Country and Time-Variation in Import Exchange Rate Pass-Through: Is It Driven By Macro Factors?

Raphael Brun-Aguerre[a], Ana-Maria Fuertes[b], Kate Phylaktis[b,*]

[a] *Economics Research Department, J.P.Morgan*

[b] *Faculty of Finance, Cass Business School, City University London*

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Abstract

A relatively large sample of developed and emerging economies is examined to explore the empirical importance of macroeconomic factors as drivers of exchange rate pass-through. The analysis supports the presence of a long run equilibrium relation between the domestic import price, the foreign effective export price and the corresponding effective exchange rate. An in-sample analysis based on panel models reveals that the pass-through elasticity is significantly linked overall to the inflation level, long-term GDP growth, relative GDP per capita, import dependence ratio and level of tariffs although there are differences across emerging and developed countries. There is also evidence for both types of countries that the direction and size of the exchange rate change matters in explaining pass-through elasticities. A univariate out-of-sample forecasting exercise confirms the predictive role of macro factors in import pass-through models and brings to the forefront the predictive role of the exchange rate volatility, import dependence ratio, tariffs and the global business cycle.

JEL classifications: C22; C51; F14; F31.

Keywords: Import Prices; Exchange Rate Pass-Through; Price Discrimination; Error Correction Model; Forecasting.

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1 Introduction

In a world of inflation targeting the impact of exchange rate fluctuations on import prices is relevant to governments as well as to producers and consumers. Domestic currency depreciations can increase import prices which, in turn, may translate into domestic price inflation. A country targeting its inflation must regularly adjust its interest rate to counteract deviations from the inflation target caused by changes in the nominal exchange rate. The higher the impact of exchange rate changes on (domestic) import prices, a phenomenon called *import* exchange rate pass-through (ERPT), the greater the level of adjustment of interest rates required to achieve the inflation target following a depreciation of the domestic currency. The extent of import ERPT a destination country experiences has implications not only for its trade balance determination but also for its choice of exchange rate regime; in fact, the "fear of floating" typically associated to poor countries and their larger degree of intervention in foreign exchange markets (among those that have adopted inflation targeting) can be theoretically linked to the fear of being subject to a relatively large degree of import ERPT. The ERPT elasticity, which represents the percentage change in local currency import prices resulting from a 1% change in the exchange rate between the importing and exporting countries, can plausibly range between 0% and 100% (complete pass-through) depending on the exporters pricing policy.[1] The Law of One Price (LOOP) states that, under the assumption of costless arbitrage, identical goods sold in different markets must have the same common-currency price which, in turn, implies that import pass-through should be complete. Export prices are set up as a markup over the producer's marginal costs. If the currency of the importing (or destination) country depreciates and the export price in foreign currency is maintained then the import price will reflect all the change in the exchange rate so there is complete ERPT (i.e. unit import ERPT elasticity). In this

[1] This paper focuses on the narrowest definition of pass-through to the prices of goods observed "at the dock", i.e. when they first arrive in the destination country, as opposed to wider definitions such as the pass-through to the price of the same imported goods at retail (store counter) or more generally to the general price level (CPI). Additional mechanisms are in place in the broader definitions (over and above the pricing policy of the exporter) such as the costs of transportation from the exporting country to the destination country, the costs of distribution and retail (including real wages and rents) that apply between the dock in the country of import and the store counter and the degree of competition among local producers. Tariffs and transportation costs are likely to be more important in emerging than developed markets and the other way round for distribution and retail costs. The low cost of retail services in poor countries, relative to the value of the product, has been used as one argument to explain higher retail pass-through in developing countries (Frankel et al., 2005). It has been argued theoretically that the degree of pass-through to consumer goods prices decreases as the level of competition among local producers increases (Bacchetta and Van Wincoop, 2002).

case, we have producer-currency pricing (PCP). On the other hand, if exporters are willing to lower their prices (hence, their markups) to offset the full exchange rate change, this implies that import prices in local currency of the importing country will be left unchanged (zero import ERPT elasticity); in other words, exporters adopt local-currency pricing (LCP). If the latter prevails, then there is an "insulation" effect by which the expenditure-switching effects of FX fluctuations are lessened. Hence, monetary policy becomes a more effective tool. The greater the ERPT the more challenging it is to keep the monetary stability and the inflation target. Hence, a deeper understanding of the ERPT mechanisms is of key importance to Central Banks for policy-making.

Using as laboratory the currency realignments of 1971, Kreinin (1978) documented various degrees of ERPT; a relatively small pass-through to US import prices at 50 percent and larger ones at 60 percent for Germany, 70 percent for Japan and 100 percent for Italy. Moreover, the currency crises (i.e. depreciations) which took place during the 1990s surprisingly did not entail high inflation rates implying that ERPT was incomplete. This apparent resilience of import prices to fluctuations in the exchange rate has been the subject of a vast theoretical and empirical literature. In a seminal paper, Dornbusch (1987) provides a theoretical model that explains incomplete pass-through in terms of the degree of market concentration, the relative market shares of domestic and foreign firms, the extent of product homogeneity and substitutability, and pricing-to-market or the practice by exporters to adjust markups in order to price discriminate across destination countries.

The present study seeks to contribute to a vast literature in several directions. First, it investigates the pass-through of exchange rate movements into a country's import prices using a relatively large sample of 19 developed markets (DMs) and 18 emerging markets (EMs) over the period 1980Q1 -2009Q3 which includes the recent global financial crisis period 2008-2009 characterized by sharp rises in FX rate volatility. The majority of previous studies have focused on either DMs (e.g., Campa and Goldberg, 2005), or EMs (e.g., Zorzi et al., 2007). Few papers have as yet considered a large cross-section of countries in their analyses including both DMs and EMs, the latter being more open economies and more prone to financial crises and large depreciations. Among these, Frankel et al. (2005) use highly disaggregated data, Choudhri and Hakura (2006) focus on the importance of the inflation history, and Golfajn and Werlang (2000) concentrate on macroeconomic factors. This previous research has documented that ERPT is higher for EMs than for DMs which is often linked to larger inflation in

the former type of countries. Our findings based on a more recent sample period and a different export proxy suggest that the ERPT elasticities are higher for DMs than for EMs both in the short-run and long-run.

Second, whereas previous studies rely on bilateral FX data and domestic price indices to proxy export prices we use not only *effective* or *trade-weighted* exchange rates but also *trade-weighted* foreign export prices. This aspect of the data can be important since ERPT estimates that are based exclusively on the exchange rate against the U.S. dollar as the base currency are likely to be biased; this is because the exchange rate captures not only the variation in the value of the domestic currency but also of the foreign one too. Effective and bilateral exchange rates will on the whole differ. Since effective and bilateral rates are associated with the same observed variation in the price level a potential bias emerges. This bias leads to an underestimation of ERPT if the fluctuations in the value of the US dollar (the base currency) are more important than those of other international currencies, and to an overestimation of ERPT if other international currencies are more volatile. Thus the use of effective exchange rates aims at providing a better insulation of the measured ERPT from shocks affecting mainly the US economy. Similarly, our effective export price measure for each (domestic) importing country is constructed using the relative bilateral trade as weighting of each (foreign) export price index instead of relying on general price indices as in previous work which should further improve the accuracy of ERPT estimates. Many previous studies proxy the foreign export price by the CPI or PPI in the currency of the country of export. For instance, Anderton (2003) computes an effective export price as a weighted average of the producer prices of 7 major euro area import suppliers. The weights are based on the share of euro area manufacturing imports accounted for by the individual country. Instead, Campa and Goldberg (2005) proxy it by a trading-partner cost index, $\frac{NEER_t \times P_t}{REER_t}$, where NEER and REER are nominal and real effective (foreign vis-à-vis domestic) exchange rates and P is the domestic import price index;[2] all these proxies reflect mainly the evolution of prices for domestic consumption or production and not the dynamics of prices for exports.

Thirdly, in measuring ERPT we allow for the possibility that exchange rate changes might not influence import prices in a linear manner. Our definition of 'nonlinear' ERPT mechanism is a broad one: any departure

[2] A bilateral exchange rate involves a currency pair, while an effective exchange rate is an index representing the weighted average value of a country's currency relative to all major trading partner's currencies. The weights are determined by the relative trade importance a home country places on the other currencies as measured by the balance of trade. A real effective exchange rate (REER) adjusts the NEER by appropriate foreign price levels and deflates it by the home country price level.

from the traditional restrictive linear modeling approach(es) that imply time-constant ERPT elasticities. Theoretical explanations exist for non-constant pass-through elasticities where the dynamics could be dictated by macroeconomic variables. For instance, competition among exporters to preserve (or gain) market share may imply that the amount of the exchange rate change that is passed onto import prices is inversely related to the exchange rate volatility level (Froot and Klemperer, 1989). Downward price rigidities may imply a discrete-type variation (or 'switch') in the ERPT elasticity over time whereby the impact of exchange rate changes on (domestic) import prices is smaller for appreciations (of the domestic currency) than for depreciations. Menu costs, which are like a fixed cost, may imply that changing invoice prices is only worthwhile if the exchange rate change is above a certain threshold and hence, the ERPT elasticity may differ for large and small exchange rate changes. The empirical literature that has considered the possibility of nonlinear responses to date is still quite scant. The majority of the few existing studies relate to either individual or a small number of countries: Webber (2000) focuses on 7 Asian countries, Herzberg et al (2003) on the UK, Marazzi et al. (2005) on the US, Khundrakpam (2007) on India, and Bussière (2007) on the G7 economies. Overall the findings are mixed.

