Does exchange rate pass-through respond to measures of macroeconomic instability?

Reginaldo Pinto Nogueira Júnior
Fundação João Pinheiro, Brazil

Miguel A. León-Ledesma
University of Kent, UK

Abstract: We argue that, theoretically, exchange rate pass-through (ERPT) into consumer prices may be nonlinear in contrast to standard linear estimates found in the literature. ERPT can be higher in periods of financial or confidence crises, when firms have no incentive to absorb cost increases in their margins. We test this hypothesis applying a logistic smooth transition (STR) model to Mexican data. Using two different measures of macroeconomic instability as transition variables, we find that ERPT does seem to increase in periods of macroeconomic distress, which highlights the importance of a stable macroeconomic environment in reducing ERPT in emerging markets.

JEL Classification: E31, E52, F41.

Keywords: Exchange rate pass-through, smooth transition regression models, emerging markets.

2 Corresponding author: Reginaldo Pinto Nogueira Júnior. Escola de Governo, Fundação João Pinheiro. Alameda das Acáias, 70. Belo Horizonte, Brazil. CEP: 31.275-150. E-mail: reginaldo.nogueira@fjp.mg.gov.br.
1. Introduction

The extent to which exchange rate changes are transmitted into prices is of utmost importance for policymakers. This effect, known as exchange rate pass-through (ERPT), influences not only current inflation, but also inflation expectations, the setting of monetary policy, and the ability of exchange rate changes to correct trade imbalances.

Various studies have shown that ERPT has declined in recent years\(^3\). The most common interpretation for this finding is that of Taylor (2000), which relates the decline to a lower inflation environment. According to this view the rate of inflation affects the persistence of costs changes, which is positively correlated with ERPT. A somewhat similar explanation argues that this finding is a corollary of credibility gains of monetary policy (see for e.g. Mishkin and Savastano, 2001; Choudhri and Hakura, 2006). Both hypotheses suggest that there might be a role for the macroeconomic environment in determining the degree of ERPT.

We analyze this corollary directly by investigating the existence of a possible link between the macroeconomic environment and the degree of ERPT. We first present a simple model where we put forward the possibility that, theoretically, ERPT may be nonlinear in contrast to linear estimates traditionally found in the literature. In particular, ERPT may be higher in periods of macroeconomic instability, such as financial or confidence crises. Then we test this hypothesis using a smooth transition regression (STR) model of ERPT for Mexican data. The case of Mexico is rather important being one of the largest emerging markets and having faced important crises in the past decades\(^4\).

There is little work on the issue of nonlinearities and asymmetries in ERPT\(^5\). In addition, the existing literature provides mixed evidence on the matter: while studies such as Herzberg, Kapetanios and Price (2003) and Marazzi et al. (2005) have not found evidence of nonlinear or asymmetric behaviour, others such as Gil-Pareja (2000) and Mahdavi (2002) have found support for nonlinear ERPT. Moreover, much of the literature has focused exclusively on asymmetries with respect to the size and direction of exchange rate changes. Hence, a further contribution of this paper is the investigation of another potential source of nonlinearity in ERPT.

\(^3\) See for e.g. Gagnon and Ihrig (2004) and Choudhri and Hakura (2006).
\(^4\) For an overview on the recent developments of the Mexican economy see Ball and Reyes (2004).
\(^5\) For a brief survey see Marazzi et al. (2005).
Our results present some evidence in favour of nonlinearities in ERPT with respect to our measures of macroeconomic instability (EMBI spreads of dollar-denominated bonds and real interest rate differentials with the United States). This finding suggests that market’s confidence in a stable macroeconomic environment plays an important role in reducing ERPT. Evidently, this conclusion does not rule out other possible sources of nonlinearities, but it does complement our understanding on ERPT dynamics in emerging markets.

The rest of the paper is structured as follows. Section 2 presents a simple model of nonlinear ERPT. Section 3 discusses our empirical methodology. Section 4 presents the results. Finally, section 5 concludes.

2. Theory

A simple theoretical model helps illustrating the reasons for the potential existence of a nonlinear ERPT that depends on the macroeconomic environment. The model we present here is very parsimonious but it suffices to illustrate the argument. We build on Korhonen’s (2005) model for ERPT into import prices, which draws on the micro-founded model of Burnstein, Eichenbaum and Rebelo.

Let us consider a foreign firm that exports its product to the domestic country. Under imperfect competition, a profit-maximizing exporter with prices set in importing country currency will set its price at time \( t \) equal to:

\[
P_t = \theta E_t C^*_t
\]

(1)

Where \( P \) is the local currency price, \( C^* \) is the exporter’s marginal cost expressed in its own currency, \( E \) is the domestic exchange rate, and \( \theta \) is a mark-up over marginal cost.

