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The Exchange Rate Pass-Through to Import and Export Prices: The Role of Nominal Rigidities and Currency Choice

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The Exchange Rate Pass-Through to Import and Export Prices: The Role of Nominal Rigidities and Currency Choice

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Abstract

Using both regression- and VAR-based estimates, the paper finds that the exchange rate pass-through to import prices for a large number of countries is incomplete and larger than the pass-through to export prices. Previous studies have reported similar results, which give rise to the puzzle that while local currency pricing is needed to account for incomplete import price pass-through, it would not imply a lower export price pass-through. Recent explanations of this puzzle have emphasized markup adjustment in response to exchange rate changes. This paper suggests an alternative explanation based on the presence of both producer and local currency pricing. Using a dynamic general equilibrium model, the paper shows that a mix of producer and local currency pricing can explain the pass-through evidence even with a constant markup. The model can also explain the observed exchange rate and inflation variability as well as the fact that the regression and VAR estimates tend to be similar.

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I. INTRODUCTION

There continues to be much interest in understanding the mechanism that determines the exchange rate pass-through to import and export prices. A key issue is what role is played by nominal rigidities and currency choice (for setting the price of traded goods) in determining the behavior of import and export prices relative to the exchange rate. This question has important implications for the international transmission mechanism and the design of optimal monetary policy in an open economy. The conventional assumption—incorporated in the Mundell-Fleming model and adopted by Obstfeld and Rogoff's seminal (1995) contribution to new open economy macroeconomics—is that prices of traded goods are sticky in the currency of the producer. Models based on the assumption of producer currency pricing (PCP) imply that flexible exchange rates are desirable in achieving relative price adjustment. Moreover, optimal monetary policy rules are inward looking in that they stabilize domestic prices and output, and do not react to international variables like the exchange rate (Corsetti and Pesenti, 2001; Clarida, Gali, and Gertler, 2002). An alternative view assumes that prices of traded goods are sticky in the currency of local consumers (e.g., Betts and Devereux, 2000). The assumption of local currency pricing (LCP) leads to very different prescriptions for monetary policy. For example, Devereux and Engel (2003) show that under LCP, there is no benefit to exchange rate flexibility and fixed exchange rates are to be preferred. If exchange rates are flexible, optimal monetary policy under LCP would react to international variables (Corsetti and Pesenti, 2005).

Both PCP and LCP hypotheses, however, have problems explaining the evidence on the exchange rate pass-through to import and export prices. In the baseline versions based on preset prices, PCP implies that the pass-through (the elasticity of the price in home currency with respect to the exchange rate) equals one for import prices and zero for export prices. These implications are reversed under LCP. Inconsistent with these predictions, the import price pass-through for OECD countries tends to be significantly different from zero and one, and on average, is close to one-half (Campa and Goldberg, 2005). Estimates of the export price pass-through also generally do not support the baseline versions of both PCP and LCP (e.g., see Bussière and Peltonen, 2008). Models based on staggered prices modify the values of pass-through predicted by PCP and LCP, but not sufficiently to conform to the evidence. For example, if prices are set according to the Calvo (1983) model, the pass-through to import prices under LCP would be positive but generally lower than the estimated value. More problematically, LCP would still imply a value of the pass-through to export price, which is too high. This inconsistency was highlighted by Obstfeld and Rogoff (2000), who pointed out that a depreciation of national currency (an increase in the exchange rate) should improve the terms of trade (the price of exports relative to imports) according to LCP, but in data, depreciation tends to be associated with a deterioration in the terms of trade (implying that the pass-through to export prices tends to be smaller than to import prices).

To explain the evidence on the pass-through to import and export prices and the behavior of the terms of trade, a number of recent studies have suggested variations of the basic model, which allow markup adjustment in response to exchange rate changes. Corsetti and Dedola (2005), for example, develop a model of price discrimination based on the assumption that sale of goods abroad requires an input of local (nontraded) distribution services. They show that even in the presence of flexible prices (in which case, currency choice does not matter), a sufficiently high distribution margin in this model can generate the import price pass-through and the exchange rate-terms of trade association observed in data.[2] Corsetti, Dedola, and Leduc (2008) extend this model to also explain exchange rate volatility and other empirical regularities. This literature suggests that markup adjustments rather than nominal rigidities and currency choice play a key role in determining the pass-through behavior.

This paper argues that the pass-through evidence can be explained by an alternative model, which assumes that there is a mix of firms using PCP and LCP in each economy. This assumption is consistent with data on the currency of invoicing of exports and imports (e.g., see Goldberg and Tille, 2005), which shows that international transactions are invoiced in national currency as well as trade partner currency or a vehicle currency like the US dollar, and the share of national currency invoicing varies across countries and industries. A hybrid model with both PCP and LCP is also found to help explain the degree of the exchange rate pass-through to various prices in non-US G6 countries (Choudhri, Faruqee, and Hakura, 2005). The literature on optimal currency choice (for pricing of export goods) suggests that the use of both PCP and LCP in an economy can arise under two types of equilibria. One possibility is an equilibrium where firms using PCP and LCP are indifferent between the two types of pricing (Devereux, Engel, and Storegaard, 2004). Another possibility is that PCP is preferred for one type of products while LCP is preferred for others (Bacchetta and van Wincoop, 2005). The key factors determining the currency choice are sensitive to how the model is specified. We do not attempt to explain the currency choice, but focus instead on whether a model with both PCP and LCP and without a markup adjustment mechanism is capable of explaining the evidence on the pass-through to trade prices for a wide range of countries. For this purpose, we estimate the exchange rate pass-through to import and export prices for a large number of countries and examine how well a Dynamic Stochastic General Equilibrium (DSGE) model with staggered prices, a mixture of both PCP and LCP and a constant markup can explain the pass-through estimates.

The standard empirical model for estimating import or export price pass-through is based on a micro or a partial-equilibrium framework, and examines the price response of an exporter to an exogenous exchange rate change. The pass-through elasticity in this model is estimated by a regression of the import/export price index on the exchange rate with additional variables included in the regression to control for the marginal cost, markup, and short-run dynamics. In a general equilibrium model suitable for analyzing the behavior of aggregate

[2] See Corsetti, Dedola, and Leduc (2007), and Gust, Leduc, and Vigfusson (2010) for alternative models of markup adjustment.

trade prices, the exchange rate as well as some control variables are endogenous, and thus the estimated coefficients would be sensitive to the combination of shocks affecting the economy and the policy regime influencing the transmission of the shocks.[3] An alternative approach uses a VAR model to estimate the exchange rate pass-through to a price as the elasticity of the price to an orthogonalized VAR innovation to the exchange rate (ordered as the first variable in the VAR). An appealing feature of this approach is that it focuses on the price response to exogenous shocks. A limitation of this methodology, however, is that the exchange rate innovation represents a composite of structural disturbances that are difficult to identify.

We first review the evidence on the short-run (one quarter) exchange rate pass-through to import and export prices for a large sample of countries for which data on import and export price indexes are available for a sufficiently long period. Using simple versions of both regression and VAR models, we find that the pass-through to import as well as export prices tends to fall in the unit interval and the import price pass-through tends to be larger than the export price pass-through. Remarkably, for both import and export prices, the regression estimates of the pass-through are similar to the VAR estimates.

To explain this evidence, we use a standard DSGE model of an open economy with a Calvo price-setting mechanism, but add a new feature that the proportion of firms using PCP (rather than LCP) can vary between zero and one. To explore the potential of nominal rigidities with a mix of PCP and LCP to explain the evidence on pass-through to trade prices, we abstract from features (such as requirements of nontraded goods in distribution) that allow the markup to vary and be a function of the exchange rate. We incorporate both nominal and real shocks in the model.

A quantitative version of the model shows that reasonable values of the frequency of price adjustment and the proportion of PCP firms can explain the average values of pass-through elasticities for import and export price levels as well as the differences between these values. The model is also able to explain the similarity of the regression and VAR estimates. We also examine if the model can explain the variation of pass-through values across countries. The model predicts that the import price pass-through increases as the proportion of foreign varieties with PCP increases, and the export price pass-through decreases as the proportion of home varieties with PCP increases. We find support for this prediction for a subset of countries in our sample for which data on currency of invoicing are available.

[3] In addition to the endogeneity issues, the estimates are subject to omitted-variable and measurement-error biases since good measures of controls are not available and they are either omitted or represented by proxies.

II. Empirical Evidence

Let $P_{M,t}$ and $P_{X,t}$ denote the import and export price indexes for a country (expressed in home currency), and S_t the nominal exchange rate (an increase representing an appreciation of the foreign currency). Using lower-case letters to denote values in logs, and assuming that these series (in logs) are nonstationary and not cointegrated,[4] we can express the regression model typically used to estimate the exchange rate pass-through to the import or export price as

$$\Delta p_{T,t} = c + a\Delta s_t + \mathbf{b}\mathbf{g}_t + e_t, \quad T = M, X, \tag{1}$$

where $\mathbf{g}_t$ is a vector of variables (possibly including lagged values of $\Delta p_{T,t}$ and $\Delta \mathrm{s}_t$) included to control for the effect of certain factors and to introduce dynamics in the relation. This model is motivated by a partial-equilibrium framework, in which the exchange rate pass-through can be defined as the price response to a 1% change in the exchange rate by an exporter who takes certain factors (such as marginal cost and demand functions) as given, and may change prices gradually in the presence of price adjustment costs or other nominal frictions. The short-run pass-through defined in this way can be measured by estimating coefficient a in (1). The pass-through is sometimes estimated by a simple regression of $\Delta p_{T,t}$ on Δs_t (i.e., by estimating (1) without $\mathbf{g}_t$). The pass-through coefficient in this case can be interpreted as the price response that includes the direct as well as the indirect effect of the exchange rate (operating through variables in $\mathbf{g}_t$).

