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Asymmetric exchange rate pass-through: Evidence from Peru

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Asymmetric exchange rate pass-through: Evidence from Peru*

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Abstract

We study the response of prices to exchange rate shocks for the Peruvian economy in a non-linear context. For that purpose we specify a Structural Vector Autorregressive model (SVAR) and compute impulse-responses functions for prices after exchange rate shocks. We follow Hamilton (2010) and Kilian and Vigfusson (2011), who explore the presence of asymmetric effects on USA output after oil prices shocks that either decrease or increase oil prices. In our setup we analyze shocks that either appreciate or depreciate the local currency under censored exchange rate changes. The results exhibit a remarkable asymmetry in the response of consumer prices and wholesale import good prices, both on impact and on propagation. In absolute value, the effect of a depreciation shock on the consumer price index after one year is about twice the size of that corresponding to an appreciation shock. Roughly speaking, the one-year passthrough to prices is 20 percent under depreciations and only 10 percent after appreciations.

JEL Classification: C32 E31 F31

Key words: Exchange rate pass-through, asymmetric impulse responses, non-linear models.

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1 Introduction

The exchange rate pass-through to prices (ERPT) never ceases to be a relevant topic. Although there exists general evidence of reduced pass-through to prices in countries with floating exchange rate regimes, the effect is still non negligible. Given the latter argument, in this paper we go one step further and ask whether the response of prices is symmetric across depreciation and appreciation episodes. This question is important since the evidence of a decreasing ERPT during the last decade is mainly based on an episode of a persistent appreciation of the domestic currency with occasional periods of depreciation, as it is depicted in Figure 1. Moreover, during the period 1993:01-2014:12 we find for the year-to-year depreciation rate that 58% of the observations correspond to a depreciation episode and 42% correspond to appreciation ones. Therefore, we have a considerable amount of information in order to conduct proper inference.

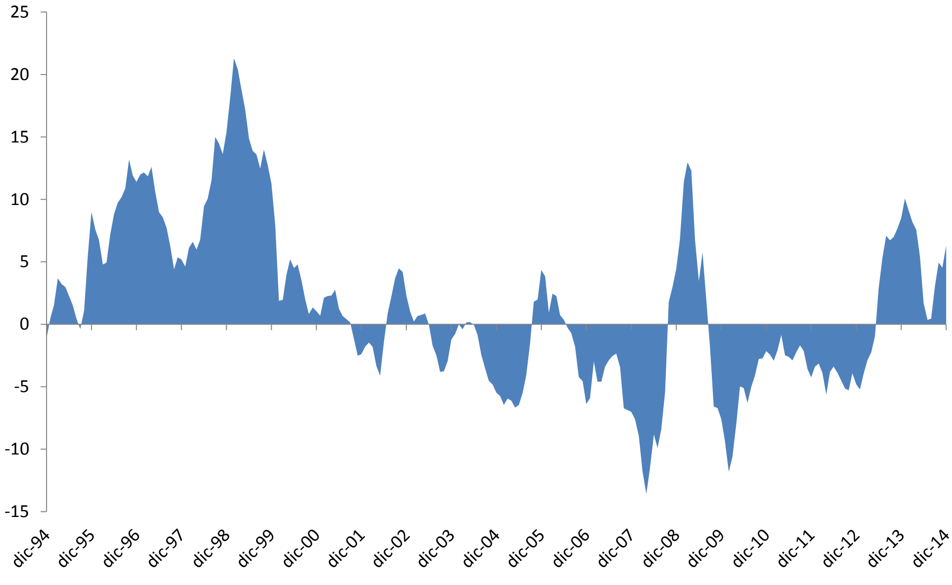

Figure 1. *Year-on-year depreciation rate of the Sol (in percentages)*

We focus on the case of Peru and we find that depreciation shocks produce stronger effects than appreciation shocks. This result is relevant in terms of monetary policy design, since we provide aggregate evidence of different behavior of prices under the two considered regimes, and price stability is the ultimate goal for a Central Bank that actively implements the Inflation Targeting scheme, such as the BCRP. Furthermore, our analysis can be easily extrapolated to similar Small Open Economies and Emerging Market ones, but we leave this for future research agenda.