Campa and Goldberg (2005) allow for continuous-type time variation in long run and short run import ERPT elasticities estimated using panel data for 23 OECD countries over the 1975Q1-2003Q4 period. They split this time period into four subperiods and for each of them they obtain, first, the (long-) short-run elasticities using linear regressions with all variables (exchange rate and prices) in first-differences. In a second, pooled time-series linear modeling approach they seek to explain the time variation in those ERPT elasticities in terms of money growth, GDP growth, inflation changes, exchange rate volatility changes and an imputed aggregate elasticity measure that reflects the changes in a country's ERPT that are due exclusively to changes in its imports composition. Their main conclusion is that the observed decline in pass-through is more a 'micro' than a 'macro' phenomenon: changes in the composition of the importing basket away from goods characterized by a high pass-through (raw-material products, particularly energy) toward goods characterized by a low pass-through (manufacturing items). They do not find any significant influence of the macroeconomic variables on the ERPT elasticities. On the other hand, Marazzi et al. (2005) provides little evidence for the view that a shift in the country composition of U.S. imports away from high-wage countries toward low-wage countries (which are

able to continue to provide goods at or near the price that prevailed before the exchange rate change) is a key element in explaining the decline in pass-through. Although for one country only, Pollard and Coughlin (2004) and Herzberg et al. (2003) are the first studies to investigate the role of the nonlinearities and asymmetries in the pass-through to, respectively, US and UK import prices; the former study concludes that the size of FX changes is more important than the direction (sign) effect whereas the latter concludes that the ERPT mechanism is linear. More recently, Bussière (2007) contributes to the ERPT literature by considering the presence of asymmetries/nonlinearities in the reaction of *both* import and export prices. The main focus of his analysis is the comparison of large versus small exchange rate changes, and appreciations versus depreciations. On the basis of first-difference models for G7 countries over the 1980Q1-2006Q4 period, evidence is provided that nonlinearities and asymmetries cannot be ignored.

We start by exploiting our large sample of 37 countries over a 30-year period (1980Q1-2009Q3) to contrast three competing linear models that have been employed in the recent ERPT literature: a first-difference regression and two distinct ECMs that incorporate the levels (or long-run equilibrium) relationship between the nominal exchange rate, import and export prices; the significance of the levels relationship is documented through several cointegration tests. The selected linear model according to the adjusted-R^2, AIC and SBC is next estimated in a rolling window approach for each country. The time-series sequence of short-run ERPT elasticities thus obtained is pooled across countries and panel regressions are estimated to ascertain the role of different potential *drivers* of ERPT variation. Plausible candidates are the domestic inflation rate, the extent of FX rate volatility, the size of the economy, the degree of trade openness, and the cyclical component of output. Lastly, we analyze the relative ability of nonlinear ERPT models to conditionally forecast out-of-sample future import price changes, an exercise which has not been attempted in previous work. This constitutes our fourth and final contribution to the literature.

2 Data and Methodology

2.1 Import price, export price and exchange rate data

The main three variables in empirical models of ERPT are the exchange rate, the local-currency (domestic) import price and the foreign-currency export price. The former is a nominal effective, or trade-weighted, exchange rate (NEER) index of the foreign currency per unit of the domestic currency and hence, an increase in the index marks an appreciation of the domestic currency. The import and export prices are proxied by customs unit value indices. For each importing or destination country in our sample, the import price (proxying the domestic price of goods and services "at the dock") is matched with an *effective* foreign export price (proxying the foreign price of goods and services coming into the country). The latter is constructed from foreign export unit value indices together with bilateral trade data; thus the export price measure used is a weighted average of foreign export unit value indices where each weight represents the relative share of each of the foreign countries' exports in the domestic importing economy. The frequency of the observations is quarterly over a maximum time span of $T = 119$ quarters from 1980Q1 to 2009Q3; the observations begin later for some countries. A total of $N = 37$ economies are included in the analysis of which 18 are identified as "emerging markets" (EMs) and 19 are "developed countries" (DMs) according to *The Economist* magazine; see **Appendix A** for more details.[3] In what follows we employ the following notation: $t = 1, ..., T$ denote quarterly observations for $i = 1, ..., N$ importing countries, $p_{i,t}$ denotes the import unit value index, $p^*_{i,t}$ is the *effective* export unit value index[4] and $s_{i,t}$ is 1/NEER with NEER defined as above (all in logarithms).

[3] The classification of EMs by *The Economist* is employed by Michigan State University for its Market Potential Index (for emerging markets) available at http://globalEDGE.msu.edu/resourceDesk/mpi/. We build individual import EPRT models for each if the 37 countries listed in Appendix Table A. However, the sample is larger since export price indices for several countries not listed in the table are also included in our sample to construct the effective export prices for those 37 importing countries. Examples are Bolivia, Ecuador, India, Indonesia, Taiwan and Russia, for which we were not able to collect import price data and so their ERPT regressions were not feasible.

[4] For each import country in our sample, $i = 1, ..., 37$, the trade-weighted export price $p^*_{i,t}$ representing the foreign price of goods and services (exports) received by the country is constructed as $p^*_{i,t} = \sum_{j=1}^{J} w^j_{i,t} p^{j*}_{i,t}$ where $j = 1, ..., J$ are the trading (exporters) partners to country i. Therefore $p^*_{i,t}$ is a proxy for the rest-of-the-world foreign export price faced by importing country i. It is constructed as a weighted average of all trade partners total export prices (*i.e.* a unique export price is available for each country and no distinction is made with respect the direction of trade). In order to reflect the country (foreign) composition of the importing basket, the foreign export prices are weighted by $w^j_{i,t} = \frac{M^j_i}{\sum_{j=1}^{J} M^J_i}$, the share of country's i imports coming from trading partner j, where M^j_i is the value of country's i imports from trading partner j. The bilateral trade figures are obtained from the IMF *Direction of Trade Statistics*.

We start by examining the time-series properties of the three main variables for the analysis, p_{it}, p^*_{it} and s_{it}. For each of them, the null hypothesis of unit root non-stationarity (against the alternative hypothesis of stationarity) is tested using the Augmented Dickey-Fuller test, and the null stationarity (against the alternative of non-stationarity) is tested using the Kwiatkowski et al. (1999; KPSS) test. We gather robust evidence that the series are I(1) non-stationary.[5] According to the LOOP for traded goods over the long-run these non-stationary variables should co-move, namely, p_{it}, p^*_{it} and s_{it} are cointegrated; thus first-difference ERPT models that exclude the long-run (i.e. level) relationship are misspecified. **Figure 1** plots the import price, export price and FX rate (all in logarithms) for 6 EMs and 6 DMs in the sample.[6] The EMs have been chosen to represent different geographical regions: Latin America (Colombia), Europe (Hungary, Turkey), Asia (Hong Kong, Pakistan), Africa (South Africa). Likewise for the DMs: Asia (Japan), European Union (France, Germany), Europe (UK) and North America (Canada, US).It is interesting to note that the NEER for all EMs is upward sloping for the 30-year sample period indicating a depreciation, while for the DMs it is downward sloping indicating an overall appreciation. Although the graphical evidence is inconclusive, it can be perceived that although the individual series are nonstationary they do not diverge too much from each other in the long run. The importance of acknowledging the presence of this long-run relationship in ERPT modeling is formally revisited in the next section.

2.2 Linear time-constant ERPT framework

Given that different empirical models have been adopted in the recent ERPT literature, it would be interesting to draw comparisons across them in a *ceteris paribus* sense. It is difficult to do so on the basis of results from existing studies because not only they differ in the cross-section and time dimensions of their samples but also in the variables used (e.g. nominal versus effective exchange rates) and often in the empirical models too. Therefore we start by contrasting the in-sample goodness of fit and out-of-sample forecast accuracy of three linear specifications (all of which imply time-invariant ERPT elasticities) on the basis of the same large sample of 37 developed and emerging markets 1980Q1-2009Q3.

[5] To save space we do not present the detailed results but they are available from the authors upon request

[6] The plots for the remaining 25 countries in our sample are available from the authors upon request.

