Currency Choice and Exchange Rate Pass-through*

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Abstract

In the open economy macro literature with nominal rigidities, the currency in which goods are priced has important implications for optimal monetary and exchange rate policy and for exchange rate pass-through. We show, using novel data on currency and prices for U.S. imports, that even conditional on a price change, there is a large difference in the pass-through of the average good priced in dollars (25%) versus non-dollars (95%). We document this to be the case across countries and within disaggregated sectors. This finding contradicts the assumption in an important class of models that the currency of pricing is exogenous. We present a model of endogenous currency choice in a dynamic price setting environment and show that the predictions of the model are strongly supported by the data.

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1 Introduction

In the open economy macro literature with nominal rigidities, the currency in which goods are priced has important implications for optimal monetary and exchange rate policy and for exchange rate pass-through. In a large class of models used to evaluate optimal policy prices are exogenously set either in the producer currency or in the local currency.[1] In these models, in the short run when prices are rigid, pass-through into import prices of goods priced in the producer's currency is 100% and it is 0% for goods priced in the local currency. However, when prices adjust there is no difference in pass-through. With firms thus exogenously assigned to currencies there can be sizeable departures from allocative efficiency. A fundamental question then is whether indeed it is the case that when prices adjust pass-through is unrelated to the currency of pricing. We address this question both empirically and theoretically in this paper.

We show, using novel data on currency and prices for U.S. imports, that even conditional on a price change, there is a large difference in the pass-through of the average good priced in dollars (25%) versus non-dollars (95%). We document this to be the case across countries and within disaggregated sectors. We then present a model of endogenous currency choice in a dynamic price setting environment and show that these findings are consistent with the predictions of the model. As firms mimic the short-run flexible price benchmark by choosing currency optimally, the deviations from allocative efficiency are less severe.

In the empirical work we evaluate pass-through in several ways. First, we construct the average monthly price change of import prices from each country into the U.S., for goods priced in dollars and in non-dollars separately. We then estimate pass-through from the exchange rate shock into prices over time. We find that there is a large difference in pass-through into U.S. import prices of the average good priced in dollars versus the average good priced in non-dollars at all horizons up to 24 months. One month after the shock pass-through is nearly zero for goods priced in dollars and nearly complete for goods priced in non-dollars, consistent with the substantial amount of nominal price stickiness in the

[1] For instance, Obstfeld and Rogoff (1995) assume producer currency pricing, Betts and Devereux (2000) and Chari, Kehoe, and McGrattan (2002) assume local currency pricing. Devereux and Engel (2003) allow prices to be exogenously set both in local and producer currencies.

data. More interestingly, pass-through 24 months after the shock is only 0.17 for dollar priced goods and 0.98 for non-dollar priced goods. The difference in pass-through therefore declines from 1 on impact to around 0.8 at 24 months.

Second, we evaluate at the good level exchange rate pass-through conditional on a price change. We document that even conditioning on a price change there is a large difference in the pass-through of the average good priced in dollars (25%) versus non-dollars (95%). This significant difference is shown to be present for individual countries and within sectors as detailed as the 10 digit level that have a mix of dollar and non-dollar pricers. We show that similar patterns also hold for exports from the U.S., even though only a neglibible fraction of exports is priced in non-dollars. Specifically, exports priced in the producer currency (dollars) exhibit 84% pass-through conditional on adjustment, while exports priced in local currency (non-dollars) pass-through only 25% conditional on a price change.

We also estimate the difference across dollar and non-dollar priced goods in pass-through after multiple rounds of price adjustment. We find that, while there remains a significant difference in pass-through, this gap is considerably smaller than conditional on the first instance of price adjustment. This arises because dollar priced goods pass-through 49% after multiple adjustments, which is twice as high as after the first adjustment.

There are other important differences between dollar and non-dollar pricers. At the 1-2 digit harmonized code level sectors that would be classified as producing more homogenous goods (following the Rauch (1999) classification) are dollar priced sectors. Sectors such as 'Animal or Vegetable Fats and Oils', 'Wood and articles of Wood' and 'Mineral Products' are dominated by dollar pricers. On the other hand there is a greater share of non-dollar pricers in the 'Footwear', 'Textiles and textile articles', 'Machinery and mechanical appliances' sectors that are in the differentiated sector. We also document that non-dollar pricers adjust prices less frequently than dollar pricers.

The evidence presented in the empirical section are new facts that need to be matched by models of open economy macroeconomics. We next examine how these facts compare with the predictions of standard open economy models. Is the evidence consistent with a model where firms choose which currency to price in? While there exist several important papers[2] in the theoretical literature on currency choice, our paper is most closely related to Engel (2006) who presented an important equivalence result between pass-through and the currency of pricing. The two main departures we make from the existing literature are, one, we consider a multi-period dynamic (as opposed to a static) price setting environment, and two, we provide conditions under which a sufficient statistic for currency choice can be empirically estimated using observable prices.

Currency choice is effectively a zero-one indexing decision of the firm's price to exchange rate shocks. If prices adjust every period, currency choice is irrelevant. However when prices are sticky, the firm can choose its currency to keep its preset price closer to it's desired price (the price it would set if it could adjust flexibly). We show that a sufficient statistic for currency choice is the average desired pass-through over the period of non-adjustment. This *medium-run* pass-through is determined by both the dynamic path of desired pass-through and the duration of non-adjustment. This result does not rely on the particular details of the economic environment or the specific source of incomplete pass-through, that is whether it is driven by variable mark-ups, imported inputs or decreasing returns to scale in production.

Intuitively, if a firm desires low exchange rate pass-through in the short run—before it has a chance to adjust prices—the firm is better off choosing local currency pricing that results in 0% pass-through in the short run. Conversely, if short-run desired pass-through is high, the firm should choose producer currency pricing that results in complete (100%) pass-through prior to price adjustment.

An important insight the model delivers is that currency choice cannot be predicted solely by long-run pass-through or desired pass-through on impact of the exchange rate shock. As a result, a firm, for instance, with a high flexible price (long-run) pass-through can well choose local currency pricing if real rigidities lead to a low desired pass-through in the medium-run. In addition we develop conditions under which medium-run pass-through can be measured using actual prices chosen by firms. This is equal to the pass-through conditional on the first adjustment to the exchange rate shock, as estimated in the empirical section.

[2]The papers in this literature include Giovannini (1988); Friberg (1998); Bacchetta and van Wincoop (2003, 2005); Devereux, Engel, and Storgaard (2004); Corsetti and Pesenti (2004); Goldberg and Tille (2005).

We also study numerically a currency choice model with two sources of incomplete long-run pass-through — demand-driven variable mark-ups and imported inputs in production. We show that a firm is more likely to select into producer currency pricing the lower the elasticity of its mark-up and the lower the share of imported inputs in its production cost. In addition we show that the main results of the analytical section hold up well to several extensions and the empirical estimate of pass-through conditional on first price adjustment robustly approximates medium-run pass-through.

The predictions of the theoretical model find strong support in the empirical evidence, as we document that pass-through conditional on adjusting prices is significantly greater for non-dollar as compared to dollar pricers. Further, as theory predicts, long-run pass-through is a less relevant statistic to evaluate currency choice with pass-through after multiple rounds of price adjustment exceeding 0.5 for dollar pricers. Consistent with theory, non-dollar pricers adjust less frequently than dollar pricers. We also observe an important sorting of goods, with dollar pricers being predominant in the homogenous sector and non-dollar pricers in the differentiated goods sector, which is consistent with endogenous currency choice.

The debate on whether currency choice is exogenous or endogenous is important as it can significantly alter our understanding of exchange rate determination and optimal exchange rate policy. It is well understood (following Caplin and Spulber (1987)) that the same median frequency of price adjustment can lead to differing outcomes for policy depending on whether frequency is chosen endogenously or exogenously. A similar reasoning applies to the currency in which prices are set. With exogenous currency choice, one obtains stark results such as the optimality of floating under PCP that ensures expenditure switching and of pegging under LCP that ensures consumption risk-sharing (Devereux and Engel, 2007). This stark difference arises because firms are forced to price in one or the other currency irrespective of their flexible price desired pass-through. Once they are allowed to choose currency optimally, they will choose it to fit their desired pass-through patterns enhancing the effective amount of price flexibility and reducing the welfare gap between floating exchange rates and pegs.

There are other important considerations that arise with endogenous currency choice in a general equilibrium environment, as pointed out in Devereux, Engel, and Storgaard (2004) and Bacchetta and van Wincoop (2005). In such an environment exchange rate volatility affects currency choice that in turn effects exchange rate volatility generating the possibility for multiple equilibria. Consequently, a country that follows more stable monetary policies will experience greater price stability as more of the exporters to that country set prices in the country's currency.

The paper proceeds as follows. Section 2 presents empirical evidence. Section 3 develops the theory and provides numerical simulations. Section 4 interprets the empirical findings through the lens of the proposed theory of currency choice, evaluates alternative interpretations and concludes. All proofs are relegated to the Appendix.

Data

We use unpublished micro data on import prices collected by the Bureau of Labor Statistics for the period 1994-2005. This data are collected on a monthly basis and contain information on import prices of a very detailed good over time, with information on the country of origin and the currency of pricing. Details regarding the underlying database are reported in Gopinath and Rigobon (2008).

In the price survey the BLS asks firms to report on the currency of denomination of the price. Gopinath and Rigobon (2008) document that prices are rigid, with a median duration of 11 months, in the currency in which they are reported as being priced in. This fact suggests that the currency information is meaningful and it is not the case, for instance, that firms price in non-dollars and simply convert the prices into dollars to report to the BLS. This would imply that dollar prices would then show a high frequency of adjustment, which is not the case.

Around 90% of U.S. imports in the BLS sample are reported as priced in dollars. This fraction however varies by country of origin. The fraction of imports in the exporter currency is, for example, 40% from Germany, 21% from Japan, 16% from U.K. and 4% from Canada. From all developing countries the share in the exporters currency is close to zero. As is well known a significant fraction of trade (40% of the BLS sample) takes place intra-firm. Since we will test theories of prices that are driven mainly by market forces we exclude intra-firm prices from our analysis.[3]

In our empirical analysis we include countries that have a non-negligible share of their exports to the U.S. priced in both dollar and non-dollar currency. This includes Germany, Switzerland, Italy, Japan, UK, Belgium, France, Sweden, Spain, Austria, Netherlands and Canada.[4] Table 1 lists the total number of goods (second column) for each country and the fraction invoiced in the exporter's currency (last column).

2 Pass-through and Currency in the Data

In this section we present empirical evidence on pass-through across dollar and non-dollar priced goods. It follows mechanically that during the period when prices do not adjust, exchange rate pass-through into the dollar price of goods is 0% for the goods priced in dollars and 100% for the goods priced in non-dollars. The question we ask is what is the difference in pass-through once prices of these goods *change*.

To address this question we present two main types of regression results. First, we present estimates from a standard aggregate pass-through regression using average monthly price changes, with the difference that we construct separate series for dollar and non-dollar priced goods. We show that the difference in pass-through into U.S. import prices of the average good priced in dollars versus the average good priced in non-dollars is large even at a two-year horizon. Second, we present micro-level regressions where we condition on the price having changed. We show that even conditional on a price change there is a large difference in the pass-through of the average good priced in dollars (25%) versus non-dollars (95%). We present both country-wise and sector-wise evidence that supports these results.

[3]For empirical evidence on the differences between intra-firm and arms-length transactions, using this data, see Gopinath and Rigobon (2008) and Neiman (2007).

[4]We used the following two formal criteria for selection: (1) a country should have at least 10 items priced in non-dollars; and (2) at least 5% of all items imported from a country should be priced in non-dollars. We made an exception for Canada, which has only 4% of exports to U.S. priced in Canadian Dollars, but this constitutes a large number of goods since Canada is an important trade partner.

2.1 Aggregate Evidence

For each country we construct two separate monthly average price change series — one including only goods that are priced in dollars and the other using only those goods priced in the exporter's currency. For most countries, for exports to the U.S., these are the only two types of pricing. Some goods are priced in a third currency, but such instances are rare. The average price change was calculated using equal weights, since we were not provided with BLS weights at the good level.

We estimate the following standard pass-through regression,

$$\Delta p_{k,t} = \alpha_i + \sum_{j=0}^{n} \beta_j \Delta e_{k,t-j} + \sum_{j=0}^{n} \gamma_j \pi_{k,t-j} + \sum_{j=0}^{3} \delta_j \Delta y_{t-j} + \epsilon_{k,t}, \tag{1}$$

where k indexes the country, Δp is the average monthly log price change in dollars, π is the monthly foreign country inflation using the consumer price index, and Δy is average GDP growth in the U.S.; n is the number of lags which varies from 1 to 24. Since the data is monthly, we include up to 24 lags for the nominal exchange rate and foreign inflation and 3 lags for GDP growth. We estimate specification (1) for the full sample of goods and for the two sub-samples of dollar and non-dollar priced goods.

The statistic of interest is the sum of the coefficients on the nominal exchange rate: $\beta(n) \equiv \sum_{j=0}^{n} \beta_j$. These coefficients reflect the impact that the current change in the exchange rate has on the price of imports over time. The objective is to compare these estimates across currencies as we increase the number of lags included in the specification from 1 to 24. Figure 1 depicts the pass-through coefficients, $\beta(n)$, from estimating a pooled regression of all countries (with country fixed effects) against the number of lags, n, on the x-axis. The line in the middle depicts the average pass-through for all goods. This measure of pass-through increases from 0.21 with one lag to 0.30 with 24 lags. The feature that at the aggregate level most of the pass-through takes place in the first two quarters and levels off soon after is consistent with the findings of Campa and Goldberg (2005) and others who have estimated pass-through into the U.S. using the BLS price index.

