

Import Price Misalignment after the Crisis: A New Keynesian Perspective

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Abstract

Current account imbalances particularly between East Asian and the U.S. have persisted into the second decade of the 21st century. A contemporary debate focuses on the role of exchange rate misalignment in explaining these imbalances. This paper focuses instead on the role of *import price* misalignment defined as the deviation of import prices from exchange-rate-adjusted marginal cost of production. We find evidence that the markups over marginal cost of U.S. import prices are historically low in recent years. The failure of import prices in reflecting U.S. dollar depreciation can prevent the expenditure switching necessary to clear imbalances. We estimate a structural New Keynesian model of local currency pricing to show that rates of import price adjustment are low, particularly for imports from Asian economies. We find industry-level evidence suggesting that this low passthrough may be explained by concentration of Asian economies on final goods which have low passthrough.

JEL Classification Numbers: E31; F31; F41

Keywords: exchange rate passthrough; import prices; local currency pricing; price stickiness; new Keynesian Phillips curve; forward-looking expectation

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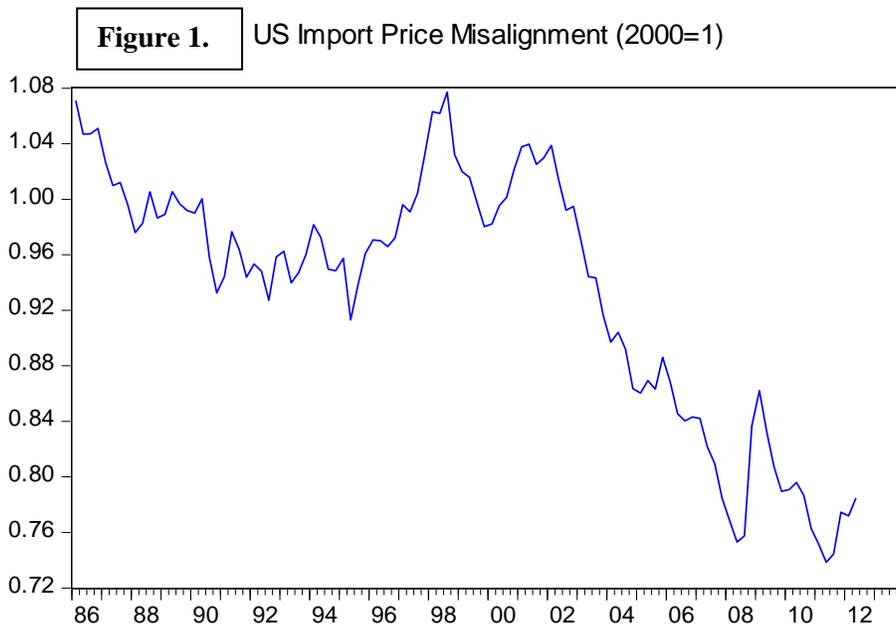
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I. Introduction

Over the last decade, imbalances in global trade have raised continuing questions about international stability (see IMF World Economic Outlook, April 2011), particularly in terms of the size of the United States trade deficit. One explanation of this is a persistent overvaluation of U.S. dollar in light of assessing exchange rates and current balance (Lee et al., 2008; and IMF Research Department, 2013). This paper will also help uncover an important source of trade imbalance.

Our focus in this paper will be on the pricing of imported goods. We can illustrate the importance of the large misalignment in U.S. import prices relative to the production costs of its trading partners. Figure 1 shows the U.S. Bureau of Labor Statistics Import Price Index for All Goods except Petroleum (source: BLS) deflated by a trade weighted index of exchange-rate-adjusted producer costs amongst 51 large trading partners (the methodology of construction of the index is described in Choi and Cook, 2013).

The import price index series begins in 1985 and stayed consistently within the 8% of the 2000 baseline level through 2003. Following that date, however, we see import prices fall persistently through the latest decade, except for rebounds after the global financial crisis.



Despite the persistent depreciation of the U.S. dollar relative to its trading partners, U.S. import prices have been very slow to adjust so that real prices of imports dropped by 25% relative to the beginning of the decade. This observation suggests that the failure of U.S. trade rebalance has not been primarily attributable to insufficient exchange rate adjustments by its trading partners but could rather stem from the pricing mechanism of exporters to the U.S.

Over recent periods, the most severe import price misalignment has been concentrated in the Pacific Rim region. We report an import price index by locality of origin (discounted by a trade-weighted, exchange-rate-adjusted index of producer prices from the region of origin). The set of countries includes China, Japan, Australia, Indonesia, Malaysia, New Zealand, and the Philippines. Dictated by data availability, the data start from the end of 2003.

Figure 2a shows that, relative to the beginning of the series, the price of imports into the U.S. from all regions has been falling relative to domestic marginal cost, especially prior to the global financial crisis of 2008. During the crisis, exchange-rate-adjusted domestic costs increased as international currencies appreciated. However, such

Figure 2a US Import Price Misalignment by Global Region of Origin (2004= 1)