The failure to find evidence of cointegration (i.e. a long-run levels relation supporting the theoretical Law of One Price) between p_{it}, p^*_{it} and s_{it} has led many studies to adopt the following first-difference specification:

$$\Delta p_{i,t} = c_i + \sum_{k=0}^{K_1} \beta_{k,i} \Delta s_{i,t-k} + \sum_{k=0}^{K_2} \gamma_{k,i} \Delta p^*_{i,t-k} + e_{i,t}, \tag{1}$$

where the short-run ERPT is given by $\beta_{0,i}$ and the long-run ERPT by $\sum_{k=0}^{K_1} \beta_k$. Thus, for instance, a value $\beta_{0,i} = 0.4$ implies that a 1 percent depreciation of the importing country's currency (i.e. $\Delta s_{i,t} > 0$) would make import prices 0.4 percent more expensive in the short-run; since this model is linear it does not accommodate asymmetries and therefore $\beta_{0,i} = 0.4$ equally implies that a 1 percent appreciation ($\Delta s_{i,t} < 0$) would reduce import prices by 0.4 percent. In our analysis, the "short run" amounts to one quarter since the frequency of the observations is quarterly. Using quarterly data for 23 OECD countries, Campa and Goldberg (2005) estimate this model for $K_1 = K_2 = 4$ additionally including as control variable the destination market's real GDP growth. Bandt et al. (2008) criticize this specification, in particular, regarding how sensitive the inferences on the long-run ERPT are to the choice for K_1. Moreover, Bandt et al. (2008) argue that the failure to support a long-run ERPT level relationship in most studies is because of the presence of structural breaks which pose a challenge for conventional cointegration tests.[7] If evidence of cointegration is found then it is, in principle, more appropriate to employ a specification such as:

$$\Delta p_{i,t} = c_i + \sum_{k=0}^{K_1} \beta_{k,i} \Delta s_{i,t-k} + \sum_{k=0}^{K_2} \gamma_k \Delta p^*_{i,t-k} + \theta_{i,1} p_{i,t-1} + \theta_{i,2} s_{i,t-1} + \theta_{i,3} p^*_{i,t-1} + \epsilon_{i,t}, \tag{2}$$

which is a parsimonious re-parameterization of the equation

$$\Delta p_{i,t} = \sum_{k=0}^{K_1} \beta_{k,i} \Delta s_{i,t-k} + \sum_{k=0}^{K_2} \gamma_k \Delta p^*_{i,t-k} + \delta_i (p_{i,t-1} - \bar{p}^{ERPT}_{i,t-1}) + \epsilon_{i,t},$$

known as error correction model (ECM) in the literature. Model (2) has the additional appealing feature over (1) that it captures the dynamics of import prices in response to the disequilibrium gap, $z_{i,t} \equiv p_{i,t-1} - \bar{p}^{ERPT}_{i,t-1}$ where $\bar{p}^{ERPT}_{i,t-1} \equiv A_i + B_i s_{i,t-1} + C_i p^*_{i,t-1}$ is the long-run (ERPT) import price level; hence, the term $\delta_i z_{it}$ represents an *error correction* mechanism. Model (2) can be estimated by OLS which yields unbiased and

[7] Another might be the use of bilateral exchange rates instead of effective exchange rates as it is argued in Section 1, or the use of unrepresentative domestic (consumer or producer) price indices expressed in foreign currency as proxies for the export price.

consistent measures of $A_i = -c_i/\theta_{i,1}, B_i = -\theta_{i,2}/\theta_{i,1}$ and $C_i = -\theta_{i,3}/\theta_{i,1}$. In this setting, the short-run and long-run ERPT measures are given, respectively, by $\beta_{0,i}$ and $-\frac{\theta_{i,2}}{\theta_{i,1}}$. Campa et al. (2008) employ:

$$\Delta p_{i,t} = a_i + \beta_i \Delta s_{i,t} + \gamma_i \Delta p^*_{i,t} + \alpha_i \Delta p_{i,t-1} + \theta_{i,1} p_{i,t-1} + \theta_{i,2} s_{i,t-1} + \theta_{i,3} p^*_{i,t-1} + \epsilon_{i,t}, \tag{3}$$

which is a more parsimonious variant of model (2) as it includes only the current first-difference of the export price and exchange rate (i.e. no lags). It also differs in that it incorporates the lagged dependent variable as regressor to capture inertia (persistence) in import price changes. In this model, the short run and long run ERPT are measured, respectively, by β_i and $-\frac{\theta_{i,2}}{\theta_{i,1}}$.

We start by estimating models (1), (2) and (3) which are then compared in terms of: *i*) the plausibility of the short-run and long-run ERPT estimates, and *ii*) goodness-of-fit criteria such as the adjusted-R^2, AIC and SBC. We employ $K_1 = K_2 = 4$ throughout.

2.3 Nonlinear time-varying ERPT framework: macroeconomic data

The main focus is on the short-run pass-through elasticity parameter to which the term "ERPT" will refer hereafter. Our definition of a nonlinear ERPT mechanism is quite accommodating: any departure from the previous linear framework where the short-run response of the import price to FX fluctuations (e.g. devaluations) is assumed constant. Thus our modelling approach allows for time variation in ERPT in either a continuous or discrete (regime-switching) fashion. To formalize our approach, let $\beta_{i,t}$ denote the time-varying ERPT elasticity (or ERPT process) which nests as a particular case the linear time-constant ERPT elasticity β_i in (3). Let $Z^1_{i,t}, ..., Z^k_{i,t}$ denote k observable stationary covariates *driving* the ERPT process deterministically as follows:

$$\beta_{i,t} = \beta_i + \lambda^1_i Z^1_{i,t-1} + ... + \lambda^k_i Z^k_{i,t-1} \tag{4}$$

where $\lambda^j_i > 0 (< 0)$ indicates that large levels in the observable process at $t-1$ drive up (down) the ERPT elasticity in the subsequent period t, and vice versa. Introduced in the linear ECM equation (3), the above

mechanism gives the more general ECM specification:[8]

$$\Delta p_{i,t} = a_i + \beta_i \Delta s_{i,t} + \lambda_i^1 Z_{i,t-1}^1 \Delta s_{i,t} + ... + \lambda_i^k Z_{i,t-1}^k \Delta s_{i,t} + \gamma_i \Delta p_{i,t}^* + \alpha_i \Delta p_{i,t-1} + \theta_{i,1} p_{i,t-1} + \theta_{i,2} s_{t-1} + \theta_{i,3} p_{i,t-1}^* + \epsilon_{i,t} \tag{5}$$

Below we describe a set of economic/financial covariates that we consider as drivers of the ERPT process:

FX rate volatility. Although the theoretical literature suggests a link between the degree of import pass-through and FX rate volatility, the direction of the relationship is not clearcut. In a highly competitive export environment, higher exchange rate volatilities are typically associated with lower ERPT. Under high export competition, a lower ERPT occurs if exporters are prepared to quote competitive prices by absorbing the fluctuations in their markup seeking to hold or increase market share (Froot and Klemperer, 1989). On the contrary, in a more monopolistic environment exporters will seek to stabilize their profit margins by setting prices in their own currency (i.e. high ERPT) and so the expected effect is positive. Higher FX volatility has been associated with higher ERPT also in Floden and Wilander (2006) where a model is put forward that accommodates local currency pricing and price-setters follow S-type adjustment rules. Another argument put forward in the theoretical literature to rationalize the direction of the nexus between FX rate volatility and ERPT is how long-lasting the exchange rate movements are perceived: if the increase in FX rate volatility is perceived as very short-lived then exporters are more likely to adjust their profit margins rather than incur the costs associated with frequent price changing. Our FX rate variability measure is the quarterly realized variance defined as the sum of daily squared logarithmic FX returns over a quarterly window, $\sqrt{\sum_{j=1}^{D} [\log(\frac{NEER_{j,t}}{NEER_{j,t-1}})]^2}$ where D is the number of days in each quarter. In order to smooth out noise we consider as explanatory variable the square root of the moving average (over the most recent year) realized variance.

Inflation. Lower levels and variability of inflation are expected in countries where the monetary authority is more vigilant and credible at fighting inflation. Importing countries with chronic high level and volatility of inflation rates are expected to experience higher ERPT as argued in Taylor (2000), Gagnon and Ihrig (2004), and Choudhri and Hakura (2006). *Inflation* is defined as the annual logarithmic change in the consumer price

[8] Our main motivation for considering one-period-lagged covariates as opposed to contemporaneous is twofold. One is to rule out endogeneity issues in subsequent regressions of the estimated ERPT elasticities on the macrovariables. The other is to shed light on the predictive role of these macrovariates in an out-of-sample sense.

index (CPI) of the importing country, $\log(\frac{CPI_{i,t}}{CPI_{i,t-4}})$.