An alternative methodology uses a VAR model to estimate the exchange rate pass-through.[5] Letting P_t and P_t^* denote the home and foreign CPI (expressed in home and foreign currency, respectively), we can express the basic VAR models (with n lags) as

$$\mathbf{h}_t = \mathbf{c} + \mathbf{D}_1\mathbf{h}_{t-1} + \mathbf{D}_2\mathbf{h}_{t-2} + \ldots + \mathbf{D}_n\mathbf{h}_{t-n} + \mathbf{e}_t, \tag{2}$$

where $\mathbf{h}_t = \left[\Delta s_t, \Delta p_{T,t}, \Delta p_t, \Delta p_t^*\right]$ for $T = M, X$, lower-case letters represent values in logs, and $\mathbf{e}_t$ is a vector of reduced-form shocks.[6] Extended versions of this model include additional variables. Based on orthogonalized impulse response functions (with Δs_t as the

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[4] This is a typical characterization of these series, which is generally supported by time-series unit root and cointegration tests (e.g., Campa and Goldberg, 2005).

[5] See McCarthy (2000) for an early use of this methodology.

[6] The basic version assumes that variables $s_t, p_{T,t}, p_t, p_t^*$ are $I(1)$ and not cointegrated. A vector error correction model can be used instead if the variables are cointegrated.

first variable), the short-run pass-through to an import or export price index is defined as the (current period) elasticity of Δp_t to an innovation in Δs_t. The exchange rate innovation is generally a combination of different shocks, but can be identified as a particular shock under certain conditions.[7]

We use both regression and VAR models to estimate the exchange rate pass-through to import and export prices for a sample that includes all countries for which quarterly trade price data are available for at least 52 quarters since 1979. For the regression model, we use the simple regression (without $\mathbf{g}_t$) as it incorporates the indirect effects and is not sensitive to the choice of controls, which vary from one model to another and are difficult to measure. The regression results are presented in Table 1 for two sets of countries. The first set includes 18 advanced countries, and the second set of 16 countries largely includes emerging economies in our sample. For advanced countries, the pass-through coefficient for the import price is, with one exception, between zero and one, is significantly different from zero for all countries, and also significantly different from one for most countries. The estimates of the export price pass-through coefficient for these countries are similar: the coefficient is between zero and one for all countries but one, and is generally significantly different from both zero and one. The pass-through coefficient for the export price tends to be smaller than the coefficient for the import price. The average value of the pass-through coefficient is 0.39 for the export price and 0.67 for the import price. The average pass-through coefficients for the two prices are significantly different from each other. The results for emerging market countries show a similar pattern. Some notable differences for this group are that the pass-through coefficients are negative for two countries and the import price pass-through is smaller than the export price pass-through for two countries. The average values of the import price pass-through coefficients is 0.63 and that for the export price is 0.53.

Table 2 shows the estimates of exchange rate pass-through to import and export prices derived from the VAR model (2) for both advanced and emerging economies.[8] The VAR-based estimates are close to the estimates based on the simple regression (see Figure 1). For the advanced country group, the exchange rate pass-through to both price indexes is above zero and below one for all countries, and is significantly different from both zero and one for most countries (i.e., the two-standard deviation band for the pass-through estimate lies in the (0,1) interval). Moreover, the import price pass-through is larger than the export price pass-through for 17 out of 18 countries and this difference is significant (with the p-value of 0.1 or less) for 11 countries. The average value of the pass-through is 0.60 for the import price and 0.39 for the export price. The results for emerging countries show similar tendencies, but as

[7] For example, Choudhri, Faruqee and Hakura (2005) use the exclusion restriction that contemporaneous information on prices is not available to participants in the financial markets to identify the exchange rate innovation as the shock to the interest parity relation.

[8] We also experimented with a VAR that includes GDP as an additional variable, but this variation did not make much difference to the results.

in the case of the regression results, the pass-through coefficients are negative and the import price coefficient is smaller than the export-price coefficient in a few cases. For this group, the average values of the pass-through for the import and export prices are 0.54 and 0.51, respectively.

It has been suggested that the pass-through, especially to import or consumer price indexes has declined recently due to inflation stabilization (Gagnon and Ihrig, 2004) or trade integration resulting from lower trade costs (Gust, Leduc, and Vigfusson, 2010). To explore whether the exchange rate pass-through to trade prices has changed over time, we also obtained regression and VAR estimates for the sub-periods 1985-1997 and 1998-2010. The pass-through estimates for the subperiods are reported in Appendix Tables 1 and 2 (for the countries which had sufficient data to conduct the estimations). These estimates are less precise, and show no clear pattern for change between the two periods. The estimates of the pass-through for both price indexes, for example, are higher for some countries but lower for others in the second relative to the first sub-period.

III. The Model

In this section, we develop a DSGE model to explain the evidence on the exchange rate pass-through to trade prices. We design our model for explaining the results for the advanced countries whose financial markets and monetary policies can be modeled by an interest rate parity relation and a Taylor-type interest rate rule. Our model, however, would also be relevant for a range of emerging countries with similar financial markets and monetary policies. To examine how well a staggered price model with a combination of PCP and LCP can by itself explain the pass-through results, we consider a simple framework without nontraded goods or distribution services and a CES index for aggregating varieties. This setup excludes adjustment mechanisms (e.g., via markup variation or distribution channels), which have been emphasized by alternative explanations of the pass-through evidence. In our basic setup, we assume that all traded goods and labor services are differentiated. The adjustment of wages and prices is based on the Calvo model. A novel feature of our model is Calvo price adjustment with a mix of PCP and LCP, and the discussion below focuses on the implications of this feature. The rest of the model is standard and is described briefly.

A. Basic Setup

For a continuum of households indexed by $i \in [0,1]$, preferences are given by

$$U_t(i) = E_t \sum\nolimits_{k=t}^{\infty} \beta^k \left(\frac{C_k(i)^{1-\chi}}{1-\chi} + \varphi \frac{L_k(i)^{1+\mu}}{1+\mu} \right), \tag{3}$$

where $C_t(i)$ and $L_t(i)$ represent the household's aggregate real consumption and labor supply. Households hold one-period domestic and foreign bonds, and holding of foreign bonds (used for international borrowing or lending) is subject to transaction cost shocks. Their budget constraint is

$$B_{t+1}(i)+S_t B^*_{t+1}(i)=(1+R_{t-1})B_t(i)+S_t(1+R^*_{t-1})B^*_t(i)X_{TC,t-1} \\ +W_t(i)L_t(i)+PF_t(i)-P_tC_t(i) \quad , \tag{4}$$

where $B_t(i)$ and $B^*_t(i)$ represent home and foreign bonds held at the beginning of period *t*; S_t is the exchange rate; R_{t-1} and R^*_{t-1} are the home and foreign interest rates, and $X_{TC,t-1}$ is the transaction cost in period $t-1$; P_t denotes the price level; $PF_t(i)$ is the household's shares of total profits, and $W_t(i)$ is the wage rate set by the household .

Optimization of utility in (3) subject to the budget constraint (4) implies the following standard conditions in symmetric equilibrium (where household index is dropped for simplicity):

$$\beta\left(\frac{E_tC_{t+1}}{C_t}\right)^{-\chi}\left(\frac{P_t}{E_tP_{t+1}}\right)=\frac{1}{1+R_t}, \tag{5}$$

$$\frac{S_t}{E_tS_{t+1}}=\frac{(1+R^*_t)X_{TC,t}}{1+R_t}, \tag{6}$$

where $X_{TC,t}$ can be interpreted as the shock to the interest parity relation.[9] The central bank sets the interest rate by the following simple rule that reacts to only the inflation rate:

$$1+R_t=(1+\bar{R})(\Pi_t/\bar{\Pi})^{\delta}X_{R,t}, \quad \delta>0, \tag{7}$$

where $\bar{R}$ is the steady-state interest rate and $\bar{\Pi}$ is the target value of $\Pi_t=P_t/P_{t-1}$ and $X_{R,t}$ is the shock to the monetary rule. Using (7) and its foreign counterpart to substitute for the home and foreign (gross) interest rates in (6), we can express the exchange rate as

$$S_t=\frac{E_tS_{t+1}(1+\bar{R}^*)(\Pi^*_t/\bar{\Pi}^*)^{\delta^*}X_t}{(1+\bar{R})(\Pi_t/\bar{\Pi})^{\delta}}, \tag{8}$$

where $X_t=X^*_{R,t}X_{TC,t}/X_{R,t}$ is a composite nominal shock.

[9] The shock to interest parity relation could also be motivated by Devereux and Engel's (2002) noise-trader model, where there is a stochastic bias in the expectations of foreign exchange dealers.

Aggregate consumption is defined as

$$C_t = \left[(1-\gamma)^{1/\eta}(C_{H,t})^{(\eta-1)/\eta} + \gamma^{1/\eta}(C_{F,t})^{(\eta-1)/\eta}\right]^{\eta/(\eta-1)}, \tag{9}$$

where $C_{H,t}$ and $C_{F,t}$ are bundles of home and foreign varieties. Assume a continuum of home firms indexed by $j \in [0,1]$, and foreign firms indexed by $j^* \in [0,1]$, and define these bundles as

$$C_{H,t} = \left(\int_0^1 C_{H,t}(j)^{(\varepsilon-1)/\varepsilon} dj\right)^{\varepsilon/(\varepsilon-1)}, \quad C_{F,t} = \left(\int_0^1 C_{F,t}(j^*)^{(\varepsilon-1)/\varepsilon} dj^*\right)^{\varepsilon/(\varepsilon-1)}, \tag{10}$$

where $C_{H,t}(j)$ and $C_{F,t}(j^*)$ represent the consumption amounts for a home and a foreign variety. Let $P_{H,t}(j)$ and $P_{M,t}(j^*)$ denote the prices of a home and an imported (foreign) variety. The demand functions for the home and foreign bundles and the varieties within each bundle can then be derived as follows:

$$C_{H,t} = (1-\gamma)\left(\frac{P_{H,t}}{P_t}\right)^{-\eta} C_t, \quad C_{F,t} = \gamma\left(\frac{P_{M,t}}{P_t}\right)^{-\eta} C_t, \tag{11}$$

$$C_{H,t}(j) = \left(\frac{P_{H,t}(j)}{P_{H,t}}\right)^{-\varepsilon} C_{H,t}, \quad C_{F,t}(j^*) = \left(\frac{P_{M,t}(j^*)}{P_{M,t}}\right)^{-\varepsilon} C_{F,t}, \tag{12}$$

with price indexes defined as

$$P_t = \left[(1-\gamma)(P_{H,t})^{1-\eta} + \gamma(P_{M,t})^{1-\eta}\right]^{1/(1-\eta)}, \tag{13}$$

$$P_{H,t} = \left[\int_0^1 P_{H,t}(j)^{1-\varepsilon} dj\right]^{1/1-\varepsilon}, \quad P_{M,t} = \left[\int_0^1 P_{M,t}(j^*)^{1-\varepsilon} dj^*\right]^{1/1-\varepsilon}. \tag{14}$$