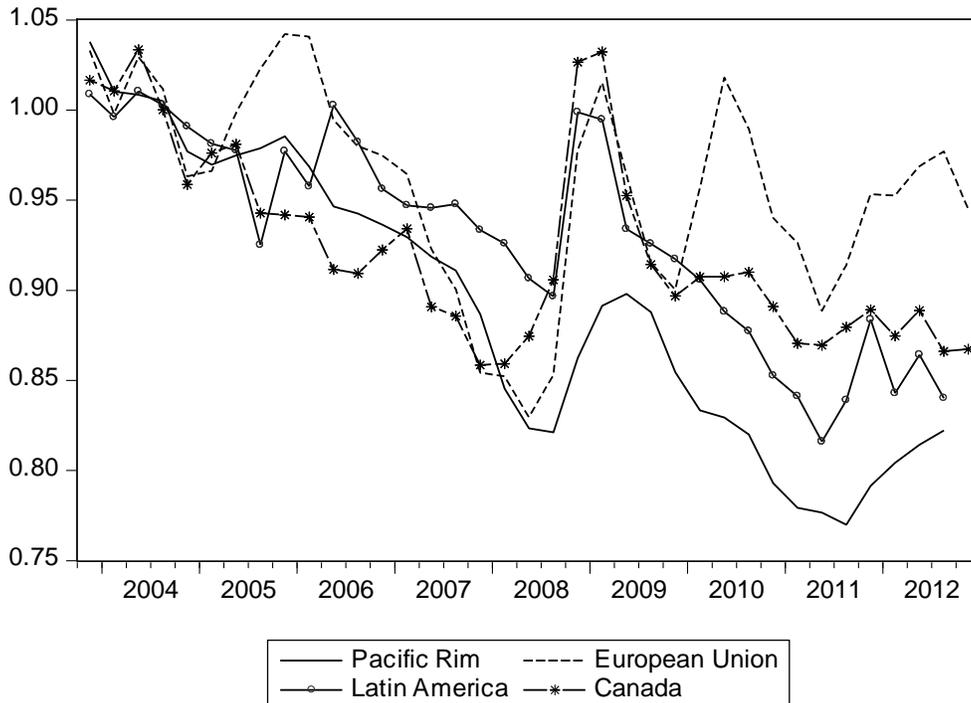
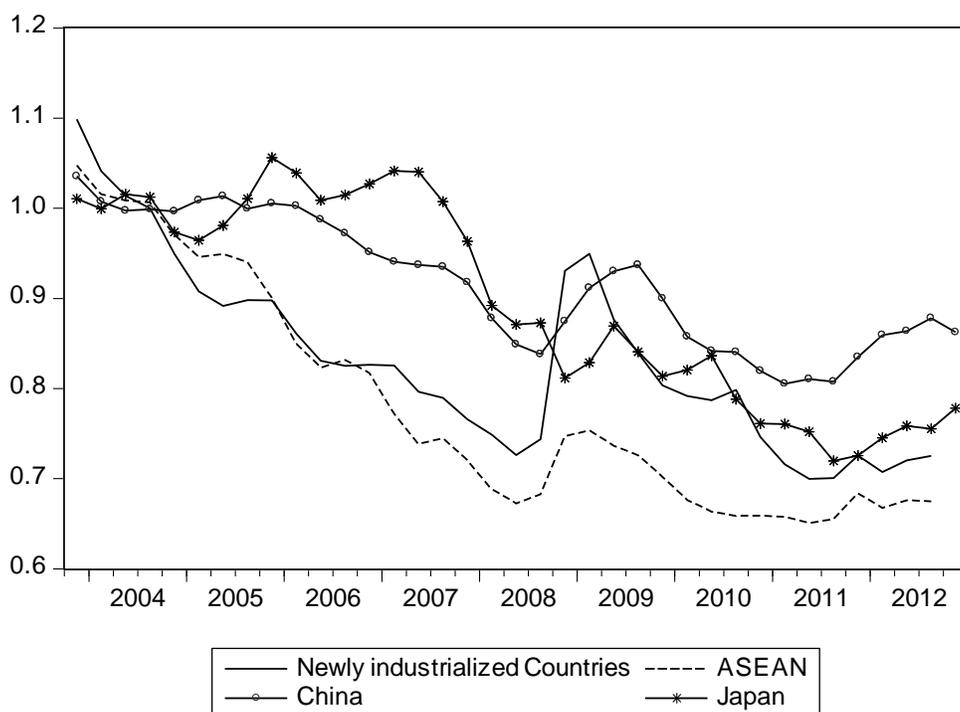


Figure 2b: Import Price Misalignment: Asian Subregions (2004 = 1)



cost increases were mildest in the Pacific Rim countries. The appreciation of emerging market currencies in the aftermath of the crisis pushed markups down. Near the end of 2011, import markups over marginal cost were nearly 25% below where they had been at the beginning of the sample, although 2012 saw some reversals.

The misalignment of some East Asian sub-regions is depicted by Figure 2b. The ASEAN countries experienced a consistent decline in import markups. By the end of the period, the import price markups for ASEAN countries are 30 per cent below their 2004 level. Qualitatively similar results are found for the Asian Newly Industrialized Countries (Hong Kong SAR, Singapore, and Taiwan), Japan and Korea. Perhaps interestingly, the mildest drop in markups occurs among Chinese exporters.

Import prices may tend to have slow exchange-rate passthrough (see Campa and Goldberg, 2005). We examine whether the U.S. import price misalignment observed the last decade, particularly amongst economies in the Asian region can be explained by this import price stickiness. To identify whether import price stickiness itself can explain import price misalignment, we use GMM techniques to estimate a New Keynesian style sticky price model in parallel to the New Keynesian Phillips curve models developed by

Galí and Gertler (1999). We can use these econometric models to identify structural parameters governing price stickiness at the aggregate, regional or industry level. This model controls for both marginal cost and forward-looking expectations.

These methods are most appropriate for this project because we are interested in distinguishing import price stickiness from potential accusations of monetary policy manipulation. As noted by Gagnon and Ihrig (2004), non-structural models of exchange rate passthrough conflate the responses of import prices and the responses of monetary policy to real exchange rates (see the Lucas critique). An optimization based model can focus specifically on pricing behavior. Conversely, likelihood based structural models such as Smets and Wouters (2002) require assumptions about monetary policy. We estimate an industry-level measure of import price stickiness that abstracts from the monetary (and other) policy responses which preclude reduced-form estimates. Our results would also shed light on the role of exchange rate adjustments in U.S. trade rebalancing after the crisis.