Size. Models of price discrimination rationalize higher pass-through in small countries in terms of the lack of local substitutes (Dornbusch, 1976). The smaller the importing economy the more likely it is that the local substitutes are sparse which, in turn, hinders the expenditure-switching effects associated with FX fluctuations; hence, the larger the power of the exporter which translates into higher ERPT. We consider as driver of ERPT at t the previous 5-year moving average annual growth in real GDP, $\log(\frac{GDP_{i,t}}{GDP_{i,t-4}})$. According to the above argument, a steady positive growth over a 5-year period implies an expansion of the domestic market which, in turn, should have a negative effect on import ERPT.

Output gap. A measure of the country-specific business cycle stage is obtained as deviations of actual real GDP from potential real GDP. A positive output gap signals that the economy is running above potential, namely, the demand is expanding and a lower ERPT may be observed in this context if export companies try to "fill the gap" (i.e. to gain market share) by absorbing the FX fluctuations in their profit margins in order to quote competitive prices. The output gap is computed as the difference between real GDP and a Hodrick-Prescott (with standard 1600 parameter for quarterly frequency) filtered real GDP series for each country, $\log(\frac{GDP_{i,t}}{GDP^*_{i,t}})$. Real GDP is the nominal value of output in national currency deflated using the GDP deflator.

Income. Another piece of evidence that would support *price discrimination* by exporting firms would be the finding that rich countries tend to experience lower import pass-through than poor countries ceteris paribus. To examine this conjecture, we define income in relative terms as the logarithmic ratio of the importing country's real GDP per capita to the world's real GDP per capita. The latter is defined as $GPC_t^{world} \equiv \sum_{i=1}^{N} GPC_{i,t}$ where N is the total cross-section (37 countries) in our sample.

Import penetration. The weight of import trades in the overall economy may influence the import ERPT elasticities.We consider as covariate the *self-sufficiency index*, sometimes also referred to as *import dependence* index, defined as the share of total domestic demand that is satisfied by imports, $\frac{M_{i,t}}{GDP_{i,t}+M_{i,t}-X_{i,t}} \times 100$, where $M_{i.t}$ and $X_{i,t}$ are, respectively, the total amount of imported and exported goods by country i and GDP is its nominal output; all of the data is in compatible current U.S. dollar units. Although some studies have tried to pin down the effect of trade openness on ERPT to the overall price level, to the best of our knowledge there

are no theoretical models as yet linking trade openness and the ERPT to import prices at the dock.[9] In a highly competitive environment, export firms may be more inclined to reduce import ERPT in the context of less import open countries in order to penetrate those markets, i.e. to gain market share. Under full ERPT, a domestic depreciation makes imports dearer thereby discouraging imports – the impact depends on the price elasticity of demand for imports. It is possible that exporting firms regard countries with higher import trade shares as being characterized by relatively small price elasticities of import demand and, in turn, subject them to higher import ERPT. Both arguments support a positive nexus between openness and import ERPT.[10]

Protectionism. An import tariff increases the cost to importing firms and puts pressure on the price of imported goods in the local markets thus acting towards lowering the quantity of goods imported ceteris paribus. Hence, import tariffs are a type of trade barrier which represent limits-to-arbitrage and thus should be associated with lower or incomplete pass-through which is a particular violation of the Law of One Price (LOOP). The higher the barrier to entry for foreign products (imposed by the government in the domestic country) the more willing the exporter may be to absorb currency fluctuations in profit margins in order to stabilize the price in the currency of the importer thus seeking to "compensate" the importing firms for the higher tariffs they have to pay. We employ an import trade "tariff barrier" index representing the level of tariff rates constructed by Gwartney et al. (2010) from World Trade Organization *World Tariff Profiles* sources. An index level of 10 indicates absent tariff barriers and the level moves toward 0 as the tariff barriers increase.[11] Therefore a positive coefficient is interpreted as consistent with the theory, namely, as the trade barriers (or barriers to arbitrage)

[9] For instance, as argued in Akofio-Sowah (2009) there is a direct and indirect link, with opposite signs, between trade openness and pass-through to the overall price level. The direct (positive) link simply comes from the fact that exchange rate fluctuations are more strongly fed into prices for more import open economies ceteris paribus. On the other hand, trade openness may be negatively correlated with inflation providing an indirect (negative) link.

[10] Following exchange rate fluctuations, if ERPT is relatively high and close local substitutes are available the substitution or expenditure-switching effects that follows will be large and hence, the import demand will be relatively elastic. If import demand is perceived as highly elastic in a given country, it is more likely that the foreign firms (exporters) are willing to 'cushion' currency depreciations with lower markups or equivalently, less ERPT; hence, for depreciations the expected relation between import demand elasticity and ERPT is negative. Under the same argument, the effect for appreciations would be the opposite: if the import demand elasticity is perceived as high or increasing, exporters could be prepared to sacrifice improved margins by increasing the pass-through too so as to make foreign goods more attractive (cheaper) to the importer seeking to gain market share; hence, for appreciations the expected relation is positive. To capture the effect of the perceived import demand elasticity on the pricing policy of exporters distinctively for appreciations versus depreciations one can interact the demand elasticity with a sign indicator variable (I^{sign}) for FX rate changes which is explained next.

[11] The data is available at annual frequency so we adopt a simple "step function" interpolation method where the four quarters of each year are set at the same value.

increase the extent of pass-through (ERPT elasticity) decreases and vice versa.

FX rate changes. The direction and the size of exchange rate changes may impart nonlinearity (*asymmetry*) of regime-switching type in the behaviour of ERPT elasticities. Two distinct regimes of import pass-through may be present corresponding, respectively, to depreciations and appreciations of the importer's currency. Under complete ERPT, depreciations (appreciations) imply one-to-one increases (decreases) in the price paid by importers for the incoming foreign products. However, a lower ERPT elasticity for appreciations than for depreciations ($\beta^{app} < \beta^{dep}$) would be consistent with both downward price stickiness and foreign firms' capacity constrains. In the short run, if foreign firms have reached full capacity it may be difficult for them to respond to upward export demand pressure that may accompany a reduction of import prices in the destination country (higher ERPT) following appreciations and thus they may opt for not passing favourable exchange rate fluctuations to the importer. On the contrary, as argued by Marston (1999) foreign firms may increase import pass-through during appreciations and reduce it during depreciations in order to quote competitive import prices if they are seeking to gain market share ($\beta^{app} < \beta^{dep}$). The majority of the few papers in the literature that examine the role of asymmetries with respect to the FX rate, define depreciations as increases in the exchange rate at t with respect to its previous level at $t-1$ (i.e. $\Delta s_t = s_t - s_{t-1} > 0$). Accordingly, we define

$$I_{i,t}^{sign} \equiv \begin{cases} 1 & \text{if} \quad \Delta s_{it} > 0 \text{ (depr.)} \\ 0 & \text{if} \quad \Delta s_{it} \leq 0 \text{ (appr.)} \end{cases}$$

and order to obtain measures of β^{app} and β^{dep} we consider two covariates simultaneously as drivers of the ERPT elasticity, $|\Delta s_{i,t}|$ and $I_{i,t}^{sign}|\Delta s_{i,t}|$, with regression coefficients, respectively, b_1 and b_2. Effectively this implies rewriting (4) as $\beta_{i,t} = \beta_i + b_1|\Delta s_{i,t-1}| + b_2 I_{i,t}^{sign}|\Delta s_{i,t-1}|$ which is particularized as $\beta_{i,t} = \beta_i + \lambda^{depr}|\Delta s_{i,t-1}|$ for FX rate depreciations with $\lambda^{depr} = b_1 + b_2$, and as $\beta_{i,t} = \beta_i + \lambda^{appr}|\Delta s_{i,t-1}|$ for appreciations with $\lambda^{appr} = b_1$.

Menu costs, which are like a fixed cost, may result in size asymmetry whereby import pass-through differs for small versus large changes in the exchange rate. The direction of the asymmetry depends on the type of invoicing followed by the exporter at the time the exchange rate change occurs. If the exporter is invoicing in his own currency, it may not be worthwhile to re-invoice following a small FX rate change which implies that its proceeds are kept intact and all of the change is born by the importer (large ERPT). On the contrary, under

local-currency invoicing again the presence of menu costs hinders the re-invoicing for small FX rate changes and hence, these translate fully into fluctuations in the exporters markup (small ERPT). In order to investigate this type of asymmetry, we define the size indicator:

$$I_{i,t}^{size} \equiv \begin{cases} 1 & \text{if} \quad \Delta s_{it} < \tau_{appr}^{-} \text{ or } \Delta s_{it} > \tau_{depr}^{+} \text{ (large)} \\ 0 & \text{if} \quad \tau_{appr}^{-} \leq \Delta s_{it} \leq \tau_{depr}^{+} \text{ (small)} \end{cases}$$

where $\tau_{appr}^{-} < 0 < \tau_{depr}^{+}$ are two threshold parameters that define, respectively, the cutoff point for large appreciations and large depreciations, namely, any positive FX rate change above τ_{depr}^{+} is regarded as a large depreciation and any negative FX rate change below τ_{appr}^{-} is deemed a large appreciation.

