(3)	(4)	(5)	(6)
Price	CPI	CPI	CPI	PPI	PPI	PPI
Trade intensity	0.056*** (0.010)	0.057*** (0.010)	0.052*** (0.010)	0.063*** (0.013)	0.062*** (0.013)	0.059*** (0.014)
Correlation of broad money growth		-0.008 (0.024)	-0.009 (0.025)		0.007 (0.027)	0.007 (0.027)
Correlation of output gaps			0.069*** (0.024)			0.039 (0.025)
N	210	210	210	200	200	200
Adjusted R^2	0.068	0.064	0.102	0.266	0.263	0.270

Note: *, ** and *** denote statistical significance at 10%, 5% and 1% levels, respectively. Estimated with period fixed effects. White robust standard errors in parentheses. Excluding Hong Kong SAR and Singapore. Coefficients are normalised by dividing the variables by their standard deviation.

Sources: authors' calculation.

It may appear surprising that even as consumer prices are affected by distribution costs and other non-tradable components, the point estimate on the trade intensity variable is only slightly higher when producer prices are used, with no statistically significant difference between the coefficients. However, as argued by Campa and Goldberg (2010), when imported inputs are used in the production of non-tradables, the sensitivity of consumer prices to exchange rate movements is enhanced. We also note that the goodness of fit of the model with producer prices is markedly higher than the one with consumer prices, as shown in the last row of Table 2.

These results should only be taken as suggestive. Reverse causality cannot be ruled out in the econometric specification, and factors other than trade intensity and our measure of commonality of monetary policy may lead to a closer co-movement of inflation rates. Indeed, the adjusted R squared values suggest that some, but clearly not all, of the variation in inflation co-movement is captured by our model. For these reasons, the use of sectoral data has benefits. It allows us to disentangle the impact of real integration in a *relative* sense – sectors that are more closely connected between different economies are also more likely to experience greater cross-border price spillovers, controlling for country-specific or common trends across time. The next sections use sectoral data that reflect variation in the Asian supply chain to investigate the link between real integration and price spillovers.

[18] Results are available upon request. While the correlation between real interest rates is highly statistically significant when used in specifications (2) and (4), part of the close correlation with inflation likely arises due to the construction of the real interest rate variable (specified as nominal interest rate less inflation).

3. Supply chain links and price spillovers into the Asia-Pacific

3.1. Details about the data

We use data from the World Input-Output Database (WIOD henceforth) to illustrate the supply chain dynamics and measure price spillovers within the Asia-Pacific region. The WIOD has been developed, among other objectives, to aid the analysis of the effects of globalisation on trade patterns (WIOD, 2012). The world input-output table is basically an extension of the national input-output tables. The national tables specify, for each industry, the use of the product, being either for industry (intermediate) or final use. Final use includes domestic final use (private consumption, government consumption, investment) and exports. The difference with the world input-output table is that the world table breaks down the use of products by their origin – each product is produced either by a domestic or a foreign industry. The world table also shows in which foreign industry the product was produced, and how the exports of a country are used, by which foreign industry or final end user.

The WIOD covers 27 EU countries and 13 other major economies, for 1995-2009. We focus our analysis on data from the Asia-Pacific region, ie China, India, Japan, Korea, Australia, Chinese Taipei and Indonesia, with all the WIOD's 40 economies included as trading partners. Table 3 shows the economies included in the database.

European Union			North & Latin America	Asia and Pacific
Austria	Germany	Netherlands	Canada	China
Belgium	Greece	Poland	United States	India
Bulgaria	Hungary	Portugal	Brazil	Japan
Cyprus	Ireland	Romania	Mexico	Korea
Czech Republic	Italy	Slovak Republic		Australia
Denmark	Latvia	Slovenia	**Other**	Chinese Taipei
Estonia	Lithuania	Spain	Turkey	Indonesia
Finland	Luxembourg	Sweden	Russia	
France	Malta	United Kingdom		

Table 3: Economies included in the world input-output database

Note: The regional classification in this study has been adapted to more closely match the BIS definition of regions.

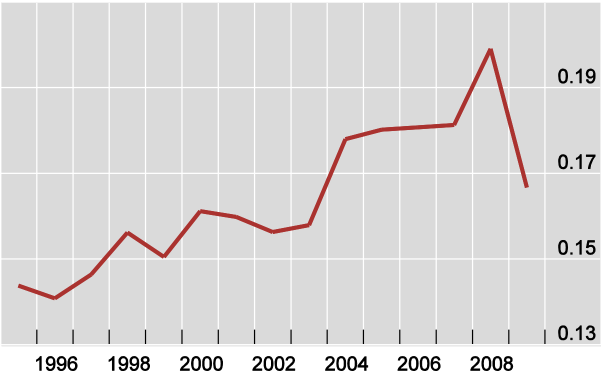
Differences in the intensity of import use across sectors play a crucial role in our analysis. Using data from the WIOD, we specify the intensity of import use as the share of intermediate imports in total output. We further assume that variable costs account for 70% of total costs, and proxy the latter by the value of total output. The cost share of imported intermediate inputs is then equal to the share of imported intermediate inputs in total variable costs.