There exists a large literature that studies the phenomenon of ERPT for the Peruvian economy. In particular, we can mention the work of Winkelried (2003), Miller (2003), Maertens Odría, Castillo, and Rodríguez (2012), Winkelried (2013), among others. The main conclusion is that the ERPT, if any, is small and has been decreasing during the last decade. Winkelried (2003) studies the ERPT asymmetry in the sense of different stages of the Business Cycle. On the other hand, this is the first paper that explores the presence of asymmetries in terms of the direction of the exchange rate variation.

The task of capturing the differences between the responses of prices during depreciation and appreciation episodes is far from being straightforward. Empirical analysis for asymmetric responses can be found in the oil price literature. In this context, Hamilton (2010) and Kilian and Vigfusson (2011) propose the use of censored variables in Structural Vector Autoregressions model (SVARs) in order to capture these asymmetries (see the complete debate in Hamilton (2009) and Kilian (2009)). In this paper we import this methodology and apply it to the ERPT context, where we take the benchmark specification of Winkelried (2013). In addition, we propose a bootstrap routine with the purpose of computing the confidence intervals for the impulse response functions, a procedure that might be useful for readers interested in the methodology.

Finally, we link our results with existing theoretical models in the literature. In this regard, the asymmetry in price setting behaviour is not new. Since Ball and Mankiw (1994), there exists a vast literature that takes into account some form of asymmetries when studying price setting[1]. In particular, we mention the work of Pollard and Coughlin

[1]See Meyer and Von Cramon-Taubadel (2004) for a literature survey.

(2004), who present a price setting model for importing good firms.

The document is organized as follows: section 2 covers the empirical analysis, section 3 discusses the main results, section 4 describes the theoretical background and section 5 concludes.

2 The Model

In this section we study the presence of asymmetries in the responses of prices to exchange rate shocks. To do that, we employ the framework described by Hamilton (2010) and Kilian and Vigfusson (2011), a framework previously used to capture the nonlinear responses after oil prices shocks. To the best of our knowledge, this is the first paper that treats the asymmetry of ERPT in this way. We start with a minimal example and then develop the current framework used for the estimation.

2.1 Non-linear models and econometrics

A non-linear relationship can be captured as follows. First, consider the regression model between y and its own past, such that

$$y_t = c + \phi_1 y_{t-1} + e_t$$

The latter model is linear, therefore the relationship between y and its past is symmetric. However, if we wanted to capture a non-linear relationship, then we specify

$$y_t = c + \phi_1 y_{t-1} + \phi_1^F F\left(y_{t-1}\right) + \widetilde{e}_t \tag{1}$$

where $F\left(.\right)$ is a well defined non-linear function. Ideally, we would like to estimate equation (1), and this is what Kilian and Vigfusson suggest. On the other hand, Hamilton

(2010) argues that it is preferable to estimate the expression

$$y_t = c + \phi_1^F F\left(y_{t-1}\right) + e_t^* \tag{2}$$

since it is a more parsimonious model. Clearly, in a single equation model this difference is neglibible. However, once we turn to a multivariate model with many lags and variables, the context changes dramatically. That is, we consider a $VAR\left(p\right)$ model

$$\mathbf{Y}_t = \mathbf{c} + \mathbf{\Phi}_1\mathbf{Y}_{t-1} + \mathbf{\Phi}_2\mathbf{Y}_{t-2} + \cdots + \mathbf{\Phi}_p\mathbf{Y}_{t-p} + \mathbf{E}_t \tag{3}$$

where $\mathbf{E}_t \sim iidN\left(0, \Omega\right)$ and where $\dim \mathbf{Y}_t = M \gg 2$. The model in this form has $K = M^2p + M$ coefficients. Furthermore, if we wanted to consider a non-linear relationship then