¿From just this aggregate measure however it is impossible to discern the role of currency. Now we consider the separate currency series. The top line depicts the pass-through for the non-dollar priced goods. The bottom line is the pass-through for the dollar priced goods. The bands represent the 95% confidence interval around the point estimate for each lag specification. As Figure 1 demonstrates, the regression using only the contemporaneous and 1 month lag of the exchange rate estimates a pass-through of close to 0 for goods priced in dollars and close to 1 for goods priced in non-dollars, consistent with the substantial nominal rigidity in the data. Further, we observe that the pass-through increases for the dollar items with the inclusion of lags. Note that the pass-through into the dollar priced goods is far more gradual than pass-through into the aggregate index. Pass-through decreases slightly for the non-dollar index. We might expect to see the two pass-through numbers converge as we get past periods of nominal rigidity. A striking feature of the plot however is that the gap between pass-through of the dollar and non-dollar index remains large and significant even 24 months out. At 24 months the pass-through is 0.17 in the dollar sub-sample and 0.98 in the non-dollar sub-sample.

In Figure 2, we replicate the aggregate regressions country by country. Notice that the aggregate level of pass-through varies substantially across countries. This can be seen from the middle line in the plots. For instance, for Germany, the pass-through is around 40 percent at all horizons; for Japan and U.K., the numbers are smaller, as they increase from 23 percent to 32 percent; while for Sweden and France pass-through is always smaller than 20 percent. The difference between the pass-through of the dollar and non-dollar price series is again striking. The exception to this is Canada where the two pass-through elasticities intersect.[5] For all other countries the differences remain large even at long horizons and for 9 out of the 12 countries the difference is significant even for the specification with 24 lags. The two exceptions, other than Canada, are Austria and Netherlands for which there is simply not enough data to statistically distinguish the two pass-through elasticities at 24 months horizon. Notice that in all other countries the confidence intervals for the dollar and non-dollar pass-through do not intersect.

An alternative empirical specification would be to use individual (good level) price

[5]Average pass-through for Canada increases from 12 percent in the short run to almost 45 percent at 24 months, though these numbers are highly imprecisely estimated. One must be cautious to interpret the evidence for Canada since the Canadian exchange rate is more likely to be driven by the price of its main export—commodities—rather than the other way round (see Chen and Rogoff (2003)).

changes for each country as the left hand side variable, without averaging across observations. So instead of one monthly observation for a country we have a panel of monthly price changes for a country. We perform this panel regression where the left hand side is the monthly change in the price of a good. We also include fixed effects for each country and primary strata (BLS defined sector code, mostly 2-4 digit harmonized trade code) pair and cluster the standard errors at this level. The impulse responses from this estimation are reported in Figure 3. As is evident the impulse responses are almost identical to those presented in Figure 1.

To summarize, aggregate pass-through even two years after an exchange rate shock is strikingly different for dollar and non-dollar priced goods, with non-dollar pass-through being much higher than dollar pass-through. Consequently, the fraction of goods that are priced in different currencies has significant predictive power for measures of aggregate import pass-through, even at very distant horizons.

Since prices change infrequently in this sample with a median duration of 11 months, aggregate price indices are dominated by unchanging prices. Increasing the horizon of estimation to several months so as to arrive at the flexible price pass-through does not solve this issue because around 30% of the goods in the BLS sample do not change price during their life, i.e. before they get replaced. Consequently, when estimating pass-through using the BLS index such prices have an impact on measured pass-through even at long horizons. In the next section we estimate micro-level regressions that condition on a price change.

2.2 Micro-level Evidence

At the good level, we estimate the following regression:

$$\Delta \bar{p}_{i,t} = \left[\beta_D \cdot D_i + \beta_{ND} \cdot (1 - D_i)\right] \cdot \Delta_c e_{i,t} + Z'_{i,t}\gamma + \epsilon_{i,t}, \tag{2}$$

where i indexes the good; $\Delta \bar{p}_{i,t}$ is the change in the log dollar price of the good, *conditional on price adjustment in the currency of pricing*;[6] $\Delta_c e_{i,t}$ is the cumulative change in the log

[6]In the BLS database, the original reported price (in the currency of pricing) and the dollar converted price are both reported. We use the latter, conditional on the original reported price having changed. Since the first price adjustment is censored from the data, we also perform the analysis excluding the first price change and find that the results are not sensitive to this assumption.

of the bilateral nominal exchange rate over the duration for which the previous price was in effect. D_i is a dummy that takes the value of 1 if the good is priced in dollars and zero if the good is priced in non-dollars. $Z_{i,t}$ includes controls for the cumulative change in the foreign consumer price level, the US consumer price level, the US GDP and fixed effects for every BLS defined primary strata (mostly 2-4 digit harmonized codes) and country pair. The coefficients of interest are β_D and β_{ND} that estimate pass-through conditional on price adjustment for dollar and non-dollar pricers respectively.

The results from estimation of specification (2) are reported in Table 2. The point estimate for pass-through and the standard error for dollar and non-dollar pricers are reported in columns 2–5. The difference in the pass-through and the t-statistic of the difference is reported in columns 6–7. The number of observations, number of goods and R-squared are reported in the remaining columns. The first row reports the results from pooling all observations. The pass-through, conditional on a price change, to the cumulative exchange rate change, is 0.24 for dollar priced goods and 0.92 for non-dollar priced goods. The difference in these pass-through estimates is large and strongly significant.[7] We estimate this specification for each country and obtain similarly that there is a sizeable difference in the point estimate of dollar and non-dollar priced goods. This difference is statistically significant for 9 out of the 11 countries, the exceptions being Spain and Canada.

One might be interested in the subset of goods for which the firm has arguably more pricing power. Accordingly, in Table 3 we repeat the analysis for the sub-sample of differentiated goods, according to the Rauch (1999) classification.[8] The average pass-through for dollar priced firms is 0.24 and it is 0.96 for non-dollar priced firms, in line with the results for the

[7]In an environment with endogenous frequency, conditioning on a price change can induce a bias in the exchange rate pass-through estimates as it generates a conditional correlation between the exchange rate and the residual even if the unconditional correlation is zero. The direction of this correlation is different under LCP and PCP — negative in the first case and positive in the second case resulting in pass-through estimates that are biased upwards for LCP firms and downwards for PCP firms. This bias therefore works against finding a result of significant difference in the pass-through's of dollar and non-dollar pricers. The fact that we find a large difference implies that either this bias is small or the gap is even larger than we estimate. More details on the magnitude of this selection bias in the simulated data are provided in footnote 26 of Section 3.2.

[8]Rauch (1999) classified goods on the basis of whether they were traded on an exchange (organized), had prices listed in trade publications (reference) or were brand name products (differentiated). Each good in our database is mapped to a 10 digit harmonized code. We use the concordance between the 10 digit harmonized code and the SITC2 (Rev 2) codes to classify the goods into the three categories. We were able to classify around 65% of the goods using this classification.

full sample of goods. This difference is also observed at the country level, with the difference in the pass-through estimates significant for all countries except Spain and Canada.

Sectoral Evidence: In Table 4 we estimate pass-through (equation 2) at the disaggregated sectoral level.[9] The last column of this table reports the fraction of goods that are priced in non-dollars for the particular sector in the BLS sample. There is substantial variation across sectors. Most sectors that would be classified as producing more homogenous goods (using the Rauch (1999) classification) are dollar priced sectors. For instance, the sector 'Animal or Vegetable Fats and Oils' is wholly dollar priced. 'Live Animals, Animal Products', 'Wood and articles of Wood', 'Vegetable Products', 'Mineral Products' are dominated by dollar pricers. On the other hand there is a greater share of non-dollar pricers in the 'Footwear', 'Textiles and textile articles', 'Machinery and mechanical appliances' sectors. As we get down to more disaggregated levels most sectors in our sample are either all dollar or non-dollar priced. At the 6 (10) digit level, where there are 2 or more observations, 57% (63%) are all dollar-priced and 5% (7%) are all non-dollar priced (the sample sizes are however very small).

For 18 of the 19 sectors that have a mix of dollar and non-dollar pricers, the pass-through of non-dollar priced goods exceeds that of dollar priced goods and this difference is significant at conventional levels for 15 sectors. For one sector, 'Live Animals, Animal Products' the difference is negative, but not statistically significant. The share of non-dollar priced goods in this sector is also small.

We next examine differential pass-through within those 10 digit classification codes that have a mix of dollar and non-dollar prices. We find that dollar pricers have a pass-through of 30%, non-dollar of 95% and the difference is highly statistically significant.[10] Lastly, in our data set there are 125 items for which the currency of invoicing changed during the life of the good. However, not all of them have a price change within each regime. When

[9]This table can be compared to Table 8 in Gopinath and Rigobon (2008). The two differences are, one, the sample is smaller because we look at a subset of countries and two, we separate goods based on their currency of pricing. We include fixed effects for each 6 digit harmonized code and country pair and the standard errors are clustered at this level.

[10]Fixed effects for each 10 digit classification code were included and standard errors clustered at the country level.

we concentrate on that smaller sub-sample, and exclude the price change that exists when the shift in the currency of invoicing takes place, for the dollar invoiced items the pass-through is roughly 50 percent. The pass-through when the good is priced in the non-dollar currency is 68 percentage points higher than when it was priced in dollars and this difference is statistically significant.

Exports: This far we have focused on U.S. imports. In the case of U.S. exports the fraction of goods priced in dollars is overall 97.3% and there is very little variation across countries, unlike for imports. For instance, for exports to the Euro area this fraction is 96.4%. When we estimate pass-through conditional on price adjustment, as in equation (2), we find that exchange rate pass-through into local currency prices of goods that are priced in dollars (producer currency pricing) is 84%, while pass-through for goods priced in the importing country's currency (local currency pricing) is 25% and the difference has a t-statistic of 3.88. These results are very similar to those observed for imports across pricing regimes. Given the small fraction that are non-dollar pricers we cannot however perform detailed sectoral and country comparisons as was done for the case of imports.

2.2.1 Life-Long Pass-through

When we condition on a price adjustment we get past the period of strict nominal rigidity for the firm. However, if there are real rigidities that arise say from strategic complementarities in pricing and if competitor firms adjust at different points in time, then it is well known that a single price adjustment will not measure the full pass-through that occurs when all firms have adjusted their price. In this subsection we estimate a measure of pass-through that incorporates multiple rounds of price adjustment and consequently allows for fuller adjustment to an exchange rate shock.

We estimate life-long pass-through using price changes over the life of the good in the sample and measure its response to cumulative exchange rate movements over this period. Specifically, we estimate

$$\Delta \bar{p}^{L}_{i,T} = \left[\beta^{L}_{D} \cdot D_i + \beta^{L}_{ND} \cdot (1 - D_i)\right] \cdot \Delta_L e_{i,T} + Z'_{i,T}\gamma + \epsilon_{i,t}, \tag{3}$$

where $\Delta\bar{p}^L_{i,T}$ is the difference between the last observed new price of the good and the first price in the sample; and $\Delta_L e_{i,T}$ is the exchange rate change over the respective period.[11] We have therefore one observation for each good that has at least one price adjustment during its life in the sample and estimate the exchange rate pass-through over the life of the good. The estimates of life-long pass-through for dollar and non-dollar priced goods are β^L_D and β^L_{ND} respectively.

The results from estimation of the life-long specification (3) are reported in Table 5. Life-long pass-through is 0.49 for dollar-priced goods and 0.98 for non-dollar priced goods and the difference is statistically significant. The difference in life-long pass-through is also large and significant in different sub-samples, including differentiated goods and goods imported separately from Euro and Non-Euro areas, in some cases exceeding 0.5. While for non-dollar priced goods, life-long pass-through is not very different from pass-through conditional on first adjustment, it is important to note that for dollar-priced goods pass-through is about twice as high over the life of the good compared to the first round of adjustment. This implies that it takes far longer than one price adjustment for these goods to attain long-run pass-through.

2.2.2 Frequency and Size of Price Adjustment

Lastly, we report on differences in the frequency of price adjustment and the size of price adjustment conditional on a price change across dollar and non-dollar priced goods. As reported in Gopinath and Rigobon (2008), non-dollar pricers have longer price durations than dollar pricers (14 versus 11 months).[12] For the sub-sample of countries used in this paper this difference remains. As Table 6 reports, the median frequency for dollar pricers (0.10) is higher than that for non-dollar pricers (0.07). In terms of duration this is a difference of around 4 months. This difference partly reflects the fact that non-dollar pricers are in the differentiated sector where price durations are longer as reported in Gopinath and Rigobon

[11]See Gopinath and Itskhoki (2008) for more on this measure of long-run pass-through and how it closely approximates true flexible price pass-through in several standard models.

[12]This drives the negative relation between frequency and pass-through conditional on the first adjustment to the exchange rate shock reported in Table 11 of Gopinath and Rigobon (2008). Non-dollar pricers have longer durations and, as we show in this paper, they have higher pass-through conditional on a price change, which drives the relation in the table.

(2008). To see if this difference persists at a more disaggregated level we restrict the sample to those 6 digit sectors that have a mix of dollar and non-dollar goods. The median frequency for the dollar pricers is 0.13 and for the non-dollar pricers is 0.08. Similarly, restricting the sample to those 10 digit classification groups that have both dollar and non-dollar pricers we find that frequency for the dollar pricers is 0.11, while for non-dollar pricers it is 0.07. The standard deviation of the frequency measure is large across all specifications. The evidence at the disaggregated level is therefore consistent with the more aggregate level evidence that dollar pricers adjust prices more frequently than non-dollar pricers.