We assume the mark-up to respond to demand pressures in the importing country. Moreover, we also assume the mark-up to depend on the importing country’s general macroeconomic stability, i.e. when the economy faces a financial or a confidence crisis, ERPT is higher. The intuition behind this hypothesis is that the firm’s decision on how much to pass-through cost changes into prices depends on its view on the importing country’s macroeconomic conditions. In periods of bad macroeconomic environment in the importer country, the exporter may decide to pass-through a larger proportion of its cost changes in view of the increased likelihood of default from the importer. In periods of good
macroeconomic conditions, the exporter may be willing to reduce the pass through in order to keep the loyalty of a stable export market. Hence, the mark-up has the following functional form:

\[ \theta_t = \theta(\rho, E^{\alpha(Z)}) \]  

(2)

Where \( \rho \) accounts for the demand pressures in the importing country, and the component \( Z \) depicts the nonlinear response to the general macroeconomic condition. We model \( Z \) in a way that high values imply a bad macroeconomic environment. In other words, \( Z \) would actually be a measure of macroeconomic instability. The function \( \omega(Z) \) can be seen as a mark-up multiplier, where firms respond more to exchange rate changes if their confidence in the economy is low. Hence, during a crisis ERPT would increase.

From (1) and (2), simple log-linearised reduced form equation for prices would be:

\[ p_t = \beta c_t + \kappa y_t + \alpha \omega_t + \alpha(Z)e_t \]  

(3)

Equation (3) states that there are two channels of ERPT. The first channel is given by \( \alpha \) and is bounded between 0 and 1. The second channel is given by the function \( \omega(Z) \), and depends on the macroeconomic environment. We will follow Korhonen (2005) and further assume that there is some threshold \( Z^* \) which divides the extreme cases of good (low) values of \( Z \) and bad (high) values of \( Z \) (macroeconomic environment).

\[ \alpha(Z) = \begin{cases} 0; & Z \leq Z^* \\ \psi > 0; & Z > Z^* \end{cases} \]  

(4)

For these two extreme cases we find two different ERPT. If the importing country faces a good macroeconomic environment, then ERPT is equal to \( \alpha \). If the importing country faces a bad macroeconomic environment, then ERPT is equal to \( \alpha + \psi \). We can see that ERPT is higher in the second case, as \( \alpha + \psi > \alpha \). Intuitively, with an unstable macroeconomic environment firms have no incentive to absorb cost increases in their margins. Hence, the model implies that perceptions about the importing country’s general macroeconomic conditions would raise ERPT in a nonlinear way.
Rewriting (3) in difference form, we have:

\[ \Delta p_i = \beta \Delta e^*_i + \kappa \Delta y_i + [\alpha + \alpha(Z)] \Delta e_i \]  

(5)

The above threshold model may be likely for one firm, but not for the aggregate of firms, as there is probably some heterogeneity across firms on their attitude towards the state of the macroeconomic environment (Korhonen, 2005). Following this, we will make use of smooth transition models instead of threshold models in our empirical application.

Although the model presented above is for import prices, we want to analyse ERPT into consumer prices in our empirical analysis, as this is the most important variable for policymakers. Taking as starting point the composition of the consumer price index (CPI):

\[ P_{CPI} = P^\phi_H P^{1-\phi}_T \]  

(6)

Where \( P_{CPI} \) is the consumer price level, \( H \) represents the non-tradable (home) sector, \( T \) the tradable sector, and \( \phi \) is a bounded parameter that shows the participation of each sector in the composition of the CPI.

From equation (6) we can derive an inflation equation for the economy, where \( \pi \) is CPI inflation:

\[ \pi = \phi \pi_H + (1-\phi) \pi_T \]  

(7)

Following the literature on inflation persistence and the importance given to its inertial behaviour, and assuming the same (one) period lag for both tradable and non-tradable sectors, we have:

\[ \pi_{(H)y} = \delta \pi_{(H)y-1} + \varphi \Delta y_t \]  

(8)

\[ \pi_{(T)y} = \delta \pi_{(T)y-1} + \beta \Delta e^*_t + \kappa \Delta y_t + [\alpha + \alpha(Z)] \Delta e_t \]  

(9)

Equation (8) states that home prices are dependent on the output gap and past inflation. Equation (9) shows the tradable sector prices, basically following equation (5) but allowing for some price inertia. Substituting (8) and (9) into (7), yields:
Finally, rearranging equation (10), we have:

$$
\pi_t = \phi(\phi(h_t) + \phi\Delta y_t) + (1-\phi)\{\phi\Delta y_t + [\Delta y_t + (1-\phi)\Delta y_t + \Delta y_t + \Delta y_t + \Delta y_t + \Delta y_t]\Delta y_t\}
$$

Equation (11) yields the basic model for estimating ERPT at the consumer prices level, and can be described as a nonlinear backward-looking Phillips curve. In the next subsection we develop this model into a proper econometric specification.