The home firm supplies its variety to the home and the foreign market and thus it output equals

$$Y_t(j) = C_{H,t}(j) + C^*_{H,t}(j), \tag{15}$$

where $C^*_{H,t}(j)$ is the foreign consumption of the home variety. Assuming that foreign consumption indexes are analogous to home indexes, we derive the foreign demand for the variety as

$$C_{H,t}^*(j) = \left(\frac{P_{X,t}^*(j)}{P_{X,t}^*} \right)^{-\varepsilon} C_{H,t}^*, \quad C_{H,t}^* = (1-\gamma^*)\left(\frac{P_{X,t}^*}{P_t^*} \right)^{-\eta} C_t^*, \tag{16}$$

where $P_{X,t}^*(j)$ is the export price of the home variety expressed in foreign currency. Price indexes $P_{X,t}^*$ and P_t^* are given by

$$P_{X,t}^* = \left[\int_0^1 P_{X,t}^*(j)^{1-\varepsilon} dj \right]^{1/1-\varepsilon}, \quad P_t^* = \left[(1-\gamma^*)(P_{F,t}^*)^{1-\eta} + \gamma(P_{X,t}^*)^{1-\eta} \right]^{1/(1-\eta)}, \tag{17}$$

where $P_{F,t}^*$ is the price of the bundle of foreign varieties in the foreign country.

The production function for a home variety is

$$Y_t(j) = A_t L_t(j), \tag{18}$$

where A_t is a stochastic productivity index that represents a real shock for the home economy; and $L_t(j)$ is a bundle of labor services defined as

$$L_t(j) = \left(\int_0^1 L_t(i,j)^{(\varepsilon-1)/\varepsilon} di \right)^{\varepsilon/(\varepsilon-1)}. \tag{19}$$

The marginal cost is the same for all home firms and equals

$$MC_t = W_t / A_t, \tag{20}$$

where $W_t = \left[\int_0^1 W(i)^{1-\varepsilon} di \right]^{1/1-\varepsilon}$ is the wage index for the bundle.

Each home firm sets the home and export prices for its variety. The export price can be set either in producer or local currency. Let ψ and $1-\psi$ represent, respectively, the proportion of home firms setting prices using PCP and LCP. Partition the unit interval such that for $j \in [0,\psi]$, export price, $P_{XP,t}(j)$, is set in home currency, and for $j \in [\psi,1]$, export price, $P_{XL,t}^*(j)$, is set in foreign currency. The export price index expressed in home currency can then be defined as

$$P_{X,t} = \left[\int_0^\psi P_{XP,t}(j)^{1-\varepsilon} dj + S_t \int_\psi^1 P_{XL,t}^*(j)^{1-\varepsilon} dj \right]^{1/(1-\varepsilon)}. \tag{21}$$

Note that the foreign-currency export price in (17) can be converted into the home-currency export price in (21) as $P_{X,t} = S_t P^*_{X,t}$.

We allow the frequency of price change to differ between the home and foreign countries, and assume that it depends on the currency in which the price is set. In each period, there is probability, $1-\theta$, that a firm will change its price set in home currency, and probability, $1-\theta^*$, that it will change its price set in foreign currency. Let $\tilde{P}_{H,t}(j)$, $\tilde{P}_{XP,t}(j)$ and $\tilde{P}^*_{XL,t}$ denote, respectively, the new prices for home sales, for exports under PCP, and for exports under LCP if a firm sets new prices in period, t. The firm chooses these prices to maximize $Z_t(j) = \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k}\left[PF_{H,k}(j) + PF_{X,k}(j)\right]$, where $PF_{H,\tau}(j)$ and $PF_{X,\tau}(j)$ are profits from home sales and exports, and $D_{t,k}$ is the discount factor. In view of the demand functions for $C_{H,t}(j)$ and $C^*_{H,t}(j)$ in (12) and (16), $PF_{H,\tau}(j)$ equals $(\tilde{P}_{H,t}(j) - MC_\tau)C_{H,\tau}(\tilde{P}_{H,t}(j)/P_{H,\tau})^{-\varepsilon}$ while $PF_{X,\tau}(j)$ equals $(\tilde{P}_{XP,t}(j) - MC_\tau)C^*_{H,\tau}(\tilde{P}_{XP,t}(j)/P_{X,\tau})^{-\varepsilon}$ for PCP and $S_\tau(\tilde{P}^*_{XL,t}(j) - MC_\tau / S_\tau)C^*_{H,\tau}(\tilde{P}^*_{XL.t}(j)/P^*_{X,\tau})^{-\varepsilon}$ for LCP. The optimal new prices can be shown to equal (for all firms changing prices in period t)

$$\tilde{P}_{H,t} = \frac{\varepsilon E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C_{H,k} P^{\varepsilon}_{H,k} MC_k}{(\varepsilon-1) E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C_{H,k} P^{\varepsilon}_{H,k}}, \tag{22}$$

$$\tilde{P}_{XP,t} = \frac{\varepsilon E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C^*_{H,k} P^{\varepsilon}_{X,k} MC_k}{(\varepsilon-1) E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C^*_{H,k} P^{\varepsilon}_{X,k}}, \quad \tilde{P}^*_{XL.t} = \frac{\varepsilon E_t \sum_{k=t}^{\infty} \theta^{*\tau-t} D_{t,k} C^*_{H,k} P^{*\varepsilon}_{X,k} MC_k}{(\varepsilon-1) E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C^*_{H,k} S_k P^{*\varepsilon}_{X,k}}. \tag{23}$$

The home price index [as defined in (14)] equals $P_{H,t} = \left[\sum_{k=0}^{\infty} (1-\theta)\theta^k \tilde{P}_{H,t-k}{}^{1-\varepsilon}\right]^{1/(1-\varepsilon)}$ and can be expressed as

$$P_{H,t} = \left[(1-\theta)\tilde{P}_{H,t}{}^{1-\varepsilon} + \theta P_{H,t-1}{}^{1-\varepsilon}\right]^{1/(1-\varepsilon)}. \tag{24}$$

Similarly, the export price defined in (21) can be written as

$$P_{X,t} = \left[\psi P^{1-\varepsilon}_{XP,t} + (1-\psi)(S_t P^*_{XL,t})^{1-\varepsilon}\right]^{1/(1-\varepsilon)}, \tag{25}$$

where $P^{1-\varepsilon}_{XP,t} = \sum_{k=0}^{\infty} (1-\theta)\theta^k (\tilde{P}_{XP,t-k})^{1-\varepsilon}$ and $P^{*1-\varepsilon}_{XL,t} = \sum_{k=0}^{\infty} (1-\theta^*)\theta^{*k} (\tilde{P}^*_{XL,t-k})^{1-\varepsilon}$. The producer- and local-currency components of the export price can also be expressed as

$$P_{XP,t} = \left[(1-\theta)(\tilde{P}_{XP,t})^{1-\varepsilon} + \theta(P_{XP,t-1})^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}, P^*_{XL,t} = \left[(1-\theta^*)(\tilde{P}^*_{XL,t})^{1-\varepsilon} + \theta^*(P^*_{XL,t-1})^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}. \tag{26}$$

Assume that ψ^* and $1-\psi^*$ proportions of foreign firms set prices using PCP and LCP, and let $\tilde{P}^*_{MP,t}(j^*)$ and $\tilde{P}_{ML,t}(j^*)$ denote the new prices for home imports under PCP and LCP. Denoting the foreign marginal cost by MC^*, we can then derive analogous relations for determining the import price index (expressed in home currency) as follows:

$$\tilde{P}^*_{MP.t} = \frac{\varepsilon E_t \sum_{k=t}^{\infty} \theta^{*k-t} D_{t,k} C_{F,k} P^{*\varepsilon}_{M,k} MC^*_k}{(\varepsilon-1) E_t \sum_{k=t}^{\infty} \theta^{*k-t} D_{t,k} C_{F,k} P^{*\varepsilon}_{M,k}}, \quad \tilde{P}_{ML,t} = \frac{\varepsilon E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C_{F,k} P^{\varepsilon}_{M,k} MC^*_k}{(\varepsilon-1) E_t \sum_{k=t}^{\infty} \theta^{k-t} D_{t,k} C_{F,k} P^{\varepsilon}_{M,k} / S_k}, \tag{27}$$

$$P_{M,t} = \left[\psi^* (S_t P^*_{MP,t})^{1-\varepsilon} + (1-\psi^*) P^{1-\varepsilon}_{ML,t} \right]^{1/(1-\varepsilon)}, \tag{28}$$

$$P^*_{MP,t} = \left[(1-\theta^*)(\tilde{P}^*_{MP,t})^{1-\varepsilon} + \theta^* (P^*_{MP,t-1})^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}, P_{ML,t} = \left[(1-\theta)(\tilde{P}_{ML,t})^{1-\varepsilon} + \theta (P_{ML,t-1})^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}. \tag{29}$$