How has the cost share of imported inputs changed over time in the Asia-Pacific economies? Graph 3 (left-hand panel) shows that, pooling the seven economies together, the cost share of imported inputs increased from roughly 14% in 1995 to 16% in 2000. This was followed by a rapid increase in the cost share from 2003 to 2008, reaching a level close to 20%, before global trade collapsed at the time of the international financial crisis. The right-hand panel shows how the cost share has evolved in the different economies. In all other economies except Indonesia, the cost share of imported inputs was higher in 2008 than in 1998. In Chinese Taipei, the cost share was close to 35% in 2008. Perhaps surprisingly, given the importance of mainland China in the global supply chains, the imported input cost share in China in 2008 was lower than in the other economies, with the exception of Japan. Nonetheless, the increase in the imported input cost share from 1998 to 2008 in China was particularly prominent, increasing two-fold during this ten-year period.

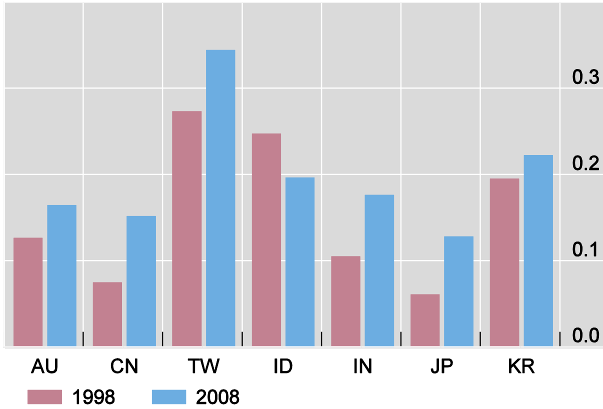
Cost share of imported inputs

Graph 3

Average imported input cost share, 7 economies pooled



Average imported input cost share, 1998 and 2008



AU = Australia; CN = China; ID = Indonesia; IN = India; JP = Japan; KR = Korea; TW = Chinese Taipei.

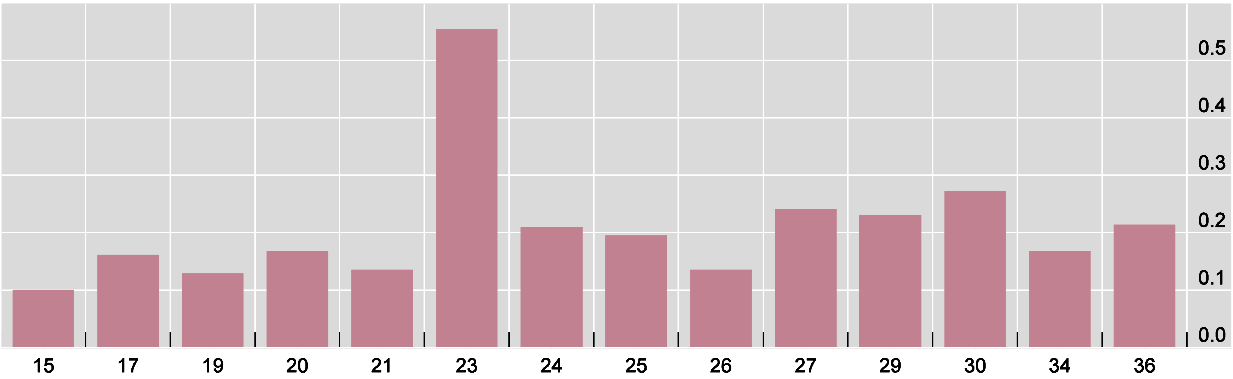
Source: Authors' calculations based on WIOD (2012).

For this group of seven Asia-Pacific economies, Graph 4 displays the sectoral imported input cost shares in 2008. The sectors follow the NACE classification. For the textiles sector, the cost share is 16%, and it exceeds 20% for the manufacturing of machinery and equipment, for example. A particularly high cost share, 55%, is recorded for the "coke, refined petroleum products and nuclear fuel" sector.

Imported input cost shares in 2008, sectoral level

7 Asia-Pacific economies pooled

Graph 4



15 = food products and beverages; 17 = textiles; 19 = leather and leather products; 20 = wood and wood products; 21 = pulp, paper and paper products; 23 = coke, refined petroleum products and nuclear fuel; 24 = chemicals and chemical products; 25 = rubber and plastic products; 26 = other non-metallic mineral products; 27 = basic metals; 29 = machinery and equipment n.e.c.; 30 = office machinery and computers; 34 = motor vehicles, trailers and semi-trailers; 36 = furniture, manufacturing n.e.c..