The above two effects can be explored by considering jointly as drivers of ERPT the four covariates, $|\Delta s_{i,t}|$, $I_{i,t}^{sign} I_{i,t}^{size} |\Delta s_{i,t}|$,$(1 - I_{i,t}^{sign}) I_{i,t}^{size} |\Delta s_{i,t}|$ and $I_{i,t}^{sign}(1 - I_{i,t}^{size})|\Delta s_{i,t}|$, with respective slope coefficients b_1, b_2, b_3 and b_4. Thus the differential effect on ERPT between depreciations and appreciations for small FX rate changes is given by b_4; the differential depreciation versus appreciation effect for large FX rate changes is given by $b_2 - b_3$; the differential large versus small change for appreciations is measured by b_3; and the effect of large versus small changes for depreciations is given by $b_2 - b_4$. The significance of the sign effect (for large changes) will be tested through a Wald test for the null hypothesis $H_0 : b_2 - b_3 = 0$ and so forth. Effectively, this model implies that the short-run ERPT mechanism may follow four distinct regimes dictated, respectively, by large appreciations, large depreciations, small appreciations and small depreciations. In order to endogenously determine the threshold parameters τ_{appr}^{-} and τ_{depr}^{+} we first subdivide the observed quarterly FX returns (pooled either across DMs or EMs) over the sample period into depreciations and appreciations, denoted $\{\Delta s_{i,t}\}^{+}$ and $\{\Delta s_{i,t}\}^{-}$, respectively. Over the sample period 1980Q1-2009Q3, the appreciations outnumbered the depreciations (57% of sampled quarters versus 43%) for the DMs and vice versa for the EMs (42% of sampled quarters versus 58%). In terms of magnitude, the largest appreciations are observed for DMs ranging between 0.002% and 312.84% whereas those for EMs are relatively milder with range (0.003%, 29.84%); the opposite is true for the depreciations in the range (0.002%, 21.991%) for DMs and (0.001%, 181.00%) for EMs. Rather than setting the thresholds τ_{appr}^{-} and τ_{depr}^{+} at some ad hoc pre-determined levels we estimate them by OLS alongside the other model parameters.[12]

[12] We conduct a grid search in the bidimensional set $S = (\{\Delta s_{i,t}\}_{0.05-0.5}^{-} \times \{\Delta s_{i,t}\}_{0.5-0.95}^{+}\})$, that is, the threshold candidates for τ_{appr}^{-} are those values between the 5th and 50th (median) percentiles of $\{\Delta s_{i,t}\}^{-}$ and the threshold candidates for τ_{depr}^{+} are the

Global economic sentiment. The analytical framework represented by equations (4)-(5) allows for the possibility that global or world factors can influence the degree of ERPT. Overall booming economic activity increases the demand for commodities which, in turn, puts upward pressure on commodity prices and possibly the ERPT. We employ as world factor the Global Purchasing Managers Index, defined as $Z_t \equiv \log(PMI_t)$, a GDP-weighted index of individual PMIs for the largest 24 developed and emerging economies covering 77% of world GDP.[13] Each individual PMI is an average of surveys for about 400 purchasing managers maintained by the Institute for Supply Management (ISM): a reading above 50 signals improving economic conditions and vice versa. It contains actual data elements but also a forward-looking confidence element thus making the PMI an extremely timely indicator of economic activity with a lot of value on Wall Street, and it is recognized as leading indicator by the Fed as borne out by its mention in the FOMC minutes.[14]

Appendix C illustrates cross-country differences in the above macroeconomic factors (with the exception of PMI which is global) on average over the entire 1980Q1-2009Q3 sample period. The graphs confirm some stylized facts such as the relatively high inflation levels and FX rate volatility of emerging markets, the soaring growth rate of China, the relatively low income levels of emerging markets, the small import dependence of the U.S., the relatively high import tariff rates of emerging markets and the hefty depreciation episodes in some emerging markets such as Brazil, Argentina, Turkey, Mexico and Venezuela.

3 Empirical Analysis

3.1 Linear Framework

The short-run (SR) and long-run (LR) import pass-through elasticities obtained from the first-difference model (1) and the two linear ECM formulations, (2) and (3), are presented in **Table 1** alongside the $\bar{R}^2$, AIC and

values between the 50th-95th percentiles of $\{\Delta s_{i,t}\}^+$. Following the least squares principle, the threshold estimates are those that minimize the residual sum of squares, i.e. $(\tau^-_{appr}, \tau^+_{depr}) = \arg\min_S RSS$.A further refinement consists of allowing the thresholds to be country-specific, ie τ_i^{depr} and τ_i^{appr} for $i = 1, ..., N$.

[13] The countries are Australia, Austria, Brazil, China, Czech Rep., Denmark, France, Germany, Greece, Hong Kong, India, Ireland, Israel, Italy, Japan, Netherlands, New Zealand, Poland, Russia, Singapore, South Africa, South Korea, Spain, Switzerland, Taiwan, Turkey, US and UK.

[14] The logarithmic quarterly change in the PMI is significantly positively correlated with quarterly commodity price inflation as measured by the first logarithmic difference in the Commodity Research Bureau (CRB) index, at 60%, and with changes in the MSCI All Country World Index (ACWI) of stock market activity at 82%.

SBC goodness-of-fit criteria. The results are obtained in two settings: using the longest available data span for each country (Table 1a) and focusing on the most recent 51 quarters spanning the 1997Q1-2009Q3 period for all countries (Table 1b); the latter setting makes the cross-country comparisons and coefficient averaging more meaningful. Each of the ERPT elasticity estimates (SR and LR) is subjected to two Wald tests for the hypotheses, respectively, of zero ($H_{0,A} : ERPT = 0$) and complete ($H_{0,B} : ERPT = 1$) import pass-through. Asterisks are used in Table 1 to denote rejections of $H_{0,A}$ whereas bold indicates rejection of $H_{0,B}$. Shaded areas are used to indicate the superior model according to each of the goodness-of-fit criteria.

In all three models, the vast majority of ERPT elasticities fall between 0 and 1.[15,16] However, the two ECM equations, (2) and (3), tend to outperform the first-differences model (1) in terms of in-sample goodness-of-fit. These findings are common to the whole unbalanced panel setting (Table 1a) and the shorter balanced panel setting (Table 1b). The results for the two ECMs are very similar both in terms of coefficient plausibility and overall goodness-of-fit but equation (3) has the advantage over equation (2) of having half less parameters to estimate, 7 versus 14. Hence, following the parsimony principle the former is preferred over the latter.

As a follow up to the above comparison, we deploy several cointegration tests: i) the time-series bounds Wald test developed in the context of equation (3) by Pesaran, Shin and Smith (2001) to test the null hypothesis of no long-run comovement ($H_0 : \theta_{i,1} = \theta_{i,2} = \theta_{i,3} = 0$), ii) Johansen's (1998) time-series sequential trace-type cointegration test based on a trivariate VECM specification, iii) Pedroni (2004), Kao (1999) and Maddala-Wu (1999) panel cointegration tests. The results are set out in **Table 2**. Shaded areas for each country are used to signify evidence of cointegration from at least one of the two time-series tests. The panel cointegration tests that jointly exploit the cross-section and time-series dimension of the data also clearly provide supportive evidence of

[15] Although it is widely accepted that implausible pass-through elasticities are those beyond [0,1], theoretically it is also possible to justify pass-through coefficients greater than 1 in terms of an amplification effect (see Knetter, 1993). Krugman (1987) justifies theoretically negative pass-through coefficients in the context of luxury goods. Feenstra (1989), Froot and Klemperer (1989) and Campa and Goldberg (2005) report ERPT estimates between -2.26 and 2.55.