We perform a similar comparison for the size of price adjustment in the reported currency of pricing, conditional on a price change. The median size of price change and the median absolute size of price change are quite similar across dollar and non-dollar pricers. The median absolute size overall for the dollar pricers is 7 percentage points while it is 6 percentage points for the non-dollar pricers. At the 6 digit level again the difference in the medians is 1 percentage point, while at the 10 digit level the difference is 2 percentage points in favor of dollar pricers. There does not appear to be a systematic difference in the size of price adjustment across dollar and non-dollar pricers. Since size is scale dependent it is difficult to infer what the size measure implies about the responsiveness to shocks, as it depends also on the size of shocks. This is unlike pass-through, which is scale independent and measures the responsiveness to shocks.

3 Currency and Prices in a Dynamic Sticky Price Model

The evidence presented in the empirical section are new facts that need to be matched by models of open economy macroeconomics. How do these facts compare with the predictions of standard open economy models? Is the evidence consistent with a model where firms choose the currency to price in? To address these questions we develop a model of endogenous currency choice in a dynamic pricing environment with nominal rigidities.

While there exist several important papers in the theoretical literature on currency choice, our paper is most closely related to Engel (2006). The two main departures from the existing literature are, one, we consider a multi-period dynamic setting as opposed to a static environment, and two, we provide conditions under which a sufficient statistic for currency choice can be empirically estimated using observable prices.

In section 3.1 we consider a general model of currency choice in an environment with Calvo (1983) staggered price setting. Using second order approximations we provide an analytical characterization of the currency choice rule. In section 3.2 we numerically analyze a specific model of incomplete pass-though and evaluate the robustness of the analytical findings to various extensions.

3.1 Analytical Model of Currency Choice

We consider a partial equilibrium environment by focusing on the pricing and currency strategies of a single firm. For concreteness, consider a firm exporting its product into the U.S. Denote by $\Pi(p_t|s_t)$ its profit function from sales in the U.S. market, where p_t is the log of the current local currency (dollar) price of the firm and s_t is a state vector that can include demand conditions, cost shocks, competitors' prices and the exchange rate.

Define the desired price of a firm as the price it would set if it could costlessly adjust its price in a given state. The desired price of the firm in local currency is:

$$\tilde{p}(s_t) = \arg\max_{p} \Pi(p|s_t). \tag{4}$$

The log of the desired price in the producer currency is then $\tilde{p}_t^* = \tilde{p}_t - e_t$, where the asterisk indicates that the variable is denominated in the producer currency and e_t is the log of the bilateral nominal exchange rate defined such that an increase in e_t corresponds to an appreciation of the exporters exchange rate. Throughout Section 3.1 we make the empirically relevant assumption that the nominal exchange rate follows a random walk and relax this assumption in Section 3.2.

3.1.1 Price Setting in Local and Producer Currency

Now assume that the firm is allowed to adjust prices each period with an exogenous probability $(1-\theta)$, as in Calvo (1983). Further, when the firm adjusts its prices it can costlessly decide whether to fix the new price in the local currency or in the producer currency.

The value to the firm that sets its price in the local currency, $V_L(p|s^t)$, is characterized by the following Bellman equation:

$$V_L(p|s^t) = \Pi(p|s_t) + \delta\theta\mathbb{E}_t V_L(p|s^{t+1}) + \delta(1-\theta)\mathbb{E}_t V(s^{t+1}), \tag{5}$$

where $s^t \equiv (s_0, s_1, \ldots, s_t)$ denotes the history of states s_{t-j} and δ is the constant discount factor.[13] The expectations are conditioned on information as of time t contained in s^t and on whether the firm adjusts its price in period $t+1$ or not. If the firm does not adjust it receives the continuation value $V_L(p|s^{t+1})$ and if it does adjust it receives $V(s^{t+1})$. The optimal price set in the local currency can be expressed as:

$$\bar{p}_{L,t} \equiv \bar{p}_L(s^t) = \arg\max_p V_L(p|s^t). \tag{6}$$

Similarly, the value to the firm of setting its price in the producer currency, $V_P(p^*|s^t)$, is given by,

$$V_P(p^*|s^t) = \Pi(p^*+e_t|s_t) + \delta\theta\mathbb{E}_t V_P(p^*|s^{t+1}) + \delta(1-\theta)\mathbb{E}_t V(s^{t+1}), \tag{7}$$

and the optimal price set in the producer currency is

$$\bar{p}^*_{P,t} \equiv \bar{p}^*_P(s^t) = \arg\max_p V_P(p^*|s^t). \tag{8}$$

During the duration of the producer-currency price, its local currency value moves one-to-one with the exchange rate and is denoted by $p_{t+\ell} = \bar{p}^*_{P,t} + e_{t+\ell}$.

The value to the firm adjusting its price in period t that incorporates the optimal currency choice is then:

$$V(s^t) = \max\left\{V_L\big(\bar{p}_L(s^t)|s^t\big), V_P\big(\bar{p}^*_P(s^t)|s^t\big)\right\}. \tag{9}$$

This maximization problem determines the currency choice of the firm when it adjusts prices in period t.

We now show the following equivalence result for optimal price setting in the local and producer currency:

[13] The variable component of the discount factor can be incorporated into $\Pi(p|s^t)$, which then should be interpreted as the profit in real discounted units. Since we later introduce first order approximations to the pricing decisions, the potential variation in the stochastic discount factor does not affect the results.

Proposition 1 *The first order approximation to optimal price setting in local and producer currency is given respectively by:*

$$\bar{p}_L(s^t) = (1-\delta\theta)\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\tilde{p}(s_{t+\ell}),$$
$$\bar{p}_P^*(s^t) = (1-\delta\theta)\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\big\{\tilde{p}(s_{t+\ell}) - e_{t+\ell}\big\},$$

which implies the following equivalence between optimal prices in the local and producer currency:

$$\bar{p}_L(s^t) = \bar{p}_P^*(s^t) + e_t.$$

The first order approximations provide certainty equivalent pricing rules. Proposition 1 shows that two otherwise identical firms, one pricing in local currency and the other in producer currency, will set the same price (in the common currency) conditional on adjustment. The firm sets its price in each currency as an expected weighted average of future desired prices in the currency of pricing. The only difference between the desired prices in each currency is the nominal exchange rate, and the expectation of this exchange rate for any future date is the current value of the exchange rate given that it follows a random walk. Proposition 1 has an important corollary: If firms are assigned the currency of pricing exogenously, their pass-through conditional already on the first adjustment should be identical regardless of the currency of pricing.[14]

3.1.2 Currency Choice

Now consider the problem of currency choice. Define the difference between the value of local and producer currency pricing by:

$$\begin{aligned}\mathcal{L}(s^t) =& V_L\big(\bar{p}_L(s^t)|s^t\big) - V_P\big(\bar{p}_P^*(s^t)|s^t\big)\\ =& \sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\left\{\Pi\big(\bar{p}_L(s^t)|s_{t+\ell}\big) - \Pi\big(\bar{p}_P^*(s^t)+e_{t+\ell}|s_{t+\ell}\big)\right\},\end{aligned} \tag{10}$$

where the final expression is obtained by iterating the value functions defined in (5) and (7). Whenever $\mathcal{L}(s^t) > 0$ the firm should choose LCP and it should choose PCP otherwise.

[14]This corollary relies on the fact that the price-setting rule in Proposition 1 is optimal independently of whether currency choice is exogenous or endogenous (see Appendix).

When making the currency choice decision, a firm compares expected profits under the two invoicing arrangements during the period of price stickiness.

To shed more light on the currency choice decision we use a second order approximation to equation (10) and arrive at the following proposition,

Proposition 2 *The second order approximation to the difference in value of LCP and PCP is:*

$$\mathcal{L}(s^t) = K(s^t)\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathrm{var}_t(e_{t+\ell})\left[\frac{1}{2} - \frac{\mathrm{cov}_t\big(\tilde{p}(s_{t+\ell}), e_{t+\ell}\big)}{\mathrm{var}_t(e_{t+\ell})}\right], \tag{11}$$

where $K(s^t) \equiv -\partial^2\Pi\big(\tilde{p}(s_t)|s_t\big)/\partial p^2 > 0$*. Therefore, the firm chooses local currency pricing when*

$$\bar{\Psi} \equiv (1-\delta\theta)^2\sum_{\ell=1}^{\infty}(\delta\theta)^{\ell-1}\ell\frac{\mathrm{cov}_t\big(\tilde{p}(s_{t+\ell}), e_{t+\ell}\big)}{\mathrm{var}_t(e_{t+\ell})} < \frac{1}{2} \tag{12}$$

and producer currency pricing otherwise.

Currency choice is effectively a zero-one indexing decision of the firm's price to exchange rate shocks. If prices adjust every period, currency choice is irrelevant.[15] However when prices are sticky, the firm can choose its currency to keep its price closer to the desired price in periods when the firm does not adjust. Accordingly, as demonstrated in equation (12), currency choice is determined by the weighted average of exchange rate pass-through coefficients into desired prices, $\mathrm{cov}_t\big(\tilde{p}(s_{t+\ell}), e_{t+\ell}\big)/\mathrm{var}_t(e_{t+\ell})$, where the weights are the product of the discount factor and the probability of non-adjustment. The particular threshold value of 1/2 arises from the second order approximation. This result generalizes the main insight of Engel (2006) in a dynamic pricing environment.[16]

We refer to $\bar{\Psi}$ as *medium-run pass-through* (MRPT). The firm chooses local currency pricing when MRPT is low and producer currency pricing when MRPT is high. Intuitively,

[15] Formally, if the firm adjusts prices every period, $\theta = 0$, and from (11), $\mathcal{L} \equiv 0$.

[16] Since in Engel (2006) firms adjust prices every period, but *before* observing the current state of the world, the equivalent to (12) dynamic expression for currency choice would be

$$(1-\delta\theta)^2\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}(\ell+1)\frac{\mathrm{cov}_{t-1}\big(\tilde{p}(s_{t+\ell}), e_{t+\ell}\big)}{\mathrm{var}_{t-1}(e_{t+\ell})} < \frac{1}{2},$$

and if the firm adjusts prices every period ($\theta = 0$), it reduces to $\mathrm{cov}_{t-1}\big(\tilde{p}(s_t), e_t\big)/\mathrm{var}_{t-1}(e_t) < 1/2$, as in Engel (2006).

if a firm desires low exchange rate pass-through in the short run—before it has a chance to adjust prices—the firm is better off choosing local currency pricing that results in 0% pass-through in the short run. Conversely, if short-run desired pass-through is high, the firm should choose producer currency pricing that results in complete (100%) pass-through prior to price adjustment.

Note that the covariance terms in (12) are not conditional on any contemporaneous variables. That is, what matters for currency choice is the unconditional covariance of exchange rate shocks and desired prices independently of whether this is a direct relationship or mediated through other contemporaneous variables such as the industry price level, costs or demand.

An important insight follows by re-writing the medium-run pass-through as a weighted average of the desired price responses to exchange rate shocks:

$$\bar{\Psi} = (1-\delta\theta)\sum_{j=0}^{\infty}(\delta\theta)^j\left[(1-\delta\theta)\sum_{\ell=1}^{\infty}(\delta\theta)^{\ell-1}\tilde{\Psi}_{\ell,\ell+j}(s^t)\right], \tag{13}$$

where

$$\tilde{\Psi}_{j,\ell}(s^t) \equiv \frac{\mathrm{cov}_t\big(\tilde{p}(s_{t+\ell}),\Delta e_{t+j}\big)}{\mathrm{var}_t(\Delta e_{t+j})}$$

is the impulse response of the desired price in period $t+\ell$ to the exchange rate shock in period $t+j$ conditional on information available at time t.[17] The term in square brackets is a weighted average of the impulse response of the desired price in period $\ell+j$ to an exchange rate shock in ℓ, where j is held fixed. For example, when $j=0$ each term is the instantaneous response to an exchange rate shock and when $j=1$ each term is the response one period after the shock. As $j\to\infty$ each term is the long-run response to the exchange rate shock. In general, therefore, MRPT and currency choice depend not just on the long-run pass-through or desired pass-through on impact, but on the entire path of the desired pass-through responses weighted by the probability of price non-adjustment. We show below that this distinction can be quantitatively important.

[17] To obtain (13), we decompose:

$$\frac{\mathrm{cov}_t\big(\tilde{p}(s_{t+\ell}),e_{t+\ell}\big)}{\mathrm{var}_t(e_{t+\ell})} = \frac{\mathrm{cov}_t\big(\tilde{p}(s_{t+\ell}),e_t+\Delta e_{t+1}+\ldots\Delta e_{t+l}\big)}{\mathrm{var}_t(e_{t+\ell})} = \frac{1}{\ell}\sum_{j=1}^{\ell}\tilde{\Psi}_{j,\ell}(s^t),$$

where the last step uses the random walk property that $\mathrm{var}_t(e_{t+\ell}) = \ell\cdot\mathrm{var}_t(\Delta e_{t+j})$ for any $\ell, j\geq 1$.

3.1.3 Structural model of incomplete pass-through

To provide further insights into the currency choice rule we need to put more structure on the determinants of pass-through. We can express the log of the desired price in local currency as:

$$\tilde{p}_t \equiv \tilde{p}(e_t, P_t|z_t) = \mu(\tilde{p}_t - P_t|z_t) + mc^*(e_t|z_t) + e_t, \tag{14}$$

where P_t is the log of the sectoral price level, μ_t is the log desired mark-up and mc^* is the log of the marginal cost in producer currency; z_t denotes all other shocks that affect mark-ups and costs, but are uncorrelated with the exchange rate and sectoral price level.

Denote by ϕ_t the elasticity of the local currency marginal cost with respect to the exchange rate: $\phi_t \equiv \partial(mc_t^* + e_t)/\partial e_t$. This elasticity can be less than 1 if it is the case that some fraction of a foreign exporters costs is set in dollars, in which case exchange rate movements only partially effect the dollar cost of the firm. For instance, ϕ can represent the constant elasticity of output with respect to domestic inputs in a Cobb-Douglas production function.