$$\begin{aligned} \mathbf{Y}_t &= \mathbf{c} + \mathbf{\Phi}_1\mathbf{Y}_{t-1} + \mathbf{\Phi}_2\mathbf{Y}_{t-2} + \cdots + \mathbf{\Phi}_p\mathbf{Y}_{t-p} \\ &\quad + \mathbf{\Phi}_1^F\mathbf{F}\left(\mathbf{Y}_{t-1}\right) + \mathbf{\Phi}_2^F\mathbf{F}\left(\mathbf{Y}_{t-2}\right) + \cdots + \mathbf{\Phi}_p^F\mathbf{F}\left(\mathbf{Y}_{t-p}\right) + \widetilde{\mathbf{E}}_t \end{aligned} \tag{4}$$

where $\mathbf{F}\left(.\right)$ is a vector-valued non-linear function. Now the model has parameters

$$K^F = 2M^2p + M \gg M^2p + M = K \tag{5}$$

Clearly, controlling for non-linearities implies a large cost in terms of degrees of freedom, which means that the model is far from being parsimonious. Once this problem is presented, then it is clear that Hamilton's suggestion is the more reasonable one. Of course, in an ideal world we would have an infinite sample and this discussion would not be relevant anymore. Because this is not the case, it is reasonable to estimate the expression

$$\mathbf{Y}_t = \mathbf{c} + \mathbf{\Phi}_1^F\mathbf{F}\left(\mathbf{Y}_{t-1}\right) + \mathbf{\Phi}_2^F\mathbf{F}\left(\mathbf{Y}_{t-2}\right) + \cdots + \mathbf{\Phi}_p^F\mathbf{F}\left(\mathbf{Y}_{t-p}\right) + \mathbf{E}_t^* \tag{6}$$

Naturally, the exclusion of the linear terms generates a potential source for the ommited variable bias. On the other hand, as it is clear in (5), the cost in terms of degrees of freedom implies that the estimation of (4) might be either unfeasible or highly unstable. As we can see, we face trade-off between consistency and parsimony. We take the lead of Hamilton and estimate the most parsimonious specification through OLS.

2.2 The SVAR model

In order to estimate the response of prices to exchange rates for the peruvian economy we specify a Structural Vector Autoregressive model (SVAR). We explore the recent literature, since it will give us light about which is the most suitable information set for the estimation. In particular, a recent paper by Winkelried (2013) explores the time variation in the ERPT for Peru, and concludes that if any, it has been decreasing during the last decade. Since this is the most recent work for ERPT, we take its information set as a benchmark. Ideally, we could use a larger information set and estimate a Factor-Augmented model (FAVAR) or even assume a more sophisticated identification scheme than the standard 'Cholesky' decomposition. However, these technical details are out of the scope of this paper, and the current specification is enough for answering our main research question.

Consider the vector $\mathbf{Y}_t = \left[\Delta RER_t, \Delta GDP_t, \Delta ER_t, \pi_t^{WM}, \pi_t^{WD}, \pi_t^{CPI}\right]'$. All variables are expressed in year-to-year changes and they are:

- ΔRER_t: Year-to-year change in Real Exchange Rate.
- ΔGDP_t: Year-to-year growth of GDP.
- ΔER_t: Year-to-year depreciation.
- π_t^{WM}: Year-to-year inflation of imported goods (wholesale price index).

- π_t^{WD}: Year-to-year inflation of domestic goods (wholesale price index).
- π_t^{CPI}: Year-to-year inflation (consumer price index).

Winkelried (2013) shows that plausible ERPT estimates can be obtained using this information set. As a benchmark, we use the same information set[2]. Furthermore, consider the nonlinear function

$$\mathbf{F}(\mathbf{Y}_t) = \begin{cases} \mathbf{Y}_t, & \text{if } \Delta ER_t > 0 \\ \mathbf{Y}_t^* & \text{otherwise} \end{cases} \tag{7}$$

where $\mathbf{Y}_t^*$ has $\Delta ER_t = 0$ and keeps the other elements of $\mathbf{Y}_t$ unchanged. That is, we include a censored variable in our VAR model. The use of censored variables in VAR models is not new, because it comes from the Oil Prices shocks literature, see Hamilton (2010) and Kilian and Vigfusson (2011). We take this lead and present the estimation results below. However, we consider separately the cases of positive ($\Delta ER_t > 0$) and negative ($\Delta ER_t < 0$) changes in the exchange rate (depreciations and appreciations, respectively), but we include the same information for the remaining variables.

