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New Keynesian Exchange Rate Pass-Through

Woon Gyu Choi and David Cook

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Prepared by Woon Gyu Choi and David Cook[1]

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Abstract

Using the theory of optimal local currency pricing, this paper constructs a structural equation to estimate the rate at which foreign producer prices pass through the local currency prices of imported goods in the U.S. This can be viewed as measuring exchange rate pass-through, in line with price stickiness in the New Keynesian Phillips curve literature. We estimate the structural equation using the generalized methods of moments for consistent estimates of exchange rate pass-through. We find that a model with a mix of local currency pricing and producer currency pricing fits the data best. The estimate of price stickiness in import prices is comparable to existing estimates of domestic price stickiness.

JEL Classification Numbers: E31; F31; F41

Keywords: pass-through effects; import prices; local currency pricing; new Keynesian Phillips curve; forward-looking expectation

Authors' E-Mail Address: wchoi@imf.org; davcook@ust.hk

[1] Woon Gyu Choi is a Senior Economist at the IMF Institute of the International Monetary Fund. David Cook is an Associate Professor of Economics at the Hong Kong University of Science and Technology.

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I. INTRODUCTION

The degree to which fluctuations in the nominal exchange rate affect the relative prices of real goods will in large part determine the effect of exchange rate policy on patterns of international trade. A large empirical literature in international macroeconomics estimates how quickly exchange rate changes translate into import prices (for example, Yang 1997; and Campa and Goldberg, 2005). This speed of adjustment is referred to as exchange rate pass-through. In this paper, we examine the speed of adjustment in the price of exports to the U.S. in response to changes in the value of the U.S. dollar. An innovation in our paper is an application of the modern sticky price theory that is embedded into the New Open Economy Macroeconomics literature (see Obstfeld and Rogoff, 1995) to create a statistical model in which exchange rate pass-through can be represented as a structural parameter immune to the Lucas critique. We then apply some of the techniques that have been developed to understand the dynamics of domestic inflation (referred to as the New Keynesian Phillips curve literature) to estimate our model and consistently identify the structural parameter which we refer to as exchange rate pass-through.

The idea that the prices of imports would incorporate exchange rate changes only slowly has a long heritage. Many theoretical models were developed to explain this particular failure of the Law of One Price. In particular, price stickiness implies that the Law of One Price fails to hold at any time, and exchange rate pass-through is imperfect.[2] In the context of sticky price models, Betts and Devereux (1996, 2000) develop the theory of local currency pricing (LCP) to explain violations of purchasing power parity. In contrast, when monopolistic producers set sticky prices in their own currency, which is the case of producer currency pricing (PCP), exchange rate changes are reflected one-for-one in foreign currency prices (perfect pass-through).

To generate richer dynamics, many sticky price models have adopted a (time-dependent) model of pricing developed by Calvo (1983) and incorporated it into the dynamic general equilibrium modeling structure (Yun, 1996). In this type of models, inflation is forward-looking and can be written as a discounted sum of future marginal costs. When real marginal costs are high or expected to be high in the future, firms have an incentive to raise prices inducing inflation. The recursive form of an inflation equation in which inflation is written as a function of current marginal cost plus expected future inflation (as a proxy for future marginal costs) has been represented as a modern alternative to the Phillips curve. Galí and Gertler (1999) and Sbordone (2002, 2005) show that this forward-looking representation of inflation can be estimated with general methods of moments (GMM) using lag variables as valid instruments. This literature can be used to estimate the rate at which sticky prices change.

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. We can decompose the real U.S. dollar marginal cost of foreign production into two parts, the real marginal cost of production in the exporting country and the

[2] In the earliest versions of the New Open Economy Macroeconomics literature, sticky prices generated a role for monetary policy. A substantial body of theory examines the welfare implications of imperfect pass-through. In some cases, LCP implies that maintaining a fixed exchange rate can be an optimal monetary policy (for example, see Devereux and Engel, 2003).

gap between import prices in the U.S. and the PCP price (that is, the exchange-rate-adjusted price of goods in the foreign country). Finding a consistent measure of marginal costs in the foreign country may be difficult. However, the theory of the New Keynesian Phillips curve implies that the measure of real marginal costs can be inferred from inflation since real marginal costs drive inflation. Therefore, the estimation of a recursive, forward-looking representation of the LCP Euler equation governing import price inflation enables us to provide a consistent estimate of the frequency with which foreign firms change their U.S. prices in response to changes in U.S. dollar marginal costs. We can interpret this structural parameter as the New Keynesian rate of exchange rate pass-through because one of the determinants of these U.S. dollar marginal costs is the exchange rate. As a byproduct of this procedure, we are able to estimate the rate at which foreign producers adjust their own home price levels without ever constructing a measure of marginal cost.

This paper offers several findings on exchange rate pass-through effects on U.S. import prices. First, we find strong evidence of price stickiness in U.S. import prices; consistent with a structural model in which prices change every 5 quarters on average. Second, a model in which some firms adopt LCP and others adopt PCP pricing fits the data most coherently. The estimated fraction of PCP price setters is small: 10 percent or less. Third, we find only weak evidence that rates of price change have slowed over time. Fourth, the degree of import price stickiness is similar to that of the domestic price stickiness of U.S. trading partners. Lastly, as in the New Keynesian Phillips curve, there is a substantial forward-looking element to the adjustment of import prices which is consistent with the idea that many firms are adjusting in an optimal manner. Also, like that literature, only a smaller fraction of firms are backward-looking.

II. The Model

We consider a model in which all firms exporting to the U.S. adjust their U.S. dollar prices infrequently and optimally according to the local currency pricing (LCP) pass-through theory. A producing firm that changes prices in the home market chooses a price to maximize the discounted sum of expected profits over the time. The firm must set producer prices in advance using the information set available at time *t*-1 (see Rotemberg and Woodford, 1997). Discounted expected profits in the domestically oriented sector can be written as:

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

where β is the discount factor of the firm's managers, κ 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

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

Linearizing the first order conditions as in Galí and Gertler (1999),

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

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., they maximize their profits from the U.S. market:

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

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 ν is the probability of adjusting the import price in each period. The optimal import price ipi_t^* that maximizes the expected profits is given by

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

We can extend the Galí and Gertler (1999) and Sbordone (2002) technique to the case of import prices and write

$$\pi_t^{IPI} = E_{t-1}\left[\frac{(1-\nu)(1-\beta\nu)}{\nu}\left(mc_t - \mu_t\right) + \beta \cdot \pi_{t+1}^{IPI}\right], \qquad (2)$$

where μ_t is the logarithm of $M_t \equiv \dfrac{IPI_t}{PPI_t / S_t}$, the mark-up of U.S. imports over their (U.S. dollar) domestic prices (that is, $\mu = \ln M$).

Combining equations (1) and (2) to eliminate mc_t, we have

$$\pi_t^{IPI} = \frac{(1-\nu)(1-\beta\nu)}{\nu} E_{t-1}\left[-\mu_t + \frac{\kappa}{(1-\kappa)(1-\beta\kappa)}\pi_t^{PPI} - \frac{\beta\kappa}{(1-\kappa)(1-\beta\kappa)}\pi_{t+1}^{PPI}\right] + \beta E_{t-1}\left(\pi_{t+1}^{IPI}\right). \quad (3)$$

Equation (3) will form the statistical model that we can use to measure structural pass-through. This equation reflects an error-correction mechanism, where μ (= ln M) is a cointegrating vector for U.S. import prices, the foreign producer price level, and the exchange rate because the ratio M converges in expectation to a steady state when firms are fully able to adjust prices.