3. Empirical model

Smooth transition regression (STR) models are a class of nonlinear models that can account for deterministic changes in parameters over time, in conjunction with regime switching behaviour. The STR model takes the following general form:

$$
y_t = \beta_1 x_t + \beta_2 x_t G(s_{t-1}, \gamma, c) + \nu_t
$$

Where, \( s_{t-1} \) is the transition variable, \( G \) is the transition function, \( \gamma \) measures the speed of transition from one regime to the other, and \( c \) is the threshold for the transition function. As discussed by van Dijk, Terásvirta and Franses (2002), the transition function \( G \) is a continuous function bounded between 0 and 1. As \( \gamma \) becomes larger, the change of the transition function becomes almost instantaneous. In this paper we use the logistic smooth transition function (LSTR), which is given by:

$$
G(s_{t-1}, \gamma, c) = \left[1 + \exp\{-\gamma(s_{t-1} - c)\}\right]^{-1}
$$

As explained by Christopoulos and León-Ledesma (2007), the LSTR specification implies that the nonlinear coefficient would take different values depending on whether the transition
variable is below or above the threshold: as \((s_t - c) \rightarrow -\infty\), the coefficient becomes \(\beta_1\); if \((s_t - c) \rightarrow +\infty\) then the coefficient is \(\beta_1 + \beta_2\); and if \(s_t = c\) it becomes \(\beta_1 + \beta_2 / 2\).

We followed the modelling approach described in Lundbergh et al. (2000), van Dijk, Terasvirta and Franses (2002) and Terasvirta (2004). The procedure was the following: first, we tested the null of linearity of a baseline linear model, if the null was not rejected, we accepted the linear model; if the null was rejected, we estimated the model for which the rejection was the strongest; then, we evaluated the estimated model for misspecification (including remaining nonlinearity), if the model failed these tests, an extended model was analysed. We applied LM\(_3\) tests with the null of linearity against LSTR nonlinearity\(^6\). After testing for linearity, we used nonlinear least squares to estimate the parameters in the model\(^7\).

The model has the following form:

\[
\pi_t = \beta_0 + \sum_{i=1}^{n} \beta_{i1} \pi_{t-i} + \sum_{i=0}^{n} \beta_{i2} \Delta \text{imp}_{t-i} + \sum_{i=0}^{n} \beta_{i3} \Delta y_{t-i} + \sum_{i=0}^{n} \beta_{i4} \Delta e_{t-i} + \left( \beta_0 + \sum_{i=0}^{n} \beta_{i5} \Delta e_{t-i} \right) G(s_t; \gamma; c) + \epsilon_t \tag{14}
\]

Where \(\pi_t\) is the inflation rate, \(\Delta \text{imp}\) is the change in import prices (in foreign currency), \(\Delta y\) is real output growth\(^8\), \(\Delta e\) is the exchange rate change, and \(\epsilon\) is an error term.

The transition variables used as measures of macroeconomic instability are the real interest rate differentials (\(\text{rids}\)) with respect to the U.S., and EMBI spreads. The use of \(\text{rids}\) as a measure of macroeconomic instability, and particularly as a leading indicator of confidence crises, has been advocated, among others, by Kaminsky, Lizondo and Reinhart (1998). Regarding the EMBI spreads, they track total returns for traded dollar denominated external debt instruments in the emerging markets. Once the debt is denominated in dollars, there is no exchange rate risk involved, thus representing a measure of “pure country risk”, which makes it our preferred measure of macroeconomic instability.

Monthly data was collected for Mexico. The period of estimation corresponds to the period 1992M1 to 2005M12. Data was obtained from the International Monetary Fund’s

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\(^6\) For a technical discussion of the test the reader is referred to van Dijk, Terasvirta and Franses (2002). We used F-versions of the LM test statistics, because these have better size properties than the chi-square variants.

\(^7\) van Dijk, Terasvirta and Franses (2002) observe that it is quite difficult to obtain an accurate estimate of \(\gamma\), which may, therefore, appear to be insignificant. This should not be interpreted as evidence of weak nonlinearity. Moreover, due to the imprecision of the estimates of the nonlinear function, we followed standard practice in the literature and first estimated \(\gamma\) and \(c\) using a grid search.