3 Results

In this section we explore the response of prices for the cases of depreciations ($\Delta ER_t > 0$) and appreciations ($\Delta ER_t < 0$) separately. We compute the Impulse Responses for the non-linear SVAR through the use of conditional forecasts as in Kilian and Vigfusson (2011). In addition, we compute the error bands using bootstrapping techniques (see Appendix B for details about the simulation). Since there is no consensus about how to sort impulse responses and compute these bands (see e.g. Inoue and Kilian (2013)), we follow the literature and plot the median values and percentiles for each horizon.

[2]See Appendix A for details about the data description.

Results exhibit a noticeable asymmetry in the response of prices, with the special case of imported good prices, but without ignoring consumer prices. Figure 2 and Figure 3 depict the main results of the paper. For comparison purposes we present the effects of an appreciation with the opposite sign, so that in both regimes we have the same shock size. Symmetry would imply that the IRFs of both regime coincide.

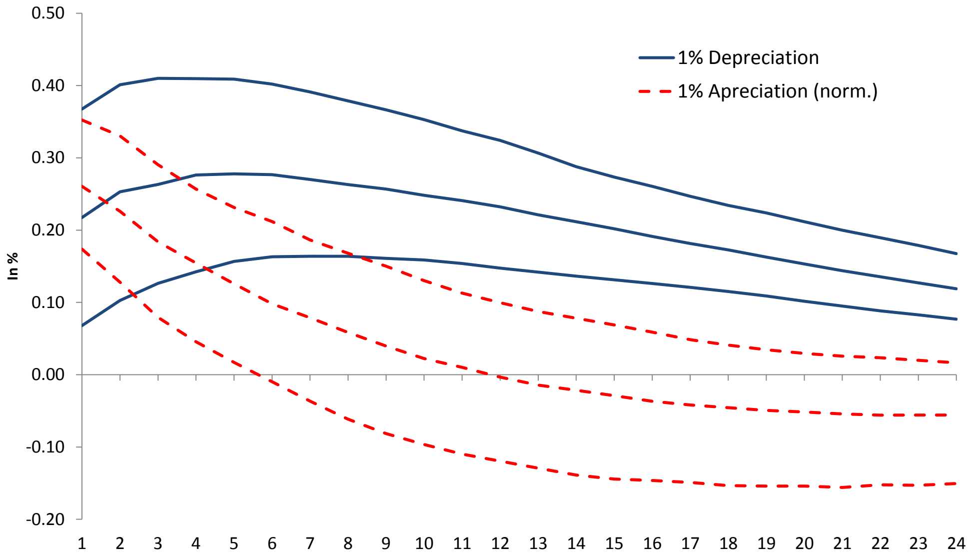

Figure 2. *Responses of Consumer Prices, 68% bands. The IRF for appreciation shock has an inverted sign.*

When there is a depreciation of 1%, CPI Inflation raises up to 0.3%, reaching this point in six months. Moreover, this effect is persistent and statistically significant for almost two years. Furthermore, wholesale prices of imported goods inflation raises up to 0.45%, reaching this maximum in three months. This effect is also persistent and significant for almost two years.

On the other hand, when there is an appreciation of 1%, CPI Inflation roses up to 0.3% as well, reaching this point in one month. Moreover, this effect is less persistent than the case of depreciation and it turns to be statistically insignificant at the 68% level of confidence after one year. Furthermore, wholesale prices of imported goods inflation does not exhibit a significant reaction in this case.