Denote by Γ_t the elasticity of the mark-up with respect to the relative price of the firm: $\Gamma_t \equiv -\partial\mu_t/\partial(p_t - P_t)$. The mark-up channel of incomplete pass-through is the classic pricing to market channel of Dornbusch (1987) and Krugman (1987). In general, Γ varies with the state s_t. Consider specific instances when Γ can be treated as a constant. Firstly, under constant price elasticity of residual demand (e.g., CES), Γ is a constant equal to 0. A more interesting case, however, is when Γ is a constant that differs from zero, that is the mark-up varies but the elasticity of the mark-up is constant.[18] To make further progress in studying the optimal currency choice we assume that the elasticity of mark-up Γ is constant. We relax this assumption in the numerical explorations of the next section. With a constant elasticity of mark-up, we can prove:

Proposition 3 *Let ϕ and Γ be constant and consider the first order approximation to price setting. Then the impulse response of desired prices, $\tilde{\Psi}_{\ell,\ell+j}(s^t)$, is independent of ℓ and s^t*

[18] If mark-up variability is demand-driven, one can write a demand system that results in constant Γ. Since $\mu \equiv \ln[\sigma/(\sigma-1)]$, where σ is the price elasticity of demand, we have $\Gamma = \varepsilon/(\sigma-1)$, where ε is the price elasticity of σ. Therefore, the appropriate demand system solves the following second order differential equation:

$$\varepsilon = \Gamma(\sigma - 1), \qquad \sigma \equiv -\partial \ln q/\partial \ln p, \qquad \varepsilon \equiv \partial \ln \sigma/\partial \ln p,$$

where Γ is an arbitrary constant measuring the elasticity of mark-up and $q(p)$ is the demand schedule.

and depends only on j, the time elapsed after the exchange rate shock. Moreover, it can be written as:

$$\tilde{\Psi}_j = \frac{\phi}{1+\Gamma} + \frac{\Gamma}{1+\Gamma} \cdot \frac{\mathrm{cov}(P_{t+j}, \Delta e_t)}{\mathrm{var}(\Delta e_t)}. \tag{15}$$

Finally, the currency choice rule can be rewritten as:

$$\bar{\Psi} = (1-\delta\theta) \sum_{j=0}^{\infty} (\delta\theta)^j \tilde{\Psi}_j < 1/2. \tag{16}$$

The interpretation of (16) is exactly the same as that of (12), except that under the assumptions of Proposition 3, medium-run pass-through ($\bar{\Psi}$) and therefore currency choice does not depend on the state, that is currency choice effectively becomes a once and for all decision.

The impulse response j periods after the shock $\tilde{\Psi}_j$ has two terms. The first term is increasing in the firm's cost sensitivity to the exchange rate shock and decreasing in the elasticity of the mark-up. It follows from (15) that the desired pass-through profile is entirely shaped by the second term that is the impulse response of the sectoral price level to the exchange rate shock interacting with the elasticity of the mark-up. If, for example, strategic complementarities in price setting across firms are strong ($\Gamma > 0$), most firms price in the local currency, and price changes across firms are not synchronized, then $\tilde{\Psi}_j$ will have an increasing profile. As a result, the medium-run pass-through, $\bar{\Psi}$, will fall short of the long-run pass-through and the gap between the two will be increasing in the extent of the nominal price stickiness quantified by θ. Consequently, we can formulate the following corollary:

Corollary 1 *Let the desired pass-through profile be increasing, $\tilde{\Psi}_j \leq \tilde{\Psi}_{j+1}$. Then the firm is more likely to chose producer currency pricing, if (1) it has longer duration of prices (higher θ) and/or (2) everywhere higher pass-through profile (higher $\tilde{\Psi}_j$ for all j).*

3.1.4 Estimation of Medium-run Pass-through

The sufficient statistic for currency choice—medium-run pass-through—is a complex expression that depends on desired prices that are not observable. We show here that, under the assumptions of Proposition 3, there is a way to estimate this statistic without knowledge of the determinants of the desired pass-through profile.

Consider a micro-level regression of a change in the dollar price of the firm conditional on price adjustment in the currency of pricing on the most recent change in the exchange rate, for both LCP and PCP firms. Specifically, we want to evaluate the coefficient in the regression of

$$\Delta\bar{p}_t \equiv \begin{cases} \bar{p}_{L,t} - \bar{p}_{L,t-\tau}, & \text{for LCP,} \\ \bar{p}^*_{P,t} + e_t - \bar{p}^*_{P,t-\tau} - e_{t-\tau}, & \text{for PCP,} \end{cases}$$

on Δe_t. Here τ denotes the most recent price duration and hence $t-\tau$ is the previous instance of price adjustment. Denote the coefficient in this regression by β_{MR}. We can prove the following result:

Proposition 4 *Under the assumptions of Proposition 3, β_{MR} equals medium-run pass-through:*

$$\beta_{MR} \equiv \frac{\mathrm{cov}\big(\Delta\bar{p}_t, \Delta e_t\big)}{\mathrm{var}(\Delta e_t)} = (1-\delta\theta)\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\tilde{\Psi}_{\ell} = \bar{\Psi}.$$

Therefore, a relevant statistic for currency choice can be estimated by using observed prices and conditioning on a price change and regressing it on the exchange rate shock. The standard concern in pass-through regressions about omitted variables is not an issue here since what matters for currency choice is the unconditional correlation between exchange rate and prices.

Finally, note that β_{MR} is the counterpart to the coefficient in the empirical regression conditional on first price adjustment (2). The only difference is that in the empirical specification we regress $\Delta\bar{p}_t$ on the cumulative change in the exchange rate ($\Delta_c e_t \equiv e_t - e_{t-\tau}$), rather than on the one-period change ($\Delta e_t \equiv e_t - e_{t-1}$). Empirically, regressing $\Delta\bar{p}_t$ on Δe_t results in very noisy estimates due to the uncertainty about the exact timing of price change within the month. In the numerical calibration of the next section we show that the two regressions indeed result in very close estimates and the empirical specification (2) provides a good approximation to β_{MR}. Moreover, we show that this specification provides accurate estimates of medium-run pass-through, $\bar{\Psi}$, even when the assumptions of Proposition 3 are not satisfied.

3.2 Numerical Simulation

In this section we study numerically a standard model of incomplete pass-through and evaluate the currency choice rule. The purpose of this exercise is twofold. First, given a specific model we relate both pass-through and the currency decision to primitives of the economic environment. Second, within a more specialized quantitative model we are able to relax the assumptions imposed in the previous section to verify the robustness of the theoretical predictions. The overall conclusion is that the main results of the analytical section hold up well to several extensions and the empirical estimate of pass-through conditional on first price adjustment robustly approximates medium-run pass-through.

3.2.1 Demand and Costs

As in the theoretical section we retain the partial equilibrium set-up and consider the problem of a single firm facing exogenously given industry price level dynamics.[19] We introduce two standard channels of incomplete exchange rate pass-through — variable mark-ups and imported intermediate inputs. We then solve for the optimal pricing and currency decisions of the firm in an environment with nominal rigidities and consider both Calvo and menu cost pricing.

We adopt the Klenow and Willis (2006) specification of the Kimball (1995) aggregator that results in a demand schedule with non-constant elasticity.[20] Specifically, the demand schedule for the firm is

$$q = q(p, P) = [1 - \varepsilon(p - P)]^{\sigma/\varepsilon}\,, \tag{17}$$

where p is the log of the firm's own price and P is the log of the industry price level that aggregates the prices of the firm's competitors. This demand specification is conveniently governed by two parameters, $\sigma > 1$ and $\varepsilon > 0$, resulting in the following non-constant price

[19] In an earlier version of this paper (Gopinath, Itskhoki, and Rigobon, 2007), we numerically solved for an industry equilibrium and endogenous sectoral price level dynamics. However, all the important insights can be obtained in this simpler environment that provides the flexibility to introduce a number of extensions.

[20] We view Kimball demand as a useful abstraction for modeling mark-up variability arising from strategic interactions between monopolistic competitors. See Yang (1997) and Atkeson and Burstein (2008) for alternative models.

elasticity of the desired mark-up[21]:

$$\tilde{\Gamma} \equiv -\frac{\partial \mu}{\partial p} = \frac{\varepsilon}{\sigma - 1 + \varepsilon \cdot (p - P)}. \tag{18}$$

In steady state, $p = P$ and hence the steady state mark-up elasticity is given by $\Gamma = \varepsilon/(\sigma - 1)$.

The log of the firm's marginal cost is given by

$$c_t = \phi e_t - a_t. \tag{19}$$

where ϕ is the sensitivity of cost to exchange rate shocks and a_t denotes the log of the idiosyncratic productivity shock. This marginal cost function can be derived from a constant returns to scale production function that combines domestic and foreign inputs, where ϕ measures the elasticity of output with respect to domestic inputs. When $\phi < 1$, this limits the desired pass-through of the firm even if the desired mark-up is constant. We label this the imported intermediate inputs channel of incomplete pass-through.

The firm faces exogenous stochastic processes (a_t), (e_t) and (P_t), and decides on the optimal currency of pricing and price setting in a given sticky price environment. Specifically, in the Calvo case, it faces an exogenous probability $(1 - \theta)$ in any given period of being able to adjust its price and currency of pricing. The decisions of the firm are governed by the Bellman equations system (5), (7), (9) of the previous section. In the Menu Cost case, the firm faces a menu cost, κ, that it needs to pay to adjust its price and currency choice. The Bellman equations system for this case is provided in the appendix.

3.2.2 Calibration and Simulation Procedure

Here we describe the calibration and leave details of the simulation procedure to the appendix. The parameter values are reported in Table 7. The model is calibrated to monthly data. We assume that both a_t and e_t follow persistent autoregressive processes. We set the

[21] The elasticity and super-elasticity (elasticity of elasticity) of this demand schedule are given by:

$$\tilde{\sigma} \equiv -\frac{\partial \ln q}{\partial p} = \frac{\sigma}{1 - \varepsilon(p - P)} \qquad \text{and} \qquad \tilde{\varepsilon} \equiv \frac{\partial \ln \tilde{\sigma}}{\partial p} = \frac{\sigma}{1 - \varepsilon(p - P)}.$$

The log desired mark-up is $\tilde{\mu} \equiv \ln\left[\tilde{\sigma}/(\tilde{\sigma} - 1)\right]$ and hence its elasticity is $\tilde{\Gamma} = \tilde{\varepsilon}/(\tilde{\sigma} - 1)$. A useful feature of this demand specification is that it converges to constant σ-elasticity demand (CES) with constant desired mark-up when $\varepsilon \to 0$.

autocorrelation of the exchange rate process to 0.986 and hence can verify the robustness of the theoretical results to the random walk assumption. The standard deviation of the exchange rate innovation is calibrated to the data for developed countries bilateral nominal exchange rates of 2.5%.

The persistence of the idiosyncratic shock process is calibrated to 0.95 to match the evidence on the persistence of firm level productivity as reported in Bils and Klenow (2004). The standard deviation of the idiosyncratic shocks is set to 8% to match the average absolute size of price adjustment in the data of 7%. The Calvo probability of adjustment is set to match the duration of prices in the sample of 9 months and the menu cost parameter, κ, is chosen to match this duration in the menu cost environment. Specifically, in the baseline calibration, the menu cost is set to equal 5% of steady state revenues conditional on price adjustment, which constitutes less than a percent of revenues on an annualized basis, well within the standard range of menu cost estimates in the literature.

The sectoral price level is assumed to follow

$$(P_t - \bar{P}) = \alpha(P_{t-1} - \bar{P}) + (1-\alpha)\bar{\phi}e_t, \qquad \bar{P} = \ln\big[\sigma/(\sigma-1)\big]. \tag{20}$$

That is, the only source of shocks to the sectoral price level is the exchange rate. The inertia in the price level is measured by $\alpha \in (0,1)$ and $\bar{\phi} \in (0,1)$ is the long-run response of the price level to the exchange rate shock.[22] Since $\sigma/(\sigma-1)$ is the steady state mark-up, $p = P$ in the steady state.

The cost and demand parameters are calibrated as follows. The steady state elasticity of demand is set to $\sigma = 5$ which corresponds to a steady state mark-up of 25%, consistent with empirical evidence and other calibrations in this literature (Klenow and Willis, 2006). In the simulations we vary the steady state super-elasticity of demand, ε, and firm's cost sensitivity to the exchange rate, ϕ. Specifically, we vary ε on $[0,8]$ and ϕ on $[0.5,1]$.[23] This variation in ε translates into the variation in steady state elasticity of mark-up (Γ) on $[0,2]$.

[22]This sectoral price level process obtains in a linearized Calvo price setting industry equilibrium model. For our simulation, we set $\alpha = 0.95$ which corresponds to a 13.5 months 'half-life' for the sectoral price level. We also set $\bar{\phi} = 0.5$ which implies a 50% long-run pass-through into the sectoral price level so that the exchange rate shock is neither purely aggregate for the industry, nor purely idiosyncratic to the firm. Note that $\bar{\phi}$ can be interpreted as cost sensitivity to the exchange rate for the average competitor of the firm.

[23]Note that we always have $\phi \geq \bar{\phi}$ so that the cost of the foreign firm is indeed more sensitive to the exchange rate than the cost of an average firm in the industry.

The pass-through of idiosyncratic cost shocks into the desired price equals $1/(1+\Gamma)$, so that when, for example, $\Gamma = 1$, the pass-through of idiosyncratic cost shocks is 50%.[24]

To solve the model, we iterate numerically the Bellman operator that yields value functions and policy functions. From the value functions we directly compute the difference in the value of LCP and PCP, as in (10), which determines currency choice exactly. From the policy function we compute medium-run pass-through, $\bar{\Psi}$, according to (12), which determines currency choice approximately as it relies on the second order approximation to the value functions. We also compute medium-run pass-through using formula (13) derived under the restrictive assumption of a constant mark-up elasticity and evaluate the error resulting from it and show that it is small. We compute the long-run pass-through as the desired pass-through given the long-run response of the price level equal to $\bar{\phi}$. Lastly, using the policy function, we simulate a time series of exchange rates and firm prices to estimate the micro-level pass-through regression similar to (2) and evaluate the properties of this MRPT estimator.