The theory constrains the dynamics of equation (3) in a number of ways. First, in this optimal price setting equation, all corrections come through changes in foreign or import prices, and the exchange rate only appears as part of the error-correction term. This is in contrast with the standard literature which essentially regresses import price inflation (or CPI inflation) on changes in the exchange rate. Second, the prices adjust geometrically so we can include only a single lead of each kind of inflation in the dynamic equation. The reduced-form exchange rate pass-through literature typically focuses on backward-looking dynamics and does not use theory to constrain the number of lags in the model. Third, the price setting firms operate under rational expectations, so the error term is uncorrelated with all variables in the information set available for price-setters at time t-1. This produces a natural set of identification conditions and allows consistent estimation with GMM. Fourth, the pass-through of real foreign marginal cost into import prices occurs at the same speed as exchange rate pass-through as implied by equation (2). Finally, the long-term pass-through is 100 percent.

III. THE DATA

To estimate the pass-through effect model given by equation (3), we need data on domestic import price inflation, foreign producer price inflation, and the relative price of imports to foreign price levels. Detailed descriptions about data are provided in the appendix.

We also need trade weights to aggregate foreign prices over U.S. trading partners. We calculate trade weights for 40 trading partners of the U.S. (Appendix C).[3] Country j's share in U.S. imports of manufactured goods at quarter t, w_t^j, is calculated as imports of manufactured goods from country j divided by the sum of imports of manufactured goods from all countries in the list of trading partners.

We measure the markup of import prices over the PCP price, M_t, using geometric weights to aggregate time series of the markup for each of the trading partners, as in Thomas and Marquez (2006). The markup for country j is defined as $m_t^j \equiv \frac{S_t^j \cdot IPI_t}{PPI_t^j}\frac{1}{PPP_{1995}}$, where S_t^j is the number of currency units in country j needed to buy 1 U.S. dollar in spot markets relative to the exchange rate in 1995 (Source: IFS), and PPI_t^j is the producer price index of manufactured

[3] U.S. trading partners comprises 40 countries as follows: Australia, Austria, Belgium, Brazil, Canada, Chile, China, Hong Kong SAR, Colombia, Costa Rica, Denmark, Finland, France, Germany, India, Indonesia, Ireland, Israel, Italy, Japan, Korea, Malaysia, Mexico, Netherlands, New Zealand, Norway, Peru, Philippines, Poland, Portugal, Russia, Singapore, South Africa, Spain, Sweden, Switzerland, Taiwan, Thailand, Turkey, and the United Kingdom.

goods for country *j* relative to the index in the base year 1995.[4] The import price index for the U.S. is from a database of OECD manufactured goods import price indices that spans the period 1975–2002. The data is extended forward in time using an index of non-petroleum import prices from BLS.[5] The variable PPP_{1995}^{j} is the purchasing power parity ratio from the Penn World Tables (Heston, Summers, and Aten, 2006) in the base year, 1995.

We create an aggregate measure of the relative prices of goods produced in the U.S. relative to its trading partners using geometric trade weights, $\widehat{w}_{t-k}^{j}$:

$$M_t = \prod_{j=1}^{40} \left(m_t^j \right)^{\widehat{w}_{t-k}^{j}}$$

where $\widehat{w}_{t-k}^{j}$ are averaged over *n* quarters and lagged by *k* quarters to eliminate endogeneity issues.

The markup does not contain any discernable secular drifts although it exhibits somewhat persistent swings over the period 1980:Q1–2005:Q4, as shown in Figure 1. The markup is measured by the logarithm of the trade-weighted index of the relative prices: that is, $\mu_t = \ln(M_t)$. An augmented Dickey-Fuller test with an intercept and four lags rejects the hypothesis of a unit root in the series at the 5 percent level. The standard deviation of this series is about 4.1 percent. Wide swings in this figure roughly correspond to fluctuations in the dollar. When the dollar is strong in the early 1980s and near the turn of the millennium, this markup is also high. This indicates that import prices are not as low as might be expected given the strong value of the dollar. In the mid-1980s and in recent periods, the markup has been low, indicating that import prices have not risen to the same degree that the dollar has fallen.

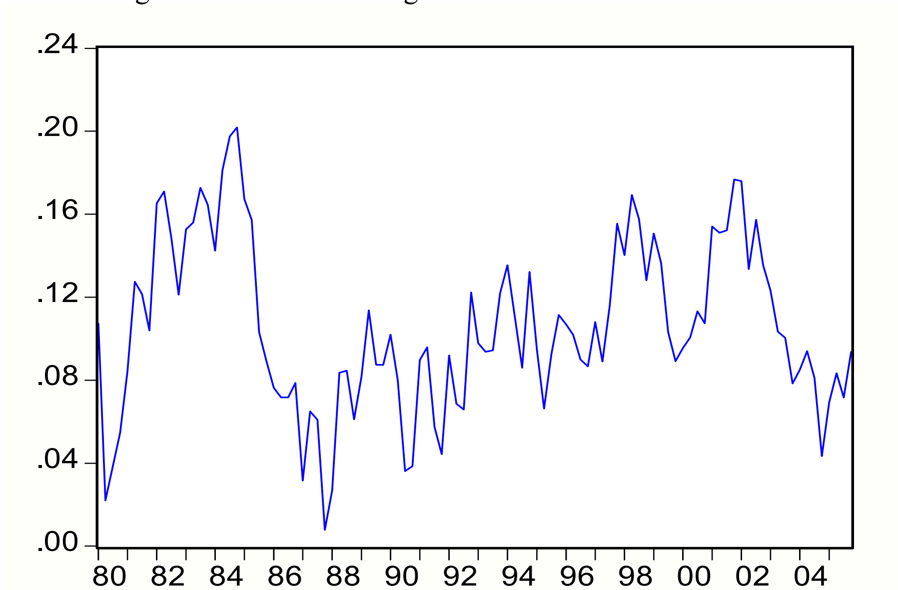
We construct aggregate measures of foreign inflation as weighted averages of the country-by-country inflation. The import-weighted average of foreign inflation is given by $\pi_t^{PPI} = \sum_{j=1}^{40} \left[\widehat{w}_{t-k}^{j} \cdot \Delta \ln(PPI_t^j) \right]$. Figure 2 shows the behavior of U.S. import price inflation (measured in U.S. dollars) and the (weighted) inflation of domestic production of U.S. trading partners (measured in foreign currency). Both were relatively high in the late 1980s and fell during the 1990s. During relatively long periods of the early and late 1980s and the 1990s, U.S. import price inflation was consistently below foreign domestic inflation.

Similarly, we construct aggregate measures of foreign interest rates and exchange rate depreciation as weighted averages of the country-by-country variables. The import-weighted exchange rate appreciation of the U.S. dollar is $ds_t = \sum_{j=1}^{40} \left[\widehat{w}_{t-k}^{j} \cdot \Delta \ln(S_t^j) \right]$; and import-weighted foreign interest rates are $i_t^* = \sum_{j=1}^{40} \left[\widehat{w}_{t-k}^{j} \cdot i_t^j \right]$ (see Appendix D for details on foreign interest rates).

