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Reginaldo P. Nogueira, Jr.
Miguel A. León-Ledesma

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DOES EXCHANGE RATE PASS-THROUGH RESPOND TO MEASURES OF MACROECONOMIC INSTABILITY?

Reginaldo P. Nogueira, Jr.*
IBMEC-MG

Miguel A. León-Ledesma
University of Kent at Canterbury

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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 (LSTR) 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 codes: E31, E52, F41
Key words: exchange rate pass-through, smooth transition regression models, emerging markets

I. 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

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* Reginaldo P. Nogueira, Jr. (corresponding author): IBMEC-MG. Rua Rio Grande do Norte, 300. Belo Horizonte, Minas Gerais, Brazil. 30.130-130. E-mail: reginaldo.nogueira@ibmecmg.br. Miguel A. León-Ledesma: School of Economics, Keynes College, University of Kent at Canterbury, United Kingdom. CT27NP. E-mail: mal@kent.ac.uk. We would like to thank, without implicating, Dimitris Cristopoulos, Mathan Satchi, John Driffill, Jorge Streb, an anonymous referee, and seminar participants at the 39th annual Money, Macro and Finance Research Group Conference, Birmingham, UK, and the 31st annual meeting of the Brazilian Econometric Society, Foz do Iguaçu, Brazil, for their insightful comments and suggestions. All remaining errors are our own.

(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.[1] 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 example 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 theoretical model where we put forward the possibility that 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. We test this hypothesis using a smooth transition regression (STR) model of ERPT for Mexican data, for the period 1992M1 to 2005M12. The case of Mexico is rather important, being one of the largest emerging market economies, and having faced important crises in the past decades.[2]

There is little work on the issue of nonlinearities and asymmetries in ERPT.[3] 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.

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

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[1] See, for example, Gagnon and Ihrig (2004) and Choudhri and Hakura (2006).

[2] For an overview on the recent developments of the Mexican economy see Ball and Reyes (2004).

[3] For a brief survey see Marazzi *et al*. (2005).

plays an important role in reducing ERPT. This is especially interesting in the case of Mexico because ERPT seems to have been particularly low post-2000, after the adoption of Inflation Targeting in that country. This is in accordance to the literature for other emerging market economies (see for example, Nogueira Jr. and León-Ledesma 2009), and reinforces the argument that the introduction of a set of policies that boost market confidence in the economy can indeed lead to lower ERPT, and hence lower costs of keeping inflation low should a depreciation episode occur. Evidently, this conclusion does not rule out other possible sources of nonlinearities, but it does complement our understanding on ERPT dynamics in emerging market economies.

The rest of the paper is structured as follows. Section II presents a simple model of nonlinear ERPT. Section III discusses our empirical methodology. Section IV presents the results. Finally, Section V concludes.

II. 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 and Juntilla's (2010) model for ERPT into import prices, which draws on the micro-founded model of Burnstein, Eichenbaum and Rebelo (2007).

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:

$$P_t = \theta_t E_t C_t^*, \tag{1}$$

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 θ is a mark-up over marginal cost.

We assume the mark-up responds to demand pressures in the importing country. Moreover, we also assume the mark-up depends 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 form:

$$\theta_t = \theta\left(y, E^{\omega(Z)}\right), \tag{2}$$

where y accounts for the demand pressures in the importing country, and can thus be proxied by aggregate output, 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), a simple log-linearised reduced form equation for prices would be:

$$p_t = \beta c_t^* + \kappa y_t + \alpha e_t + \omega(Z) e_t. \tag{3}$$

Equation (3) states that there are two channels of ERPT. The first channel is given by α 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 and Juntilla (2010) 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).

$$\omega(Z) = \begin{cases} 0; Z \leq Z^* \\ \psi > 0; Z > Z^* \end{cases}. \tag{4}$$

For these two extreme cases we find two different ERPT. If the importing country faces a good macroeconomic environment, then ERPT is equal to α. 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:

$$\Delta p_t = \beta \Delta c_t^* + \kappa \Delta y_t + [\alpha + \omega(Z)] \Delta e_t. \tag{5}$$

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 in their attitude towards the state of the macroeconomic environment (Korhonen and Juntilla 2010). 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):

$$P_{CPI} = P_H^{\phi} P_T^{1-\phi}, \tag{6}$$

where P_{CPI} is the consumer price level, H represents the non-tradable (home) sector, T the tradable sector, and ϕ 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 π is the log-difference of the price level:

$$\pi = \phi\pi_H + (1-\phi)\pi_T. \tag{7}$$

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:

$$\pi_{(H)t} = \delta\pi_{(H)t-1} + \varphi\Delta y_t, \tag{8}$$

$$\pi_{(T)t} = \delta\pi_{(T)t-1} + \beta\Delta c_t^* + \kappa y_t + [\alpha + \omega(Z)]\Delta e_t. \tag{9}$$

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:

$$\pi_t = \phi[\delta\pi_{(H)t-1} + \varphi\Delta y_t] + (1-\phi)\{\delta\pi_{(T)t-1} + \beta\Delta c_t^* + \kappa\Delta y_t + [\alpha + \omega(Z)]\Delta e_t\}. \tag{10}$$

Finally, rearranging equation (10), we have:

$$\pi_t = \delta\pi_{t-1} + [(1-\phi)\kappa + \phi\varphi]\Delta y_t + (1-\phi)\beta\Delta c_t^* + (1-\phi)[\alpha + \omega(Z)]\Delta e_t. \tag{11}$$

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.