3.2.3 Simulation Results

The results described correspond to Calvo pricing unless otherwise stated. First we examine how currency choice, medium-run pass-through (MRPT) and long-run pass-through (LRPT) respond to variation in the mark-up elasticity Γ (driven by variation in ε) and the cost sensitivity to the exchange rate ϕ. In Figure 4 we set $\phi = 0.75$ and vary Γ on $[0, 2]$. We observe that both MRPT and LRPT are declining in Γ with the wedge between the two increasing in Γ as firms put more and more weight on the sectoral price level in their own pricing decisions. The dashed vertical line separates the regions of local and producer currency pricing computed based on the *value function*. Note that the threshold of 1/2 for MRPT provides an accurate approximation for the currency choice rule. At the same time,

[24] Note that in our calibration we need to assume neither very large menu costs, nor very volatile idiosyncratic shocks, as opposed to Klenow and Willis (2006). There are a few differences between our calibration and theirs. First of all, our baseline value for super-elasticity of demand is $\varepsilon = 3$ and we almost never need ε greater than 5, as opposed to their baseline value of 10. In addition, they assume a much less persistent idiosyncratic shock process and match the standard deviation of relative prices rather than the average absolute size of adjustment. The interaction of these two deviations appears to drive the differences in the results.

LRPT stays above 1/2 for all values of Γ.

Figure 5 carries out a similar exercise, but now holds $\Gamma = 0.75$ (i.e., $\varepsilon = 3$) constant and varies ϕ on $[0.5, 1]$. As ϕ increases both MRPT and LRPT increase with the gap between them staying roughly constant. The dashed vertical line again separates the regions of local and producer currency pricing. At $\phi = 0.77$, the firm is indifferent between the currency of pricing and its MRPT is very close to 0.5, while its LRPT is quite a bit higher, equal to 67%.

The overall conclusions from Figures 4–5 are the following: First, pass-through is increasing in cost sensitivity to the exchange rate (ϕ) and decreasing in mark-up variability (Γ) making local currency pricing more appealing when ϕ is low and Γ is high. Second, comparing MRPT with the threshold of 1/2 indeed provides an accurate criterion for currency choice, while LRPT may not be a very useful measure for this purpose.

This is further evident in Figures 6 and 7. In Figure 6 we solve for the combination of Γ's and ϕ's for which the firm is indifferent between local and producer currency pricing using the value function. We do this for both Calvo and the menu cost model and plot the resulting relationship between ϕ and Γ. In both cases a firm chooses PCP if it has high ϕ or low Γ.

In Figure 7 we take the combinations of Γ and ϕ for which the firm is indifferent between local and producer currency pricing based on the value function (depicted in Figure 6) and plot for these parameters the corresponding MRPT and LRPT as a function of Γ (solid lines). Note that MRPT remains very close to 0.5.

The theoretical MRPT was estimated using equation (12) without restricting $\tilde{\Gamma}$ to be constant. We now examine how the value of MRPT would differ if we computed it under the assumption of Proposition 3 that $\tilde{\Gamma}$ is constant at its steady state value of Γ. This is also reported in Figure 7 as the dashed lines (MRPT$'$ and LRPT$'$) computed using Proposition 3, to be compared to the solid lines. As is evident the two lines are very close to the correct theoretical MRPT and LRPT respectively. This justifies our assumption in the theoretical section that a constant Γ is a useful approximation point for empirical work.

The results this far relate to theoretical estimates of MRPT constructed from desired prices. In Proposition 4 we pointed out that MRPT can be measured using actual price changes by estimating the price response conditional on first adjustment to the exchange rate shock. We now use the simulated data to estimate this regression in order to evaluate the quality of this estimator of MRPT. The results are reported in Figure 8. The dashed line is the theoretical MRPT, while the other lines are coefficients from different specifications of the pass-through regression conditional on price adjustment. We compute both the estimate from the regression on a one-period exchange rate change ($\Delta e_t = e_t - e_{t-1}$), as suggested by Proposition 3, and on the cumulative exchange rate change ($\Delta_c e_t = e_t - e_{t-\tau}$, where τ is the price duration), as was done in the empirical work. Finally, we plot coefficients both corrected and uncorrected for mean reversion bias in the exchange rate.[25] The overall conclusion that emerges from this figure is that all empirical specifications provide accurate approximations to the true theoretical MRPT and the potential biases are not important quantitatively.[26] In fact, in our simulation the coefficient from regression (2), on cumulative exchange rate change and uncorrected for mean reversion bias, provides the most accurate approximation to the theoretical MRPT.

To summarize, the numerical simulation verifies the robustness of the theoretical results to relaxing a number of assumptions. Firstly, we verify that the second order approximation to the value function of Proposition 2 is accurate and medium-run pass-through indeed accurately predicts currency choice. Secondly, empirical estimates of pass-through conditional on price adjustment indeed approximate well medium-run pass-through even in the environment with variable mark-up elasticity, where the assumption of Proposition 3 does not hold. Thirdly, the results are robust to mean reversion in the exchange rate. Lastly, the main results extend to the menu cost model of price stickiness.

[25]Mean reversion in the exchange rate leads firms to adjust by less in their currency of pricing. This leads PCP firms to have higher exchange rate pass-through relative to identical LCP firms and the size of this bias is given by $\delta\theta(1-\rho)/(1-\delta\theta\rho)$, where ρ is the autocorrelation of the exchange rate process. For empirically reasonable value of $\rho \in (0.98, 1)$, the wedge in the pass-through between PCP and LCP firms does not exceed 10 percentage points and is small relative to the documented empirical differences.

[26]This exercise was done for the data generated from the Calvo model. In the menu cost model there is in addition the selection bias discussed in footnote 7. Consistent with that discussion we find that this reduces the estimated gap in MRPT between local and producer currency pricers in the model generated data. It turns out that this bias is small to moderate when we use the specification with the cumulative exchange rate change. Specifically, with the selection bias, the pass-through estimate for LCP firm is 10-15 percentage points higher than for otherwise identical PCP firm.

4 Discussion: Linking Theory and Empirical Evidence

A main conclusion of the previous section is that in an environment with endogenous currency choice medium run pass-through that can be approximated empirically by regression (2) should be lower for goods priced in the local currency as compared to goods priced in the producer currency. In the empirical section we documented this to be robustly the case across various sub-samples of the data. Tables 2, 3 and 4 all strongly support the endogenous currency choice model's prediction.

A second insight of the theoretical section is that currency choice is closely tied to medium-run pass-through and not to long-run pass-through. LRPT could be above 0.5 but nevertheless if MRPT is below 0.5 the firm chooses local currency pricing. In the life-long regressions estimated in Section 2.2.1 the pass-through estimates are almost twice as high as pass-through conditional on first adjustment, with some point estimates exceeding 0.5. This is consistent with an important role played by real rigidities in pricing that have effects significantly past the period of nominal rigidity.

A third insight, stated in Corollary 1, is that even when firms have the same desired pass-through profiles they will choose different currencies to price in if the frequency with which they adjust differ. Firms that adjust less frequently are more likely to price in the producer currency. This is consistent with what we observe in the data for goods within very narrow classifications. As reported in Section 2.2.2, even within 10 digit classifications that have a mix of dollar and non-dollar pricers, non-dollar pricers adjust prices less frequently than dollar pricers.

The standard international macro model assumes CES demand with constant mark-ups and desired pass-through equal to 1. In this environment firms are exogenously specified to be local currency pricers or producer currency pricers. Therefore, pass-through differs in the short-run, but once firms adjust prices the pass-through is the same. In Proposition 1 we showed more generally that firms facing similar demand and cost conditions set the same price once they adjust. This implies that pass-through conditional on first adjustment should be the same. Clearly, the data strongly contradicts this theoretical description of the relation between currency of pricing and pass-through.[27]

Let us consider departures from the theoretical specification that allow for exogenous currency choice, yet obtain different pass-through conditional on price adjustment. One possibility is that price setting is backward-looking in the sense that even when the firm adjusts it's price, it keeps it close to the previous price. This generates differential pass-through conditional on adjusting prices. This particular explanation with exogenous currency choice requires that long-run pass-through is the same regardless of the currency of pricing. The evidence in Table 5 suggests that LRPT is significantly different for LCP and PCP firms. One could however argue that since goods are around for a limited period and get substituted often, we are not quite capturing the appropriate long run. The median good in this sample has a life of 3-4 years. This explanation therefore would imply that the effects of nominal and real rigidities remain well past this horizon.

A second piece of evidence against exogenous currency of pricing is that we observe sorting in the data. That is there are sectors that are dominated by dollar pricers and those by non-dollar pricers. So there is the interesting question of why certain sectors such as 'Animal or Vegetable Fats and Oils' are dominated by dollar pricers, while 'Machinery and Mechanical Appliances' by non-dollar pricers. Further, there are interesting differences even within narrow sectors that have a mix of dollar and non-dollar pricers. Non-dollar pricers tend to adjust prices less frequently than dollar pricers. These features are consistent with a model of endogenous currency choice, but less obvious in an environment with exogenous currency choice.

A second possible hypothesis is that suppose for exogenous institutional reasons sectors with identical long-run pass-through get sorted into local and producer currency pricing bins. In this case, and if there are significant real rigidities in pricing that generate delayed price adjustments, a single firm in each sector will adjust its price by small amounts in the currency of pricing. In this case, firms in the sector dominated by producer currency pricers will have higher MRPT than firms in the sector dominated by local currency pricers,

[27] The result of the proposition is exact when the nominal exchange rate follows a random walk. If the exchange rate mean reverts, we pointed out that for empirically reasonable values of mean reversion the wedge in pass-through conditional on adjusting prices falls far short of the empirical wedge of 70 points.

even if eventually LRPT is the same for all firms in both sectors. However the fact that even within narrow 10 digit sectors there is a mix of dollar and non-dollar pricers limits the generality of this explanation. It is hard to think of institutional reasons why at this level of disaggregation there is exogenous sorting of firms.

While several of our findings are consistent with an environment of endogenous currency choice and these findings rule out certain models of currency and pass-through, there is more work to be done. The ideal test for whether endogenous currency choice is the over-riding mechanism would be to relate firm-level characteristics shaping mark-up variability and cost sensitivity to the currency choice decision. This requires extensive firm-level information that can be hard to come by, as is well known in the pass-through literature. Using aggregated data, Bacchetta and van Wincoop (2005) and Goldberg and Tille (2005) present some evidence consistent with endogenous currency choice.

A strength of our analysis is that we provide a simple sufficient statistic for currency choice that summarizes complicated pieces of information. In this sense it is less sensitive to details about difficult to measure elements of the pass-through environment. Further research on the underlying determinants will be clearly valuable though beyond the scope of the current paper.

Appendix

A Proofs of Results for Section 3.1

Proof of Proposition 1: From the Bellman equations (5) and (7) and using the Envelope Theorem, we have the first order conditions for price setting by an LCP and PCP firm respectively:

$$\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\Pi_p\big(\bar{p}_{L,t}|s_{t+\ell}\big)=0,$$
$$\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\Pi_p\big(\bar{p}^*_{P,t}+e_{t+\ell}|s_{t+\ell}\big)=0,$$

where the subscript denotes the respective partial derivative. Note that, with Calvo pricing, preset prices affect the value function only through profits in the states in which they are effective, which is determined exogenously and, hence, these optimality conditions do not depend on whether currency choice is exogenous or endogenous.

First consider the LCP case. We take a first order Taylor approximation to the marginal profit state by state around that state's optimal price:

$$\begin{aligned}\Pi_p\big(\bar{p}_{L,t}|s_{t+\ell}\big)&=\tilde{\Pi}_{pp}(s_{t+\ell})\,[\bar{p}_{L,t}-\tilde{p}_{t+\ell}]+\mathcal{O}(\bar{p}_{L,t}-\tilde{p}_{t+\ell})^2\\&=\tilde{\Pi}_{pp}(s_t)\,[\bar{p}_{L,t}-\tilde{p}_{t+\ell}]+\mathcal{O}(\bar{p}_{L,t}-\tilde{p}_{t+\ell})^2+\mathcal{O}\big(\|s_{t+\ell}-s_t\|\cdot(\bar{p}_{L,t}-\tilde{p}_{t+\ell})\big),\end{aligned}$$

where $\tilde{p}_{t+\ell}\equiv\tilde{p}(s_{t+\ell})$, $\tilde{\Pi}_{pp}(s_{t+\ell})\equiv\Pi_{pp}(\tilde{p}_{t+\ell}|s_{t+\ell})$, $\mathcal{O}(\cdot)$ denotes the same order of magnitude, $\|\cdot\|$ is some norm in a vector space. The derivation of this approximation uses the first order optimality of the desired price $\tilde{p}_{t+\ell}$ that implies $\Pi_p(\tilde{p}_{t+\ell}|s_{t+\ell})=0$ and the fact that $\tilde{\Pi}_{pp}(s_{t+\ell})=\tilde{\Pi}_{pp}(s_t)+\mathcal{O}(\|s_{t+\ell}-s_t\|)$ that follows from the smoothness of the profit function and the desired price in the state of the economy (i.e., this requires convexity and continuous second derivatives of $\Pi(\cdot)$, which we assume hold).