[16] The model in first-differences, equation (1), gives 4 implausible elasticities whereas the two linear ECMs give 8 each. The implausible coeffcients from (1) are: 1 significantly negative for Norway (short-run ERPT) and 3 significantly above unity for South Africa (SR and LR) and Belgium-Luxembourg (LR). The implausible coefficients from (2) are: 4 significantly negative for, respectively, Norway (SR), Chile (SR), China (LR) and US (LR) and 4 significantly above unity for, respectively, South Africa (SR and LR), Pakistan (LR) and Turkey (LR). The implausible coeffcients for (3) are: 4 significantly negative for, respectively, Argentina (LR), Norway (SR), China (LR) and Chile (LR) and 4 significantly above unity for, respectively, South Africa (SR), South Africa (LR), Pakistan (LR) and Ireland (LR).

cointegration.[17] The upshot is that the three variables (p_{it}, p^*_{it}, s_{it}) move together in the long-run and therefore the linear ECM equation (3) is a reasonable baseline framework to estimate ERPT elasticities. We discuss next the time-constant ERPT elasticity estimates from this model.[18]

As shown in Table 1a, the zero short-run ERPT hypothesis is rejected in 11 out of 18 EMs and virtually in all, 20 out of 21, DMs; complete pass-through is also rejected in the short run for the vast majority of EMs (15) and DMs (12). Hence, the evidence corroborates partial short-run import ERPT for most countries. Unsurprisingly, in the long run it is more difficult to reject the hypothesis of complete pass-through. The short and long run ERPT elasticities for the US are relatively small in line with many previous studies (e.g. Frankel et al., 2005; Campa and Goldberg, 2005) and adding support to the consensus that exporters to the US market are more prepared to "cushion" themselves the exchange rate fluctuations instead of passing them on. This could be due to the size and hence, pricing power, of the US market

The simplest way to exploit the panel structure of the sample while allowing for full heterogeneity (i.e. all the model parameters are country-specific) is to average the ERPT elasticities across countries; Pesaran and Smith (1995) call this the mean group (MG) estimator. We report the panel MG estimates and corresponding tests of zero and unit pass-through separately for the EMs and DMs. Quite clearly, the hypothesis $H_{0A} : ERPT = 0$ is strongly rejected in the short- and long-run for both groups of countries. Another common thread among EMs and DMs is that the hypothesis of complete pass-through ($H_{0B} : ERPT = 1$) is rejected only in the short-run. These findings are common to the unbalanced and balanced panels.

[17] To preserve space, for the time-series Johansen tests we only report results for the H_0 that there are 0 cointegrating relations; we observe that for all the countries where this H_0 is rejected, the subsequent test statistic for the H_0 that there is at most 1 cointegrating relation is either not rejected or rejected marginally at the 10% level. Pedroni's and Kao's panel tests have in common that they are residual-based in the same spirit of the Engle-Granger two-step approach for time-series. Among the different panel test statistics proposed by Pedroni, the two reported allow for heterogeneity under the alternative hypothesis. The Fisher-type test can be interpreted as a pooled Johansen test.

[18] The comparison of the three linear specifications (1), (2) and (3) was also conducted on the basis of an out-of-sample forecasting exercise. A rolling window estimation approach is adopted to generate 20 one-step-ahead out-of-sample forecasts corresponding to the 2004Q3-2009Q3 period. See **Appendix Table B1**. The same exercise is repeated over a longer out-of-sample period 1999Q4-2009Q3 (40 quarters) in **Appendix Table B2**. In order to increase the reliability of the model estimates, we constrain this second exercise to the 28 countries (9 emerging and 19 developed) for which the length of the estimation (rolling) window is at least 75 quarters. The two forecast evaluation criteria adopted, root mean squared error (RMSE) and mean absolute error (MAE), indicate a clear dominance of the two ECM models over the first-difference model and the ECM specification (3) outperforms the ECM specification (2) in producing more marked reductions in RMSE and MAE. Hence, it seems reasonable to conclude that the inclusion of the long-run (cointegration) relationship and the lagged dependent (import price change) variable as regressors affords improvements in forecast accuracy. On this basis and the parsimony principle, model (3) is favoured over (2).

The balanced-panel MG estimates in Table 1b suggest that the short-run ERPT is lower than the long-run ERPT, at 0.494 versus 0.717 for EMs and at 0.744 versus 1.210 for DMs; although there are individual country exceptions. There are also cases where the long-run ERPT elasticity is lower than the short-run ERPT. This anomaly which has also been documented in other studies could be reflecting exporters' overreactions to exchange rate changes.[19] The ERPT elasticities (both short-run and long-run) seem larger for DMs than for EMs, a novel result in the light of the previous literature. It is interesting to note as well that both the long-run and short-run elasticities from this (balanced) shorter panel starting in 1997Q1 are higher than those from the unbalanced panel (Table 1a) where for a majority of countries the data starts in 1980Q1; the balanced panel MG estimates of the long-run ERPT are, in fact, closer to unity both for EMs and DMs. These findings challenge the conclusion that pass-through rates are declining over time due to more stable macroeconomic conditions of the importing countries (Taylor, 2000; Goldfajn and Werlang, 2000).

3.2 Drivers of country and time-variation in ERPT elasticities

In this section we seek, first, to uncover the most significant observable drivers of the import pass-through using a two-step analysis. Second, the linear ECM specification (3) will be generalized to accommodate time-varying/nonlinear behaviour according to each of the macroeconomic drivers, and a forward looking out-of-sample forecasting exercise for import prices will be conducted.

3.2.1 Two-step regression analysis

In order to investigate the presence of time-variation in ERPT elasticities and, in turn, the role of different drivers to explain it, we adopt the following two-step approach.[20] First, we obtain a time series of short-run

[19] For instance, Campa and Goldberg (2005; Table A1, p.690) report several such cases at disaggregate level in various sectors, e.g. food, energy and raw materials.

[20] We considered an alternative approach to these panel regressions as follows. First, we compute averages (over time) of the betas and the macroeconomic drivers as detailed in **Appendix C**. Second, a cross-section regression is estimated to explain the average betas on the basis of the average macro factors. The results uncovered only two factors, GDP growth and the output gap, as significant explanatory variables of the country variation in ERPT elasticities; the $\bar{R}^2$ is 11.61% and 6.26%, respectively. In both cases the slope coefficient is negative in line with the theoretical arguments presented in Section 2.3. However, this pure cross-section approach has two drawbacks. One is that by averaging the data over time it masks the power of the macro factors in explaining the time-evolution in ERPT elasticities, for instance, it does not us to explore the presence of sign and size asymmetries with respect to the time-evolution of the FX rate. Another is that it does not permit us to control for the presence of unobserved country heterogeneities (possibly correlated with the regressors) which could introduce biases.

ERPT elasticities, $\{\hat{\beta}_{i,t}\}_{t=1}^{T_1}$, by estimating the linear ECM equation (3) sequentially using rolling-windows (Step 1). The estimated ERPT elasticities $\hat{\beta}_{i,t}$ are then regressed (Step 2) on the observable drivers lagged one quarter, i.e. $\hat{\beta}_{i,t} = f(Z_{i,t-1}) + \varepsilon_{it}$. This approach is similar in spirit to that adopted in Campa and Goldberg (2005). However, they only obtained four ERPT elasticity estimates per country by dividing the sample period into 4 subperiods.

In the Step 1 regressions there is a trade-off between the length of the estimation (rolling) windows to obtain the $\{\hat{\beta}_{i,t}\}_{t=1}^{T_1}$ sequence, T_0, and the sample size for the Step 2 regressions, T_1, We primarily adopt $T_1 = 21$ which implies a rolling-window length of $T_0 = 99$ quarters for the vast majority of countries; thus the number of quarterly ERPT elasticities available for the regressions is 21 pertaining to the 6-year period 2004Q3-2009Q3. As a robustness check we also consider $T_1 = 41$ quarters corresponding to the 11-year period 1999Q3-2999Q3 which implies a rolling window length $T_0 = 79$ quarters.

We start by plotting in **Figure 2** the rolling short-run ERPT elasticity estimates over the 2004Q3-2009Q3 period ($T_1 = 21$ quarters) for the same 6 EMs (Figure 2a) and 6 DMs (Figure 2b) as in Figure 1.[21] A virtually monotonic downward trend is observed in some of the ERPT elasticities (e.g. Germany, US, Turkey), a virtually monotonic increase (e.g. Hungary, Pakistan and South Africa), a convex U-shape (e.g. Colombia, Hong Kong) and a monotonic decrease (increase) until the onset of the recent financial crisis with a reversal for France (Japan). **Figure 3** plots the ERPT elasticities over the recent 6-year period 2004Q3-2009Q3 averaged separately for DMs and EMs; in addition to the IMF classification we present the results for two slightly different country lists that are provided by MSCI Barra and by the FTSE Group (and Standard and Poor's).[22] All graphs suggest that the gap between the average ERPT elasticity between DMs and EMs has narrowed down over the last 6 years of the sample. This figure reveals that on average the import pass-through of DMs is higher than that for EMs. In a paper by Coulibaly and Kempf (2010) based on a relatively large sample of 27 emerging countries ending recently in 2009Q1 it is shown that inflation targeting in emerging countries

[21] The complete set of plots is available from the authors upon request.

[22] The classification of some of the countries is not unambiguous. Hong Kong, Singapore and Israel are categorized as DMs by the MSCI, FTSE and S&P but they are considered EMs by the IMF's World Economic Outlook October 2008 Report, and by *The Economist* magazine and J.P.Morgan as of today. South Korea is classified as DM by the FTSE and S&P but as EM by the MSCI, IMF and *The Economist*. For our current sample of countries the classification by the IMF and *The Economist* coincide, and that of the FTSE and S&P are also identical.

has helped to reduce the pass-through to import prices, in particular, the contribution of exchange rate shocks to price fluctuations in emerging markets declines after adopting inflation targeting. Hence, the finding that the average ERPT is higher for DMs than EMs in Figure 3 can be explained by the fact that the proportion of inflation targeters over this recent 6-year period is higher in the EM group (14 out of 18, or 78%) than in the DM group (7 out of 19, or 37%). The figure also reveals that the EMs pass-through is not declining over time. Frankel et al. (2005) also provide some evidence that FX pass-through to price of imports on the dock for developing countries has actually gone up and they provide a potential rationale in terms of declining transportation costs. The figure suggests also that the pass-through for DMs is gradually declining which in line with Campa and Goldberg's (2005) findings albeit on the basis of a different annual disagreggated (at product level) dataset over a less recent 1990-2001 period.