A large number of papers argue that the slow pace of exchange rate passthrough can be explained by strategic competition between foreign export firms, which reduces incentives to change prices quickly in response to shocks. Gust et al. (2010) develop a model in which greater trade integration increases strategic interaction. Marazzi and Sheets (2006), using data through 2004, provide evidence that industries faced with increasing entry of Chinese firms experienced declines in exchange rate passthrough.²

II. Pricing Model

We consider a model in which some firms exporting to the U.S. adjust their U.S. dollar prices infrequently and optimally according to the local currency pricing (LCP) theory (Betts and Devereux, 2000). A producing firm chooses producer prices in the home market to maximize the discounted sum of expected profits.³

² Existing studies suggest that low exchange rate passthrough is associated with low average inflation levels (Taylor, 2000; and Devereux and Yetman, 2010). Also, declining exchange rate passthrough could be attributed to changes in data measurement and model specifications (Hellerstein, Daly, and Marsh, 2006).

³ Choi and Cook (2013) develop passthrough equation based on the modification of the local currency pricing model of Betts and Devereux (2000) to allow for strategic interaction as in Kimball (1995).

Discounted expected profits in the domestically oriented sector can be written as:

$$\max_{ppi} E_t \left[\sum_{j=t}^{\infty} (\kappa\beta)^j PPI_j^{\frac{1}{1-\zeta}} Q_j^{\frac{\zeta}{\zeta-1}} \left\{ ppi_t^{\frac{\zeta}{\zeta-1}} - ppi_t^{\frac{1}{\zeta-1}} MC_j \right\} \right],$$

where β is the discount factor of the firm's managers, $(1-\kappa)$ is the probability of adjusting the producer price in each period, ppi is the firm's producer price, PPI is the aggregate producer price index, Q is output, and MC is nominal marginal cost in terms of home currency. The optimal producer price ppi_t^* that maximizes the expected profits is given by

$$ppi_t^* = \frac{1}{\zeta} \frac{E_t \left[\sum_{j=t}^{\infty} (\kappa\beta)^j \cdot PPI_j^{\frac{1}{1-\zeta}} \cdot Q_j^{\frac{\zeta}{\zeta-1}} \cdot MC_j \right]}{E_t \left[\sum_{j=t}^{\infty} (\kappa\beta)^j \cdot PPI_j^{\frac{1}{1-\zeta}} \cdot Q_j^{\frac{\zeta}{\zeta-1}} \right]}. \quad (1)$$

The dynamics of prices are given by:

$$PPI_t = \kappa \cdot PPI_{t-1} + (1-\kappa) \cdot ppi_t^*. \quad (2)$$

Linearizing the dynamics as in Galí and Gertler (1999) and Sbordone (2002),

$$\pi_t^{PPI} = E_t \left[\frac{(1-\kappa)(1-\beta\kappa)}{\kappa} mc_t + \beta\pi_{t+1}^{PPI} \right], \quad (3)$$

where $mc_t = \log(MC_t / PPI_t)$, and π_t^{PPI} is the inflation rate in foreign prices.

Similarly, when importing firms choose their prices for imports into the U.S. under LCP, they maximize their profits from the U.S. market:

$$\max_{ipi} E_t \left[\sum_{j=t}^{\infty} (\nu\beta)^j IPI_j^{\frac{1}{1-\zeta}} IM_j^{\frac{\zeta}{\zeta-1}} S_j \left\{ ipi_t^{\frac{\zeta}{\zeta-1}} - ipi_t^{\frac{1}{\zeta-1}} MC_j / S_j \right\} \right],$$

where ipi is the firm's import price measured in U.S. dollars, IPI is the aggregate import price index, IM is the volume of imports, S_t is the spot exchange rate with the U.S. dollar, and $(1-\nu)$ is the probability of adjusting the import price in each period. The optimal import price that maximizes the expected profits of an LCP firm, ipi_t^{LCP} , is given by

$$ipi_t^{LCP} = \frac{1}{\xi} \frac{E_t \left[\sum_{j=t}^{\infty} (\nu\beta)^j IPI_j^{\frac{1}{1-\xi}} \cdot IM_j^{\frac{\xi}{1-\xi}} \cdot S_j \cdot (MC_j / S_j) \right]}{E_t \left[\sum_{j=t}^{\infty} (\nu\beta)^j IPI_j^{\frac{1}{1-\xi}} \cdot IM_j^{\frac{\xi}{1-\xi}} \cdot S_j \right]}. \quad (4)$$

The price dynamics of LCP firms are:

$$IPI_t^{LCP} = \nu \cdot IPI_{t-1}^{LCP} + (1-\nu) \cdot ipi_t^{LCP}. \quad (5)$$

Define the law of one price (LOOP) gap, M_t , as the ratio of the U.S. dollar price of imports to the producer price of foreign goods converted into U.S. dollar:

$M_t \equiv IPI_t / (PPI_t / S_t)$. Under the LOOP, M_t equals one. Under local currency pricing, the LOOP gap will dissipate over time for LCP firms but only through the process of price adjustment.

In parallel with equation (3),

$$\pi_t^{LCP} = E_t \left[\frac{(1-\nu)(1-\beta\nu)}{\nu} (mc_t - \mu_t) + \beta \cdot \pi_{t+1}^{LCP} \right], \quad (6)$$

where μ_t is the logarithm of M_t and π_t^{LCP} is the log first difference of IPI_t^{LCP} . The New Keynesian Phillips curve implies that firms have an incentive to raise prices when real marginal costs are high or expected to be high, inducing inflation. There exists an exact parallel to the New Keynesian Phillips curve in the LCP literature. Import price inflation will be high when the real U.S. dollar marginal cost of foreign production is high or expected to be high. The U.S. dollar foreign marginal cost is broken into two parts: the real marginal cost and the LOOP gap.

We assume that some fraction, λ , of firms are PCP firms and price their goods in their home currency. PCP firms are randomly distributed and adjust their home prices as other firms do. While PCP firms have sticky prices in their own currency, their invoice prices in U.S. dollars adjust automatically as exchange rates change. We assume that PCP firms pass changes in producer prices and exchange rates into changes in import good prices with a one-period lag: $\pi_t^{PCP} = \pi_{t-1}^{PPI} - ds_{t-1}$.