We then combine this approximation with the first order condition above, and obtain the following first order approximation for the preset price:

$$\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\left\{\bar{p}_{L,t}-\tilde{p}_{t+\ell}\right\}=0+\alpha^2,$$

where one can verify that

$$\alpha\equiv\mathcal{O}_p\left(\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\|s_{t+\ell}-s_t\|\right),$$

with $\mathcal{O}_p(\cdot)$ denoting the same order of magnitude in the probabilistic sense. After rearranging, this delivers the first claim of the proposition for LCP firms. Analogous steps result in a similar result for PCP firms:

$$\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\left\{\bar{p}^*_{P,t}+e_{t+\ell}-\tilde{p}_{t+\ell}\right\}=0+\alpha^2.$$

Subtracting the expression for PCP from that for LCP firms, multiplying through by $(1-\delta\theta)$ and rearranging, we have:

$$\bar{p}_{L,t}=\bar{p}^*_{P,t}+(1-\delta\theta)\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_te_{t+\ell}+\alpha^2.$$

When exchange rate follows a random walk, $\mathbb{E}_t e_{t+\ell} = e_t$ and we have $\bar{p}^*_{P,t} + e_t - \bar{p}_{L,t} = \alpha^2$, which completes the proof of the proposition.[28]

Finally, we assess the magnitude of the error of approximation α. Denote by $\sigma_s \equiv \mathcal{O}_p(\|\Delta s_t\|)$ the magnitude of the standard deviation of the innovation to the state of the economy. Allowing for a unit root in $\{s_t\}$, we have $\mathcal{O}_p(\|s_{t+\ell} - s_t\|) = \ell\sigma_s$. Therefore,

$$\alpha = \sigma_s \textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^\ell \ell = \sigma_s \delta\theta/(1-\delta\theta)^2.$$

That is, even in an environment with an integrated state of the economy, the error of approximation still has the order of the standard deviation of the one-period innovation to the state vector, as long as the duration of prices is finite (more accurately, as long as $\delta\theta < 1$). ■

Proof of Proposition 2: Recall that the difference in the value to the firm of LCP and PCP is:

$$\mathcal{L}_t = \textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^\ell \mathbb{E}_t \Big\{ \Pi\big(\bar{p}_{L,t}|s_{t+\ell}\big) - \Pi\big(\bar{p}^*_{P,t} + e_{t+\ell}\,|s_{t+\ell}\big) \Big\}.$$

We take a second order Taylor expansion of the profit differential state by state around the respective desired price:

$$\Pi\big(\bar{p}_{L,t}|s_{t+\ell}\big) - \Pi\big(\bar{p}^*_{P,t} + e_{t+\ell}\,|s_{t+\ell}\big) = \frac{1}{2}\tilde{\Pi}_{pp}(s_{t+\ell})\Big[\big(\bar{p}_{L,t} - \tilde{p}_{t+\ell}\big)^2 - \big(\bar{p}^*_{P,t} + e_{t+\ell} - \tilde{p}_{t+\ell}\big)^2\Big] + \alpha^3,$$

where we use the same notation as in the Proof of Proposition 1. Using again the fact that $\tilde{\Pi}_{pp}(s_{t+\ell}) = \tilde{\Pi}_{pp}(s_t) + \alpha$, we can rewrite:

$$\mathcal{L}_t = \frac{1}{2}\tilde{\Pi}_{pp}(s_t)\sum_{\ell=0}^{\infty}(\delta\theta)^\ell\, \mathbb{E}_t \Big\{ \big(\bar{p}_{L,t} - \tilde{p}_{t+\ell}\big)^2 - \big(\bar{p}^*_{P,t} + e_{t+\ell} - \tilde{p}_{t+\ell}\big)^2 \Big\} + \alpha^3.$$

Expanding the expression inside the brackets, we obtain:

$$\begin{aligned}\big(\bar{p}_{L,t} - \tilde{p}_{t+\ell}\big)^2 - \big(\bar{p}^*_{P,t} + e_{t+\ell} - \tilde{p}_{t+\ell}\big)^2 &= \big(\bar{p}_{L,t} - \bar{p}^*_{P,t} - e_{t+\ell}\big)\big(\bar{p}_{L,t}\bar{p}^*_{P,t} + e_{t+\ell} - 2\tilde{p}_{t+\ell}\big)\\ &= -\big(e_{t+\ell} - e_t + \alpha^2\big)\big(\bar{p}_{L,t} + \bar{p}^*_{P,t} + e_{t+\ell} - 2\tilde{p}_{t+\ell}\big),\end{aligned}$$

where the second equality used the equivalence result of Proposition 1. Substituting this into the expression for $\mathcal{L}_t$, we have

$$\mathcal{L}_t = -\frac{1}{2}\tilde{\Pi}_{pp}(s_t)\sum_{\ell=0}^{\infty}(\delta\theta)^\ell \operatorname{cov}_t\big(e_{t+\ell}, e_{t+\ell} - 2\tilde{p}_{t+\ell}\big) + \alpha^3,$$

[28] If exchange rate follows a first order autoregressive process with autocorrelation ρ, we have $\mathbb{E}_t e_{t+\ell} = \rho^\ell e_t$ and therefore:

$$\bar{p}^*_{P,t} + e_t - \bar{p}_{L,t} = \frac{\delta\theta(1-\rho)}{1-\delta\theta\rho}e_t + \alpha^2.$$

Mean reversion in the exchange rate introduces a wedge between PCP and LCP pricing proportional to $(1-\rho)$. We discuss the magnitude of this wedge and its implications for the measurement of pass-through in Section 3.2.

where we use the fact from Proposition 1 that

$$\textstyle\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\mathbb{E}_t\left\{\bar{p}_{L,t}+\bar{p}^*_{P,t}+e_{t+\ell}-2\tilde{p}_{t+\ell}\right\}=\alpha^2.$$

Expanding the covariance term immediately results in expression (11) of Proposition 2. Finally, by the random walk property of the exchange rate, we have $\mathrm{var}_t(e_{t+\ell})=\ell\mathrm{var}_t(\Delta e_{t+j})$ for any $\ell,j>0$. Therefore, $\mathcal{L}_t<0$ is approximately equivalent to (12). ■

Proof of Proposition 3: With constant elasticity of mark-up and constant marginal cost sensitivity to the exchange rate, the expression for the desired price is:

$$\tilde{p}_t=-\Gamma\big(\tilde{p}_t-P_t\big)+\phi e_t+\xi_t,$$

where ξ_t combines a constant and all shocks to mark-ups and marginal cost orthogonal to the exchange rate. This expression can be rewritten as

$$\tilde{p}_t=\frac{\phi}{1+\Gamma}e_t+\frac{\Gamma}{1+\Gamma}P_t+\frac{\xi_t}{1+\Gamma}$$

and the impulse response of $\tilde{p}_{t+\ell}$ to e_{t+j} conditional on information as of period t is

$$\tilde{\Psi}_{j,\ell}(s^t)=\frac{\phi}{1+\Gamma}+\frac{\Gamma}{1+\Gamma}\frac{\mathrm{cov}_t(P_{t+\ell},\Delta e_{t+j})}{\mathrm{var}_t(\Delta e_{t+j})},$$

as stated in the proposition. Assuming that the sectoral price level follows a linear process:[29]

$$(P_t-\bar{P})=\alpha(P_{t-1}-\bar{P})+(1-\alpha)\bar{\phi}e_t+\xi_{P,t},$$

we have

$$\frac{\mathrm{cov}_t(P_{t+\ell},\Delta e_{t+j})}{\mathrm{var}_t(\Delta e_{t+j})}=\big(1-\alpha^{\ell-j+1}\big)\bar{\phi}.$$

Therefore, indeed, $\tilde{\Psi}_{j,\ell}(s^t)$ depends only on $\ell-j$ and does not depend on s^t. Moreover, the medium-run pass-through equals

$$\bar{\Psi}=\frac{\phi}{1+\Gamma}+\frac{\Gamma\bar{\phi}}{1+\Gamma}(1-\delta\theta)\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\big(1-\alpha^{\ell+1}\big)=\frac{\phi}{1+\Gamma}+\frac{\Gamma\bar{\phi}}{1+\Gamma}\frac{1-\alpha}{1-\delta\theta\alpha},$$

which is increasing in θ. ■

[29] This process can be shown to be the sectoral equilibrium outcome in a linearized Calvo model with α determined by the primitives of the model.

Proof of Proposition 4: First, we show that

$$\frac{\operatorname{cov}\big(\Delta\bar{p}_t,\Delta e_t\big)}{\operatorname{var}(\Delta e_t)}=\frac{\operatorname{cov}\big(\bar{p}_t,\Delta e_t\big)}{\operatorname{var}(\Delta e_t)}.$$

This follows from

$$\operatorname{cov}\big(\bar{p}_{t-k},\Delta e_t\big)=\mathbb{E}\big\{\bar{p}_{t-k}\cdot\Delta e_t\big\}=\mathbb{E}\left\{\mathbb{E}_{t-k}\big\{\bar{p}_{t-k}\cdot\Delta e_t\big\}\right\}=0,$$

where we have use the law of iterated expectations and the fact that $\mathbb{E}_{t-k}\Delta e_t=0$. Next, it follows from Proposition 1, that

$$\operatorname{cov}\big(\bar{p}_t,\Delta e_t\big)=(1-\delta\theta)\sum_{\ell=0}^{\infty}(\delta\theta)^{\ell}\frac{\operatorname{cov}(\mathbb{E}_t\tilde{p}_{t+\ell},\Delta e_t)}{\operatorname{var}(\Delta e_t)}.$$

Note that

$$\operatorname{cov}\big(\mathbb{E}_t\tilde{p}_{t+\ell},\Delta e_t\big)=\mathbb{E}\left\{\Delta e_t\cdot\mathbb{E}_t\tilde{p}_{t+\ell}\right\}=\mathbb{E}\left\{\mathbb{E}_t\left\{\tilde{p}_{t+\ell}\cdot\Delta e_t\right\}\right\}=\operatorname{cov}(\tilde{p}_{t+\ell},\Delta e_t),$$

again using the law of iterated expectations and the random walk property of the exchange rate. Proposition 4 follows immediately. ■

B Details for the Numerical Simulation of Section 3.2

The profit of the firm in local currency is given by

$$\Pi(p_t|P_t,e_t,a_t)=\big[\exp(p_t)-\exp(\phi e_t-a_t)\big]\cdot q(p_t,P_t).$$

For the Calvo case, we iterate the Bellman operator defined by (5), (7) and (9) to solve for the value functions. For the Menu Cost case, the Bellman equations system is give by

$$\begin{aligned}
V_L(p|s^t)&=\Pi(p|s_t)+\mathbb{E}_t\max\{V_L(p|s^{t+1}),V(s^{t+1})-\kappa\},\\
V_P(p^*|s^t)&=\Pi(p^*+e_t|s_t)+\mathbb{E}_t\max\{V_P(p^*|s^{t+1}),V(s^{t+1})-\kappa\},\\
V(s^t)&=\max\left\{V_L\big(\bar{p}_L(s^t)|s^t\big),V_P\big(\bar{p}^*_P(s^t)|s^t\big)\right\},
\end{aligned}$$

where the policy functions are given by:

$$\begin{aligned}
\bar{p}_L(s^t)&=\arg\max_p V_L(p|s^t),\\
\bar{p}^*_P(s^t)&=\arg\max_{p^*} V_P(p^*|s^t).
\end{aligned}$$

Additionally, the policy function specifies in which states the firm adjusts its price and what currency it chooses.

In both cases, the state vector contains the previous price of the firm, the idiosyncratic productivity shock, the sectoral price level and the nominal exchange rate: $s_t=(p_{t-1},a_t,P_t,e_t)$. The expectations in the Bellman operators are with respect to (a_t, P_t, e_t). All three follow exogenous first order processes with $\{a_t\}$ independent from $\{e_t, P_t\}$. We iterate the Bellman equations on a discreet grid.[30]

After obtaining the policy functions, we simulate in the given partial equilibrium environment the firm's price for $T = 12,000$ periods, where the period is calibrated to be a month. Using the policy function and the simulated data, we study the currency choice and pass-through patterns for different values of primitive parameters. We determine currency choice by evaluating $\mathcal{L} = V_L(\bar{p}_L) - V_P(\bar{p}^*_P)$.[31] For Figures 6 and 7 we find the set of parameter values $\{(\phi, \Gamma)\}$ for which $\mathcal{L} = 0$ at an initial state, s_0, at which $p = P = \bar{P}$.

We calculate MRPT according to its definition (16) by simulating a path of desired prices, $\tilde{p}_t$, for $10,000$ realizations of the exchange rate, sectoral price level and idiosyncratic shocks and estimate $\mathrm{cov}_0(\tilde{p}_t, e_t)/\mathrm{var}_0(e_t)$. We compute LRPT using the definition of desired price (equation 4). Specifically, for multiple realizations of the shocks (e_t, a_t) and conditioning on the long-run response of the sectoral price level (equal to $\bar{\phi} e_t$), we compute $\tilde{p}_t$ and evaluate the regression coefficient of desired prices on the exchange rate, which defines our long-run pass-through measure.

In addition, we compute MRPT and LRPT under the restrictive assumption that mark-up elasticity is constant at its steady state level Γ. As suggested by Proposition 3, these measures are:

$$\mathrm{MRPT}' = \frac{\phi}{1+\Gamma} + \frac{\Gamma\bar{\phi}}{1+\Gamma}\frac{1-\alpha}{1-\alpha\delta\theta} \qquad \text{and} \qquad \mathrm{LRPT}' = \frac{\phi}{1+\Gamma} + \frac{\Gamma\bar{\phi}}{1+\Gamma}.$$

Lastly, we compute the regression-based estimates of MRPT. We have a simulated series of firm's prices and exchange rates. Using these series, we estimate the micro-level pass-through regression conditional on price adjustment, equivalent to our empirical specification (2). We estimate both specifications, with the one-period exchange rate change ($\Delta e_t = e_t - e_{t-1}$) and the cumulative exchange rate change ($\Delta_c e_t = e_t - e_{t-\tau}$, where τ is the previous price duration) as right hand side variables. In addition, we estimate the mean reversion bias corrected estimates given by:

$$\hat{\Psi}_L^{b.c.} = \hat{\Psi}_L \frac{1-\delta\theta\rho_e}{1-\delta\theta} \qquad \text{and} \qquad \hat{\Psi}_P^{b.c.} = 1 - \big(1-\hat{\Psi}_P\big)\frac{1-\delta\theta\rho_e}{1-\delta\theta},$$

where ρ_e is the autocorrelation of the exchange rate process. Note that given the same underlying MRPT, the cumulative bias of PCP and LCP MRPT is

$$\hat{\Psi}_P - \hat{\Psi}_L = \frac{\delta\theta(1-\rho_e)}{1-\delta\theta\rho_e},$$

independently of the value of MRPT (see the Proof of Proposition 1).