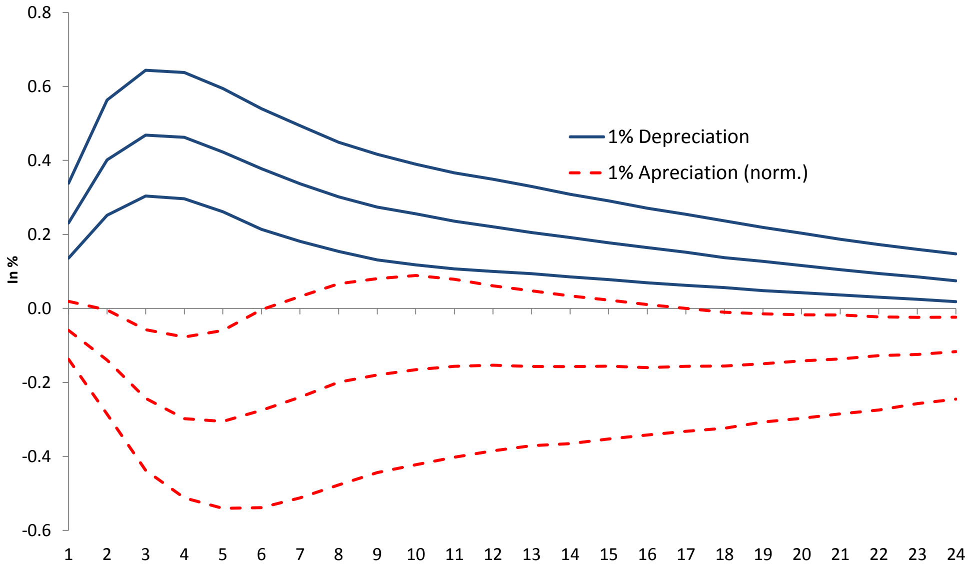

Figure 3. *Responses of Wholesale Prices of Imported Goods, 68% bands.The IRF for appreciation shock has an inverted sign.*

Overall, we observe substantial differences in the transmission of exchange rate shocks to prices, both on impact and propagation. Prices exhibit hump-shaped responses, meaning that the ERPT, if any, is not complete neither on impact, nor after many periods. Moreover, the responses are larger and more persistent under depreciations than appreciations.

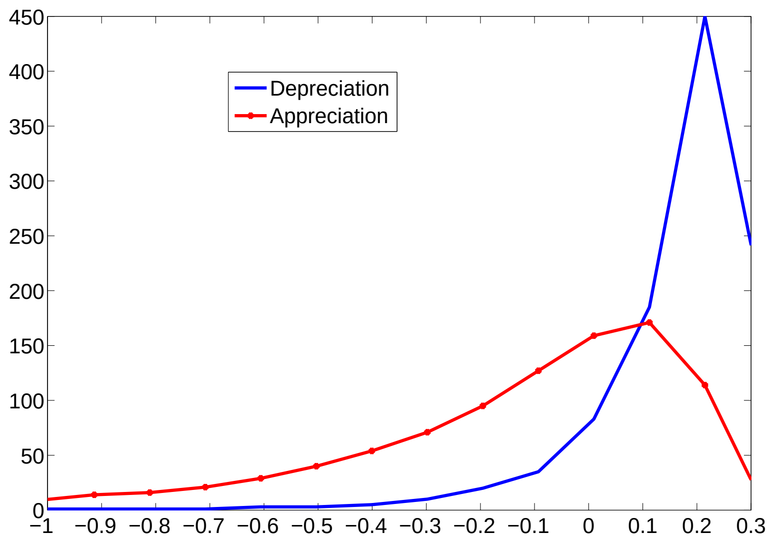

Figure 4. *Exchange Rate Pass-Through to Prices after one year*

The results can be explained by the simple framework outlined in Figure 5. There we see that the difference between depreciation and appreciation shocks do not only rely on the downward rigidity of prices which would directly make appreciation shocks weaker than depreciation shocks. We also need to take macroeconomic effects into account. Two possible aggregate demand effects arise due to exchange rate shocks. The traditional expenditure switching effect and wealth effects due to financial dollarization of assets and liabilities of households and firms. The wealth effect of a depreciation shock is negative and the wealth effect of an appreciation shock is positive. We postulate that the extent of the wealth effect due to appreciations has been much stronger than the extent of wealth effect due to depreciations. This has been so because the effect of depreciation shocks over aggregate demand conditions are softened by macroeconomic policy.