The demand for labor is affected by the price dispersion under the Calvo model. To derive labor demand, first use (15) and (18) to obtain $L_t = \frac{1}{A_t} \int_0^1 \left(C_{H,t}(j) + C^*_{H,t}(j) \right) dj$, and then use Calvo staggered price adjustment to express this relation as

$$L_t = \frac{1}{A_t} \begin{bmatrix} C_{H,t} \sum_{k=0}^{\infty} (1-\theta)\theta^k (\tilde{P}_{H,t-k} / P_{H,t})^{-\varepsilon} \\ + C^*_{H,t} \sum_{k=0}^{\infty} (1-\theta)\theta^k \left(\psi (\tilde{P}_{XP,t-k} / P_{X,t})^{-\varepsilon} + (1-\psi)(\tilde{P}^*_{XL,t-k} / P^*_{X,t})^{-\varepsilon} \right) \end{bmatrix}. \tag{30}$$

In each period, there is probability, $1-\theta_W$, that a household will change its wage. Let $\tilde{W}_t(i)$ denote the new wage for a household setting a new wage in period, t. The household chooses the wage to maximize lifetime utility (3) subject to the budget constraint (4) and labor demand, $L_t(i) = L_t(W_t(i)/W_t)^{-\varepsilon}$. This wage equals

$$[\tilde{W}_t(i)]^{\varepsilon\mu+1} = \frac{\varphi\varepsilon \sum_{k=t}^{\infty} (\theta_W \beta)^{k-t} (L_k)^{1+\mu} W_k^{\varepsilon(1+\mu)}}{(\varepsilon-1) \sum_{k=t}^{\infty} (\theta_W \beta)^{k-t} \lambda_k L_k W_k^{\varepsilon}}. \tag{31}$$

The wage index [defined in (20)] is then determined as

$$W_t = \left[(1-\theta_W) \tilde{W}_t^{1-\varepsilon} + \theta_W W_{t-1}^{1-\varepsilon} \right]^{1/(1-\varepsilon)}. \tag{32}$$

The shocks in the model are assumed to follow the following autoregressive processes:

$$X_t = \left[1-\kappa\left(S_t B_t^* / P_t\right)^2\right] X_{t-1}^{\rho_X} e^{v_{X,t}}, \quad A_t = \bar{A}^{1-\rho_A} A_{t-1}^{\rho_A} e^{v_{A,t}}, \tag{33}$$

where $\kappa > 0$, $0 < \rho_X < 1$, $0 < \rho_A < 1$, $\bar{A}$ is the steady state value of the productivity index, and $v_{X,t}$ and $v_{A,t}$ are white noise disturbances. Note that the presence of the expression, $\kappa\left(S_t B_t^* / P_t\right)^2$ in the nominal shock process ensures convergence to a unique steady state with zero net foreign real assets ($S_t B_t^* / P = 0$).

B. Key Relations

A quantitative version of the model is analyzed in the next section. Here, we briefly discuss key relations that determine the exchange rate pass-through to trade prices. We simplify these relations by using a log-linear approximation around the steady state. Letting lower-case letters denote values in logs, express the linear versions of (28), (29) and (27) as

$$\Delta p_{M,t} = \psi^* (\Delta p_{MP,t}^* + \Delta s_t) + (1-\psi^*)\Delta p_{ML,t}, \tag{34}$$

$$\Delta p_{MP,t}^* = (1-\theta^*)(\tilde{p}_{MP,t}^* - p_{MP,t-1}^*), \quad \Delta p_{ML,t} = (1-\theta)(\tilde{p}_{ML,t} - p_{ML,t-1}), \tag{35}$$

$$\begin{aligned} \tilde{p}_{MP,t}^* &= \mu + (1-\beta\theta^*)\sum\nolimits_{k=0}^{\infty} (\beta\theta^*)^k E_t mc_{t+k}^*, \\ \tilde{p}_{ML,t} &= \mu + (1-\beta\theta)\sum\nolimits_{k=0}^{\infty} (\beta\theta)^k E_t (mc_{t+k}^* + s_{t+k}), \end{aligned} \tag{36}$$

where $\mu = \ln\dfrac{\varepsilon}{\varepsilon-1}$ is the markup in logs. Next use (36) to express (35) as[10]

$$\Delta p_{MP,t}^* = \beta E_t \Delta p_{MP,t+1}^* + \lambda^* \tilde{mc}_{P,t}^*, \quad \Delta p_{ML,t} = \beta E_t \Delta p_{ML,t+1} + \lambda \tilde{mc}_{L,t}^*, \tag{37}$$

where $\lambda^* = \dfrac{(1-\theta^*)(1-\beta\theta^*)}{\theta^*}$, $\lambda = \dfrac{(1-\theta)(1-\beta\theta)}{\theta}$, and $\tilde{mc}_{P,t}^* = w_t^* - a_t^* - p_{MP,t}^* + \mu$ and $\tilde{mc}_{L,t}^* = w_t^* - a_t^* - (p_{ML,t} - s_t) + \mu$ are indexes based on real marginal costs for firms using PCP and LCP. Finally, substituting the values of $\Delta p_{MP,t}^*$ and $\Delta p_{ML,t}$ in (34), we obtain

$$\Delta p_{M,t} = \beta E_t \Delta p_{M,t+1} - \psi^* \beta E_t \Delta s_{t+1} + \psi^* \lambda^* \tilde{mc}_{P,t}^* + (1-\psi^*)\lambda \tilde{mc}_{L,t}^* + \psi^* \Delta s_t. \tag{38}$$

Similarly, we can use linear versions of (23), (25) and (26) to derive

[10] This derivation is standard for a Calvo model of price adjustment (e.g., see Yun, 1996).

$$\Delta p_{XP,t} = \beta E_t \Delta p_{XP,t+1} + \lambda \tilde{mc}_{P,t}, \quad \Delta p^*_{XL,t} = \beta E_t \Delta p^*_{XL,t+1} + \lambda^* \tilde{mc}_{L,t}, \tag{39}$$

$$\Delta p_{X,t} = \beta E_t \Delta p_{X,t+1} - (1-\psi)\beta E_t \Delta s_{t+1} + \psi\lambda \tilde{mc}_{P,t} + (1-\psi)\lambda^* \tilde{mc}_{L,t} + (1-\psi)\Delta s_t, \tag{40}$$

where $\tilde{mc}_{P,t} = w_t - a_t - p_{XP,t} + \mu$, and $\tilde{mc}_{L,t} = w_t - a_t - (p^*_{XL,t} + s_t) + \mu$.

Relations (38) and (40) identify the key channels that transmit the effects of different shocks to log differences of import and export prices. Note that the partial effect of the log difference of the exchange rate depends on the proportions of firms using PCP in the home and foreign economies, ψ and ψ^*. An increase in ψ^* would strengthen the partial effect of the exchange rate on the import prices while an increase in ψ would weaken the partial effect on the export price.

To examine the behavior of the exchange rate, we linearize (8) to obtain

$$s_t = E_t s_{t+1} + \bar{r}^* - \bar{r} + \delta^*(E_t \pi^*_t - \bar{\pi}^*) - \delta(E_t \pi_t - \bar{\pi}) + x_t \tag{41}$$

where $\bar{r}^* = \ln(1+\bar{R}^*)$, $\bar{r} = \ln(1+\bar{R})$, $\pi^*_t = \ln(\Pi^*_t)$, $\bar{\pi}^* = \ln(\bar{\Pi}^*)$, $\pi_t = \ln(\Pi_t)$, and $\bar{\pi} = \ln(\bar{\Pi})$

From (33), the linear versions of the time series processes for the stochastic variables are: $x_t = \rho_X x_{t-1} + v_{X,t}, a_t = \rho_A a_{t-1} + v_{A,t}$, and $a^*_t = \rho_{A*} a^*_{t-1} + v_{A*,t}$. The pass-through elasticities for the import and export prices depend on the impact of nominal and real disturbances, $v_{X,t}, v_{A,t}, v_{A*,t}$, on s_t (and hence Δs_t) through (41) and on $\Delta p_{M,t}$ and $\Delta p_{X,t}$ via (38) and (40). The presence of the expected future values in these relations magnifies this impact (given the persistence in the stochastic variables). Both the nominal and real disturbances affect the exchange rate and trade prices, but the presence of x_t in (41) and Δs_t in (38) and (40) suggest that the impact of the nominal disturbance would be stronger. The relative strength of the effect of nominal and real disturbances is examined in the next section.

IV. Quantitative Analysis

A. Key Determinants of the Pass-Through to Trade Prices

Our model suggests that the direct and indirect effects of the exchange rate on trade prices depend on the shares of home and foreign firms engaged in PCP (ψ, ψ^*) and the frequency of price and wage change (θ, θ_W). These parameters could play an important role in determining the degree of pass-through. Moreover, the persistence of exchange rate shocks (as reflected in ρ_x) and the interest rate response to inflation in the home and foreign economies (δ, δ^*) influence the effect of nominal shocks on the exchange rate, and these parameters could also be significant determinants of the pass-through. We use numerical analysis of the model to explore the influence of these parameters on the pass-through to trade prices.