Total import price inflation is the weighted average of inflation based on PCP and LCP:

$$\pi_t^{IPI} = \lambda \cdot \pi_t^{PCP} + (1-\lambda) \cdot \pi_t^{LCP}. \quad (7)$$

Combining equations (6) and (7), we have

$$\pi_t^{IPI} - \lambda \cdot \pi_t^{PCP} = E_t \left[(1-\lambda) \frac{(1-\nu)(1-\beta\nu)}{\nu} (mc_t - \mu_t) + \beta \cdot (\pi_{t+1}^{IPI} - \lambda \cdot \pi_{t+1}^{PCP}) \right], \quad (8)$$

Or:

$$E_{t-1} (\pi_t^{IPI} - \beta \pi_{t+1}^{IPI}) = \lambda \cdot E_t (\pi_t^{PCP} - \beta \cdot \pi_{t+1}^{PCP}) + (1-\lambda) E_t \left[\frac{(1-\nu)(1-\beta\nu)}{\nu} (mc_t - \mu_t) \right]. \quad (9)$$

In theory, equation (9) offers a moment condition whose parameters could be estimated with GMM. However, real marginal costs, mc_t , of all foreign trading partners of the U.S. may be difficult to obtain. Fortunately, the New Keynesian Phillips curve model of domestic price stickiness allows us to infer the level of foreign marginal costs from data on foreign prices. Combining equations (3) and (9) to eliminate mc_t , we have:

$$\begin{aligned} E_t (\pi_t^{IPI} - \beta \pi_{t+1}^{IPI}) &= \lambda \cdot E_t (\pi_t^{PCP} - \beta \cdot \pi_{t+1}^{PCP}) \\ &+ (1-\lambda) E_t \left[\frac{(1-\nu)(1-\beta\nu)\kappa}{\nu(1-\kappa)(1-\beta\kappa)} (\pi_t^{PPI} - \beta \cdot \pi_{t+1}^{PPI}) - \frac{(1-\nu)(1-\beta\nu)}{\nu} \mu_t \right]. \end{aligned} \quad (10)$$

III. Regional Level Import Price Models

A. Data

The Bureau of Labor Statistics constructs a set of regional level import price indices for a number of relevant regions. Estimating a regression model for the import price dynamics for a number of regions requires some additional series. We construct a set of producer price indices for 51 major trading partners of the U.S. running from the early 1980s through as far into 2012 as data are available. The indices concentrate on quarterly producer price indices for domestic manufactured goods though in some cases we use broader producer price indices as proxies. We also get quarterly data on average quarterly spot foreign exchange rates from IMF International Financial Statistics. We combine these into trade-weighted regional indices for producer price inflation, PCP pricing and misalignment.

Each region will be associated with K countries that are located in the region in our sample of 51. For each of $k = 1, \dots, K$ countries we get quarterly imports from country k into the United States as quarterly imports from IMF Direction of Trade Statistics. We measure the fraction of imports which are manufactured goods using annual SITC 2 level data from OECD International Trade by Commodity Statistics database. We measure manufactured imports from each period as the level of imports multiplied by the fraction which are manufactured. We define $imports_t^k$ as the Hodrick-Prescott filtered level of this product. For each region R , with K countries we define the weight as

$w[R]_t^k = imports_t^k / \sum_{k=1}^{K[R]} imports_t^k$. Using lagged weights to minimize endogeneity, we

aggregate import price inflation over countries as $\pi_t^{PPI} = \sum_{k=1}^{K[R]} [w[R]_{t-8}^k \cdot \pi_t^{PPI,k}]$ where

$\pi_t^{PPI,k} \equiv \Delta \ln(PPI_t^k)$, $ds_t = \sum_{k=1}^{K[R]} [w[R]_{t-8}^k \cdot \Delta \ln(S_t^k)]$, S_t^k is an index of currency units in

country k needed to buy 1 U.S. dollar in spot markets (source: IMF IFS), and PPI_t^k is the producer price index of domestic manufactured goods for country k .

For country k within region R , we construct a country specific measure of misalignment, $M[R]_t^k \equiv \frac{S_t^k \cdot IPI_t^R}{PPI_t^k} \frac{1}{PPP_{BY}^k}$ where PPP for the base year, PPP_{BY}^k , is a trade-weighted average of the relative prices in 2005 of representative manufactured goods, clothing and capital goods, from the World Bank International Comparison Project data for base year $BY = 2005$. The indices, IPI_t^R , S_t^k , and PPI_t^k are rebased relative to the base year. To calculate regional misalignment of U.S. import prices, we aggregate as in Thomas and Marquez (2006), $M_t^R \equiv \prod_{k=1}^{K[R]} (M_t^k[R])^{w[R]_{t-8}^k}$. We then represent the logarithm of the LOOP gap as: $\mu_t = \ln M_t^R$.

B. Empirical Model

We assume for the sake of simplicity that $\beta = 1$. The equation for passthrough into import prices can be simplified for estimation using GMM as the following empirical specification:

$$\left[\pi_t^{IPI} - \pi_{t+1}^{IPI} \right] = \alpha_0 + \alpha_1 \cdot \mu_t + \alpha_2 \cdot \left[\pi_t^{PPI} - \pi_{t+1}^{PPI} \right] + \alpha_3 \cdot \left[\pi_t^{PCP} - \pi_{t+1}^{PCP} \right] + \varepsilon_{t+1} \quad (11)$$