[4] The foreign price indices are from various sources and are outlined by county in Appendix A. In many cases, the coverage of the manufactured goods producer price index is incomplete and the data is extended back in time using broader producer price indices, wholesale price indices, or GDP deflators.

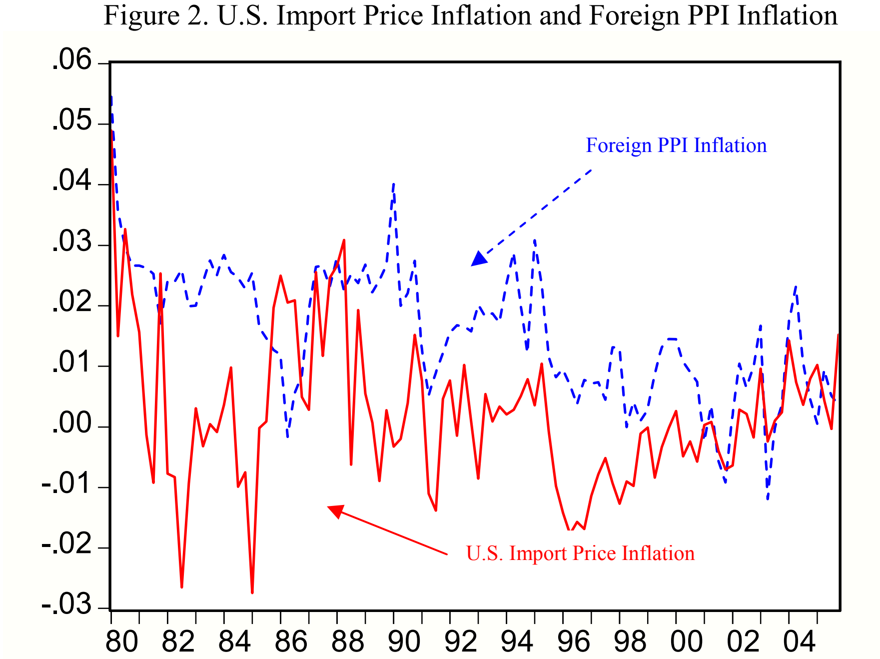
[5] The import price deflator can be obtained (http://www.bls.gov/mxp/home.htm) for the period 1985–present.

Figure 1. The Trade-Weighted Index of the Relative Prices



Notes: This figure depicts the logarithm of the trade-weighted index of the relative prices (μ_t) for the period 1980:Q1–2005:Q4.

Figure 2. U.S. Import Price Inflation and Foreign PPI Inflation



Notes: This figure compares inflation in the U.S. import price for manufactured goods with trade weighted foreign producer price inflation for the period 1980:Q1–2005:Q4.

IV. THE ESTIMATED RESULTS

A. Defining Exchange Rate Pass-through

The standard exchange rate pass-through literature typically regresses import price inflation or possibly CPI inflation on changes in the exchange rate and some variables meant to control for the marginal cost of production along with lags (for example, Yang 1997; Smets and Wouters, 2002; Campa and Goldberg, 2005; and Farugee, 2006). The rate of exchange rate pass-through can then be measured as a function of the dynamic correlation of exchange rate changes and domestic inflation at different horizons. Existing studies suggest cross-sectional differences on exchange rate pass-through, which depends on price setting behavior (for example, Dornbusch, 1987; Knetter, 1993; Devereux, Engel, and Storgaard, 2004) or concern about market share (Froot and Klemperer, 1989).

We define exchange rate pass-through in a slightly different way as determined by the model. In the context of an LCP model, the speed with which exchange rate changes begin to affect import prices depends in part on the frequency of price changes by firms. However, the degree to which firms change prices also depends on their expectations of the future dynamics of the exchange rate because they are also forward-looking agents. From this perspective, standard estimates of the exchange rate pass-through are subject to the Lucas critique: that is, they are not parameters that policy-makers can take as invariant to changes in exchange rate policy.

We will describe exchange rate pass-through as being represented by a structural parameter. We consider the degree of pass-through as the fraction of firms that change their prices in response to exchange rate pass-through, (1-v). At horizon j, the fraction of firms that have not adjusted prices will be (1-v^j).

B. Specification and Benchmark Regressions

The theory implies a rational expectations model of the form:

$$\pi_t^{IPI} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI} + \alpha_3\pi_t^{PPI} - \alpha_2\alpha_3\pi_{t+1}^{PPI}\right] \tag{4}$$

where coefficients $\alpha_1, \alpha_2, \alpha_3 > 0$ are functions of the subjective discount factor and the structural parameters of price setting. Specifically, $\alpha_1 = \frac{(1-\nu)(1-\beta\nu)}{\nu}, \alpha_2 = \beta$, and $\alpha_3 = \frac{(1-\nu)(1-\beta\nu)}{\nu}\cdot\frac{\kappa}{(1-\kappa)(1-\beta\kappa)}$. Coefficient α_1 represents the effect of exchange rate pass-through. The smaller is α_1, the slower is exchange rate pass-through: that is, an increase in the producer price relative to the import price lead to a smaller current increase in import inflation as α_1 declines. Also, a measure of $\alpha_3 < 1$ indicates that exchange rate pass-through is slower than foreign countries' domestic price adjustment.

Table 1. Estimation Results of the Pass-Through Effect Model

A. Benchmark							
Parameters	α_1	α_2	α_3	ν	κ	β	J-Test
	0.069*** (0.024)	1.114*** (0.104)	–0.033 (0.138)	0.754*** (0.019)	—	1.114*** (0.104)	18.15 [0.513]
Wald Test: $\alpha_4 = -\alpha_2 \times \alpha_3$			0.198 [0.657]				
B. No Current Information: drop contemporaneous π_t^{IPI} and π_t^{PPI}							
Parameters	α_1	α_2	α_3	ν	κ	β	J-Test
	0.078*** (0.027)	0.988*** (0.120)	–0.099*** (0.178)	0.785*** (0.018)	—	0.988*** (0.120)	18.06 [0.385]
Wald Test: $\alpha_4 = -\alpha_2 \times \alpha_3$			1.295 [0.255]				
C. Full Current Information: add contemporaneous μ_t, ds_t, i_t^*, and i_t^{FF}							
Parameters	α_1	α_2	α_3	ν	κ	β	J-Test
	0.050*** (0.018)	0.555*** (0.082)	0.213*** (0.065)	0.915*** (0.015)	0.763*** (0.055)	0.555*** (0.082)	20.87 [0.589]
Wald Test: $\alpha_4 = -\alpha_2 \times \alpha_3$			10.368 [0.017]				
D. Only Included Instruments: drop all lags of μ_t, ds_t, i_t^*, and i_t^{FF}							
Parameters	α_1	α_2	α_3	ν	κ	β	J-Test
	0.059* (0.035)	1.339*** (0.182)	–0.321 (0.367)	0.682*** (0.057)	—	1.339*** (0.182)	9.33 [0.968]
Wald Test: $\alpha_4 = -\alpha_2 \times \alpha_3$			0.004 [0.949]				
E. Short List Instruments: include only two lags of all instruments							
Parameters	α_1	α_2	α_3	ν	κ	β	J-Test
	0.072*** (0.036)	1.172*** (0.129)	–0.139 (0.287)	0.731*** (0.034)	—	1.172*** (0.129)	13.49* [0.061]
Wald Test: $\alpha_4 = -\alpha_2 \times \alpha_3$			0.635 [0.425]				
F. Two-Stage Least Squares Method							
Parameters	α_1	α_2	α_3	ν	κ	β	
	0.076* (0.042)	1.060*** (0.208)	–0.049 (0.342)	0.765*** (0.040)	—	1.060*** (0.208)	
Wald Test: $\alpha_4 = -\alpha_2 \times \alpha_3$			0.427 [0.513]				