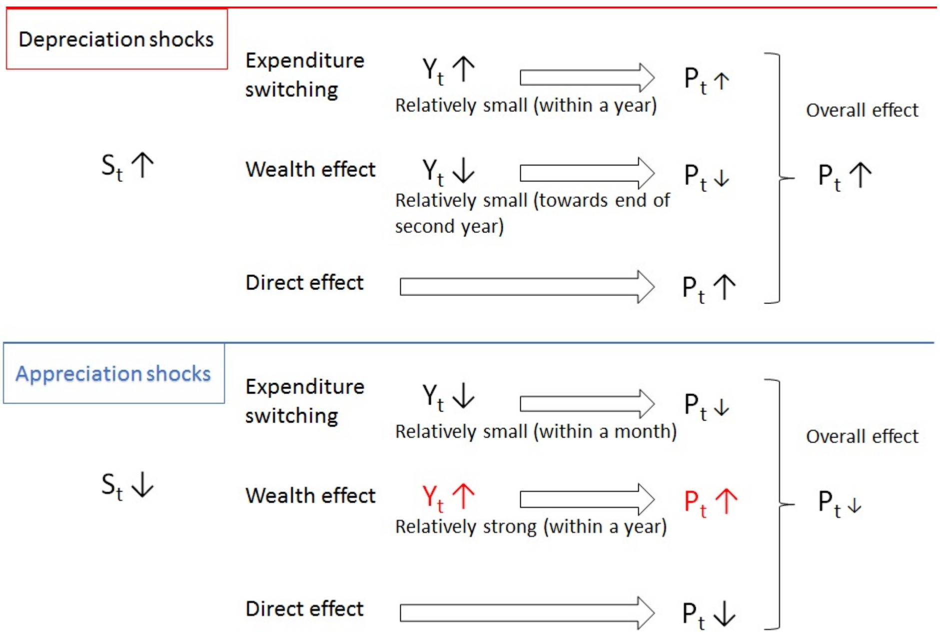

Figure 5. *Macro effects of exchange rate shocks*

Therefore, it is the extend of the positive wealth effect of an appreciation shock that may explain an upward bias in prices when the exchange rate appreciates. Overall, the effect on prices is dampened by the two countervailing forces: prices falling due to the exchange rate appreciation versus prices rising due to positive wealth effects in aggregate demand. Figure 6 shows that a depreciation shock has the strongest positive effect on GDP during month 4 and 5. At the end of the second year the effect is inverted in sign. This means that the expenditure switching effect seems to be more important in the short run while a weak negative wealth effect kicks in in the long run. Conversely, the effects of an exchange rate appreciation shock (inverted in sign) show a strong positive wealth effect within a year. This result helps explains the overall asymmetric exchange rate pass-through to prices.

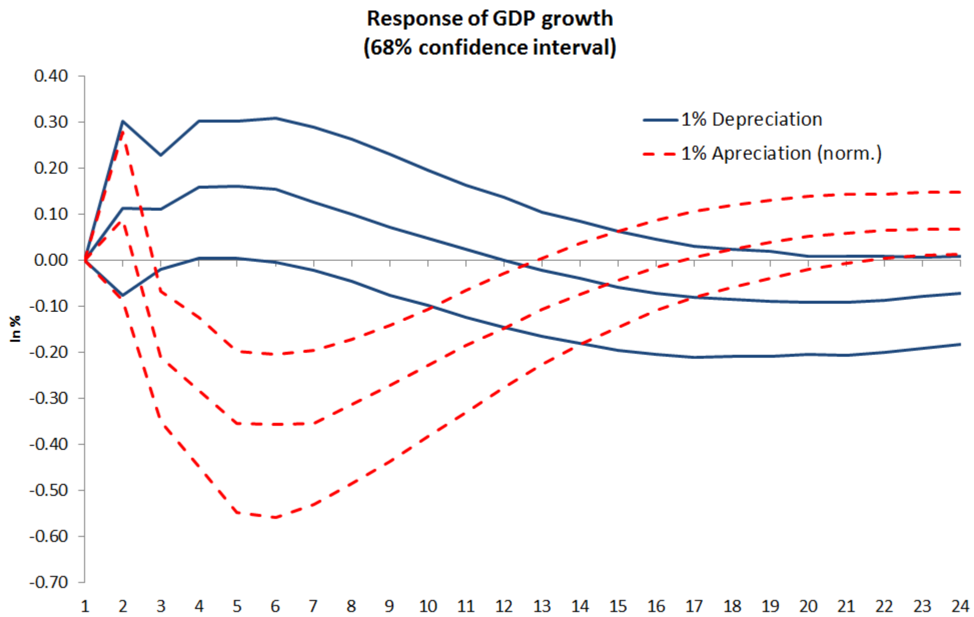

Figure 6. *Responses of GDP, 68% bands.The IRF for appreciation shock has an inverted sign.*