[30] The step of the grid for individual price p_t is no greater than 0.5% and no greater than 0.2% for the sectoral price level. There are 15 points on the grid for the idiosyncratic shock and 31 for the nominal exchange rate.

[31] There are currency switches over time but they are rare for most values of the parameters and would be absent altogether if there was a small fixed cost of currency switching.

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Figures and Tables

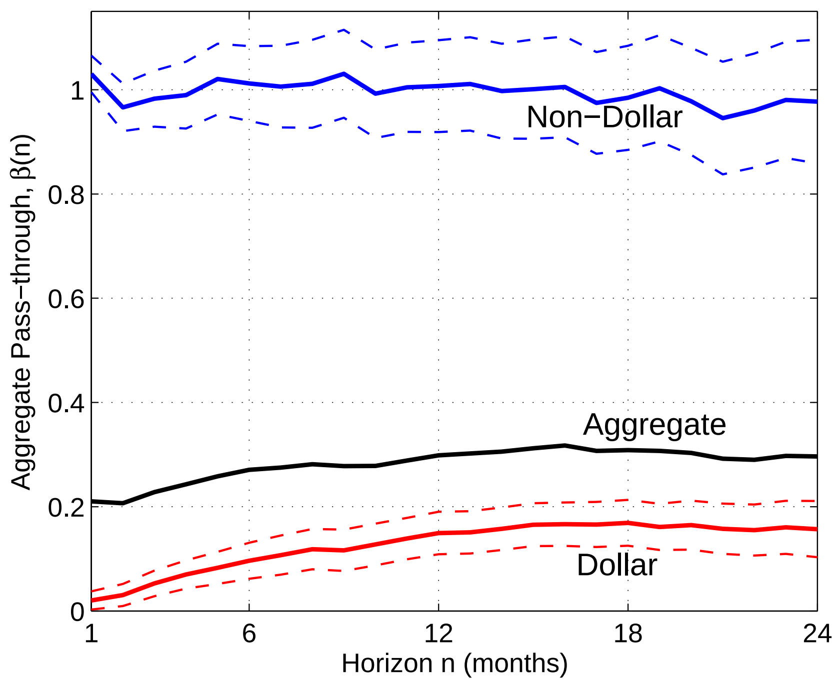

Figure 1: Aggregate ERPT at different horizons by currency

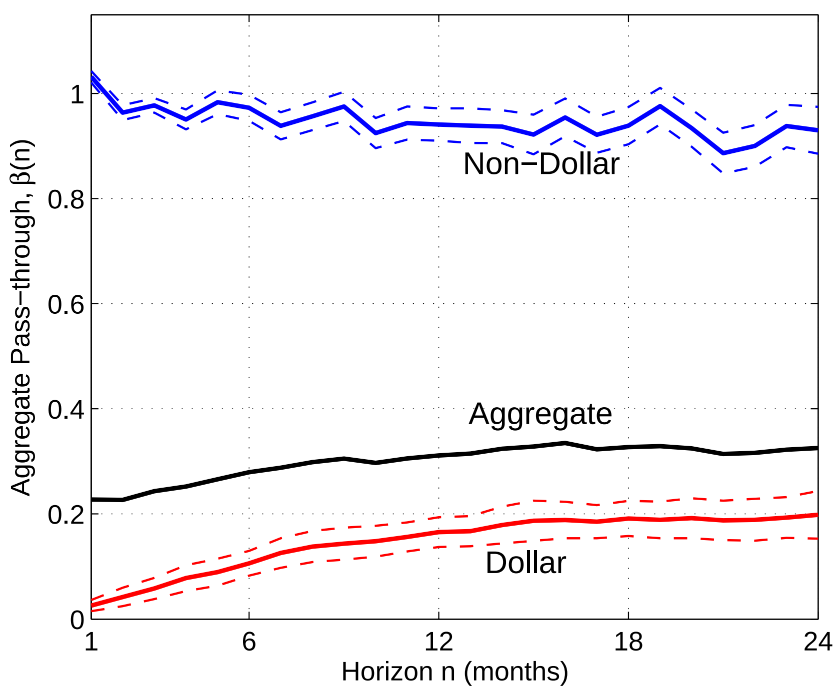

Figure 3: Aggregate ERPT at different horizons by currency: Panel specification

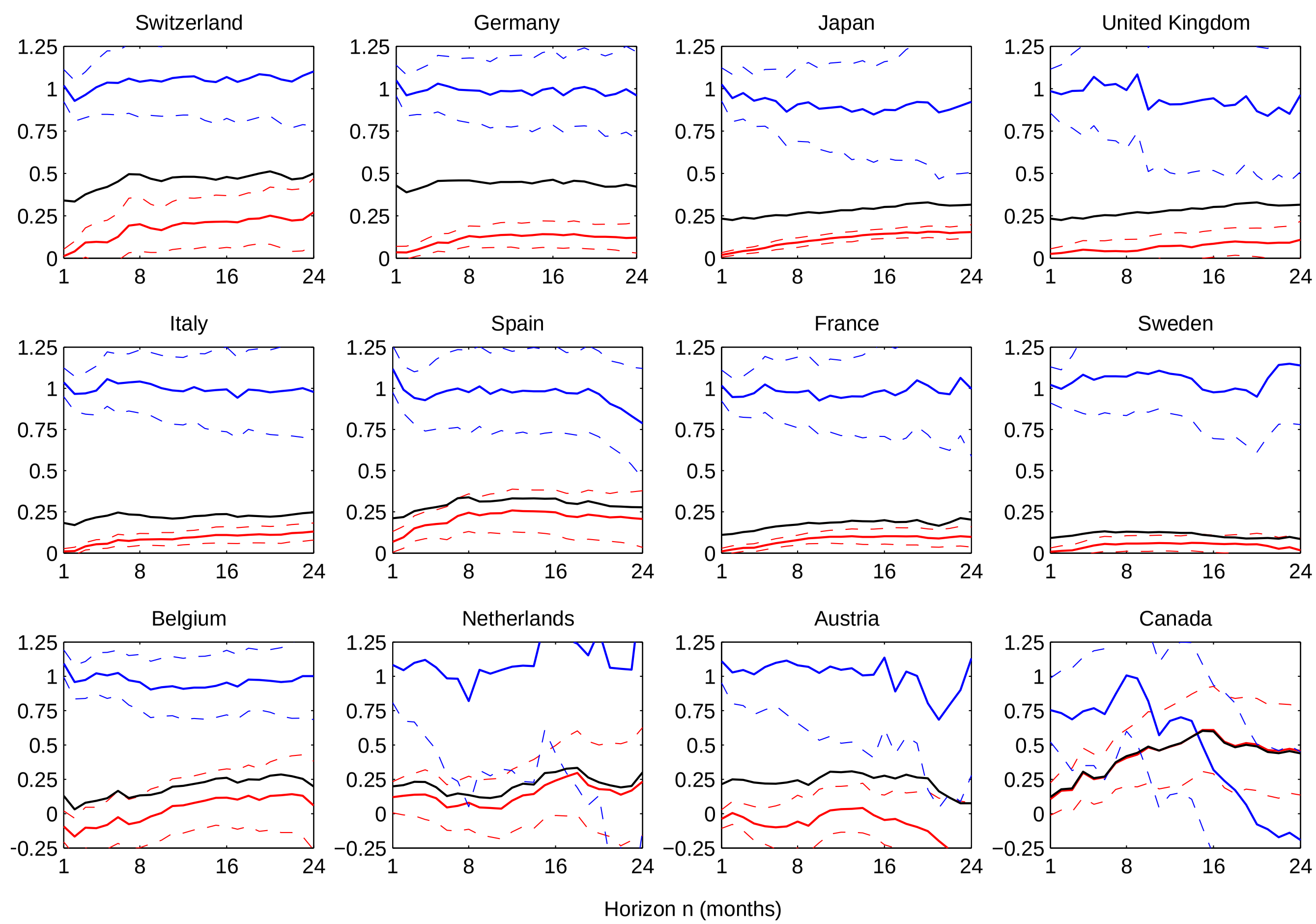

Figure 2: Aggregate ERPT at different horizons by currency from specific countries

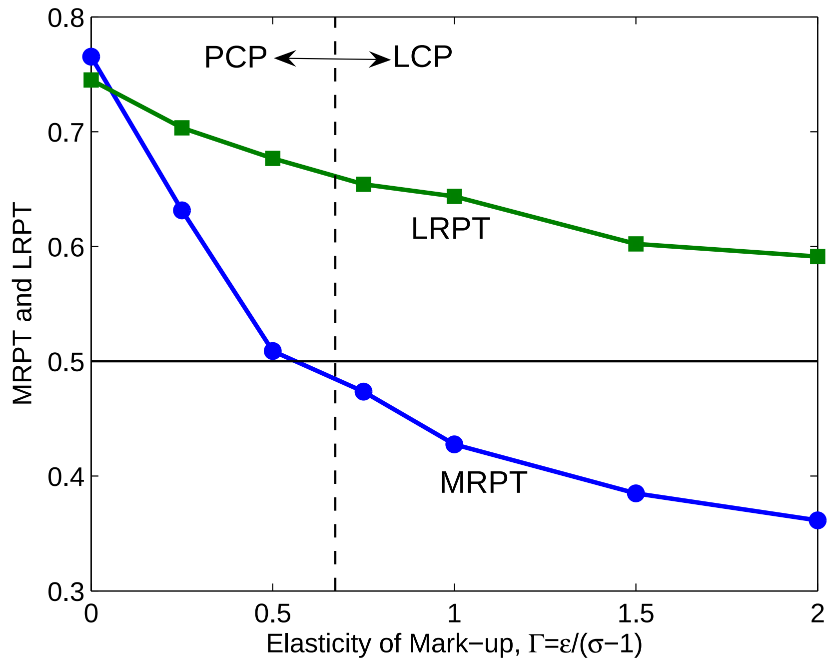

Figure 4: Currency Choice, MRPT and LRPT as a function of mark-up elasticity, Γ

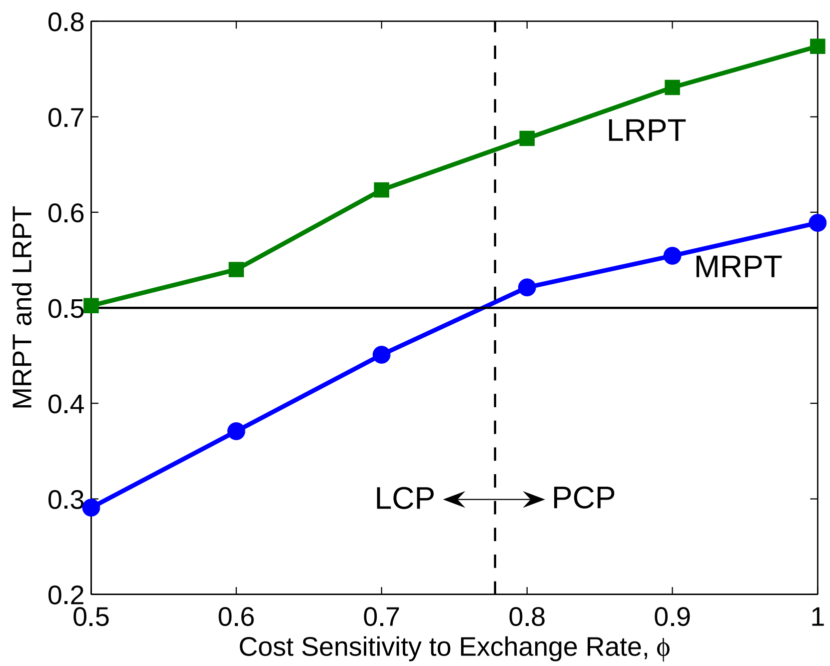

Figure 5: Currency Choice, MRPT and LRPT as a function of cost sensitivity to the exchange rate, ϕ