Notes: This table shows the GMM estimation results of a structural model given by equation (4) for 1980:Q1–2005:Q4 using different sets of instruments. The benchmark instrument set includes contemporaneous import inflation, π_t^{IPI}; contemporaneous foreign producer price inflation, π_t^{PPI}; along with 4 lags of π_t^{IPI}, π_t^{PPI}, the markup (μ_t), the appreciation rate of the U.S. dollar (ds_t), the foreign interest rate (i_t^*), and the U.S. Federal funds rate. We are able to calculate a real estimate of κ only in the case where the estimate of $\alpha_3 > 0$ (that is, panel C). Adjustments were made to the benchmark instrument set for panels B–E as indicated above. The J-test is Hansen's overidentification test for all instruments (with p-values in square brackets) which is only valid for optimal GMM. The Wald test is a parameter restriction test for the null hypothesis that $\alpha_4 = -\alpha_2 \times \alpha_3$ (with p-values in square brackets). Standard errors are reported in parentheses. In panel F, the benchmark instrument set is used, and Newey-West corrected standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

We estimate the above model with the GMM method. Our "benchmark" instrument set includes contemporaneous import inflation, π_t^{IPI}; contemporaneous foreign producer price inflation, π_t^{PPI}; along with 4 lags of π_t^{IPI}, π_t^{PPI}, the markup (μ_t), the appreciation rate of the U.S. dollar (ds_t), the foreign interest rate (i_t^*), and the U.S. Federal funds rate. This instrument list exploits the fact that the predetermined price level is "known" by price setters and is thus in the information set at time *t*-1. We use a one-step Newey-West estimator of the covariance matrix constructed with pre-whitened residuals.

The estimated results of regression (4) are reported in Table 1. The coefficients on μ_t and π_{t+1}^{IPI} (α_1 and α_2) are positive and significant at the 1 percent level. This allows us to construct estimates of the structural parameters ν and β. The estimate of the subjective discount factor, $\beta = 1.114$, is within a standard deviation of its standard estimate (say, $\beta = 0.99$) at a quarterly frequency. The estimate of the probability of no import price adjustment, $\nu = 0.754$, is consistent with price changes of imports occurring on average once per year. This is similar to estimates of Galí and Gertler (1999), as to the adjustment of domestic prices in the U.S. However, the estimate of α_3 is negative and insignificant which makes it impossible to develop a parameter estimate of foreign price stickiness, κ. The insignificance of α_3 suggests no evidence of pass-through of foreign marginal cost into U.S. import prices. The Hansen's J-statistic test of the overidentification conditions of the model, which are not rejected at the 10 percent level, suggests that our instruments are valid.

C. Robustness Checks: Estimation Methods

We also report estimated results under slightly different assumptions on the information set to check for robustness. Panel B shows the estimated model after dropping contemporaneous information on π_t^{IPI} and π_t^{PPI} from the list of instruments. The results are very similar to the baseline results. In panel C, we show the results when we use contemporaneous information about all instruments in our list. This generates some strikingly different results. First, all of the parameters of the model are positive and significant at the 1 percent level. Crucially, this allows us to find an estimate of the degree of foreign price stickiness, κ, and indicates that foreign PPI inflation can control for foreign marginal cost. Here, the estimate of the degree of price stickiness of imports is $\nu = 0.915$ indicating import prices change on average every 10 quarters. The estimate degree of price stickiness in the foreign domestic price setting is very reasonably estimated at $\kappa = 0.786$, consistent with price changes every 4–5 quarters. However, the estimate of the subjective discount factor is $\beta \approx 0.56$ is very small for a quarterly discount rate and the hypothesis that $\beta = 0.99$ can easily be rejected at any reasonable critical value. The bias in the estimate of β makes us somewhat skeptical of this specification.

Panel D reports the estimation result when we set the instrument list to include only those variables which actually appear in the equation (contemporaneous levels of π_t^{IPI} and π_t^{PPI} and lags of π_t^{IPI}, π_t^{PPI}, and μ_t). Here, the coefficient estimates are similar to those in the Benchmark results with the frequency of import price changes estimated at 3–4 quarters. The standard errors are larger when fewer instruments are used. We, again, cannot find evidence that foreign inflation is connected to import price inflation in this specification. Panel E shows estimates when only 2 lags of each instrument (along with contemporaneous levels of π_t^{IPI} and π_t^{PPI}). Here, the estimates again are very similar to the Benchmark estimates. Disturbingly, the J-test of the overidentifying assumptions is rejected at the 10 percent level. Lastly, panel F shows the 2SLS estimates using the benchmark instrument set (with Newey-West corrected standard errors) which are again similar to the GMM estimates in Panel A.

D. Robustness Checks: Alternate Specifications

Table 2 reports estimation results with some alternative specifications of the model. The chief problem with the Benchmark specification is that the dynamics of foreign inflation seem unrelated with U.S. import price inflation. One potential reason is that with coefficient α_2 being estimated as so close to one, the foreign price inflation series appears in near first difference form. We relax the restrictions on the parameter on future foreign producer price inflation in the statistical model

$$\pi_t^{IPI} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI} + \alpha_3\pi_t^{PPI} - \alpha_4\pi_{t+1}^{PPI}\right] \tag{5}$$

The results are reported in panel A of Table 2. Relaxing the constraint that $\alpha_4 = \alpha_2 \times \alpha_3$ does not change our main results. Both α_3 and α_4 are statistically insignificant, indicating little relationship between foreign producer price inflation and domestic import inflation, either contemporaneously or with a lead. The relaxation of the constraint does not have a strong effect on the other variables we measure. In fact, a Wald test of the hypothesis $\alpha_4 = \alpha_2 \times \alpha_3$ is not rejected at the 10 percent level.

To further test the model, we estimate a modified version of the benchmark model by including a lagged term of foreign producer price inflation.

$$\pi_t^{IPI} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI} + \alpha_3\pi_t^{PPI} - \alpha_2 \cdot \alpha_3\pi_{t+1}^{PPI}\right] + \alpha_5\pi_{t-1}^{PPI}\,. \tag{6}$$

Panel B reports the results of the GMM estimation of equation (6). The estimated coefficient on lagged foreign producer price inflation is near 0 and is statistically insignificant. Apparently, there is little connection between foreign price inflation and the import price inflation at leads or lags. We also see that the inclusion of the lagged term changes little about the other benchmark results: α_1 and α_2 are both significantly positive and consistent with an average frequency of price change of about one year and an economically reasonable subjective discount factor.