4 Theoretical Explanations for asymmetries in price setting behavior

Having observed the marked asymmetry in price responses, conditional on being in either a depreciation or an appreciation episode, one may wonder whether there is a rational and fully grounded explanation for this results. In this section we briefly review the existing models that treat this topic. One limitation of the models presented below is the fact that they are static. Therefore, the theoretical background that we will provide only explains the impact effect, but not the propagation of the responses observed in the data.

4.1 Ball and Mankiw's model

This is the first approach to the theme. The authors analyze the price setting problem in the context where there exists shocks to the firm's desired relative price and where prices are set for two periods (even periods). However, the firm can choose to reset its price in an odd period by paying a menu cost, which is internalized by the firm.

The crucial point in this model is the fact that desired prices also vary with a positive inflation rate. Therefore, if there is a negative shock to the price, then this is compensated by the positive inflation. As a result, the firm will be less inclined to reset its price. On the other hand, if there is a positive shock to the price, the opposite will occur. As it can be noticed, this asymmetry will disappear in a zero inflation context. Furthermore, the size of the shock also matters for the final decision. For example, a big negative shock is less likely to be compensated than a small one.

4.2 Pollard and Coughlin's model

The authors present a simple framework for exploring the effects of changes in the value of foreign currency to prices. In particular, they focus their attention on importing good firms. The idea is that exchange rates affect prices through the costs channel. That is, changes in costs will affect markups and therefore the price setting decision. They find that, in cases where firms can use both domestic and imported inputs, a shock in exchange rates will alter the profit margins. How much? It depends on the elasticity of factor prices to exchange rates and depends on the elasticity of markups to prices. The final decision of price resetting depends on the firm. Crucially, firms face strategic complementarities, and they will modify their prices according to their best response correspondence given the price of the competitor. As a result, if the exchange rate shock increases profit margins, then the firm could maintain its price and get additional profits or strategically cut it to gain some market share. On the other hand, if the exchange rate shock decreases profit margins, then the firm could raise its price to restrain it or strategically maintain its price to keep its market share constant.

We interpret the decisions related with market share as the ones that depend on the demand elasticity of each market and the market power of the firm. In addition, as it is somewhat obvious, monopolistic competitors will be more resilient to cut their prices, but will accept increases in their prices.

Pollard and Coughlin (2004) also mention that one possibility of incomplete ERPT is the fact that firms may switch between domestic and imported inputs, depending on their price. Therefore, if there is an exchange rate shock that affects price factors, firms could re-compose their input bundle such that the effect on the margin is, at least, partially mitigated. Thus, this reduces the possibility of price adjustments. Overall, we can notice that there are many reasons of why the ERPT may be asymmetric, so that our empirical evidence is somewhat plausible.

4.3 Remarks

Clearly, in line with Ball and Mankiw (1994) and Pollard and Coughlin (2004), our empirical results say that firms react differently when there are positive shocks to prices than negative ones. The observed asymmetry might be caused by one of the motives mentioned above. However, these are static models without intertemporal decisions. Therefore, the next step in our research would be to write down a dynamic model capable of replicate these facts.

5 Concluding Remarks

The central question of this paper is whether prices respond in a symmetric fashion to exchange rate shocks. We have found aggregated evidence for the 'no' using statistical methods, i.e. an asymmetric ERPT that depends on whether the current episode is a depreciation or an appreciation. This result is relevant in terms of monetary policy design, since we provide aggregate evidence of different behavior of prices under the two considered regimes, and price stability is the ultimate goal for a Central Bank that actively implements the Inflation Targeting scheme. Furthermore, our analysis can be easily extrapolated to similar Small Open Economies and Emerging Market ones, but we leave this for future research agenda.