Source: Authors' calculations based on WIOD (2012)

3.2. Framework relating import prices to production costs

In the presence of imported intermediate goods, fluctuations in the prices of imports that are themselves driven by exchange rate movements affect the domestic cost of production and, ultimately, the prices that domestic producers and exporters charge (eg Amiti et al., 2012). We next lay down a parsimonious model to describe these relations, motivating our following empirical analysis.

Denoting the domestic price that firm *n* charges for its good by $p^D_{n,t}$, the marginal cost of producing one unit of this good by $c^D_{n,t}$, and the mark-up of firm *n* by $\pi^D_{n,t}$, it holds that

$$p^D_{n,t} = \pi^D_{n,t} * c^D_{n,t} \qquad . \qquad (3)$$

The markup $\pi^D_{n,t}$ itself is a function that depends on, among other things, the per-unit cost of the firm. Denoting percentage changes of any variable by a hat, it then follows that

$$\hat{p}^D_{n,t} = \hat{\pi}^D_{n,t} + \hat{c}^D_{n,t} = \left(\frac{\partial \pi^D_{n,t}}{\partial c^D_{n,t}} \frac{c^D_{n,t}}{\pi^D_{n,t}} + 1\right) \hat{c}^D_{n,t} + \hat{\epsilon}^\pi_{n,t} \quad , \qquad (4)$$

ie the percentage change in the firms' price is equal to the change in its markup and the change in its costs. Since the markup itself is a function of marginal costs, the change in the price is then equal to $\frac{\partial \pi^D_{n,t}}{\partial c^D_{n,t}} \frac{c^D_{n,t}}{\pi^D_{n,t}} + 1$ multiplied by any cost change plus $\hat{\epsilon}^\pi_{n,t}$, which capture other fluctuations in markups that are uncorrelated with cost changes.

Costs are composed of local costs that are paid in local currencies ($c^L_{n,t}$, equal to the cost per unit $p^L_{n,t}$ multiplied by the quantity consumed $q^L_{n,t}$) and imported intermediate inputs ($c^I_{n,t}$, equal to the cost per unit $p^I_{n,t}$ multiplied by the quantity consumed $q^I_{n,t}$):

$$c^D_{n,t} = c^I_{n,t} + c^L_{n,t} = p^I_{n,t} * q^I_{n,t} + p^L_{n,t} * q^L_{n,t} \quad . \qquad (5)$$

If the input quantities do not change, and allowing for firm-specific cost shocks $\hat{\epsilon}^{C}_{n,t}$, we have

$$\hat{c}^{D}_{n,t} = \theta^{I}_{n,t} * \hat{p}^{I}_{n,t} + \left(1 - \theta^{I}_{n,t}\right) * \hat{p}^{L}_{n,t} + \hat{\epsilon}^{C}_{n,t}, \qquad (6)$$

where $\theta^{I}_{n,t}$ is the imported input cost share equal to $\frac{c^{I}_{n,t}}{c^{I}_{n,t}+c^{L}_{n,t}}$. It thus holds that

$$\hat{p}^{D}_{n,t} = \left(\frac{\partial \pi^{D}_{n,t}}{\partial c^{D}_{n,t}} \frac{c^{D}_{n,t}}{\pi^{D}_{n,t}} + 1\right) \theta^{I}_{n,t} * \hat{p}^{I}_{n,t} + \left(\frac{\partial \pi^{D}_{n,t}}{\partial c^{D}_{n,t}} \frac{c^{D}_{n,t}}{\pi^{D}_{n,t}} + 1\right)\left(1 - \theta^{I}_{n,t}\right) * \hat{p}^{L}_{n,t} + \hat{\epsilon}_{n,t} \ . \qquad (7)$$

We observe changes in both traded and nontraded input costs and $\theta^{I}_{n,t}$ in our data. We thus estimate markup elasticities in the data. Amiti et al. (2012) derive a theoretical measure of how the firm-specific markup evolves in the presence of imported input use and pricing to market.

Before presenting the estimation results, we describe the construction of the *IIPI*, the import price index that is denoted by $\hat{p}^{I}_{n,t}$ above. Each importing industry uses inputs from a variety of other sectors. We thus construct a sector's *IIPI* as the weighted average of all import price indices (*IPI*) of country *c* using the relative importance of imported inputs from sector *k* as a fraction of all inputs used by sector *s* as weights. That is, we construct:

$$I\widehat{IP}I^{I}_{c,s,t} = \frac{Inputs_{c,s,k,t}}{\sum_{k \epsilon S} Inputs_{c,s,k,t}} \widehat{IP}I^{I}_{c,k,t} \ , (8)$$

where *k* denotes the sector supplying the inputs and *s* the sector using them.