With these results in place, we fit now the Step 2 regression of the ERPT elasticities, $\hat{\beta}_{i,t}$, $t = 2004Q3, ..., 2009Q3$ obtained in Step 1 on the one-quarter lagged macroeconomic drivers outlined in the previous section (denoted $Z^j_{i,t-1}, j = 1, .., k$). **Table 3** presents the (un-weighted) averaged pairwise correlations between the different ERPT driver candidates considered in the regressions which are relatively small thus ruling out multicollinearity issues. The average correlations conceal a large degree of country heterogeneity, for instance, the correlation between the output gap and PMI ranges from a minimum of -36.26 (China), followed by -17.95 (Norway) and -17.11 (Czech Rep.), to a maximum of 49.23 (Turkey) followed by 49.92 (US) and 49.23 (Turkey). **Table 4** presents the estimation results using pooled fixed effects (FE) and random effects (RE) panel estimators.[23] Two sets of results are reported labelled as "2-way" and "1-way": the former pertain to models that accommodate both *unobserved* country and time heterogeneities whereas the latter refer to unobserved heterogeneity of cross-section type only. First, we carry out a full pooling across all ($N = 37$) countries and accordingly, the panel estimates provide measures of the nexus between import ERPT and macroeconomic drivers on average across emerging and developed economies. Second, we introduce a country dummy (EM) which equals 1 or 0 depending on whether the country is classified as EM or DM. This country dummy is interacted which each of the macro-

[23] Three countries are excluded from these panel regressions because we do not have data for some of the macrovariables; hence, the panel estimation results are for 34 countries over 21 quarters (714 observations). For Belgium imports and exports data is missing from 1997Q1. For Pakistan we were unable to find nominal and real GDP data. For Venezuela, nominal GDP is missing.

economic drivers which will enable comparisons across the two set of countries. The total number of pooled observations $T_1N = 20 \times N$ is 740. Inferences are based on White cross-section covariances that are robust to cross-equation (contemporaneous) correlation and heteroskedasticity; see Wooldridge (2002; p.148-153). The GLS covariance matrix for the composite error in the random effects model is based on the quadratic unbiased Swamy-Arora estimator which uses residuals from the within (fixed effect) and between (means) regressions.

The above models require the estimation of the threshold parameters (τ^-_{appr},τ^+_{depr}) alongside the remaining coefficients which is conducted via a grid search over the pooled quarterly FX rate changes; there is a total of 43% depreciations and 57% appreciations over the 2004Q3-2009Q3 period. The threshold estimates at $(-5.3\%, 3.9\%)$ roughly correspond, respectively, to the 8th percentile of the distribution of appreciations and the 73th percentile of depreciations. The interpretation is that any quarterly FX rate fall below 5.3% and rise above 3.9% are deemed, respectively, a "large" appreciation and depreciation on average across all countries over the 2004Q3-2009Q6 period.[24] The 2-way specification that controls for unobserved country- and time-heterogeneity and the 1-way (without the global PMI factor) specification that only controls for unobserved country effects are quite close indirectly suggesting that the *unobserved* time heterogeneity plays a relatively minor role in explaining the time-variation of ERPT elasticities; this is also borne out by the variance decomposition of the composite error term $\nu_{it} \equiv \delta_i + \gamma_t + \epsilon_{it}$ in the random effects models which suggests that 98% of the total error variance is accounted by the unobserved country-specific error component ($\sigma^2_\delta/\sigma^2_\gamma = 0.98$). The coefficient estimates from the random and fixed effects approaches are quite close both quantitatively and qualitatively; moreover, the Hausman test is only marginally in favour of the fixed effects specification.[25] On the right-hand-side of Table 4 we report the estimation results for panel models that include an emerging country dummy (equal to 1 if the importing country is EM; 0 if DM) interacted with each of the regressors; the Hausman test strongly suggests at the 1% level (statistic 64.88; p-value 0.00) that the random effects estimates are inconsistent so for space considerations we only report the fixed effects estimates.

[24] These are obtained by minimizing the residual sum of squares of the 1-way FE model that includes the country EM dummy. The remaining models produced virtually identical estimates.

[25] The Hausman test statistic applied to the 1-way (unobserved country-effects) model is significant albeit only at the 10% level (statistic 19.39; p-value=0.080) suggesting that there may be correlation between those unobserved effects and some of the included regressors.

The power of the fixed effects models to explain the ERPT process is quite high but a large amount of it can be ascribed to unobserved country-effects.[26] The adjusted-R^2 of models with macrovariables only (i.e. ignoring unobserved country or time effects) is 13.23% (33.16%) for the model without (with) an emerging country dummy and the corresponding F-tests are also strongly significant at the 1% level. Hence, there is considerable unobserved cross-section heterogeneity in the time-varying behaviour of ERPT. **Figure 4** plots the estimated fixed effects (as deviations from the average) that represent unobserved or hidden country-specific factors (i.e. over and above the macrovariables discussed above) that influence upwards or downwards the ERPT elasticity. For the DMs, the most negative effects are observed for New Zealand and the US suggesting that these two countries have the strongest pricing power as import destination markets *ceteris paribus* whereas Ireland and The Netherlands lie at the other extreme with the highest import ERPT effects. The implication is that for similar levels of the included country-specific macrovariables (FX rate volatility, inflation and so forth) the import ERPT elasticity of the United States is -0.59 units below average and that of The Netherlands is 0.26 units above average. For the EMs, China followed by Singapore exhibit the largest market power as importing markets and so their import ERPT elasticity is, respectively, 0.47 and 0.42 units below average, whereas South Korea and South Africa lie at the other extreme with unobserved effects that push the import ERPT elasticity up, respectively, by 0.62 and 0.69 units relative to average. Possible hidden factors that may lie behind the estimated fixed effects and, in turn, can explain country heterogeneities in ERPT over and above macrofactors are the credibility of the monetary policy stance and country differences in the composition of the import bundle. **Figure 5** represents the estimated time effects alongside the cyclical Global PMI evolution. The correlation between the two time series at 41.63% (p-value=0.0605) is significant at the 10% suggesting that the estimated time effects reflect partly the cyclical evolution of the overall economy as measured also by the PMI. The 1-way models with a global business cycle (PMI) indicator provide a better fit than the 2-way models.

We now discuss the main findings with respect to the role of the macro factors. On the basis of the above considerations, the inferences are based on the 1-way fixed effects model with a global business cycle

[26] The reported R^2 and F statistics for the fixed effects models are based on the difference between the RSS from the estimated model, and the RSS from a single-constant-only specification, not from a fixed-effect-only specification. Therefore these statistics are typically large because they reflect the explanatory power of the entire specification, including the estimated fixed effects.

(PMI) covariate although many of the inferences are robust across models. Several macrovariables play a role as significant drivers of ERPT elasticities confirming different hypotheses stemming from the theoretical literature. Inflation has a significantly positive effect on import pass-through and the effect is more clearly perceived for importing EMs. Another strong driver is the country's wealth (rich versus poor) as measured by the importing country's GDP per capita relative to "world" GDP per capita which is also more noticeable for EMs: the coefficient is significantly negative suggesting pricing-to-market behaviour of exporters, namely, poor countries are subject to larger import pass-through then rich ones. GDP growth is also a significant driver of pass-through for EMs with a negative sign suggesting that an increase in GDP of the import country over five consecutive years is on average followed by a decrease in import pass-through: as the domestic economy grows in size it is more likely that substitutes become available and hence, this makes exporters more prone to cushion FX rate changes against their profit margins. The import dependence index which may act as proxy for the price elasticity of imports (as perceived by exporters) is significantly and positively linked with the ERPT for EMs.

The output gap has a significantly negative effect although only for DMs implying that growth of the importing country over its potential can be perceived as an opportunistic "gap to fill" by exporting firms which may then reduce the ERPT (accepting lower margins) in an attempt to gain export market share. Likewise, tariffs play a significant role as pass-through drivers for DMs and the positive sign suggests that the higher the import tariffs the lower the level of pass-through; hence, tariffs (a type of trade barrier) act as a disincentive to arbitrage which, in turn, leads to lower import pass-through in line with the Law of One Price theory. The global economic sentiment indicator (PMI) appears significant although only for DMs. This suggests that the global cyclical (expansions versus recessions) evolution is relatively unimportant in explaining time variation in import pass-through for EMs. The positive sign suggests that the ERPT evolution for DMs has also a cyclical component — the pass-through increases in expansions and decreases in recessions; booming economic activity is likely to be reflected in higher commodity prices which increases production costs and puts upward pressure in the import ERPT if the destination market is developed. The volatility of the FX rate on average over the previous year has a positive coefficient for all model specifications suggesting a positive influence on ERPT but the effect is statistically insignificant throughout.