As a more general check, we also estimate a hybrid pass-through effect model with forward- and backward-looking expectations given by

$$\pi_t^{IPI} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI} + \alpha_3\pi_t^{PPI} - \alpha_2 \cdot \alpha_3\pi_{t+1}^{PPI}\right] + \alpha_6\pi_{t-1}^{IPI} \tag{7}$$

Table 2. Estimating the Pass-through Effect Model: Alternative Specifications

A. Equation (5): unrestricted coefficient on π_{t+1}^{PPI}					
Parameters	α_1	α_2	α_3	α_4	J-Test
	0.069***	1.087***	–0.032	–0.064	18.65
	(0.024)	(0.108)	(0.141)	(0.183)	[0.412]
B. Equation (6): with a lagged PPI inflation, π_{t-1}^{PPI}					
Parameters	α_1	α_2	α_3	α_5	J-Test
	0.072***	1.086***	–0.056	0.024	17.72
	(0.023)	(0.105)	(0.145)	(0.057)	[0.474]
C. Equation (7): with a backward-looking term, π_{t-1}^{IPI}					
Parameters	α_1	α_2	α_3	α_6	J-Test
	0.072***	1.015***	–0.076	0.075	18.39
	(0.022)	(0.111)	(0.125)	(0.060)	[0.430]
D. Equation (8): imposing $\kappa\left[(1-\kappa)(1-\beta\kappa)\right]^{-1}=12$					
Parameters	α_1	α_2			J-Test
	0.015	1.219***			18.21
	(0.010)	(0.061)			[0.573]
E. Equation (9): imposing $\kappa\left[(1-\kappa)(1-\beta\kappa)\right]^{-1}=12$; $\alpha_1 \neq \alpha_8$					
Parameters	α_1	α_2	α_8		J-Test
	0.069***	1.114***	–0.003		18.15
	(0.024)	(0.104)	(0.011)		[0.513]

Notes: This table shows the GMM estimation results of a structural model given by equation (4) for 1980:Q1–2005:Q4. In all panels, the instrument set includes contemporaneous import inflation, π_t^{IPI}; contemporaneous foreign producer price inflation, π_t^{PPI}; along with 4 lags of π_t^{IPI}, π_t^{PPI}, the markup (μ_t), the appreciation rate of the U.S. dollar (ds_t), the foreign interest rate (i_t^*), and the U.S. Federal funds rate. The J-stat is Hansen's overidentification test for all instruments (with p-values in square brackets). Standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

This equation is analogous to a hybrid New Keynesian Phillips curve à la Galí, Gertler, and Lopez-Salido (2005), which nests the pure forward-looking model as a special case. As shown in panel C, we find no evidence that the lagged variable enters significantly. The inclusion changes neither the basic result that α_1 and α_2 are both positive and statistically significant nor the estimate of the frequency of import price changes.

To assess whether our results are biased by the lack of significance of the foreign producer price inflation term, we re-estimate the model after calibrating the term $\frac{\kappa}{(1-\kappa)(1-\beta\kappa)}=12$, which would be consistent with $\kappa=0.75$ and $\beta=1$. The regression will then be of the form:

$$\pi_t^{IPI} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI} + \alpha_1 \cdot 12 \cdot \pi_t^{PPI} - \alpha_2 \cdot \alpha_1 \cdot 12 \cdot \pi_{t+1}^{PPI}\right] \tag{8}$$

We report the results in panel D. The estimate of ν is close to the benchmark level with a fairly narrow standard error. However, the coefficient α_1 is very small and only marginally significant. This model, by imposing the constraint that foreign producer price inflation passes through into import price inflation proportionately to the markup of import prices over the PCP price, results in a very low level of exchange rate pass-through. Further, we can reject the hypothesis that β=0.99 at the 5 percent level.

Panel E reports the results from a specification that relaxes the assumption that foreign producer price inflation passes through into import price inflation at the same rate as exchange rates. We estimate the specification

$$\pi_t^{IPI} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI} + \alpha_8 \cdot 12 \cdot \pi_t^{PPI} - \alpha_2 \cdot \alpha_8 \cdot 12 \cdot \pi_{t+1}^{PPI}\right] \tag{9}$$

Not surprisingly, after dropping the proportional assumption we return to our benchmark results in terms of α_1 and α_2 but find a negative estimate of α_8. The hypothesis that $\alpha_1 = \alpha_8$ was easily rejected at the 5 percent level, suggesting that the impact on domestic import inflation of foreign price inflation appears to be quite different from that of the nominal exchange rate.

We view the negative estimate of the coefficient on foreign inflation in LCP models (Tables 1 and 2) as a result of ignoring the role of PCP firms. In the following subsections, the coefficient estimates on foreign inflation become positive (and often significant) in the mix of LCP and PCP models, which enables us to report significant estimates of structural parameters.

E. Pass-through Effect Model with A Mix of LCP and PCP

The above models assume that all firms exporting to the U.S. adjust their U.S. dollar prices infrequently and optimally according to the local currency pricing (LCP) pass-through theory. We relax this assumption by modeling a world in which some fraction of firms (that is, PCP firms) price their goods in their home currency based on changes in aggregate prices.

In producer currency pricing (PCP), firms have sticky prices in their own currency while their invoice prices in U.S. dollar adjust automatically as exchange rates change. We assume that a fraction, λ, of firms set prices in their home currency and pass on the prices into import goods with a one-period lag.

$$\pi_t^{IPI,PCP} = \pi_{t-1}^{PPI} - ds_{t-1}$$

We assume that these firms are randomly distributed and their home prices adjust the same as other firms. For the LCP firms, we model the inflation in the standard way

$$\pi_t^{IPI,LCP} = \alpha_0 + E_{t-1}\left[-\alpha_1\mu_t + \alpha_2\pi_{t+1}^{IPI,LCP} + \alpha_3\pi_t^{PPI} - \alpha_2 \cdot \alpha_3\pi_{t+1}^{PPI}\right] \tag{10}$$

Total import price inflation can be written as $\pi_t^{IPI} = \lambda \cdot \pi_t^{IPI,PCP} + (1-\lambda) \cdot \pi_t^{IPI,LCP}$.

$$\begin{aligned}
\pi_t^{IPI} &= \lambda \pi_t^{IPI,PCP} + (1-\lambda) \cdot \left\{ \alpha_0 + E_{t-1}\left[-\alpha_1 \mu_t + \alpha_2 \pi_{t+1}^{IPI,LCP} + \alpha_3 \pi_t^{PPI} - \alpha_2 \cdot \alpha_3 \pi_{t+1}^{PPI} \right] \right\} \\
&= \lambda \pi_t^{IPI,PCP} + (1-\lambda) \cdot \left\{ \alpha_0 + E_{t-1}\left[-\alpha_1 \mu_t + \alpha_2 \frac{\pi_{t+1}^{IPI} - \lambda \pi_t^{IPI,PCP}}{(1-\lambda)} + \alpha_3 \pi_t^{PPI} - \alpha_2 \cdot \alpha_3 \pi_{t+1}^{PPI} \right] \right\} \\
&= \lambda \pi_t^{IPI,PCP} + \alpha_2 E_{t-1}\left[\pi_{t+1}^{IPI} - \lambda \pi_{t+1}^{IPI,PCP} \right] + (1-\lambda) \cdot \left\{ \alpha_0 + E_{t-1}\left[-\alpha_1 \mu_t + \alpha_3 \pi_t^{PPI} - \alpha_2 \cdot \alpha_3 \pi_{t+1}^{PPI} \right] \right\}
\end{aligned} \tag{11}$$

where $\pi_t^{IPI,PCP} = \pi_{t-1}^{PPI} - ds_{t-1}$. We estimate equation (11) with GMM using the benchmark set of instruments and report the results in Table 3.