\(^8\) We have opted to estimate the model using output growth instead of some measure of output gap in order to avoid using ad-hoc de-trending processes that might eliminate valuable information from the data. Nevertheless, we have also estimated the model using an HP-Filtered output gap, obtaining similar results.
International Financial Statistics. Inflation is the change in the Consumer Price Index. Exchange rate change data is the change of the national currency per unit of dollar. A positive variation means depreciation of the national currency, and a negative one means appreciation. As a proxy of monthly output growth we have used the rate of growth of the Industrial Production Index. Data on import prices is the change in the series of International Commodities Price Index. To construct the rids we used data on money market rates for Mexico and the U.S. CPI inflation was then used to obtain the real rates from the nominal rates collected. Regarding the data on EMBI spreads, it was only available for the period after 1995M1; therefore the estimation using this data has a shorter time period. With the exception of the data on rids and EMBI spreads, which are already normalized, the data used was transformed to logs. The changes refer to the 12-months differences.

4. Results

In our theoretical model we discussed the possibility that the degree of ERPT may be dependent upon the country’s general macroeconomic stability: in periods when the economy faces a confidence crisis, ERPT would be expected to increase, in opposition with periods of macroeconomic stability when ERPT would be expected to decline. In theory both rids and EMBI spreads should provide some proxy of the risks perceived by the market with respect to the general economic condition.

Table 1 shows the linearity tests using up to three lags of rids and EMBI spreads as possible transition variables. We found evidence of nonlinear response of ERPT with respect to both variables, which is consistent with our initial hypothesis.

<table>
<thead>
<tr>
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<th>rids_{t-1}</th>
<th>rids_{t-2}</th>
<th>rids_{t-3}</th>
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<tr>
<td>0.002</td>
<td>0.000</td>
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<tr>
<td>EMBI^{+}_{t-1}</td>
<td>EMBI^{+}_{t-2}</td>
<td>EMBI^{+}_{t-3}</td>
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<td>0.001</td>
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Notes: The numbers are p-values of F variants of an LM3 test of linearity against LSTR nonlinearity.

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9 Data on EMBI spreads was obtained from http://www.cbonds.info, and refers to the last day of each month.
10 Standard unit-root tests were not able to reject the null of non-stationarity for the (log) levels of the variables, but rejected the null for the 12-months differences.
Below we present the results of the estimations of the nonlinear models. Regarding the results, * denotes significance at the 10% level, and ** denotes significance at the 5% level; Sigma is the standard error of the regression; AIC is the Akaike Information Criteria; AR(4) is an autocorrelation test with 4 lags; and RNL is a LM-test for remaining nonlinearity in the model (with the null of no remaining nonlinearity). We also present the graphs of the transition functions and transitions variables over time.

The results using rids as transition variable are:

$$\pi_t = 0.001 + 1.379^{**} \pi_{t-1} - 0.493^{**} \pi_{t-2} + 0.086 \pi_{t-3} - 0.011 \Delta \gamma_t - 0.003 \Delta \rho_{t}^{imp} + 0.040^{**} \Delta \epsilon_t - 0.010 \Delta \epsilon_{t-1} + 0.016^{*} \Delta \epsilon_{t-2} - 0.018 \Delta \epsilon_{t-3} - 0.007 \Delta \epsilon_{t-4} + (-0.002^{**} + 0.001 \Delta \epsilon_t + 0.033^{**} \Delta \epsilon_{t-1} - 0.036^{**} \Delta \epsilon_{t-2} + 0.098^{**} \Delta \epsilon_{t-3} - 0.083^{**} \Delta \epsilon_{t-4})G(rid_{t-1}, \gamma, c) + \nu_t$$

$LSTR : G(rid_{t-1}, \gamma, c) = [1 + \exp(-99(rid_{t-1} - 6.873))]^{-1}$

$R^2 = 0.999; \text{Sigma} = 0.0036; AIC = -11.174; AR(4) = 0.503; RNL = 0.152$

**Graph 1**

Transition function and transition variable (rids)
The results using EMBI+ spreads as transition variable are:

$$\pi_t = 0.002^{**} + 1.322^{**} \pi_{t-1} - 0.428^{**} \pi_{t-2} + 0.058 \pi_{t-3} + 0.018 \Delta y_t - 0.009^{**} \Delta p_t^{imp} + 0.007 \Delta e_t + 0.002 \Delta e_{t-1} + 0.014 \Delta e_{t-2} - 0.016 \Delta e_{t-3} + 0.012 \Delta e_{t-4} + \left(-0.006 + 0.047^{**} \Delta e_t + 0.027 \Delta e_{t-1} - 0.037 \Delta e_{t-2} + 0.099^{**} \Delta e_{t-3} - 0.102^{**} \Delta e_{t-4}\right) G(EMBI_{t-1}, \gamma, c) + \nu_t$$