3.3. Results

In the following, we examine the extent to which domestic producer prices react to changes in imported input costs. A crucial role is played by the importance of imported inputs as a fraction of a sector's total variable costs. In all specifications of the upper panel of Table 4, the dependent variable is defined as the monthly change in the (log of the) sectoral producer price index of the importer. All models include sector fixed effects to capture any cross-sectoral trends, such as structural transformation, that could have occurred during the sample period.

Specification (1) in Table 4 documents that, on average, domestic prices are positively correlated with the import price index. In this most parsimonious specification, the only independent variable is the change in the imported intermediate goods price index, *IIPI*. The coefficient is estimated at 0.21, which if interpreted causally would imply that a 1% increase in the imported intermediate goods price index is associated with a 0.21% increase in domestic producer prices.

Specification (2) documents that the imported intermediate goods price index is more correlated with domestic prices in sectors and countries that more intensively use imported intermediate inputs. This specification adds the interaction of the change in the sectoral import price index with the sector-importer-specific cost share. The estimated specification is of the form:

$$\hat{p}^{D}_{s,t,t-1} = k^{D}_{s} + \alpha I\widehat{IP}I_{s,t,t-1} + \beta \theta^{I}_{s,t} I\widehat{IP}I_{s,t,t-1} + \gamma \theta^{I}_{s,t} + \hat{\epsilon}_{s,t} \ . (9)$$

The interaction coefficient, β, is estimated to be 0.34, while the main effect, α , is 0.13. As we include an interaction variable, $\theta_{s,t}^{I}$ must also be included directly. We note that because of the inclusion of fixed effects, γ captures the impact of changes in imported inputs-intensity on prices rather than the level.

Panel A	(1)	(2)	(3)		(4)
	Dependent variable is monthly % change in sectoral PPI				
Δ Importer IIPI	0.211*** (0.0113)	0.128*** (0.0181)	–0.160*** (0.0577)		
Δ Trade-weighted exchange rate (sector specific weights)					–0.0462*** (0.0164)
Cost change from IIPI = Δ IIPI * cost share		0.342*** (0.0585)	0.673*** (0.209)		
Cost change from exchange rate = Δ exr * cost share					0.198*** (0.0619)
Imported input cost share		0.0127 (0.0103)	0.0215** (0.0107)		0.0218** (0.0106)
Observations	5,996	5,996	5,996		5,996
R-squared	0.055	0.061			0.002
Number of panelvar	40	40	40		40
Panel B **Variables**			(3.1) Δ IIPI	(3.2) Δ IIPI * cost share	
Δ Trade-weighted exchange rate (sector specific weights)			0.279*** (0.0146)	–0.00273 (0.00443)	
Cost change from exchange rate = Δ exr * cost share			0.0208 (0.0586)	0.301*** (0.0178)	
Imported input cost share			0.0282*** (0.0105)	0.00723** (0.00319)	
Observations			5,996	5,996	
R-squared			0.114	0.084	
Number of panelvar			48	48	

Notes: Panel A: dependent variable is the change in the importer's PPI. Panel B: first-stage regressions: dependent variables are the change in the importer's IIPI and the cost change from IIPI, respectively. Δ is a first-difference operator. Dummy variables' estimates not shown. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Estimation results, spillovers to producer prices from imported inputs.

To gauge the magnitude of the uncovered effect in Specification (2), consider two sectors with input intensities of 0.04 and 0.49, respectively. Such intensities correspond to the 5th and 95th percentiles of the imported intermediate input cost share in our sample. For the sector that uses fewer imported intermediate inputs, a 1% increase in *IIPI* is associated with a 0.14% increase in domestic producer prices (0.14% = 0.13% + 0.04*0.34%). In contrast, the same 1% increase in *IIPI* is associated with a 0.30% increase in domestic producer prices when half the sector's costs stem from imported intermediate inputs (0.30% = 0.13%

+ 0.49*0.34%). The spillover rate into domestic prices is thus well over twice as large in the sector that uses imports more intensively.

Specification (3) uses exchange rate movements as drivers of the *IIPI*, thereby addressing the problem that import prices and domestic prices might co-react to common cost shocks. This specification presents the results of a two-stage least squares estimation relating first the exchange rate to changes in the *IIPI* and then the projected changes in the *IIPI* on changes in domestic prices. Because we want to instrument for both the change in the *IIPI* as well as for the interaction of the *IIPI* with the imported intermediate cost share, we instrument with the exchange rate, as well as with the exchange rate interacted by the imported intermediate cost share. The exchange rate is specified as units of domestic currency per unit of foreign currency, so that an increase denotes a depreciation of the importer's currency.