The fraction of firms that operate as PCP firms is estimated as very small as indicated by the λ estimate, which is about 6 percent and significant at the 1 percent level. The inclusion of the $\pi_t^{IPI,PCP}$ and $\pi_{t+1}^{IPI,PCP}$ terms has little effect on the estimates of α_1 and α_2, which are still consistent with an economically reasonable subjective discount factor and import price changes with a frequency between 4 and 5 quarters. The coefficient on foreign price inflation, α_3, is positive but not significant. We are able to estimate a parameter of probability of foreign producer price stickiness of $\kappa \approx 0.4$ although its standard errors are so large that it is not statistically different from 0 or 1.

We split the sample in two parts, one running through 1991 and the other from 1992 to 2005, for two reasons. The first reason concerns data quality. To construct our measure of foreign producer price inflation, we had to make compromises with data quality in order to extend our measure back through 1980 (see the appendix). The data after 1992 should be more accurate. The second reason pertains to a recent literature on structural change in import price pass-through. Marazzi et al. (2006) argue that import price pass-through has declined significantly over time, partly attributable to a change in pricing strategy after the 1997–98 East Asian crises. Our structural model estimations help us identify the reasons for changes in reduced form regressions. We estimate the sub-sample regressions by GMM with contemporaneous measures of π_t^{IPI} and π_t^{PPI} and only two lags of each of the instruments owing to the shortness of the sample period.

In both sub-samples, we find considerably more evidence for the pass-through of foreign prices into U.S. import inflation as shown by positive estimates of α_3. Panel B shows the parameters from the regression for 1980–1991. The estimates of α_1 and α_2 are both positive and of a size consistent with the benchmark regressions. The estimated frequency of import price changes is on average once per 5 quarters. The average frequency of foreign producer price changes is estimated at around once per 4 quarters. Interestingly, in this period, the fraction of PCP firms is near zero (λ=0.029), and we cannot statistically reject the hypothesis that λ=0 at the 10 percent level.

Table 3. Estimating the Pass-through Effect Model: A Mix of LCP and PCP

A. Full Sample 1980:Q1–2005:Q4				
Coefficient Estimates	α_1	α_2	α_3	J-test
	0.067***	1.032***	0.036	16.40
	(0.023)	(0.119)	(0.127)	[0.565]
Structural Parameters	v	κ	β	λ
	0.784***	0.420	1.032***	0.066***
	(0.023)	(0.432)	(0.119)	(0.018)
B. Sub-Sample 1980:Q1–1991:Q4				
Coefficient Estimates	α_1	α_2	α_3	J-test
	0.080*	0.945***	0.609	9.65
	(0.049)	(0.277)	(0.415)	[0.140]
Structural Parameters	v	κ	β	λ
	0.796***	0.749***	0.945***	0.029
	(0.052)	(0.060)	(0.277)	(0.050)
C. Sub-Sample 1992:Q1–2005:Q4				
Coefficient Estimates	α_1	α_2	α_3	J-test
	0.024	1.069***	0.534***	10.67
	(0.018)	(0.104)	(0.134)	[0.099]
Structural Parameters	v	κ	β	λ
	0.839***	0.802***	1.069***	0.135***
	(0.047)	(0.053)	(0.104)	(0.018)

Notes: This table reports the GMM estimation results of a pass-through model with a mix of local currency pricing (LCP) and producer currency pricing (PCP), structural equation (11) for different sample periods. For the full sample, the benchmark instrument set is used. For the sub-sample regressions (panels B and C), we use contemporaneous measures of π_t^{IPI} and π_t^{PPI} and only two lags (rather than four lags) of each of the variables included in the benchmark instrument list. The J-test is Hansen's overidentification test for all instruments (with p-values in square brackets). Standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

In the 1992–2005 period (panel C), we find a much larger fraction of PCP firms, as suggested by the λ estimate of 13.5 percent. In this sample period, we estimate a frequency of import price change that is slightly slower than that during the earlier sample period: price changes occurring on average about once per 6 quarters. However, the increase in the stickiness of import prices is neither sizable nor significant. In both sub-samples, the estimate of the subjective discount factor is within a standard error of an economically reasonable number. It should be noted that the model restricts the data along a number of dimensions, and in the latter sample period the overidentifying restrictions are marginally rejected at the 10 percent level.

F. Regional Models and Country-Specific Exports

The BLS collects import price indices at finer levels of aggregation based on the location of the originating country. From 1991, BLS has reported import price indices for manufactured goods from industrial countries and from developing countries. We sub-divide our list of 40 countries into two groups and construct aggregates of foreign producer prices, foreign interest rates and exchange rates that correspond to each country group.[6] Table 4 reports the estimated results of a mix of LCP and PCP model for both country groups. Note that the source for the import prices in this table (the BLS) is different from that for the previous import prices (a splice of OECD and BLS data). Again, owing to the shortness of the sample period, all regressions are estimated with contemporaneous π_t^{IPI} and π_t^{PPI} and 2 lags of each of the instruments.

In panel A, we report the results for the industrial countries. All parameter estimates are significant at the 1 percent level including α_3, the coefficient on foreign producer price inflation. The estimate of the subjective discount factor is low compared to standard estimates, and the hypothesis that β=0.99 can be rejected at the 5 percent level. The coefficient on the markup of import prices over the PCP price (α_1) is low although precisely estimated. Accordingly, the price stickiness is extreme, with the probability of both import prices and foreign prices remaining fixed in any given quarter above 0.9, while a substantial fraction of firms act according to PCP with $\lambda \approx 16$ percent. The overidentifying restrictions are rejected at the 10 percent level.

Among the developing countries, price stickiness is even more extreme. In this sample, the coefficient on the markup term (α_1) is negative indicating no convergence of import prices to the PCP price. As a result, we cannot estimate a parameter of price stickiness. We also find a very small (6.9 percent), but statistically significant, fraction of manufacturing imports are set according to PCP in developing countries.

Finally, we estimate the partial PCP model for manufacturing imports from Canada and Japan. The BLS produces import price series for goods from these countries, again beginning in 1991. Since this exercise involves data from a single country, we do not need to aggregate prices, exchange rates, or interest rates over countries. The results are reported in panels C and D of Table 4. We find very similar results for both countries: (i) very low and statistically insignificant coefficients on the markup of import prices over marginal cost are accompanied by fairly sluggish import prices (with ν being close to 0.9): the average price change frequency is about 10 quarters; (ii) positive estimates of the impact of foreign producer prices on import inflation (α_3) render foreign price adjustment substantially sluggish (with κ around 0.85); (iii) a small but statistically significant fraction of firms (around 7 percent) follow producer currency pricing; and (iv) the estimated subjective discount factor is not significantly different from the standard estimate (β=0.99).