We calibrate our model as follows. We normalize the initial prices and home income to equal one. We set the share of imports in income (γ) equal to the median import share for advanced countries in our sample, which equals 0.3 (and is slightly lower than the average share of 0.32).We let the quarterly discount factor (β) equal 0.99, the coefficient of risk aversion (χ) equal 2.0, and the elasticity of labor supply ($1/\nu$) equal 0.5. We choose a value of 2.0 for the elasticity of substitution between home and foreign goods (η), and a value of 6.0 for the elasticity of substitution between varieties (ε). The values of these parameters are similar to the ones used by other studies. Our sensitivity analysis indicates, moreover, that variations in these values have little effect on the pass-through results. For productivity shocks, we follow Corsetti, Dedola, and Leduc (2008) and let the autoregressive coefficients (ρ_A and ρ_{A^*}) equal to 0.95, standard errors of white noise disturbances (v_A and v_{A^*}) equal to 0.007, and the correlation coefficient for these disturbances equal to 0.25. For the nominal shock, we initially also let the standard error of the white noise disturbance (v_X) equal 0.007, but assume a more moderate persistence (let $\rho_X = 0.8$). For parameters that are expected to have a significant influence on the pass-through, we start with plausible baseline values and later consider the effect of variations in these values. The baseline values chosen for these parameters are $\psi = \psi^* = 0.5$, $\theta = \theta_W = 0.75$, and $\delta = \delta^* = 0.5$. We assume that the home economy is small and treat the foreign (rest of the world) real wage and inflation (W^*/P^* and Π^*) as exogenous.[11]

To examine the model dynamics underlying the relation between the exchange rate and trade prices, Figure 2 shows the dynamic response (over 20 quarters) of the log differences of the exchange rate and import and export prices ($\Delta s, \Delta p_M, \Delta p_X$) to a one-standard-deviation increase in each shock (v_X, v_A, v_{A^*}) . The figure illustrates two important features of the dynamic adjustment of these variables. First, the effect of each shock is concentrated in the first quarter. Second, the first-quarter effect of the nominal shock is much larger than that of real shocks for each variable. As discussed further below, these features suggest that the exchange rate innovations in the VAR model would be dominated by the nominal shock. Moreover, as changes in the exchange rate and trade prices largely reflect the effect of

[11] We also assume a very small value for the transactions cost parameter (κ), which ensures convergence of the economy to a unique steady state with zero net foreign assets but has a negligible effect on the dynamics of the model.

current rather than past innovations, the regression estimates of the exchange rate pass-through would be similar to the VAR estimates.

To explore how different shocks may affect the short-run pass-through, Table 3 shows the current-period elasticity of the import and export prices to the exchange rate resulting from a one-standard-deviation increase in each shock, one at a time. For the nominal shock, the pass-through elasticity is 0.57 for the import price and 0.40 for the export price in the baseline case. For the real shocks, the pass-through elasticity is very different and negative for either import or export price: for example, the elasticity is -0.22 for the export price in the case of home productivity shock, and is -0.17 for the import price in case of the foreign productivity shock. The table also shows the elasticities generated by a simultaneous one-standard-deviation increase in all three shocks. As we would expect from the impulse response functions in Figure 2 (which show the nominal shock to dominate the movements in the exchange rate and trade prices), the elasticity for the composite shock is not much different than the elasticity for only the nominal shock. The results in Table 3 suggest that while the pass-through is very different for nominal and real shocks, its value is likely to be largely determined by nominal shocks in the presence of both types of shocks.

We next examine the sensitivity of pass-through elasticities to variations in the baseline values of key parameters. Given the importance of the nominal shock in influencing the pass-through, we focus on the elasticities determined by this shock. Figure 3(a) illustrates the effect of variations in the foreign PCP share on the import price elasticity, and Figure 3(b) illustrates the effect of variations in the home PCP share on the export price elasticity.[12] As Figure 3(a) shows, the import price pass-through is low (below 0.2) if the foreign PCP share is zero. The pass-through increases as the PCP share increases and reaches one when the PCP share equals one. In contrast, the export price pass-through in Figure 3(b) is high (above 0.8) at the zero value of the home PCP share, decreases as PCP share increases and falls to a low value (below 0.1) when the PCP share equals one. These figures suggest that the pass-through for the import price would exceed that for the export price if both home and foreign PCP shares are greater than a value just below 0. 5.

The influence of other parameters on the pass-through elasticities is explored in Table 4 for the case of a one-standard deviation nominal shock. We first examine the effect of smaller wage-price stickiness. We increase θ_W to 0.667 and θ to 0.5, as suggested by estimates based on US data.[13] Higher frequency of wage and price change has a significant impact on

[12] Home PCP share is kept at the baseline value of 0.5 in Figure 3(a). Similarly, the foreign PCP share equals 0.5 in Figure 3(b). However, the relations in these figures would essentially remain the same in the symmetric case where the home PCP share equals the foreign PCP share.

[13] Our baseline values of the wage-price stickiness parameters imply that the average duration of fixed wage or price equals one year, which is not too long in view of estimates for the Euro area (Smets and Wouters, 2003).

(continued…)

the pass-through: the import price pass-through increases from 0.57 to 0.73 while the export price pass-through decreases form 0.41 to 0.33. The effect of greater wage-price flexibility is thus similar to increasing home and foreign PCP shares. We also explore how the degree of persistence in the exchange rate shock and the interest rate reaction to inflation influences the exchange rate pass-through. We varied ρ_x from 0.6 to 0.95 and δ from 0.1 to 1.0. The table shows that the effect of the lower or higher persistence of the exchange rate shock or less or more aggressive reaction to inflation does not have a large effect on the pass-through values. Thus, our numerical analysis of the short-run pass-through generated by the nominal shock suggests that PCP shares and wage-price stickiness are the major source of variation in the pass-through for import and export prices.

Finally, we evaluate the ability of the model to match the regression and VAR estimates of the pass-through and to explain certain empirical regularities concerning the exchange rate and inflation behavior. We focus on regularities about exchange rate volatility and inflation variability, which have received considerable attention. We examine the performance of the model in matching the data for advanced countries that are more likely to conform to the interest parity and Taylor rule assumptions. Stochastic simulation of the model is used to generate series on $\Delta s_t, \Delta p_{M,t}, \Delta p_{X,t}$ and Δp_t for a period as long as our sample (127 quarters).[14] The simulated data were used to compute the standard deviations of Δs_t and Δp_t, and to estimate the pass-through coefficients using both the simple regression and VAR models.[15] Table 5 first presents the results for the baseline values of the parameters. As compared to the average values for advanced countries, the baseline model yields estimates of the pass-through coefficients that are not too different from the average estimates for advanced countries. Interestingly, the regression and VAR estimates derived from the model are very close to each other. The baseline model implies greater exchange rate variability and smaller inflation variability than in the data. However, the performance of the model in matching the variability and pass-through statistics can be improved by simply adjusting the stochastic process for the nominal shock and the PCP shares. For example, Variant 1 of the model decreases the standard deviation of v_X from 0.007 to 0.0025, increases ρ_X from 0.8 to 0.95, increases ψ from 0.5 to 0.55, and decrease ψ^* from 0.5 to 0.45. This variant produces values for the standard deviations and the pass-through coefficients that match the values from data very well.

Estimates for the United States by Christiano, Eichenbaum, and Evans (2005) suggest that the average duration of a contract is 3 quarters for wages and 2 quarters for prices (i.e. $1-\theta_W = 1/3$ and $1-\theta = 1/2$).

[14] The series were, in fact, generated for 227 quarters, but the values for the first 100 quarters were dropped.

[15] The VAR model used for the simulated data is the same as that used for actual data except that it does not include Δp_t^*, which is constant in our model. We did experiment with introducing a shock to this variable in the model, but it did not affect the results much.

To examine the relative importance of nominal and real shocks in determining the different statistics in the table, two further variants are considered. Variant 2 represents the case where the economy is subject to only nominal shocks, and is the same as Variant 1 except that the standard deviations of v_A and v_{A*} are set equal to zero. The absence of real shocks does not much affect either the standard deviations or the pass-through coefficients. In Variant 3, the economy faces only real shocks, and the standard deviation of v_X is set equal to zero in this variant while keeping all other parameter values the same as Variant 1. The absence of nominal shock leads to major differences in results. The standard deviations for both the exchange rate appreciation and inflation decrease substantially. Moreover, the VAR pass-though estimates are very small and much lower than the regression estimates. Thus the model does not produce sufficient exchange rate and inflation variability or account for the similarity of the regression and VAR estimates without the presence of nominal shocks.

The effects of changes in home and foreign PCP shares are examined in the last two variants in the table. Variant 4 increases ψ to 0.8 and lowers ψ^* to 0.2 and represents the case of a country (like the United States) with high home PCP share and a low foreign PCP share. As expected, this change lowers both import and export price pass-through coefficients. The decrease in the coefficients is accompanied by a relatively small decrease in the exchange rate and inflation variability. Variant 5 represents the opposite case of a low home and a high foreign PCP share ($\psi = 0.2, \psi^* = 0.8$). Both the pass-through and variability measures increase in this case. A number of studies have noted that inflation or exchange rate stability is associated with low pass-through coefficients. The above simulations suggest that such association can result from cross-country differences in home and foreign PCP shares.

B. Currency of Invoicing and the Pass-Through

Our numerical analysis, based on both impulse response functions and simulated time series, suggests that wage-price stickiness, the mix of nominal and real shocks and the home and foreign PCP shares are key determinants of the pass-through to import and export prices. As currency of invoicing data can be used to measure of PCP shares, we briefly explore in this section some empirical evidence on the importance of these shares in explaining cross-country variation in the pass-through values. We utilize data on currency of invoicing from Kamps (2006). This source provides data on shares of invoicing in exporter's currency and in US dollars for both imports and exports. The data are available only for a few years for a subset of our sample consisting of 15 countries. Another limitation of the data is that the coverage and the collection method vary from one source to another. Nevertheless, this data provide rough measures of the PCP shares in home and foreign economies and enable us to test one implication of the model.

Table 6 presents results of OLS regressions explaining VAR estimates of the import and export pass-through by an index of the home or foreign PCP share based on invoicing-currency data. First, we examine the influence of the share of exporter currency invoicing in exports as a measure of the home PCP share. Consistent with the model, this measure has a negative and significant effect on the export price pass-through. Next, we explore the effect of the share of exporter currency invoicing in imports as a measure of the foreign PCP share. This index has a positive effect on the import price pass-through (as predicted by the model), but this effect is insignificant. One problem with this measure of the foreign PCP is that it includes invoicing in Euro, which is also the home currency for a large number of countries in our sample.[16] As an alternative measure of the foreign PCP, we use the share of US dollar invoicing in imports (for US, we use one minus the US dollar share).[17] The US dollar measure is found to have a positive and significant effect on the import price pass-through. The empirical evidence based on currency invoicing thus lends support to the model's predictions about the effect of PCP shares on the pass-through to the export and import prices.