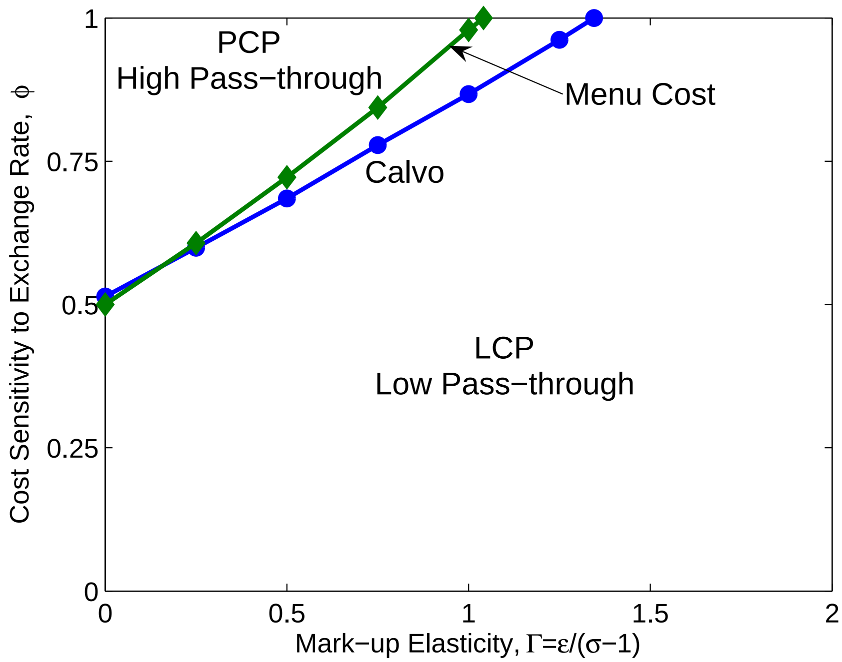

Figure 6: Currency Choice in the (Γ, ϕ)-space

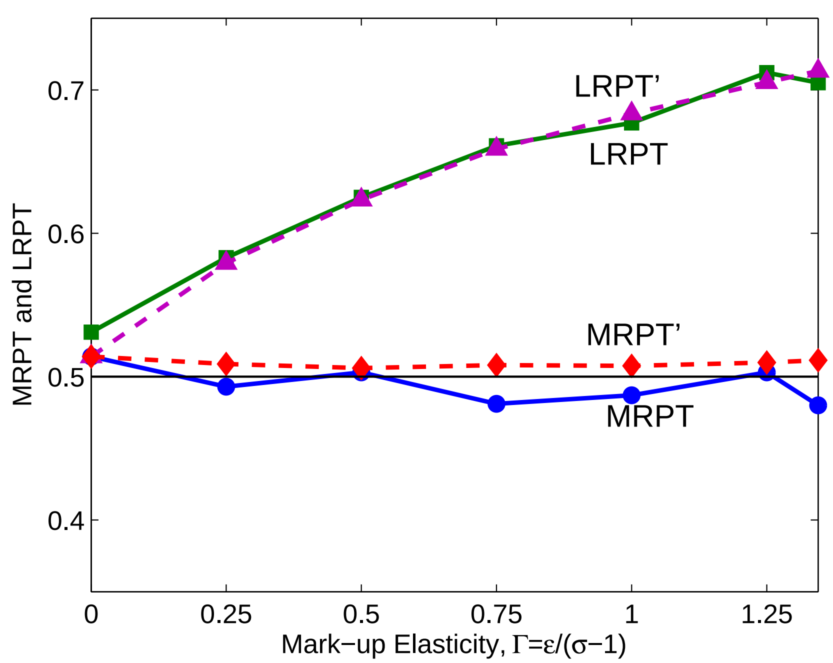

Figure 7: MRPT and LRPT for the firm indifferent between LCP and PCP, as a function of Γ (MRPT$'$ and LRPT$'$ are computed under the restrictive assumption that $\tilde{\Gamma} \equiv \Gamma = const$)

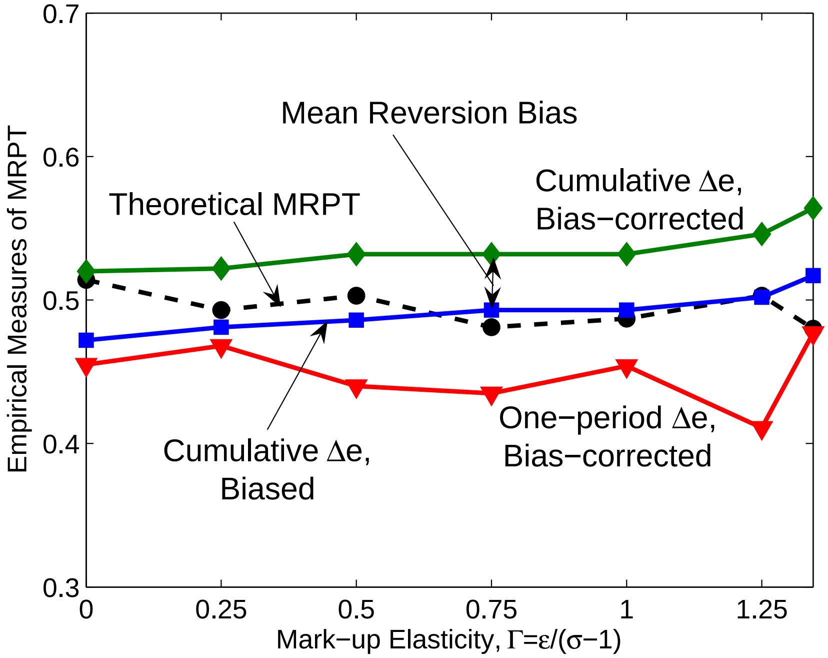

Figure 8: Empirical MRPT: Pass-through Conditional on Price Adjustment

Country	N	$Frac_{ND}$
Germany	1,156	0.40
Switzerland	220	0.38
Italy	1,112	0.22
Japan	2,553	0.21
UK	719	0.19
Belgium	122	0.17
France	582	0.13
Sweden	181	0.12
Spain	261	0.11
Austria	93	0.10
Netherlands	148	0.10
Canada	1,906	0.04

Table 1: Number of Goods and Fraction Non-Dollar Priced

	Dollar		Non-Dollar		Difference		N_{obs}	N_{goods}	R^2
	β_D	s.e.(β_D)	β_{ND}	s.e.(β_{ND})	$\beta_{ND}-\beta_D$	t-stat			
All Countries	0.24	0.03	0.92	0.04	0.68	13.89	35,622	6,637	0.11
Germany	0.31	0.07	0.87	0.10	0.56	4.71	3,339	801	0.22
Switzerland	0.24	0.11	0.96	0.18	0.72	3.36	556	130	0.36
Italy	0.21	0.06	0.84	0.13	0.63	4.31	2,030	744	0.18
Japan	0.23	0.04	0.96	0.06	0.73	10.77	6,297	1,733	0.14
UK	0.19	0.11	0.74	0.17	0.55	2.92	2,654	541	0.17
Belgium	0.01	0.07	0.98	0.14	0.98	6.05	544	97	0.41
France	0.26	0.07	1.03	0.10	0.77	6.11	1,384	425	0.21
Sweden	0.28	0.14	0.94	0.21	0.66	2.73	639	160	0.25
Spain	0.46	0.14	0.83	0.27	0.37	1.19	758	164	0.20
Netherlands	0.21	0.09	0.89	0.35	0.67	1.85	881	126	0.08
Canada	0.22	0.13	0.66	0.45	0.44	1.00	16,428	1,654	0.03

Table 2: Pass-through Conditional on Price Adjustment: Dollar vs. Non-Dollar Pricers

	Dollar		Non-Dollar		Difference		N_{obs}	N_{goods}	R^2
	β_D	s.e.(β_D)	β_{ND}	s.e.(β_{ND})	$\beta_{ND}-\beta_D$	t-stat			
All Countries	0.24	0.04	0.96	0.06	0.72	10.20	13,578	3,191	0.15
Germany	0.44	0.10	0.92	0.12	0.48	3.09	1,848	489	0.24
Switzerland	0.10	0.15	0.99	0.31	0.89	2.77	358	78	0.38
Italy	0.23	0.08	0.81	0.10	0.58	4.81	1,029	409	0.18
Japan	0.19	0.04	0.98	0.10	0.81	7.31	2,725	838	0.16
UK	0.32	0.19	0.89	0.16	0.56	2.14	848	277	0.16
Belgium	0.01	0.07	0.98	0.14	0.98	6.05	544	97	0.41
France	0.29	0.12	1.17	0.14	0.88	4.89	534	178	0.23
Sweden	0.44	0.15	1.43	0.11	0.99	6.01	350	95	0.33
Spain	0.53	0.11	0.73	0.15	0.20	0.92	458	97	0.17
Netherlands	0.17	0.22	1.19	0.03	1.01	4.33	319	36	0.07
Canada	-0.06	0.12	0.51	1.16	0.57	0.50	4,945	619	0.06

Table 3: Pass-through Conditional on Price Adjustment: Differentiated Goods

Category	Harm. Code	Dollar β_D	Dollar s.e.(β_D)	Non-Dollar β_{ND}	Non-Dollar s.e.(β_{ND})	Difference $\beta_{ND}-\beta_D$	Difference t-stat	N_{obs}	N_g	N_g^{ND}/N_g	R^2
Live Animals; Animal Products	01-05	0.18	0.16	−0.11	0.53	-0.28	−0.53	1,841	170	0.06	0.12
Vegetable Products	06-14	0.04	0.27	1.10	0.10	1.06	4.00	746	87	0.07	0.07
Animal or Vegetable Fats and Oils	15	0.65	0.32	—	—	—	—	145	18	0.00	0.12
Prepared Foodstuffs	16-24	0.24	0.06	0.83	0.25	0.59	2.27	1,668	426	0.07	0.12
Mineral Products	25-27	0.96	0.19	1.14	0.35	0.18	0.50	4,588	310	0.05	0.03
Products of chemical and allied industries	28-38	0.27	0.08	0.64	0.25	0.37	1.38	2,291	456	0.08	0.27
Plastics and Rubber articles	39-40	0.21	0.07	0.53	0.11	0.32	2.57	896	219	0.15	0.22
Raw Hides leather articles, furs etc.	41-43	-0.15	0.14	0.91	0.03	1.06	7.40	158	54	0.30	0.71
Wood and articles of wood	44-46	-0.08	0.17	0.85	0.00	0.94	5.39	4,850	307	0.01	0.03
Pulp of wood/other fibrous cellulosic material	47-49	0.26	0.11	1.02	0.20	0.76	2.85	482	293	0.05	0.57

Table 4: Pass-through Conditional on Price Adjustment: Sectoral Evidence (continued on the next page)

Category	Harm. Code	Dollar		Non-Dollar		Difference		N_{obs}	N_g	N_g^{ND}/N_g	R^2
		β_D	s.e.(β_D)	β_{ND}	s.e.(β_{ND})	$\beta_{ND}-\beta_D$	t-stat				
Textile and textile articles	50-63	0.41	0.14	0.94	0.12	0.54	2.85	482	175	0.18	0.57
Footwear, headgear etc.	64-67	0.45	0.16	0.97	0.07	0.52	4.61	161	72	0.58	0.56
Miscellaneous manufactured articles	68-70	0.19	0.17	1.06	0.24	0.87	2.83	460	214	0.19	0.32
Precious or semi precious stones etc.	71	0.24	0.13	2.03	0.74	1.79	2.48	1,882	171	0.05	0.09
Base metals and articles of base metals	72-83	0.21	0.04	1.35	0.36	1.15	3.19	3,693	614	0.13	0.27
Machinery and mechanical appliances etc.	84-85	0.22	0.05	0.90	0.06	0.67	9.51	5,943	1775	0.20	0.26
Vehicles, aircraft etc.	86-89	0.17	0.07	0.93	0.10	0.76	6.31	2,337	662	0.13	0.13
Optical, photographic etc.	90-92	0.22	0.07	1.09	0.20	0.88	4.18	928	317	0.27	0.37
Arms and ammunition	93	0.08	0.16	1.03	0.09	0.95	5.00	109	47	0.15	0.31
Articles of stone, plaster etc.	94-96	0.54	0.09	0.76	0.12	0.23	1.53	394	148	0.15	0.29

Table 4: Pass-through Conditional on Price Adjustment: Sectoral Evidence (continued from the previous page)

	Dollar		Non-Dollar		Difference		N_{goods}	R^2
	β^L_D	s.e.(β^L_D)	β^L_{ND}	s.e.(β^L_{ND})	$\beta^L_{ND} - \beta^L_D$	t-stat		
All Countries	0.49	0.06	0.98	0.06	0.49	5.84	6,643	0.37
Euro Area	0.42	0.09	0.95	0.08	0.53	4.54	2,374	0.49
Non Euro Area	0.56	0.09	0.96	0.12	0.40	2.88	4,269	0.32
Differentiated Goods	0.52	0.10	1.07	0.08	0.55	4.45	3,193	0.38
Euro Area	0.50	0.11	0.99	0.10	0.48	2.99	1,264	0.48
Non Euro Area	0.54	0.15	1.12	0.20	0.58	2.34	1,928	0.32

Table 5: Life-Long Pass-through: Dollar vs. Non-Dollar Pricers

	Dollar Pricers			Non-Dollar Pricers		
	All	6-digit	10-digit	All	6-digit	10-digit
Frequency						
Median	0.10	0.13	0.11	0.07	0.08	0.07
Mean	0.24	0.17	0.18	0.11	0.12	0.12
SD	0.24	0.17	0.20	0.15	0.14	0.16
Size						
Median	0.01	0.02	0.02	0.02	0.02	0.02
Mean	0.01	0.02	0.02	0.02	0.02	0.02
SD	0.14	0.11	0.10	0.13	0.12	0.12
Abs. Size						
Median	0.07	0.08	0.08	0.06	0.07	0.06
Mean	0.10	0.10	0.10	0.09	0.10	0.08
SD	0.12	0.09	0.09	0.12	0.11	0.11

Table 6: Frequency and Size of Price Adjustment: Dollar vs. Non-Dollar Pricers

Parameter	Symbol	Value	Source
Discount Factor	δ	$0.94^{1/12}$	Monthly data
St.dev. of e_t	σ_e	2.5%	Data
Persistence of e_t	ρ_e	0.99	Data
St.dev. of a_t	σ_a	8.0%	Absolute Size of Price Adjustment
Persistence of a_t	ρ_a	0.95	Bils and Klenow (2004)
Inertia in P_t	α	0.95	
Long-run response of P_t	$\bar{\phi}$	50%	
Cost sensitivity	ϕ	0.75	Input-Output Tables
Calvo parameter	θ	0.875	Price Durations
Menu Cost	κ	5%	Price Durations
Demand elasticity	σ	5	Steady State Mark-up
Demand super-elasticity	ε	3	

Table 7: Parameter Values