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[6] The industrial country group comprises Australia, Austria, Belgium, Canada, Hong Kong SAR, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Korea, Netherlands, New Zealand, Norway, Portugal, Singapore, South Africa, Spain, Sweden, Switzerland, Taiwan, and the United Kingdom. The developing country group is confined to Brazil, Chile, China, Colombia, Costa Rica, India, Indonesia, Israel, Malaysia, Mexico, Peru, Philippines, Poland, Russia, Thailand, and Turkey, considering their shares in U.S. imports.

Table 4. Regional Pass-through Effect Models: A Mix of LCP and PCP

A. Industrial Countries				
Coefficient Estimates	α_1	α_2	α_3	J-test
	0.023*** (0.007)	0.642*** (0.130)	0.507*** (0.098)	11.55* [0.073]
Structural Parameters	v	κ	β	λ
	0.947*** (0.010)	0.910*** (0.021)	0.642*** (0.130)	0.164*** (0.014)
B. Developing Countries				
Coefficient Estimates	α_1	α_2	α_3	J-test
	–0.003 (0.011)	1.111*** (0.094)	0.569*** (0.086)	7.84 [0.25]
Structural Parameters	v	κ	β	λ
	—	—	1.111*** (0.094)	0.069*** (0.017)
C. Canada				
Coefficient Estimates	α_1	α_2	α_3	J-test
	0.003 (0.017)	1.128*** (0.117)	0.400*** (0.066)	10.35 [0.111]
Structural Parameters	v	κ	β	λ
	0.872*** (0.060)	0.855*** (0.110)	1.128*** (0.117)	0.075*** (0.026)
D. Japan				
Coefficient Estimates	α_1	α_2	α_3	J-test
	0.012 (0.012)	0.969*** (0.146)	0.291 (0.240)	9.00 [0.174]
Structural Parameters	v	κ	β	λ
	0.912*** (0.033)	0.841*** (0.083)	0.969*** (0.146)	0.073*** (0.008)

Notes: This table reports the GMM estimation results of a pass-through model with a mix of local currency pricing (LCP) and producer currency pricing (PCP), equation (11), for the industrial and the developing country groups in panels A and B, respectively. Panels C and D show the GMM estimation results of the model for manufacturing imports from Canada and Japan, respectively. In all regressions, we use contemporaneous measures of π_t^{IPI} and π_t^{PPI} and only two lags of each of the variables included in the benchmark instrument list. The J-test is Hansen's overidentification test for all instruments (with p-values in square brackets). Standard errors are reported in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

V. CONCLUDING REMARKS

In this paper, we estimate a structural parameter that is crucial for determining the effectiveness of exchange rate policy: the frequency with which sticky-price firms are able to change prices in response to changes in exchange rates or marginal costs. Our estimate of pass-through effect of about 7 percent per quarter suggests that a depreciation of the U.S. dollar would not increase the U.S. import price in dollars substantively. We also find that the stickiness of import prices is slightly increased over time but the increase is neither sizable nor significant.

As noted by Taylor (2000), there exists a potential complementarity between monetary stability and policy effectiveness. Since price setting and pass-through effects could be endogenous (Devereux and Yetman, 2003; Devereux, Engel, and Storgaard, 2004), higher exchange rate volatility could induce an increase in import price pass-through by reducing domestic price inertia.[7] That is, higher exchange rate volatility reduces monetary stability but increases import price pass-through effects. Pass-through effects in the U.S. would be more susceptible to a change in exchange rate volatility because our findings provide evidence on forward-looking expectations in the determination of pass-through effects. We acknowledge that our pass-through effect models based on the optimization framework with the New Keynesian Phillips curve treat the import price setting parameter (v) as a time-invariant, deep parameter, but this parameter itself could be chosen endogenously for optimization when the degree of monetary stability could vary.

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[7] If local currency pricing remains to hold for some reasons (for example, concern about market share), however, large and frequent exchange rate changes may result in increased exchange rate volatility without curving import demand pressures since local currency pricing eliminates the pass-through from changes in exchange rates to consumer prices (Devereux and Engel, 2002).

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Appendix: Data Descriptions

A. Foreign Producer Prices

The following countries have OECD manufacturing PPI's for the whole period 1975–2005: Canada, Japan, Germany, Korea, U.K., France, Ireland, India, Switzerland, Indonesia, Netherlands, Spain, Australia, Austria, Finland, Denmark, Turkey, and New Zealand. The data for Norway, Turkey, and Belgium are available only from 1977, 1979, and 1980, respectively. The same dataset has data for Mexico and Italy from 1981 and Sweden from 1982. Data on Portugal, Russia, and Poland are available only from 1990, 1992, and 1996, respectively.

The data on Mexico is extended to 1980 with a Banco de Mexico series from 1980 to 1981 with a PPI excluding petroleum and services (source: CEIC). Mexican data is extended to 1975 with quadratically interpolated annual GDP deflator data from IMF International Financial Statistics (hereafter IFS). The data on Sweden is extended to 1975 using the monthly PPI/WPI from IFS. On a quarterly basis, the Italian and Portuguese data are extended using the GDP deflator from IFS to 1980 and 1977, respectively. Quarterly data on PPI/WPI are used to extend the data to 1982 for Poland. Polish data is extended to 1980 with quadratically interpolated data from IFS. Annual GDP deflators from UN National Accounts are quadratically interpolated to extend Belgian, Italian, Norwegian, Polish, Portuguese, Russian and Turkish data to 1975.

Monthly data on PPI or WPI are available from IMF for Brazil, Chile, Costa Rica, Israel, Peru, Singapore, South Africa and Thailand for the entire period. The data for Peru, Malaysia, and Philippines are available from 1980, mid-1986, and 1993, respectively. Monthly data are converted to a quarterly frequency by averaging. The Philippines data is extended to 1981 using the Philippines quarterly GDP deflator from IFS. A quarterly average of a discontinued WPI series for Metro Manila from Philippines National Statistical Office via CEIC is used to extend the Philippines data through 1976. Peruvian, Malaysian, and Filipino data is extended to 1975 with quadratically interpolated annual data on GDP deflators from IFS. For Colombia, the GDP deflator is used to infer the level of the PPI in the 1st quarter of 1994, which is not available.

A PPI series on industrial products for China is available from China's National Bureau of Statistics from October 1996. To extend this data, we calculate an equal weighted average of inflation from a Raw Materials Price Index for Textile Materials and Construction Materials. This inflation series is used to chain the deflator to January 1990. A quarterly average of this series is taken. An annual GDP Deflator from UN National Accounts is quadratically interpolated to extend the data to 1975. The manufacturing WPI for Taiwan (from Directorate-General of Budget, Accounting and Statistics in Taiwan via CEIC) available from 1981 is extended with a broader WPI from 1975. The data on quarterly GDP deflator for Hong Kong SAR is available over the period from Hong Kong Bureau of Census and Statistics.