V. Concluding Remarks

The debate on the nature of nominal rigidities in open economies has focused on whether firms use PCP or LCP. In view of the limitations of either PCP or LCP to account for the evidence on exchange rate pass-through to import and export prices, recent studies have emphasized the role of markup adjustment in explaining the pass-through evidence. This paper argues that even with no markup adjustment, a hybrid model with an appropriate mix of PCP and LCP can fit the data on not only pass-through elasticities but also measures of exchange rate and inflation variability. The model also accounts for the finding that pass-through estimates based on regression equations tend to be similar to the ones derived from VAR models

Studies based on models with variable markup (e.g., Corsetti, Dedola, and Leduc, 2008) can explain low pass-through elasticities and exchange rate volatility with minimal nominal rigidities and without the presence of nominal shocks. These studies, however, do not address the question of why regression estimates of the pass-through elasticities tend to be close to VAR estimates. Our model explains the similarity of the two types of estimates, and interestingly, nominal rigidities play a major role in this explanation. We show that in the presence of significant wage-price stickiness, short run changes in the exchange rate and trade prices are determined largely by current innovations to shocks, and thus regression and VAR estimates of the pass-through do not differ much. We also show that the presence of nominal shocks is essential in our model not only in explaining this result, but also in accounting for the pattern of pass-through elasticities for import and export prices and the observed exchange rate and inflation variability.

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[16] The currency invoicing data are available for selected years in early 2000's after the monetary union in Europe.

[17] In addition to most imports from US, a large proportion of imports from non-US countries is also invoiced in US dollars (Goldberg and Tille, 2005).

As well, the model is consistent with the evidence that a depreciation of the home currency would worsen the terms of trade, and thus changes in the exchange rate would bring about appropriate relative price adjustment in the transmission of shocks. The benefits of exchange rate flexibility would, therefore, be realized even in the hybrid case where both PCP and LCP are used.

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Table 1. OLS Regressions: Impact of Exchange Rate Changes on Trade Prices, 1979-2010

	Import Price Pass-Through			Export Price Pass-Through		
	Log-change of NEER	No. of Obs.	R-squared	Log-change of NEER	No. of Obs.	R-squared
Advanced economies						
United States	0.387 (0.107)***###	113	0.164	0.152 (0.090)*###	109	0.063
United Kingdom	0.350 (0.052)***###	127	0.310	0.210 (0.066)***###	127	0.104
Belgium	0.622 (0.197)***#	71	0.135	0.548 (0.146)***###	71	0.196
Denmark	0.953 (0.133)***	127	0.452	0.592 (0.097)***###	127	0.297
France	0.554 (0.205)***##	120	0.093	0.300 (0.126)**###	120	0.072
Germany	0.686 (0.117)***###	127	0.247	0.202 (0.036)***###	127	0.210
Italy	0.748 (0.121)***##	126	0.228	0.427 (0.084)***###	126	0.203
Netherlands	1.045 (0.223)***	111	0.148	0.821 (0.235)***	111	0.102
Norway	0.491 (0.076)***###	127	0.203	-0.233 (0.266)###	127	0.007
Sweden	0.402 (0.091)***###	127	0.267	0.303 (0.044)***###	127	0.304
Switzerland	0.633 (0.113)***###	121	0.275	0.212 (0.080)***###	121	0.056
Canada	0.605 (0.065)***###	127	0.396	0.288 (0.138)**###	127	0.064
Japan	0.982 (0.095)***	127	0.560	0.528 (0.038)***###	127	0.653
Finland	0.581 (0.076)***###	125	0.208	0.461 (0.081)***###	125	0.209
Ireland	0.808 (0.081)***##	126	0.457	0.772 (0.097)***##	126	0.400
Spain	0.863 (0.150)***	127	0.231	0.410 (0.078)***###	127	0.142
Australia	0.658 (0.035)***###	97	0.748	0.442 (0.064)***###	107	0.299
New Zealand	0.698 (0.072)***###	127	0.509	0.626 (0.083)***###	127	0.390
Average	0.67(0.03)			0.39(0.03)		
Emerging market economies						
South Africa	0.302 (0.082)***###	108	0.2551	0.452 (0.06)***###	108	0.2343
Argentina	0.91 (0.046)***#	64	0.8915	0.879 (0.051)***##	69	0.7057
Colombia	0.491 (0.06)***###	123	0.059	0.503 (0.158)***###	123	0.0367
Brazil	0.997 (0.022)***	127	0.9118	0.986 (0.027)***	127	0.8693
Mexico	-0.021 (0.015)###	126	0.028	-0.160 (0.073)** ###	126	0.072
Jordan	0.733 (0.282)***	81	0.1815	0.554 (0.211)***##	81	0.0748
Hong Kong	0.311 (0.069)***###	127	0.2514	0.146 (0.0655)**###	127	0.0655
Republic of Korea	0.85 (0.097)***	127	0.5216	0.73 (0.096)***###	127	0.5705
Pakistan	0.618 (0.203)***#	125	0.1046	0.352 (0.137)**###	125	0.0663
Singapore	-0.18 (0.188)###	127	0.0119	-0.368 (0.199)*###	127	0.0373
Thailand	0.82 (0.107)***#	127	0.3552	0.769 (0.187)***	127	0.0282
Peru	0.785 (0.122)***#	71	0.402	0.773 (0.311)**	71	0.116
Hungary	0.606 (0.182)***##	127	0.187	0.586 (0.11)***###	127	0.359
Poland	0.888 (0.104)***	99	0.022	0.872 (0.066)***#	99	0.023
Turkey	0.955 (0.055)***	115	0.834	0.913 (0.024)***###	115	0.868
Chile	0.976 (0.106)***	59	0.659	0.475 (0.204)**#	59	0.112
Average	0.63(0.03)			0.53(0.04)		

Trade prices for advanced economies come from the OECD's Monthly International Statistics database. Trade prices for the emerging market economies come from the IMF's International Financial Statistics database and the nominal effective exchange rate data comes from the IMF's Information Notice System. An * indicates if the coefficient is significantly different from 0. A # indicates if the pass-through coefficient is significantly different from 1. ***, **, and * denote the 1, 5, and 10 percent levels, respectively. ###, ##, and # denote the 1, 5, and 10 percent levels, respectively. Robust standard errors are reported in parentheses. Countries are classified as advanced and emerging market economies based on the classification in the IMF's World Economic Outlook publications. A constant is included in the OLS regressions of each country. The estimated coefficients are not reported here.

Table 2. VAR: First Quarter Trade Price Response to a One Percent Change in the Exchange Rate, 1979–2010

	Import Price Response		Export Price Response		p-value for t test of difference in trade
Country	Estimate	Std. error	Estimate	Std. error	price responses
Advanced economies					
United States	0.38	(0.08)	0.17	(0.05)	0.02
United Kingdom	0.37	(0.05)	0.25	(0.05)	0.10
Belgium	0.50	(0.18)	0.48	(0.13)	0.92
Denmark	0.69	(0.10)	0.42	(0.08)	0.03
France	0.30	(0.14)	0.11	(0.09)	0.25
Germany	0.61	(0.09)	0.17	(0.03)	0.00
Italy	0.62	(0.11)	0.33	(0.06)	0.02
Netherlands	0.93	(0.20)	0.61	(0.18)	0.24
Norway	0.57	(0.09)	0.06	(0.26)	0.06
Sweden	0.39	(0.06)	0.30	(0.04)	0.19
Switzerland	0.52	(0.08)	0.27	(0.07)	0.03
Canada	0.59	(0.07)	0.39	(0.09)	0.08
Japan	0.90	(0.09)	0.55	(0.05)	0.00
Finland	0.63	(0.10)	0.57	(0.09)	0.69
Ireland	0.70	(0.08)	0.78	(0.09)	0.53
Spain	0.76	(0.14)	0.40	(0.09)	0.03
Australia	0.63	(0.06)	0.46	(0.07)	0.06
New Zealand	0.65	(0.07)	0.61	(0.07)	0.66
Average	0.60	(0.03)	0.39	(0.02)	0.00
Emerging market economies					
South Africa	0.31	(0.05)	0.38	(0.08)	0.43
Argentina	0.94	(0.10)	0.87	(0.09)	0.59
Colombia	0.51	(0.15)	0.76	(0.20)	0.32
Brazil	0.86	(0.08)	0.99	(0.09)	0.28
Mexico	-0.01	(0.01)	-0.15	(0.06)	0.02
Jordan	0.43	(0.24)	0.03	(0.20)	0.20
Hong Kong	0.24	(0.03)	0.10	(0.03)	0.00
Republic of Korea	0.87	(0.08)	0.70	(0.08)	0.14
Pakistan	0.45	(0.18)	0.44	(0.13)	0.97
Singapore	-0.22	(0.13)	-0.33	(0.15)	0.61
Thailand	0.87	(0.11)	0.60	(0.43)	0.53
Peru	0.80	(0.13)	1.34	(0.33)	0.13
Hungary	0.52	(0.10)	0.53	(0.08)	0.93
Poland	0.31	(1.61)	0.52	(1.57)	0.92
Turkey	0.94	(0.07)	0.89	(0.07)	0.62
Chile	0.86	(0.11)	0.52	(0.15)	0.07
Average	0.54	(0.10)	0.51	(0.11)	0.84

Trade prices come from the OECD's Monthly International Statistics database for advanced economies and the IMF's International Financial Statistics database for the emerging market economies. All other variables come from the IMF's Information Notice System database.

Table 3. Pass-Through Elasticities for Different Shocks

Shocks	Pass-Through Elasticities	
	Import Price	Export Price
Nominal Shock	0.572	0.404
Home Productivity Shock	0.346	-0.221
Foreign Productivity Shock	-0.167	0.404
Composite Shock	0.523	0.381

In the case of nominal and home and foreign productivity shocks, the elasticities are based on a one-standard-deviation increase in each shock. The elasticities for the composite shock represent the effect of a simultaneous one-standard-deviation increase in all shocks.