Theory suggests signs for the variables, $\alpha_1 < 0; \alpha_2, \alpha_3 > 0$ and a mapping between $\{\alpha_1, \alpha_2, \alpha_3\}$ and structural parameters $\{\nu, \kappa, \lambda\}$. To consistently estimate regression (11), we use instrumental variables for $\left[\pi_t^{PPI} - \pi_{t+1}^{PPI} \right]$ and μ_t , because the former is not known at time $t-1$ and because of concerns about measurement error. We construct a period by period inflation forecast for each country k , $\hat{\pi}_{t+1}^{PPI,k}$. The forecast is the weighted average of out-of-sample, one-step-ahead forecasts from rolling regressions. The forecasting variables vary from country to country depending on availability of forecast power. We use $\pi_t^{PPI} - \hat{\pi}_{t+1}^{PPI}$, the difference between current weighted PPI inflation and a weighted forecast of future inflation, as an instrument for $\left[\pi_t^{PPI} - \pi_{t+1}^{PPI} \right]$,

where $\hat{\pi}_t^{PPI} = \sum_{k=1}^{K[R]} \left[w[R]_{t-8}^k \cdot \pi_t^{PPI,k} \right]$. Since all of the elements, including the parameters of

the forecasting regression, needed to construct $\hat{\pi}_{t+1}^{PPI.k}$ are known at time t , this is a valid instrument. We use lagged misalignment, μ_{t-1} , as an instrument for contemporary misalignment, μ_t .

C. Results

i. New Keynesian Exchange Rate Passthrough: Asian Region

Table 1 reports just identified regression estimates for a number of Asian regions. Depending on data availability, the sample length ranges between 20 years and slightly less than 10 years (i.e., since end-2003). The $R = Pacific\ Rim$ economies which include major Asian economies plus Australia and New Zealand. We report the estimates in Column A along with Newey-West standard errors. We find the estimated parameters are all of the sign indicated by theory. Coefficients α_2 and α_3 are statistically significant at the 1 percent level. The adjusted R^2 indicates this model explains a large share of the variance in the dynamics of import price inflation. Given the small sample size, we should be concerned about weak instruments. We report the first-stage Cragg-Donald test for weak instruments from the regression. This statistic, near 15, exceeds the Stock-Yogo 5% critical value of the hypothesis that the bias in the just identified GMM/2SLS estimates is larger than 10%.

We also report implied estimates of the structural parameters along with Newey-West standard errors. The structural parameters indicate that Pacific Rim imports display extremely slow passthrough. The parameter estimate of ν could be interpreted within the Calvo pricing framework as indicating an average time between price changes of nearly 5 years. A variety of previous studies find evidence of slow import price passthrough; one interpretation is deviation from the CES demand function which generates strategic reasons for low passthrough. However, the estimate of passthrough seems to be particularly small. We also estimate relatively sticky prices. The parameter estimate for κ indicates domestic price change occurring every 10 quarters, lower than typical estimates for the EU in tradition of Galí and Gertler (1999). We also estimate a fraction

of producer currency pricing firms of about 6%, which is consistent with BLS estimates that 5–10% of imports are invoiced in foreign currencies.

We also report estimates for some sub-regions including $R =$ Newly industrialized countries, ASEAN countries, China, and Japan. Data on the NICs and Japan date from the 4th quarter of 1992. In each of these cases, our estimates of the structural parameters have the correct sign with the single exception of α_2 , estimated to be negative for Japan. In all cases, the estimate of α_1 is negative, but small and not statistically significantly different from zero. The other parameters are significant at the 1% critical value with the exception of α_2 for China which is positive but insignificant. The R^2 is large, ranging between roughly 15% in China and Japan and approximately 45% for the NICs and ASEAN. In each case, the Cragg-Donald statistic exceeds the Stock-Yogo critical value, indicating that weak instruments aren't a great concern for the IV regressions.

We find estimates of very sharp import price stickiness in Asian countries. The implied estimates of ν range between 0.920 and 0.975, consistent with price changes, on average, of from every 12.5 quarters to every 40 quarters. Perhaps interestingly, we estimate the adjustment of import prices to be fastest for Chinese firms and slowest for the NICs. No measure of domestic price stickiness can be measured for Japan. China is estimated to have price stickiness similar to that observed for the U.S. by the general literature. Other regions have stickier prices. Finally, except for Japan, most of the regions have estimates of the share of producer currency pricing that is on the high side of likely levels.

ii. Pooled Regression

The available sample period for many regions is short. We estimate regression (11) using pooled data from seven regions including the NICs, ASEAN, China, Japan, Canada, the European Union, and Latin America. For each of the non-Asian economies, we report the estimates from the pooled regression in Table 2. Import price data from the NICs, Japan, Canada, and the EU begin at the end of 1990; those from Latin America begin in 1997; while those from ASEAN and China begin only in 2003. Therefore, the

pool is unbalanced. We allow all of the regions to have individual estimates of α_0 to control for fixed-region effects.

The joint estimates of $\{\alpha_1, \alpha_2, \alpha_3\}$ are all of the correct sign and are statistically significant at the 1% significant level. The J -test of the over-identification conditions (i.e., that all regions have the same coefficient estimates, $\{\alpha_1, \alpha_2, \alpha_3\}$) does not reject the model at even the 10 percent significance level. Estimates of the structural parameters of the Calvo model indicate a high level of price stickiness. The implied level of $\nu = .932$ is consistent with prices changing on average once every 14 quarters. The associated domestic price stickiness, $\kappa = 0.877$, in U.S. trading partners would be associated with price adjustment every 8 quarters.

We can also estimate a pooled specification that allows for different parameters for Asian regions and for non-Asian regions. We report the parameter estimates in columns B and C. Again we find all the parameters have the correct sign and are significant at the 1% level. The J -test again fails to reject the over-identification conditions at the 5% level. We test the joint hypothesis that parameters $\{\alpha_1, \alpha_2, \alpha_3\}$ in column B are equal to those in column C. A Wald test rejects the hypothesis that the parameters are the same at the 1% significance level. In addition, we are able to reject each of the individual hypotheses that any of the three parameters are equivalent across groups.