Panel B presents the two first-stage estimations for the two endogenous variables, the change in the *IIPI* and the latter interacted with the imported intermediate cost share. In Specification (3.1), the dependent variable is the change in the *IIPI*. The independent variables include the change in the trade-weighted exchange rate (*exr*), the imported intermediate cost share and the interaction of the two. As expected, the change in the *IIPI* is strongly correlated with the exchange rate, while the interaction of the exchange rate with the imported intermediate cost share has no explanatory power. Indeed, there are arguably no *a priori* reasons to believe that the sensitivity of the import price index itself to the exchange rate depends on the cost share of imports.

In Specification (3.2), the dependent variable is the change in the *IIPI* interacted with the imported intermediate cost share. The independent variables are the same as in (3.1). In this first-stage estimation, the interaction of the exchange rate with the imported input cost share is strongly correlated with the dependent variable, while the exchange rate itself has no explanatory power.[19]

The results from the second stage estimation are presented in Panel A. These reveal a negative coefficient for the main effect (–0.16) and a positive interaction coefficient of 0.67, both with high statistical significance. Repeating the previous back-of-the-envelope example for the two sectors at 5^{th} and 95^{th} percentiles of imported input intensity; a 1% increase in the *IIPI* is now associated with a –0.13% decrease in domestic prices in the low import-intensive sector (–0.13% = –0.16% + 0.04*0.67%). And, a 1% increase in the *IIPI* is associated with a 0.17% increase in domestic prices in the sector that intensively uses imported inputs (0.17% = –0.16% + 0.49*0.67%).

How should one interpret the sizeable differences between Specifications (3) and (2)? They likely reflect the fact that import prices and producer prices react to the same, possibly global, shocks. We thus continue to instrument for changes in the *IIPI* with the exchange rate in the remainder of the analysis. But instead of following the approach in (3), we present reduced-form estimations as in (4), where the independent variables are the exchange rate, the imported intermediate cost share, and the exchange rate interacted with the cost share. These estimates capture the implicit notion that the exchange rate drives changes in import

[19] The fact that each of the two endogenous variables is correlated with a different instrument implies that the second-stage estimation is well identified: the Cragg-Donald Wald F-statistic is 249, far exceeding the 10% critical value of 7.03 (10% is the lowest critical value calculated for this specification by Stock and Yogo (2005)).

prices, while these import prices affect domestic prices, especially so in sectors that intensively use imported inputs.[20]

Importantly, using the reduced-form approach allows us to extend the sample. In some countries, sector-specific import price indices are not published, while sector-specific producer prices are available (China, India). For the other five economies (Australia, Chinese Taipei, Indonesia, Japan and Korea), data on the PPI are available for a longer time period than the IIPI.

Our baseline model, using the reduced form approach, is shown as specification (5) of Table 5. This is basically a re-estimation of (4) in the previous table, but including all country-sector combinations for which a PPI is available. This increases the sample from 40 to 89 country-sector combinations and from 5996 to 11990 monthly observations. In our baseline estimation, the exchange rate itself is not statistically significant. However, the interaction of the exchange rate with the imported input cost share is highly statistically significant and has an economically sizeable coefficient of 0.16.

An obvious question is how robust the results are to the inclusion of economy- or time-specific trends. As will be shown next, an important role is played by the intensity with which the different sectors use imported inputs. This variation is mostly cross-sectional (ie cross-industry and cross-country), so our findings are not driven by aggregate patterns.

Specification (6) adds the importer's CPI inflation rate as a control variable, to capture the possibility that economy-wide demand pressures could correlate with the increase in producer prices. Specifications (7) to (9) add different sets of time fixed effects to the estimation. In particular, in (7), we add time (month) fixed effects to capture general trends. In (8), we include time-country-fixed effects in order to absorb all variation in the data that is common within a country. Finally, Specification (9) adds industry-time fixed effects that absorb all variation across sectors.

The results in Table 5 show that our results are not driven by aggregate common trends across the economies (7), neither by country-specific demand shocks or inflation patterns (6) nor any other country-specific patterns (8). Finally, the results are not driven by sector-specific fluctuations (9).

[20] The relationship between the coefficients in (3) and (4) is the following. In the second stage estimation (3.1), a 1% exchange rate depreciation implies a 0.279% increase in the IIPI. In the first stage estimation, this 0.279% higher IIPI is associated with a –0.160*0.279% = –0.0446% lower PPI. Also the effect of a changing exchange rate in the other first-stage estimation has to be taken into account (resulting in an impact of the exchange rate on the domestic PPI of –0.00273%*0.673 = –0.0018%). In the reduced-form estimation, these two effects are estimated directly and added together (up to a rounding error: –0.0462 = –0.0446–0.0018).