Table 4. Wage-Price Stickiness, Exchange Rate Persistence, Inflation Reaction and the Pass-Through

	Import Price Pass-Through	Export Price Pass Through
Baseline values	0.572	0.419
Less stickiness ($\theta = .5, \theta_W = .667$)	0.727	0.332
Lower persistence ($\rho_x = 0.6$)	0.553	0.438
Higher persistence ($\rho_x = 0.95$)	0.599	0.402
Weaker reaction ($\delta = 0.1$)	0.564	0.420
Stronger reaction ($\delta = 1.0$)	0.583	0.417

The pass-through values represent elasticities generated by the nominal shock.

Table 5. Stochastic Simulations

	Standard Dev.		Import Price PT		Export Price PT	
	Δs_t	π_t	OLS	VAR	OLS	VAR
Data						
Adv. Coun.	0.0245	0.0086	0.67	0.60	0.39	0.39
Model						
Baseline	0.0329	0.0075	0.607	0.614	0.466	0.463
Variant 1	0.0233	0.0083	0.655	0.611	0.398	0.405
Variant 2	0.0238	0.0079	0.625	0.615	0.418	0.400
Variant 3	0.0040	0.0015	0.341	0.072	0.324	0.050
Variant 4	0.0227	0.0072	0.437	0.424	0.230	0.204
Variant 5	0.0263	0.0094	0.883	0.873	0.667	0.700

Variant1: Adjusted nominal shock process and PCP shares
($\rho_X = 0.95, std.dev.(v_X) = 0.0025$, $\psi = 0.55, \psi^* = 0.45$);

Variant 2: Only nominal shocks ($std.dev.(v_A) = 0$, $std.dev.(v_{A*}) = 0$);

Variant 3: Only real shocks ($std.dev.(v_X) = 0$);

Variant 4: Higher home and lower foreign PCP share ($\psi = 0.8, \psi^* = 0.2$); and

Variant 5: Lower home and higher foreign PCP share ($\psi = 0.2, \psi^* = 0.8$).

Table 6. Invoicing Currency Shares and the Pass-Through

	Export Price Pass-Through	Import price Pass-Through	
Constant	0.615 (0.074)*	0.461 (0.186)*	0.306 (0.145)*
Exp. curr. share in exports	-0.005 (0.0016)*		
Exp. curr. share in imports		0.002 (0.0027)	
US dollar share in imports			0.008 (0.0035)*
No. of Obs.	15	13	13
R-squared	0.440	0.053	0.287

Standard errors are shown in brackets. * Indicates significance at the 5% level or less.

Figure 1. Regression and VAR Estimates of the Pass-Through

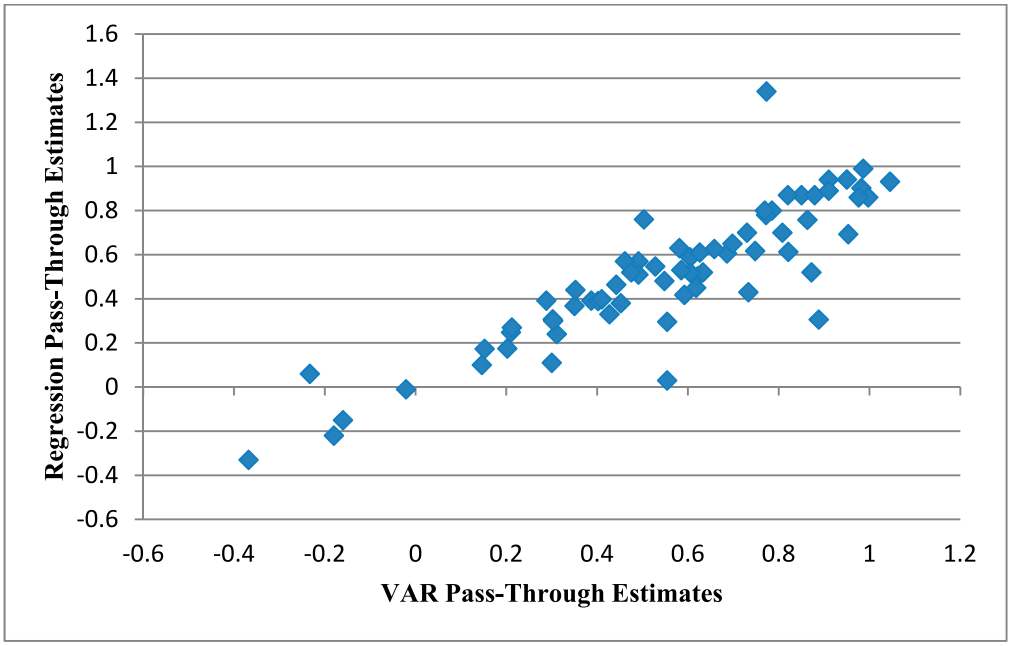
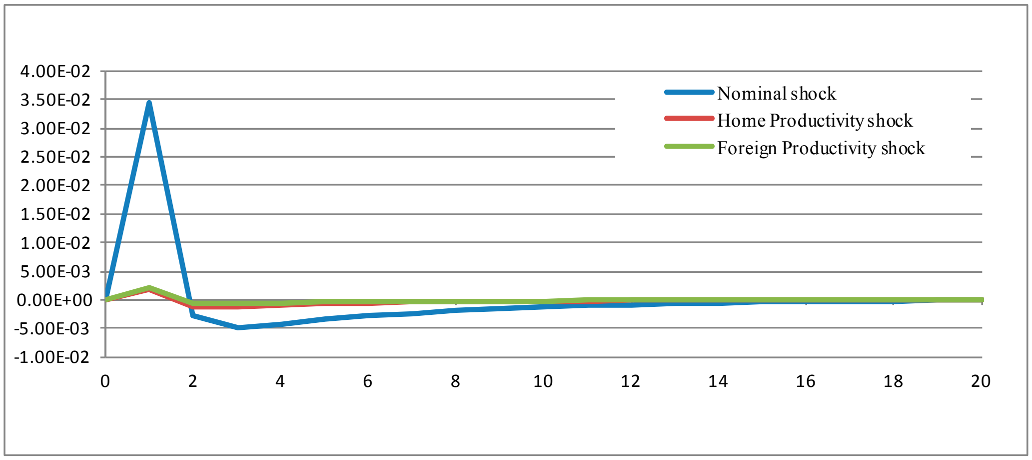
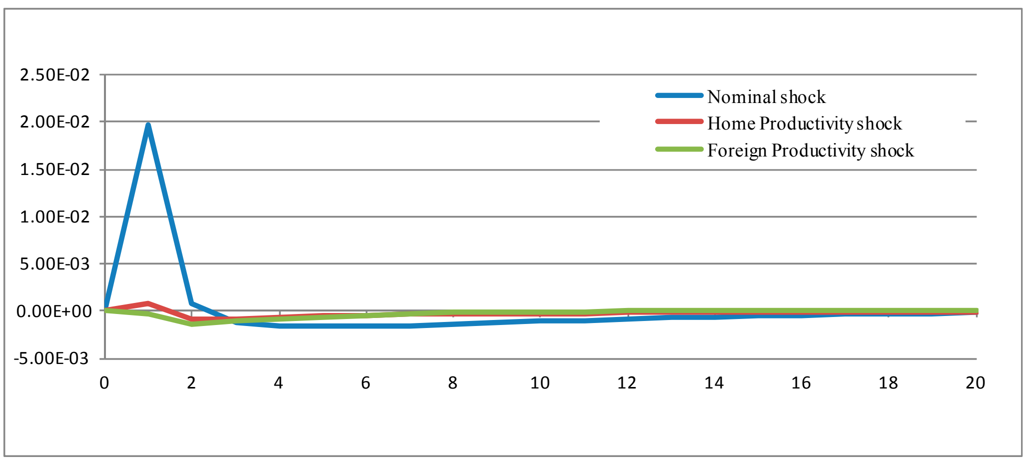