B. Exchange Rates

Monthly exchange rates are obtained for all the countries from IFS except for Russia (before mid-1992) and Taiwan. For Russia, annual exchange rate series are taken from United Nations national accounts and quadratically interpolated to a monthly frequency. Taiwanese and

U.S. exchange rates with the H.K dollar are obtained from HKMA and the cross-rates are used to calculate a monthly exchange rate. For Poland, the series on exchange rates begins only in 1978. Information from the UN National Accounts data indicated that the zloty rate was fixed to the dollar for 1975–1978, so this rate was used.

C. Trading Partners' Exports of Manufactured Goods to the U.S.

For each country, we obtain data on exports to the U.S. by SITC code from Feenstra et al. (2005) at 5-year intervals for 1975–2000. Totals of manufacturing exports to the U.S. for each country are defined as the total of exports that are not coded as SITC 2 or 3 (that is, raw materials). Quarterly data are obtained from IMF Direction of Trade Statistics (DOTS) with the exception of Taiwan which is obtained from U.S. Customs from 1987. To obtain Taiwan exports to the U.S. before 1987, annual data from Feenstra et al. (2005) is split evenly between quarters. Countries are ranked by size of manufacturing exports in each of 1975, 1980, 1985, 1990, 1995, and 2000. Any country that ranks in the top 30 exporters of manufactured goods to the U.S. in any of those years is included in the list. To estimate exports to the U.S. at quarterly frequency, we interpolate the fraction of each country's exports of manufactured goods to the U.S. for each of the years examined into quarterly fractions, and then multiply these fractions by DOTS data to get an estimate of the bilateral trade in manufactured goods on a quarterly basis.

D. Foreign Interest Rates

Our primary data source for interest rates is IFS. Short-term rates are available in a number of categories. Our first choice of data for any country is money market rates. The IFS has a full set of money market rates (1975–2006) for the U.K., Italy, Germany, Norway, Canada, Japan, Australia, South Africa, Spain, Malaysia, and Singapore. Austria and the Netherlands have their own money market rates to January 1999, and Belgium to February 1999. These series are extended to the end of the sample period using Eurozone money market rates. If IFS money market data are not available for some countries for some periods, then our next option is to fill in the gap with government bill rates (a first choice), deposit rates (a second choice), or discount rates (a third choice) from IFS.

- France and Ireland: The IFS data on France's money market rates ends in March 1999, and Eurozone money market rates are used to extend the data to the end of sample. A gap in the series for March–May 1986 is filled with government bill rates from IFS. The IFS money market data for Ireland runs from 1975 to September 2006. Eurozone money market money market rates are used to extend the data to the end of sample. In the Irish data, we fill in several one- or two- month gaps in 1976 with quarterly averages, and a five month gap for January–May 1978 with the discount rate.
- Portugal and Sweden: For Portugal, we use the discount rate for January 1975–November 1975, a deposit rate for December 1975–December 1982, money market rates for January 1983–March 2000, and Eurozone money market rates thereafter. For Sweden, money market data from IFS ends in December 2004. The data is extended to July 2006 using government bill rates and then to the end of period using the discount rate.
- New Zealand, the Philippines, and Turkey: For New Zealand, discount rates are used for January 1975–December 1978, and government bill rates for January 1979–December

1984. For the Philippines, the discount rate is used for 1975 while government bill rates are used for 1976. For Turkey, the discount rate is used for January 1975–August 1985, and government bill rates for September 1985–March 1986. Money market rates are then used to complete the sample to the end of period for these countries.

- Columbia, Costa Rica, and Peru: The interest rate for Columbia is the discount rate for 1975–1985, the deposit rate for January 1986–February 1995, and the money market rate thereafter. For Costa Rica, we use the discount rate from IFS for 1975–1981 and a deposit rate from IFS afterwards. The interest rate for Peru is the discount rate for 1975–1987, the deposit rate for January 1988–August 1995, and the money market rate thereafter.
- Switzerland and Denmark: The discount rate is used for Switzerland for January–August 1975 and for Finland for January 1975–November 1977. The money market rate for both countries is used for subsequent periods. For Denmark, the money market rate is used for the whole sample, but the discount rate is used to fill in a gap for January–March 2001.
- Korea and Indonesia: For Korea, we use deposit rates for January 1975–July 1976, and money market rates thereafter. The interest rate in Indonesia is also based on deposit rates for January 1975–November 1984 and on money market rates thereafter. We fill a gap in money market rates during the first 4 months of 1986 through linear interpolation.

For a few countries, we complete the sample with non-IMF data. For example, the IFS series on the money market rate of India runs from January 1975 to May 1998. The series is extended to the end of sample using the Mumbai Stock Exchange listed interbank rate (from CEIC). The Israel interest rate is constructed from IFS from January 1979 through the end of sample. We use a lending rate for January 1979–February 1982, the discount rate for March 1982–April 1984, a deposit rate for May–December 1984, and the money market rate thereafter. For 1975–1978, however, we use an overdraft rate from the Bank of Israel website. For Taiwan, which is not a member of the IMF, all of the data comes from CEIC with a discount rate for January 1975–September 1985 and a money market rate thereafter.

Finally, for countries for which we could find no short-term interest rate for some part of the sample, we produce an artificial interest rate. The artificial interest rate is set equal to 2 percent plus a measure of expected inflation. Note that, with the exception of Russia and briefly Brazil, the use of the artificial interest rate is restricted to pre-1980 data.

- Mexico, Chile, and Thailand: For Mexico, we use the artificial rate for 1975, a deposit rate from IFS for 1976–1977, a government bills rate from IFS for January 1978–March 1981, and the IFS data money market rate thereafter. For Chile, we use the artificial rate for 1975–1976, a deposit rate from IFS for January 1977–November 1999, and an IFS money market rate from December 1999 through the end of sample. For Thailand, we use the artificial rate for 1975, a mortgage lending rate from IFS for 1976, and the money market rate thereafter. For Mexico, Chile, and Thailand, we measure expected inflation by the ex post year-on-year monthly inflation using the CPI from IFS.
- Hong Kong: The artificial rate is used for January 1975–August 1976. The measure of expected inflation is measured by the ex post year-on-year monthly inflation calculated with the CPI from Bureau of Census and Statistics. This series, however, begins in July 1975. For January-June 1975, the average year-on-year inflation for July 1974–July 1975 is used.

- Poland, Russia, and China: For Poland, the artificial rate is used for calendar years 1975–1978, a mortgage lending rate from IFS for January 1979–November 1989, a deposit rate from IFS for December 1989–November 1990, and a money market rate from December 1990 to the end of sample. For Russia, the artificial rate is used for January 1975–June 1995, and a money market rate from IFS thereafter. For China, the artificial rate is used for 1975–1979, and a deposit rate from IFS from 1980 to the end of sample. For Poland and Russia, the measure of expected inflation is the inflation rate for the year calculated with the annual final consumption deflator from UN National Accounts. For China, expected inflation is the CPI inflation rate for the year from the National Statistics Office.
- Brazil: The expected inflation is calculated with annual CPI Inflation from the Long-term Financial Database (http://www.globalfindata.com). We use the artificial rate for January 1975–November 1982, a deposit rate from IFS for December 1982–December 1994, and a government bills rate thereafter. However, some gaps (July 1998–February 1999; July 1999; and November 2002–January 2003) are filled in with an interbank rate from CEIC.