The asymmetric impact of FX appreciations versus depreciations is noticeable for both DMs and EMs. The direction (sign) of the relationship depends on whether the FX rate change is small or large.[27] For EMs, large appreciations are more strongly passed into import prices than large depreciations and vice versa for small exchange rate changes. For DMs, appreciations are also more strongly passed than depreciations but only if they are small; the opposite effect is found for large exchange rate changes.[28] Intuitively the contrast in the direction of the sign asymmetry between EMs and DMs may be explained by the willingness of exporting firms to gain market share in emerging markets. If appreciation of the importer's currency on quarter t-1 increases the ERPT elasticity on the subsequent quarter t this will make foreign products cheaper for the importer at the expense of a markup loss for the exporter which could be "compensated" by a future gain in market share.

Another hypothesis that follows from theoretical menu cost and hysteresis models is that import prices more strongly take on the FX rate fluctuations when these are large in magnitude. The panel model estimates in Table 4 support this conjecture for DMs in the context of depreciations and for EMs in the context of appreciations. For DMs, the size asymmetry effect is only present for depreciations; intuitively this means that, following depreciations of the importer's currency, exporting firms are more willing to bear them by reducing profit margins when they are small ones than if they are large; for small depreciations the menu type costs involved in changing import prices (and risk of subsequent loss of market share) outweigh the markup loss involved in leaving prices unchanged.[29] For EMs there is an interesting contrast between appreciations (large ones are passed more strongly than small ones) and depreciations (large ones are passed less strongly than small ones); however, both effects can act towards making exporting firms gain emerging market share.

[27] We also considered simplified model versions where the explanatory variables are simply $|\Delta s_{i,t}|$, $I_{i,t}^{sign}|\Delta s_{i,t}|$ and $I_{i,t}^{size}|\Delta s_{i,t}|$, which amounts to imposing the restriction that the extent of the sign asymmetry for small changes is equal to that for large changes, and the extent of the size asymmetry for appreciations is the same as that for depreciations. The goodness-of-fit of the models was somewhat worse that of the more general models reported in Table 4. For instance, the 1-way FE model (with emerging country dummy) has an $\bar{R}^2 = 98.606\%$.

[28] On the basis of a product-disaggregated dataset Marston (1999) provides evidence that the pass-through is larger for appreciations than depreciations but he does not examine the extent and direction of the sign asymmetry for small versus large FX rate changes.

[29] For EMs the results suggest that sign asymmetry is only apparent for small changes and size asymmetry is only apparent for depreciations. In the context of the sign and size asymmetry model formulation outlined in the previous section this implies $b_2 = b_3 = 0$ and hence, the two prevailing asymmetry effects are given by $(\beta^{dep} - \beta^{app})_{\text{SMALL}} = b_4$ and $(\beta_{\text{LARGE}} - \beta_{\text{SMALL}})^{dep} = -b_4$. This is in line with the results in Table 4 since a Wald test suggests that -5.4198 is statistically insignificant from -5.8590.

3.2.2 Out-of-sample forecasting

With the above analysis in place, we turn now attention to the forecast improvement afforded by the model inclusion of observable macroeconomic covariates as drivers of ERPT. The benchmark forecasting model is the baseline linear ECM equation (3) that portrays a time-constant ERPT. This model is generalized to accommodate time-variation in the ERPT elasticity as in (5) according to each of the observable drivers discussed previously and its forecasting power is examined. The forecasting exercise is conducted on the basis of individual time-series models (per country) in order to allow for full heterogeneity in the coefficient estimates. Moreover, each of the macroeconomic drivers is examined at a time in order to save degrees of freedom and also to enable comparisons of their relative predictive power. A rolling window approach is adopted to generate 20 conditional one-quarter-ahead forecasts over the 2004Q4-2009Q3 period; the fixed length of the rolling estimation window is 99 quarters.[30] In order to mitigate the bias introduced by Jensen's inequality, the log import price forecasts are transformed into level forecasts using the bias-corrected transformation $\hat{p}_{it} = \exp(\log \hat{p}_{it} + \frac{1}{2}\hat{\sigma}^2)$ where $\hat{\sigma}^2$ is the residual variance of the model at hand instead of the simple exponential transformation. The average forecast losses corresponding to each of the models are reported in **Table 7** according to the two distinct loss functions implicit, respectively, in the root mean squared error (RMSE) and mean absolute error (MAE) criteria. The table presents also, in those cases where a forecast gain is observed, the magnitude of the forecast error reduction in percent relative to the linear benchmark and the average and cumulative forecast error reduction (over countries) for each of the macroeconomic covariates.

The out-of-sample forecast analysis suggest that there is some predictive content in each of the drivers but there are important differences across countries. Particularly for the EMs, the import tariff stands out by bringing the largest average and cumulative forecast error reduction followed closely by the global PMI, volatility

[30] The first and last estimation windows correspond, respectively, to the period 1980Q1-2004Q3 and 1984Q4-2009Q2. The sample size of the rolling estimation window is fixed at 99 quarters *maximum*; the effective sample size is slightly smaller for some of the models due to the moving averaging and first differencing involved in some of the macro covariates. In some cases it is notably smaller as dictated by the data availability for some of the macro covariates. For instance, in the models that include the realized FX volatility (one-year moving average) as driver, the initial estimation window size is 39 since the series start in 1993Q4 and the last estimation window size contains 58 observations. Even more constrained are the models that include tariffs as driver since the data is available from 2000: the initial window contains 18 quarters only and the last window contains 37 quarters. See Appendix A for more details.

of the FX rate and import penetration index; at the other extreme, the less salient predictors are the sign and size asymmetry and the output gap. It is interesting to note that these results corroborate the well-known fact that in-sample and out-of-sample performance do not necessarily coincide (c.f. Table 4 and Table 5). For instance, the in-sample exercise suggested that the FX volatility is not a significant driver of ERPT whereas the out-of-sample forecasting exercise suggests that it provides important forecast error reductions in the majority of EMs on average at 16.46% and 14.53% according to the RMSE and MAE, respectively. Likewise for the global PMI which was suggested as a marginally significant driver of ERPT in the in-sample analysis. For the DMs, the import tariff and global PMI also bring the largest forecast error reduction. Overall, the degree of reduction in RMSE and MAE suggests that the merit of allowing for time-variation in ERPT for short term (one-quarter-ahead) forecasting purposes is slightly more prominent for the emerging economies than for the developed ones; in particular, the linear forecasting models enhanced with tariffs and the global PMI afford decreases in the RMSE for 14 emerging countries but only for 8 developed economies. The forecast error reduction associated with the size asymmetry of the FX rate change is larger than that of the sign asymmetry, for both DMs and EMs.

4 Conclusions

Exchange rate pass-through into prices has been the subject of a vast literature. This paper contributes to it by providing a systematic investigation of the empirical importance of macroeconomic factors in explaining country and time-variation in the extent of import exchange rate pass-through. It departs from existing studies in that it exploits a large and relatively recent sample 1980Q1-2009Q3 with a large cross-section of both emerging (18) and developed (19) countries. Moreover, our export price proxy is a trade-weighted average of export unit value indices to reflect the bilateral trade composition instead of relying on producer or consumer type indices in foreign currency. In contrast with existing studies, our analysis of import pass-through is based not only on typical in-sample statistical measures but also on an out-of-sample conditional forecasting exercise.

Time-series and panel tests provide reasonably strong evidence of cointegration, namely, they confirm the presence of a long-run equilibrium relation between the domestic import price, foreign effective export price and

the corresponding effective exchange rate. Rolling-window estimates do not support the view that the import pass-through is gradually falling: for some countries (particularly, developed) it has decreased over time from the year 2004 until present time whereas for others (particularly, emerging) it has increased instead. Overall the in-sample analysis suggests that several macroeconomic factors play a role in explaining time-variation in the ERPT and there are significant sign and size asymmetries in the response of the ERPT elasticity to the change in the FX rate; but there are notable differences across emerging and developed countries. An unanimous result in the out-of-sample forecasting exercise is that tariffs and the global business cycle evolution contain large predictive power for both emerging and developed economies. The volatility of the FX rate and the global business cycle evolution were suggested as weak drivers by the in-sample statistical analysis but these macro factors stand out in terms of out-of-sample predictive power. The opposite is true for the sign and size asymmetries. Therefore the findings also corroborate the well-known fact that in-sample statistical significance (or lack thereof) is not always tantamount to out-of-sample forecast (in)accuracy .

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