	(5)	(6)	(7)	(8)	(9)	(10)
	Dependent variable is monthly % change in sectoral PPI					
Trade-weighted exchange rate change (sector specific weights)	–0.0138 (0.0114)	–0.0221* (0.0113)	–0.00402 (0.0114)	–0.00869 (0.0117)	–0.00403 (0.0114)	
Cost change from exchange rate = Δ exr * cost share	0.157*** (0.0460)	0.165*** (0.0457)	0.131*** (0.0453)	0.169*** (0.0449)	0.137*** (0.0456)	
Change in importing country's seasonally adjusted CPI		0.683*** (0.0531)				
6-month change in importer exchange rate						0.0857*** (0.0277)
6-month cost change from exchange rate = Δ exr * cost share						0.626*** (0.119)
Additional time fixed effects	-	-	T-FE	T-FE by country	T-FE by industry	T-FE
Observations	11,990	11,990	11,990	11,990	11,990	11,750
R-squared	0.002	0.015	0.064	0.118	0.074	0.083
Number of panelvar	89	89	89	89	89	89

Notes: dependent variable is the change in the importer's PPI. Specification (7) includes time fixed effects; (8) includes time-country fixed effects; (9) includes industry-time fixed effects. Dummy variables' estimates not shown. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 5: Estimation results, baseline reduced form model and additional time fixed effects.

We note that all the models discussed so far relate monthly changes in import prices to contemporary changes in domestic producer prices. In the presence of nominal price stickiness of any sort, it is likely that changes in the costs of imported intermediate goods affect the prices that domestic producers charge only with a lag. In (10), we add five monthly lags of both the exchange rate change and the interaction of the exchange rate change with the imported intermediate goods cost share. We consider a regression of the form:

$$\hat{p}^D_{s,t,t-1} = k^D_s + \sum_{z=0}^{5} \alpha_z \widehat{exr}^{TW}_{s,t-z,t-(z+1)} + \sum_{z=0}^{5} \beta_z \theta^I_{s,t-z} \widehat{exr}^{TW}_{s,t-z,t-(z+1)} + \sum_{z=0}^{5} \gamma_z \, \theta^I_{s,t-z} + \hat{\epsilon}_{s,t} \quad (10)$$

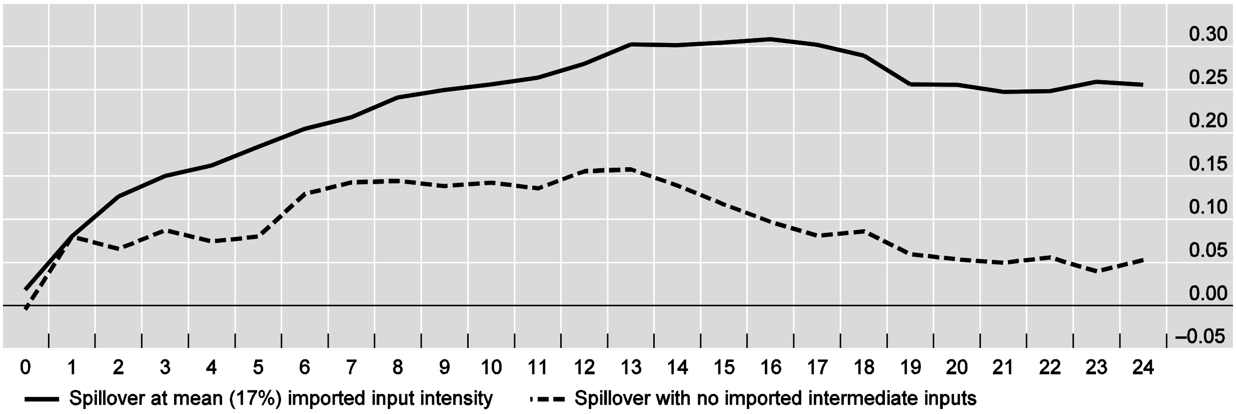
Specification (10) in Table 5 reports the sum of the 6 coefficients $\sum_{z=0}^{5} \alpha_z$, $\sum_{z=0}^{5} \beta_z$, and the corresponding standard errors.

Over a 6-month horizon, imported input use can explain over one half of the correlation between import and producer prices for the mean sector in our data. The interaction coefficient is 0.63, while the main effect is estimated to be statistically significant and positive at 0.09. We note that in our sample, the mean sector has a cost share of 0.17. For this sector, the estimated rate of spillovers is equal to 0.09 + 0.17*0.63 = 0.20, of which 0.11 = 0.17*0.63, or over one half, can be attributed to the imported input cost channel.

The spillover to producer prices, derived from the long-run specification (10), can also be displayed graphically (Graph 5). We estimate the model for horizons from 0 to 24 months. The dashed line represents the estimated cumulative main effect, ie it is equal to $\sum_{z=0}^{k} \alpha_z$, with k taking values from 0 to 24. The solid line is equal to the estimated effect for the median sector in our dataset, ie $\sum_{z=0}^{k} \alpha_z + 0.17 * \sum_{z=0}^{k} \beta_z$.