$LSTR: G(EMBI_{t-1}, \gamma, c) = \left[1 + \exp \left(-4^* (EMBI_{t-1} - 760.8^{**}) \right) \right]^{-1}$

$R^2 = 0.999; Sigma = 0.0035; AIC = -11.174; AR(4) = 0.336; RNL = 0.921$

**Graph 2**

Transition function and transition variable
The estimated nonlinear models pass the diagnostic tests of no remaining nonlinearity and autocorrelation, and provide a good fit to the data. As expected there is a positive relationship between ERPT and our measures of macroeconomic instability, which can be verified by the fact that the sum of the nonlinear exchange rate coefficients is positive. Under both specifications, estimated long-run ERPT\(^ {11} \) is around 1, i.e. there is complete pass-through, when the transition function \( G \) equals 1, but is in the 0.4 to 0.75 range when \( G \) equals zero (the smaller long-run ERPT was estimated in the specification that uses EMBI spreads as transition variable, and, hence, has a shorter sample period). Therefore, the results suggest that there is an important effect of indicators of macroeconomic instability on the ERPT. Moreover, the results represent sensible estimates for ERPT in Mexico over the period analysed, as the literature has usually found higher rates of pass-through for this country than for most emerging markets.

Turning our attention to the graphs, both specifications tell a similar story: the transition function is higher, i.e. closer to 1, basically after the collapse of the Peso in 1995, and around the Russian and Brazilian crises, in late 1998 and the beginning of 1999, which is consistent with our initial hypothesis that ERPT should be higher during periods of confidence crises. It is worth-noting that the threshold values are quite high (6.9% for \( rids \) and 761 basis points for EMBI spreads), which is a sign of the general weakness of macroeconomic fundamentals in Mexico during most of the 1990s. Nevertheless, analysing the graphs of the transition variables we can observe that both \( rids \) and EMBI spreads have been falling in the past few years, particularly post-1999, when Mexico adopted an Inflation Targeting regime. After the year 2000, with the consistent drop of the transition variables, the transition functions are very close to 0, and hence ERPT is substantially lower. In this sense, our model suggests that the adoption of a sounder set of policies may have an important role in reducing ERPT, and hence in decreasing the costs of maintaining inflation stability after depreciations. This is an important finding for countries that have been historically subject to sudden-stops of foreign capital flows, and large exchange rate pressures.

In summary, the combined evidence of the nonlinear models using EMBI and \( rids \) as transition variables provides some evidence in favour of the argument put forward by Mishkin

\(^{11}\) As long-run ERPT we refer to the cumulative effect of a change in the exchange rate on consumer prices until this effect has died-out. This is a standard procedure in the literature on ERPT (see for e.g. Gagnon and Ihrig, 2004). Long-run ERPT was computed as: $LR = \sum_{i=0}^{n} \beta_{4,i} \Delta e_{r,i} + \left( \sum_{i=0}^{n} \beta_{4,i}^{*} \Delta e_{r,i} \right) G(\xi, \gamma; c) / \left[ 1 - \sum_{i=1}^{n} \beta_{1,i} \pi_{r,i} \right]$.  

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and Savastano (2001), Choudhri and Hakura (2006), Gagnon and Ihrig (2004) and others, that policy credibility may influence ERPT. This appears to be the case for Mexico.

5. Conclusions

We have analysed the role of nonlinearities in exchange rate pass-through (ERPT) into consumer inflation for an emerging market economy. In our approach, this nonlinearity appears as a consequence of macroeconomic instability, rather than asymmetries in terms of sign and size of exchange rate changes as in previous literature. We presented this argument in a simple mark-up model of import prices. Under bad economic conditions, firms would have no incentive to absorb cost increases in their margins which would thus lead to a higher ERPT. From this model, we derived an empirical nonlinear ERPT model using smooth transition regressions. The model was applied to Mexican data for the 1992M1-2005M12 period.

Our findings suggest that ERPT does seem to depend on our measures of macroeconomic instability (EMBI spreads of dollar denominated bonds and real interest rate differentials with the U.S.). That is, ERPT appears to be highly nonlinear and dependent on measures of market confidence. In other words, economic crises brought about by poor macroeconomic policies may lead to an increase in ERPT. On the other hand, a more stable environment could account for a decline in ERPT. In this sense the adoption of sounder policies in emerging markets may be an effective tool to reducing ERPT.

References