The estimated degree of import and domestic price stickiness is much larger in the four Asian regions than in the non-Asian regions. Import prices from Asia would be estimated to adjust on average roughly every 20–25 quarters ($\nu^{Asia} = .956$) while import prices from non-Asia adjust once every 8–9 quarters ($\nu^{NonAsia} = .883$). We can reject the hypothesis that the structural parameter $\nu^{Asia} = \nu^{NonAsia}$ at the 1% significance level. Domestic price stickiness is also higher in Asia than in the non-Asian regions. Domestic prices in Asia adjust, on average, once every 10–11 quarters ($\kappa^{Asia} = .905$) while domestic prices in non-Asian counterparts adjust roughly every 6 quarters ($\kappa^{NonAsia} = .837$). Again we can reject the hypothesis $\kappa^{Asia} = \kappa^{NonAsia}$ at the 1% significance level. The fraction of producer currency price setting is between 5% and 10% in both

groups of countries, though it is slightly larger amongst Asian countries (the hypothesis that $\lambda^{Asia} = \lambda^{NonAsia}$ can be rejected at the 5% significance level).

IV. Industry Level Data

Our finding in the above section is that Asian economies have generally had stickier import prices than other regions of the world. We would argue this explains the extreme import price misalignment from this region of the world. In the next section we explore whether this is a matter of country-characteristics specific to the Asian region. An alternative hypothesis is that the mix of goods exported to the U.S. from Asia happens to include goods with relatively slow pass-through. We explore this hypothesis by examining passthrough at the individual level.

A. Industry-Level Misalignment

The BLS constructs price indices for HS 2-digit-industry-level imports which we designate IPI_t^j for each of $j=1, \dots, J (= 36)$ manufacturing industry. For every country $k = 1, \dots, 51$, we construct an industry-trading-partner-specific measure of misalignment,

$$M_t^{j,k} \equiv \frac{S_t^k \cdot IPI_t^j}{PPI_t^k} \frac{1}{PPP_{BY}^k}. \text{ Both } S_t^j \text{ and } PPI_t^j \text{ are rebased relative to the base year.}$$

We aggregate these into industry specific using industry-country trade weights. We get annual industry-level import values by country from HS 1988 code from OECD International Trade by Commodity Statistics database. Annual data are converted to quarterly and Hodrick-Prescott filtered. The country-industry trade weight is

$$w_t^{j,k} = \text{imports}_t^{j,k} / \sum_{k=1}^{51} \text{imports}_t^k. \text{ To calculate misalignment of import prices for industry } j, \text{ we aggregate, } M_t^j \equiv \prod_{k=1}^{K=51} (M_t^{j,k})^{w_t^{j,k}}.$$

Looking across these 36 industries, we examine the percentage increase in import price misalignment. We measure markup declines since 2002 by examining for each industry j , $\ln(M_{2012}^j / M_{2002}^j)$, where M_{year}^j is the average quarterly level for *year*. The average level of $\ln(M_{2012}^j / M_{2002}^j)$, is -0.124 with a standard deviation of 0.34 , indicating

that depreciation is occurring on average across sectors with a great deal of variance. Gopinath and Rigobon (2008) find that final goods have less frequent price changes than intermediate materials. Bricongne et al. (2010) classify different sectors into categories of demand with a broad set of industries concentrated in upstream stages of the production process categorized at intermediate materials. We break down our goods into two basic categories *Intermediate*—exactly following Bricongne et al. (2010)—and *Final* goods being all others. The goods fall evenly into the two categories. Here we see an immediate difference. For *Final* goods, the average level of import markup decline, $\ln(M_{2012}^j / M_{2002}^j) = -.231$, (indicating a 23% decline in markups) with a relatively smaller standard deviation of 0.179; while in the upstream *Intermediate* materials category, the average decline in markups is close to zero, $\ln(M_{2012}^j / M_{2002}^j) = -.018$, with a much larger spread (standard deviation of 0.428).

This outcome might be explainable with a closer look at some of the industries that had large increases in their markups, $\ln(M_{2012}^j / M_{2002}^j) > .29$. All of these are clearly resource intensive industries HS71: Pearls, precious stones, metals, etc; HS72: Iron and steel; HS28: Inorganic chemicals, precious metal; HS74: Copper and articles thereof and HS27: Mineral fuels, oils, etc. Given the commodity heavy nature of these industries, it may not be surprising that U.S. import prices have risen (relative to general producer prices among exports) along with the relative prices of these commodities. However, not all intermediate materials have appreciated. Some intermediate materials industries particularly organic materials related to lumber seem to have seen large declines in $\ln(M_{2012}^j / M_{2002}^j)$ thus the variance might be high.

To assess the significance of various factors determining the degree of depreciation, we examine a cross-sectional regression of markup changes, $\ln(M_{2012}^j / M_{2002}^j)$, on some various factor observed in Table 3. Regressing $\ln(M_{2012}^j / M_{2002}^j)$ on a dummy variable for intermediate industries, the coefficient reported in column A is positive and significant at the 10% level (p -value = .06). The difference as suggested above between types of industries is in line with a 20% decline in relative markups for final goods. Column B reports the results of a regression of

$\ln(M_{2012}^j / M_{2002}^j)$ on the average share of imports accounted for by the ASEAN+3 countries (including greater China) over the years 2000 and 2001. The coefficient is negative but statistically insignificant at the 10% level. However, the coefficient *is* economically significant. The coefficient is near -.33. Given the gap between the least ASEAN + 3 dominated industry and the most ASEAN+3 dominated industry is almost 90%, this variable could be associated with a substantial decline of markups. However, when we regress the variable on both ASEAN+3 share and the Intermediate Materials dummy, we find the coefficient on the former is substantively much closer to zero and statistically insignificant, while the coefficient on the latter continues to be economically and statistically significant. We also run a similar regression using China share instead of the ASEAN+3 share. The results are similar except that we find a positive coefficient on China share.