Figure 2. Impulse Response Functions

(a) Response of Δs_t



(b) Response of $\Delta p_{M,t}$



(c) Response of $\Delta p_{X,t}$

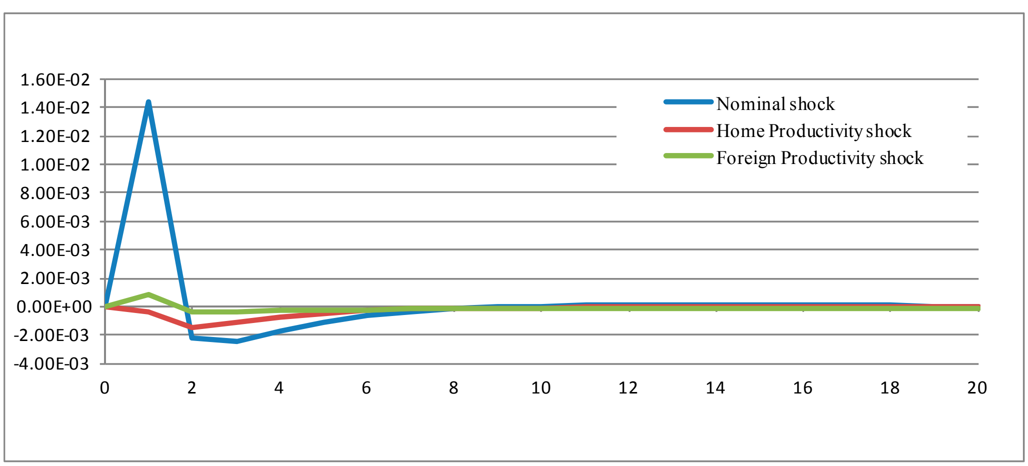
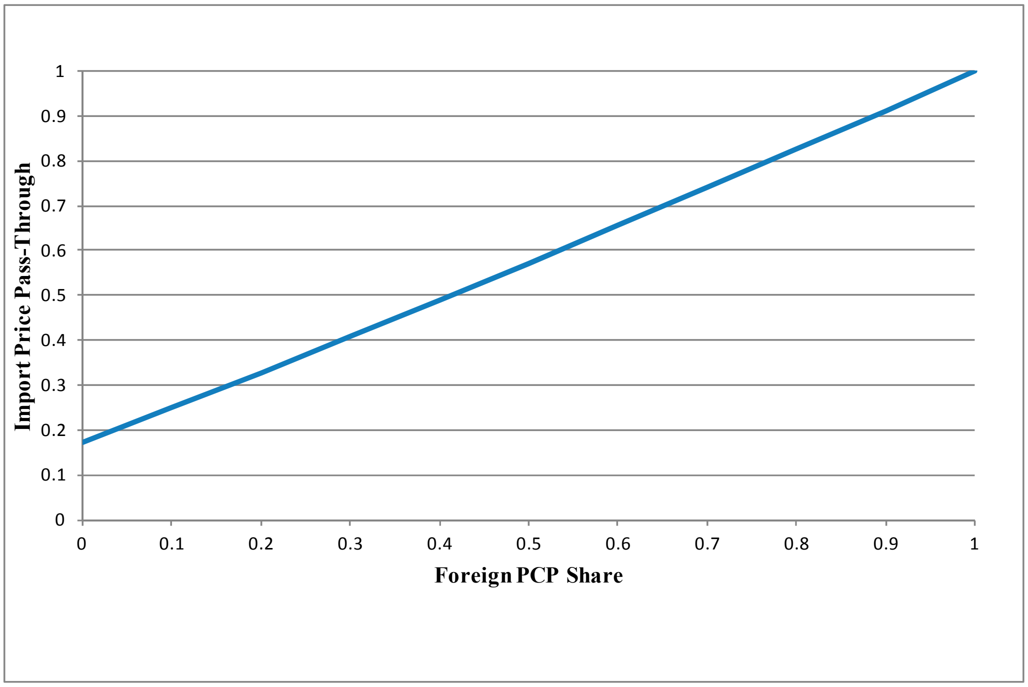
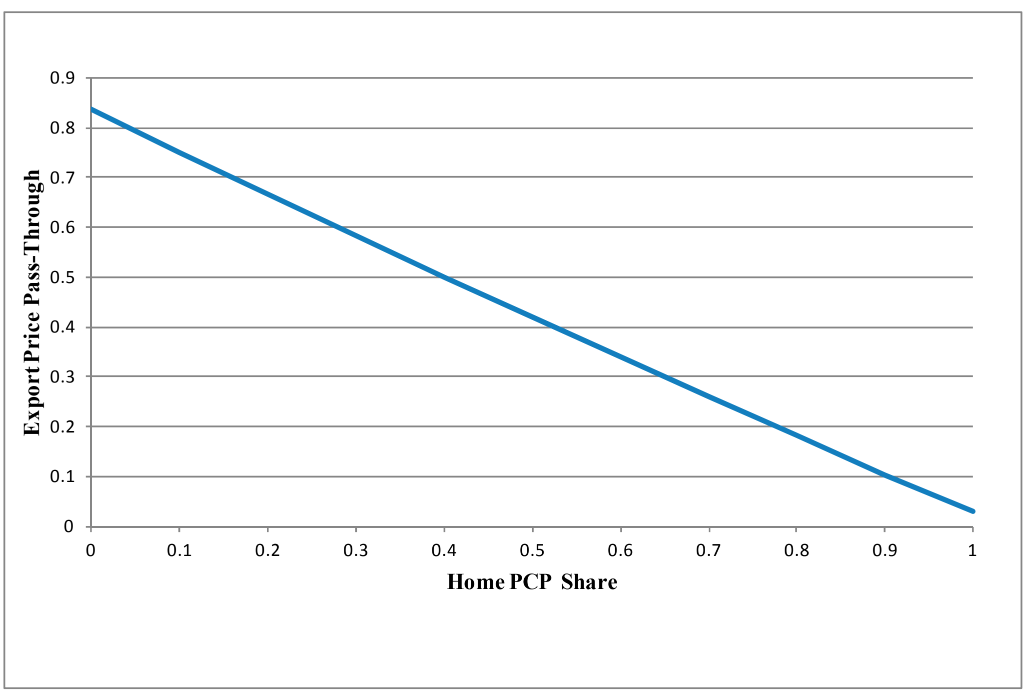


Figure 3. PCP Shares and the Pass-Through

(a) Foreign PCP Share and the Import Price Pass-Through



(b) Home PCP Share and the Export Price Pass-Through



Appendix Table 1: VAR: First Quarter Trade Price Response to a One Percent Change in the Exchange Rate 1985–1997

Country	Import Price Response		Export Price Response		p-value for t test of differences in trade price responses
	Estimate	Std. error	Estimate	Std. error	
Advanced economies					
United States	0.22	0.08	0.08	0.063	0.18
United Kingdom	0.41	0.08	0.29	0.086	0.29
Belgium	0.78	0.28	1.08	0.197	0.38
Denmark	0.38	0.12	0.13	0.114	0.13
France	0.05	0.13	0.07	0.114	0.92
Germany	0.40	0.11	0.16	0.029	0.04
Italy	0.52	0.12	0.25	0.072	0.05
Netherlands	0.66	0.22	0.32	0.225	0.29
Norway	0.77	0.23	-0.03	0.443	0.11
Sweden	0.49	0.08	0.43	0.073	0.56
Switzerland	0.57	0.11	0.24	0.068	0.01
Canada	0.42	0.15	0.15	0.145	0.19
Japan	1.06	0.15	0.63	0.081	0.01
Finland	0.64	0.12	0.56	0.108	0.63
Ireland	0.78	0.16	0.88	0.169	0.67
Spain	0.57	0.20	0.52	0.114	0.82
Australia	0.70	0.08	0.43	0.082	0.02
New Zealand	0.49	0.10	0.45	0.128	0.82
Average	0.55	0.036	0.37	0.037	0.00
Emerging market economies					
South Africa	0.20	(0.07)	0.38	(0.17)	0.33
Argentina	0.01	(0.24)	0.40	(0.54)	0.51
Colombia	0.30	(0.46)	0.61	(0.56)	0.67
Brazil	0.87	(0.12)	0.99	(0.12)	0.50
Mexico	0.03	(0.01)	-0.02	(0.08)	0.53
Hong Kong	0.14	(0.04)	0.02	(0.03)	0.01
Republic of Korea	0.72	(0.11)	0.37	(0.13)	0.04
Pakistan	-0.11	(0.33)	-0.01	(0.24)	0.79
Singapore	-0.12	(0.16)	-0.40	-(0.05)	0.09
Thailand	0.77	(0.14)	-0.54	(0.91)	0.16
Peru	0.83	(0.18)	0.61	(0.28)	0.51
Hungary	0.73	(0.20)	0.60	(0.16)	0.60
Poland	1.21	(3.15)	1.49	(3.07)	0.95
Turkey	1.07	(0.12)	0.97	(0.11)	0.55
Average	0.47	(0.86)	0.39	(0.89)	0.95

Trade prices come from the OECD's Monthly International Statistics database for advanced economies and the IMF's International Financial Statistics database for the emerging market economies. All other variables come from the IMF's Information Notice System database.

Appendix Table 2: VAR: First Quarter Trade Price Response to a One Percent Change in the Exchange Rate 1998–2010

Country	Import Price Response		Export Price Response		p-value for t test of differences in trade price responses
	Estimate	Std. error	Estimate	Std. error	
Advanced economies					
United States	0.64	0.14	0.424	0.079	0.19
United Kingdom	0.39	0.10	0.468	0.115	0.60
Belgium	0.54	0.23	0.428	0.155	0.68
Denmark	0.46	0.16	0.496	0.154	0.86
France	-0.02	0.27	-0.178	0.172	0.61
Germany	0.07	0.13	0.056	0.039	0.92
Italy	0.74	0.25	0.330	0.100	0.12
Netherlands	0.17	0.49	0.591	0.381	0.50
Norway	0.55	0.09	0.300	0.392	0.53
Sweden	0.29	0.10	0.251	0.063	0.73
Switzerland	0.42	0.11	0.223	0.125	0.22
Canada	0.68	0.10	0.537	0.178	0.47
Japan	0.80	0.17	0.499	0.090	0.11
Finland	0.44	0.23	0.789	0.199	0.24
Ireland	0.55	0.11	0.756	0.184	0.34
Spain	0.82	0.30	0.783	0.222	0.93
Australia	0.63	0.10	0.871	0.189	0.25
New Zealand	0.65	0.11	0.624	0.113	0.89
Average	0.49	0.048	0.46	0.044	0.63
Emerging market economies					
South Africa	0.38	(0.08)	0.74	(0.12)	0.01
Argentina	0.95	(0.12)	0.86	(0.11)	0.58
Colombia	0.72	(0.09)	0.95	(0.16)	0.23
Brazil	0.91	(0.15)	1.13	(0.20)	0.39
Mexico	-0.17	(0.04)	-0.67	(0.14)	0.00
Jordan	1.45	(0.51)	0.69	(0.39)	0.23
Hong Kong	0.27	(0.08)	0.19	(0.07)	0.43
Republic of Korea	0.67	(0.16)	0.80	(0.14)	0.52
Pakistan	1.34	(0.31)	0.56	(0.19)	0.03
Singapore	-0.39	(0.35)	-0.46	(0.31)	0.88
Thailand	1.17	(0.18)	0.83	(0.14)	0.13
Peru	1.09	(0.18)	1.48	(0.35)	0.33
Hungary	0.75	(0.09)	0.74	(0.09)	0.92
Poland	0.60	(0.13)	0.43	(0.14)	0.36
Turkey	0.90	(0.10)	0.84	(0.09)	0.67
Chile	0.90	(0.12)	0.51	(0.15)	0.05
Average	0.72	(0.05)	0.60	(0.05)	0.09

Trade prices come from the OECD's Monthly International Statistics database for advanced economies and the IMF's International Financial Statistics database for the emerging market economies. All other variables come from the IMF's Information Notice System database.