A Data Description

Data is taken from the website of the Central Reserve Bank of Peru (BCRP) and from the National Institute of Statistics (INEI) for the period of January 1992 to December 2014. All variables were expressed in year-to-year percent changes. These variables are:

- RER_t: Bilateral Real Exchange Rate Index (2009=100).
- GDP_t: Gross Domestic Product Index (2007=100) [3].
- ER_t : Nominal Exchange Rate (S/. per US$).
- WMP_t: Wholesale Prices of Imported Goods Index (1994=100).
- WDP_t: Wholesale Prices of Domestic Goods Index (1994=100).
- CPI_t: Consumer Price Index of Lima (2009=100).

B Computation of Impulse responses in the non-linear model

The computation of responses in a nonlinear context is far from being straightforward. In particular, since we use censored variables, the responses must be consistent with this assumption. In this regard, Kilian and Vigfusson (2011) suggest the use of conditional forecasts for a given horizon through a bootstrap procedure. We describe our loop below for the interested readers.

0. Recover the residuals of the model and substract the mean $\widehat{\mathbf{E}}^* = \left\{\widehat{\mathbf{E}}_t^* - (1/T)\sum_{t=1}^{T}\left(\widehat{\mathbf{E}}_t^*\right)\right\}_{t=1}^{T}$.
1. Set the numbers N, L and M, the horizon h and the size of the shock δ.

[3] Growth rates for dates before 2003 were recovered using the index of base 1994=100

2. For each $n = 1, \ldots, N$ do

a. Draw $\widetilde{\mathbf{E}}^* = \left\{\widetilde{\mathbf{E}}_t^*\right\}_{t=1}^{T}$ as random draws with replacement from $\widehat{\mathbf{E}}^*$.

b. Simulate $\widetilde{\mathbf{Y}}$ using $\widetilde{\mathbf{E}}^*$ and equation (2).

c. Estimate the model using the new data $\widetilde{\mathbf{Y}}$ and get the Cholesky decomposition matrix $\widetilde{\mathbf{P}}$.

d. For each $m = 1, \ldots, M$ do

- **i.** Draw $\mathbf{U}_{1,T^*}$ from $N(0, I)$ and create $\mathbf{U}_{2,T^*}$ such that the third component is δ and the others are equal to $\mathbf{U}_{1,T^*}$.

- **ii.** For each $l = 1, \ldots, L$ do

  - ***.** Draw $r \sim U(1, T)$ and take $T^* = round\,(r)$.

  - ****.** Compute recursive forecasts for h periods starting from $\widetilde{\mathbf{Y}}_{T^*}$. Use $\widetilde{\mathbf{E}}_{1,T^*}^* = \widetilde{\mathbf{P}}\mathbf{U}_{1,T^*}$, $\widetilde{\mathbf{E}}_{2,T^*}^* = \widetilde{\mathbf{P}}\mathbf{U}_{2,T^*}$ and equation (2). Call them $\widetilde{\mathbf{Y}}_1^l = \left\{\widetilde{\mathbf{Y}}_{1,t}^l\right\}_{t=T^*+1}^{T^*+h}$ and $\widetilde{\mathbf{Y}}_2^l = \left\{\widetilde{\mathbf{Y}}_{2,t}^l\right\}_{t=T^*+1}^{T^*+h}$.

- **iii.** Take the averages $\widetilde{\mathbf{Y}}_1^m = \sum_{l=1}^{L} \widetilde{\mathbf{Y}}_1^l$ and $\widetilde{\mathbf{Y}}_2^m = \sum_{l=1}^{L} \widetilde{\mathbf{Y}}_2^l$.

e. Take the averages $\widetilde{\mathbf{Y}}_1^n = \sum_{m=1}^{M} \widetilde{\mathbf{Y}}_1^m$ and $\widetilde{\mathbf{Y}}_2^n = \sum_{m=1}^{M} \widetilde{\mathbf{Y}}_2^m$.

f. Compute the difference $IRF_n = \widetilde{\mathbf{Y}}_1^n - \widetilde{\mathbf{Y}}_2^n$.

3. Collect all $IRF = \{IRF_n\}_{n=1}^{N}$ and compute percentiles.

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