One interpretation is that, although Asian dominated industries tended to have a bigger drop in markups, this was due more to the concentration of Asian exporters (to the U.S.) being concentrated in non-intermediate materials areas. ASEAN+3 exporters had an average share across industry types of 34%, but on average fewer than 20% of intermediate materials sectors and close to an average of 50% of other types of sectors including final goods.

B. Industry-Level Import Price Stickiness

For each of $j = 1, \dots, 36$ industries, we also use country-industry-level trade weights and country-level aggregate data to construct industry-level measures of foreign producer price inflation and exchange rate depreciation: $\pi_t^{PPI,j} = \sum_{k=1}^K [w_{t-8}^{k,j} \cdot \pi_t^{PPI,k}]$ and $ds_t^j = \sum_{k=1}^K [w_{t-8}^{k,j} \cdot \Delta \ln(S_t^k)]$. Setting $\mu_t^j = \ln M_t^j$, we estimate industry-level price stickiness using the New Keynesian import equation under the assumption for the sake of simplicity that $\beta = 1$. This equation can be estimated using GMM with the following specification:

$$\left[\pi_t^{IPI,j} - \pi_{t+1}^{IPI,j} \right] = \alpha_0 + \alpha_1^j \cdot \mu_t^j + \alpha_2^j \cdot \left[\pi_t^{PPI,j} - \pi_{t+1}^{PPI,j} \right] + \alpha_3^j \cdot \left[\pi_t^{PCP,j} - \pi_{t+1}^{PCP,j} \right] + \varepsilon_{t+1} \quad (12)$$

We use μ_{t-1}^j , and $\pi_t^{PPI,j} - \hat{\pi}_{t+1}^{PPI,j}$ as instruments with $\pi_t^{PPI,j} = \sum_{k=1}^K [w_{t-8}^{k,j} \cdot \pi_t^{PPI,k}]$.

We can use an estimate of $\alpha_1^j > 0$ to calculate an industry-level estimate of New Keynesian price stickiness ν^j . However, we find that in 12 of 36 industries our estimate of $\alpha_1^j < 0$, as opposed to the expected sign from passthrough theory. Nine of these twelve negative results are in commodity-intensive industries such as lumber and basic metals related production or fuels. One conjecture is that these commodity-driven sectors lack the degree of product differentiation based market power necessary for sticky price setting behavior. Hence we will concentrate further analysis on the 24 industries with $\alpha_1^j > 0$. Our estimates of ν^j range from 0.697 (HS28: Inorganic Chemicals) to 0.988 (HS84 Machinery) with a median level of .947 indicating a typically high degree of price stickiness.

We report some potential determinants of the degree of price stickiness. We regress the cross-section $\ln \nu^j$ on some industry level determinants and show robust standard errors in Table 4. Given such a small sample size, such results might be merely indicative at this stage. The first finding reported in Table 4, column A is that the intercept coefficient from a regression of $\ln \nu^j$ on a dummy variable indicating an intermediate materials industry is negative and statistically significant. Relative to the mean level of .95 (roughly consistent with a price change every 20 quarters), the typical coefficient for an intermediate materials industry might be closer to .88, (roughly consistent with an average time between price changes occurring every 8 quarters). Column B adds a variable measuring the share of the industry imports coming from the ASEAN plus 3 countries in the years 2000–2001, while column C adds a variable representing China’s share of exports. We find (in contrast to Bergin and Feenstra, 2009) that the degree of price stickiness at the industry level does not depend on the share from Asia once very simple industry characteristics are taken into account.

We consider some other industry-level characteristics of exporters. We examine whether the share of the industry that comes from a single country can affect import price

stickiness. Column D reports a regression including the share from the leading partner in the years 2000–2001. The coefficient is statistically insignificant at the 5% level.

Considering that low inflation could be associated with low passthrough (Devereux and Yetman, 2010), we regress $\ln \nu^j$ on industry class and average producer price inflation in the weighted average trading partner, $\pi_t^{PPI,j}$, in annualized percentage terms. As reported in column E, we find the coefficient is negative, indicating that high inflation countries do have less sticky prices. However, the coefficient is not significant at the 10% level. Among the industries included, the span from minimum to maximum of $\pi_t^{PPI,j}$, is 0.8% on the low side and 5.8% on the high side, so the coefficient would be associated with a moderately substantial difference in price stickiness over that span. Column F reports a similar regression substituting the average depreciation for the average producer price inflation.

C. Sticky Prices and Import Price Misalignment

In Table 3, column E, we also consider whether sticky prices are associated with the degree of misalignment. We regress $\ln(M_{2012}^j / M_{2002}^j)$ on $\ln \nu^j$ and the average annual percentage depreciation rate of the currency over the period 2000–2011, $\overline{ds}_{2000-11}^j$. What we find is that the industries that had the greatest degree of price stickiness, (i.e., high levels of $\ln \nu^j$) also had the most severe misalignment. The coefficient on $\ln \nu^j$ is significant at the 1% level. Conversely, the coefficient on exchange rate depreciation is neither economically nor statistically significant. This indicates that the degree of misalignment is better explained by the degree of price-stickiness than the actual size of the change in exchange rates. The R^2 is remarkably high.

V. Conclusion

We consider two possibilities for U.S. import price misalignments. The first possibility is that the historic level of misalignment is simply a function of the well-known slow passthrough of exchange rates into import prices (Campa and Goldberg, 2005; and Gust et al., 2010). The second is that the country-level behavior of trading partners has changed. We propose that the industries with low have not adjusted quickly

to the run up in currency values of U.S. trading partners. This rather than monetary policy per se may explain the degree of misalignment.