Spillover into producer prices and imported input intensity

7 economies pooled | Graph 5



Note: The graph displays the increase in producer prices, resulting from a 1% depreciation in the trade-weighted exchange rate of the importing sector.

Source: Authors' calculations

Graph 5 shows that a 1% depreciation of the trade-weighted exchange rate leads to a 0.25% increase in producer prices for the mean importing sector, *ceteris paribus*, when nine months have passed from the exchange rate movement. After that, the impact stabilises, fluctuating between 0.25% and 0.30% at longer horizons.[21] In contrast, for a sector that uses no imported inputs, the impact on producer prices reaches 0.15% after eight months, followed by stabilisation and a gradual decline to only 0.05%. Again, these results emphasise the importance of the relative intensity with which the different sectors use imported inputs.

3.4. Asymmetries and nonlinearities

The specifications so far have assumed linear and symmetric spillovers from exchange rate fluctuations to producer prices. Next, we investigate whether the impacts on producer prices vary for small and large exchange rate movements, and for positive and negative exchange rate fluctuations, respectively.

We first evaluate whether increases in the costs of intermediate imported inputs have stronger impacts on producer prices than do decreasing costs. A reason for this is the inherent asymmetry in the profit function, whereby firms are more averse to goods being underpriced than overpriced. The maximum loss from underpricing could be unbounded, whereas the maximum loss from overpricing is limited by zero profits (see Devereux and Siu, 2007). As a result, it is likely that increases in costs of intermediate inputs are passed through more fully to prices than decreasing costs.

Specifications (1) and (2) in Table 6 split the sample into appreciation episodes of the foreign exchange rate (resulting in decreasing costs of imported inputs) and depreciations

[21] These are only the point estimates. Taking into account estimation uncertainty, the impact on producer prices for the mean sector is between 0.15 and 0.36 after two years, using 99% confidence intervals.

(increasing import costs), respectively. This sample split documents that domestic prices do react strongly to increasing foreign input prices as a result of exchange rate depreciation, whereas no significant relation is found when foreign currencies become less expensive. The finding that producer prices only react to increasing costs of import prices emerges also at a 6-month horizon. Here, we split the sample into periods in which the 6-month cumulative change was below (3) or above (4) the median. Again, we find that the 6-month cumulative effect of the cost change resulting from exchange rate movements is only significant for above median cost changes, reflecting exchange rate depreciation (positive cost changes).

Is the response of domestic prices more pronounced when the magnitude of the exchange rate movement is large? In the presence of menu costs, whereby changing the price is costly, firms may allow their mark-ups to absorb the effects of small changes in costs, keeping final prices constant. However, large changes in costs, possibly through sizeable shocks to the exchange rate, are more likely to lead to changes in final prices.

Specification (5) includes only those exchange rate movements that fall between the 25th and 75th percentiles, ie it includes only small fluctuations. In this specification, the interaction of exchange rates and the imported input cost share is not statistically significant. In contrast, in the sample of large movements (Specification 6; exchange rate movement either below 25th or above 75th percentile), the latter interaction is statistically significant at the 5% level. Also, when evaluated at the 6-month horizon, we find that there is no effect of the exchange rate interacted with the imported input cost share when the preceding 6-month exchange rate change falls between the 25th and 75th percentiles (7). In contrast, there is a strong effect on prices in the sample where the preceding 6-month exchange rate change is either below the 25th or above the 75th percentile (8).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	(1) – (4) Asymmetric PPI Response?				(5) – (8) Nonlinear PPI Response?			
	Negative exr change	Positive exr change	Below median 6m exr change	Above median 6m exr change	Exr change between 25^{th} and 75^{th} percentile	Exr change below 25^{th} or above 75^{th} percentile	6m exr change between 25^{th} and 75^{th} percentile	6m exr change below 25^{th} or above 75^{th} percentile
	Dependent variable is monthly % change in sectoral PPI							
Trade-weighted exchange rate change (sector specific weights)	–0.0111 (0.0254)	–0.0129 (0.0251)			0.0784 (0.0607)	–0.00557 (0.0134)		
Cost change from exchange rate = Δ exr * cost share	–0.134 (0.0873)	0.372*** (0.0799)			–0.117 (0.253)	0.125** (0.0527)		
6-month change of importer exchange rate			0.306*** (0.0686)	–0.131** (0.0562)			0.490*** (0.150)	0.116*** (0.0319)
6-month cost change from exchange rate = Δ exr * cost share			–0.473 (0.290)	1.708*** (0.186)			–1.113** (0.595)	0.592*** (0.136)
Time fixed effects	T-FE	T-FE	T-FE	T-FE	T-FE	T-FE	T-FE	T-FE
Observations	5,915	6,075	5,875	5,875	5,996	5,994	5,876	5,874
R-squared	0.077	0.098	0.104	0.133	0.073	0.089	0.084	0.115
Number of panelvar	89	89	89	89	89	89	89	89