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Table 1. Regression Results for Passthrough into U.S. Import Prices

This table reports parameter estimates from just identified GMM estimates of equation (11). Reported parameters include linear coefficients α_1 - α_3 ; and theoretical coefficients ν and κ . Heteroskedasticity and auto-correlation consistent standard errors are reported in parenthesis. We report the Cragg-Donald weak-instruments-test statistic along with Stock-Yogo 5% critical values. The results are reported (by column) for an import price index for manufactured goods from (A) Newly Industrialized Countries; (B) ASEAN countries; (C) China; (D) Japan; and (E) Pacific Rim countries. Symbols *, **, and *** indicate significance at the 10%, 5%, 1% levels, respectively.

	Pacific Rim (A)	NICs (B)	ASEAN (C)	China (D)	Japan (E)
α_1	-0.003 (.005)	-0.001 (.001)	-0.002 (.004)	-0.007 (.008)	-0.005 (.004)
α_2	0.258*** (.059)	0.096*** (.038)	0.517*** (.076)	0.085 (.097)	-0.049 (.057)
α_3	0.061*** (.022)	0.138*** (.042)	0.120*** (.027)	0.092*** (.03)	0.057*** (.011)
Adj. R^2	0.601	0.450	0.457	0.173	0.158
Num. Obs.	34	86	34	35	87
Cragg-Donald	15.48	17.88	11.98	9.576	42.79
		Critical	Value	7.03	
ν	.944*** (.036)	.974*** (.031)	.952*** (.046)	.920*** (.042)	.934*** (.021)
κ	.895*** (.065)	.922*** (.011)	.934*** (.061)	.770*** (.132)	---

Table 2. Pooled Regression Results

This table reports parameter estimates from just identified GMM estimates of equation (11). Parameters reported include the GMM estimates of linear coefficients α_1 - α_3 ; and identifiable parameters λ , ν , and κ . Heteroskedasticity and auto-correlation consistent standard errors are reported in parenthesis. The results are reported (by column) for (A) a pool of import prices from seven regions with each assumed to have a common values of parameters; and the same pool with separate parameters for (B) a set of four Asian regions including ASEAN countries, Newly Industrialized economies, China, and Japan; and (C) a set of three non-Asian regions including the European Union, Latin America, and Canada. Symbols *, **, and *** indicate significance at the 10%, 5%, 1% levels, respectively.

	Joint Estimation (A)	Asian Regions (B)	Non Asian Regions (C)
α_1	0.005*** (.001)	0.002*** (.001)	0.016*** (.003)
α_2	0.272*** (.019)	0.197*** (.017)	0.465*** (.047)
α_3, λ	0.083*** (.004)	0.093*** (.006)	0.072*** (.008)
<i>N·T</i>	474	474	
<i>J Stat</i>	18.547	15.993	
<i>d.f.</i>	18	15	
<i>5% C.V.</i>	28.869	24.996	
ν	0.932*** (0.005)	0.956*** (0.008)	0.883*** (0.008)
κ	0.877*** (0.009)	0.905*** (0.017)	0.883*** (0.008)

Table 3. Industry-Level Misalignment: $\ln(M_{2012}^j / M_{2002}^j)$

This table reports estimates from cross sectional regression of the percentage change in import price markups $\ln(M_{2012}^j / M_{2002}^j)$ for 36 two-digit industries on a set industry specific variables. The regressions include one with a dummy variable for the 18 industries classified as intermediate materials, Column (A); one with the fraction of industry imports in 2000-1 from the Greater ASEAN +3 region, Column (B); one with both an intermediate materials dummy and an ASEAN+3 share, Column (C); one with both an intermediate materials dummy and the share of industry imports from China in 2000–2001 (D). The fifth includes a regression on the parametric estimate of price stickiness and the average depreciation rate of the trading partner over the period 2000–2011. All standard errors are heteroskedasticity robust. Symbols *, **, and *** indicate significance at the 10%, 5%, 1% levels, respectively.

	Independent Variable $\ln(M_{2012}^j / M_{2002}^j)$				
	(A)	(B)	(C)	(D)	(E)
Constant	-0.231*** (.042)	-0.010 (.132)	-0.196 (.148)	-0.252*** (.082)	-0.423*** (.049)
Intermediate Dum	0.213** (.109)		0.192** (.111)	0.228*** (.112)	
Asean+3 Share 2000		-0.331 (.242)	-0.069 (.25)		
China Share 2000				0.087 (.19)	
$\ln(v^j)$					-2.331*** (.711)
Average Appreciation 2000–2011					-0.019 (.026)
N	36	36	36	36	24
R^2	.100	.055	.102	.102	.633

Table 4. Industry-Level Import Price Stickiness: $\ln(v^j)$

The table reports estimates from cross sectional regression of the log of the New Keynesian parameter of import price stickiness on some determinants. The regressions include one with a dummy variable for the industries classified as intermediate materials, Column (A); one with both an intermediate materials dummy and with the fraction of industry imports in 2000–2001 from the Greater ASEAN +3 region, Column (B); one with both an intermediate materials dummy and the share of industry imports from China in 2000–2001 Column (C); one with both an intermediate materials dummy and the share of industry imports from the single leading import source in 2000–2001 (D); and one with both an intermediate materials dummy and a trade weighted average of annual inflation in the source country over the period (D). The fifth includes a regression on the parametric estimate of price stickiness and the average depreciation rate of the trading partner over the period 2000–2011. All standard errors are heteroskedasticity robust. Symbols *, **, and *** indicate significance at the 10%, 5%, 1% levels, respectively.

	Independent Variable $\ln(v^j)$				
	(A)	(B)	(C)	(D)	(E)
Constant	-0.051*** (.007)	-0.053** (.026)	-0.047** (.011)	-0.023 (.027)	-0.031 (.043)
Intermediate Dum	-0.072** (.035)	-0.071** (.029)	-0.074 (.035)	-0.078 (.035)	-0.063 (.031)
Asean+3 Share 2000		0.004 (.050)			
China Share 2000			-0.012 (.030)		
Dominant Producer Share 2000				-0.079 (.074)	
Average Inflation 2000–2011					-0.009 (.019)
N	24	24	24	24	24
R^2	.208	.208	.209	.228	.221