Notes: dependent variable is the change in the importer's PPI. Specification (1) includes appreciation periods of the exchange rate; (2) includes depreciation periods of the exchange rate; (3) includes periods when the 6-month cumulative change was below median; (4) includes periods when the 6-month cumulative change was above median; (5) uses a sample with exchange rate movements between 25th and 75th percentiles; (6) uses a sample with exchange rate movements below 25th and above 75th percentiles; (7) uses a sample with exchange rate movements between 25th and 75th percentiles; (8) uses a sample with exchange rate movements below 25th and above 75th percentiles. Dummy variables' estimates not shown. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Asymmetries and nonlinearities in the impact on producer prices

3.5. Some policy implications

The previous results show significant spillovers, especially over time, from costs of imported inputs to domestic producer prices in economies participating in the Asian supply chain. However, little so far has been said about the implications of the findings for monetary policy. Should monetary policymakers care about the behaviour of sectoral producer prices? After all, as the changes in sectoral prices represent relative price shifts, their effect on aggregate inflation may be muted.

As noted by Ball and Mankiw (1995), an important factor is the distribution of relative price changes. The authors show that in the model with menu costs, firms will change their prices in response to large shocks but not in response to small ones. Then, the size of shocks and the distribution of relative price changes matters for aggregate inflation dynamics. When there is a large negative shock and the distribution of relative prices is skewed to the left, prices fall; in the opposite case they rise. Empirical evidence supporting the model's predictions is provided eg by Sekine (2009) and Auer and Fischer (2010).[22] To the extent that increasing real integration will add to the sensitivity of Asian economies to large relative price shocks that affect the distribution of relative price changes, they are of relevance for policymakers.

Recall that in Section 2 we found a positive correlation between trade intensity and price co-movement at the level of aggregate inflation, both in terms of producer and consumer prices. The sectoral analysis can then be seen as identifying one possible mechanism by which increased trade intensity leads to a closer co-movement of prices. It confirms that the results are not driven by common trends, but by the intensity with which different economies and sectors use imported intermediate inputs. Taken together, the results show that the economies and sectors more closely connected by trade experience greater price spillovers, even at the level of aggregate inflation.

4. Conclusion

When economies are becoming more globalised, disinflationary pressures from lower-cost imports are likely to contribute to lower domestic inflation, but this effect should be temporary. Once the economies are already closely interconnected (globalised), they are exposed to cross-border demand and supply shocks and the volatility of the domestic inflation rate could increase. We examine one such case of close interconnectedness, ie that of economies participating in the manufacturing supply chains in the Asia-Pacific region. In the sectoral analysis, we use the novel World Input-Output Database.

We have documented that inflation rates measured both in terms of consumer and producer prices share a common stochastic trend in the recent decades as opposed to the 1980s, which we interpret as a sign of increased co-movement of inflation in the region. Moreover,

[22] Sekine (2009) shows that global shocks to two relative prices – wage costs and import prices – account for an important share of disinflation in the OECD countries. Similarly, Auer and Fischer (2010) examine the distribution of sectoral price changes in the United States that result from comparative advantage-induced supply shocks in the emerging economies. They show that the distribution of the shocks has been highly left-skewed, implying that the distribution of price changes would have been more right-skewed if the shocks had not taken place. Using the result by Ball and Mankiw (1995), this implies that the supply shocks resulting from globalisation lowered aggregate inflation in the United States.

we find that both headline inflation rates and producer prices move more closely together between those Asian economies that trade more with one another, ie that share a higher degree of trade intensity.

In the sectoral analysis, we have shown that the share of imported intermediate inputs in a sector's total costs is roughly 17% on average, for the seven Asia-Pacific economies. When prices of intermediate imports change by 1%, domestic producer prices change cumulatively by close to 0.3% during two years.[23] An important role is played by the intensity with which different sectors use imported intermediate inputs. This variation is mostly cross-sectional (cross-sector and cross-economy) and therefore our results are not driven by common trends. This pinpoints the mechanism by which increased trade intensity leads to greater co-movement of prices.

In sum, real integration through the supply chain matters for domestic price dynamics in the Asia-Pacific region. They are also a concern for policymakers. Increased interconnectedness is likely to lead to greater sensitivity of aggregate inflation rates to costs of imported inputs, especially when the cost changes are large in magnitude.

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Appendix

Start date of estimation sample in Table 2, model with broad money stock:

CN	1996:Q1
IN	1991:Q2
ID	1981:Q1
KR	1980:Q1
MY	1985:Q4
PH	1981:Q1
TH	1996:Q1
JP	1980:Q1
AU	1980:Q1
NZ	1989:Q1

For the country-pair of Philippines and Indonesia, data for 1996 are missing.