III. Empirical model

According to Clifton, Leon and Wong (2001), 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:

$$y_t = \beta_1 x_t + \beta_2 x_t . G(s_{t-i}, \gamma, c) + \upsilon_t, \tag{12}$$

where, s_{t-i} is the transition variable, G is the transition function, γ 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, Terasvirta and Franses (2002), the transition function G is a continuous function bounded between 0 and 1. As γ 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:

$$G(s_{t-i}, \gamma, c) = \left[\left(1 + \exp\left\{ -\gamma (s_{t-i} - c) \right\} \right)^{-1} \right]. \tag{13}$$

As explained by Christopoulos and León-Ledesma (2007), the LSTR specification implies that the nonlinear coefficient takes different values depending on whether the transition variable is below or above the threshold: as $(s_t - c) \rightarrow -\infty$, the coefficient becomes β_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 follow the modelling approach described in Lundbergh et al. (2000), van Dijk, Terasvirta and Franses (2002) and Terasvirta (2004). The procedure is the following: first, test the null of linearity of a baseline linear model; if the null is not rejected, accept the linear model, otherwise estimate the model for which rejection is strongest; then, evaluate the estimated model for misspecification (including remaining nonlinearity); if the model fails these tests, an extended model is analysed. We applied LM_3 tests with the null of linearity against LSTR nonlinearity.[4] After

[4] 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.

testing for linearity, we used nonlinear least squares to estimate the parameters in the model. [5]

The model has the following form:[6]

$$\pi_t = \beta_0 + \sum_{i=1}^{n} \beta_{1,i}\pi_{t-i} + \sum_{i=0}^{n} \beta_{2,i}\Delta p_{t-i}^{imp} + \sum_{i=0}^{n} \beta_{3,i}\Delta y_{t-i} + \sum_{i=0}^{n} \beta_{4,i}\Delta e_{t-i} + \left(\beta_0^* + \sum_{i=0}^{n} \beta_{4,i}^*\Delta e_{t-i}\right) .G(s_t;\gamma;c) + \varepsilon_t, \tag{14}$$

where π is the inflation rate, Δ^{imp} is the change in import prices (in foreign currency) and can thus be seen as imported inflation, Δy is real output growth[7], Δe is the exchange rate change, and ε is an error term.

The transition variables used as measures of macroeconomic instability are the real interest rate differentials (*rids*) with respect to the U.S., and EMBI+ spreads. The use of *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 from the IMF's IFS database. The period of estimation corresponds to 1992M1 to 2005M12. Inflation is the change in the Consumer Price Index. Exchange rate data is the change of the national currency per unit of dollar. A positive variation means depreciation of the national currency. 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

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[5] van Dijk, Terasvirta and Franses (2002) observe that it is quite difficult to obtain an accurate estimate of γ, 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 γ and c using a grid search.

[6] As mentioned before, the model can be described as a nonlinear backward-looking Phillips curve. For another example of nonlinear Phillips curves, see Clifton, Leon and Wong (2001).

[7] 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.

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[8]. 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.

Table 1. Unit root tests

	Lags	ADF	KPSS	DF-GLS
Exchange rate (Δe)	10	0.00^{**}	0.26	-2.00^{**}
Domestic Inflation (π)	5	0.09^{*}	0.71^{*}	-3.20^{**}
Output (Δy)	4	0.00^{**}	0.25	-2.45^{**}
Imported inflation (Δp^*)	2	0.12	0.16	-3.86^{**}

Notes: Numbers of lags determined by the Schwarz Information Criterion. For the ADF test the numbers are the p-values under the null of non-stationarity. For the KPSS test the numbers are the LM-statistics under the null of stationarity. For the DF-GLS test the numbers are the t-statistics under the null of non-stationarity. A time-trend was included in the test equation for inflation in all three tests. ** denotes significance at the 5% level. * denotes significance at the 10% level.

Unit root tests rejected non-stationarity for the 12-months differences (see Table 1). In spite of a vast empirical literature on this topic, the matter of whether these variables are cointegrated is open to dispute. We thus opted to follow standard practice in the literature and estimated the model in differences (e.g., Choudhri and Hakura, 2006; Ca'Zorzi, Hahn and Sanchez, 2007; Gagnon and Ihrig, 2004). Moreover, our choice also reflects the fact that the analysis focuses on short-term dynamics as opposed to long-term equilibrium relationships between the variables, as well as taking into account the short sample period under consideration.

IV. 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 is expected to increase, in opposition with periods of macroeconomic stability when ERPT is 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 2 shows the linearity tests using up to three lags of *rids* and EMBI+ spreads as possible transition variables. We find evidence of nonlinear response of ERPT with respect to both variables, which is consistent with our initial hypothesis.

[8] Data on EMBI+ spreads refers to the last day of each month.

Table 2. Linearity tests

$rids_{t-1}$	$rids_{t-2}$	$rids_{t-3}$
0.002	0.000	0.000
$EMBI+_{t-1}$	$EMBI+_{t-2}$	$EMBI+_{t-3}$
0.001	0.000	0.000

Notes: The numbers are p-values of F variants of an LM_3 test of linearity against LSTR nonlinearity.

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:

$$\begin{aligned}&\pi_t = 0.001 + 1.379^{**}\pi_{t-1} - 0.493^{**}\pi_{t-2} + 0.086\pi_{t-3} - 0.011\Delta y_t - 0.003\Delta p_t^{imp} + \\&0.040^{**}\Delta e_t - 0.010\Delta e_{t-1} + 0.016^{*}\Delta e_{t-2} - 0.018\Delta e_{t-3} - 0.007\Delta e_{t-4} + \\&\left(-0.002^{**} + 0.001\Delta e_t + 0.033^{**}\Delta e_{t-1} - 0.036^{**}\Delta e_{t-2} + 0.098^{**}\Delta e_{t-3} - 0.083^{**}\Delta_{t-4}\right)\\&.G(rid_{t-1}, \gamma, c) + \upsilon_t\end{aligned} \tag{15}$$

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

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

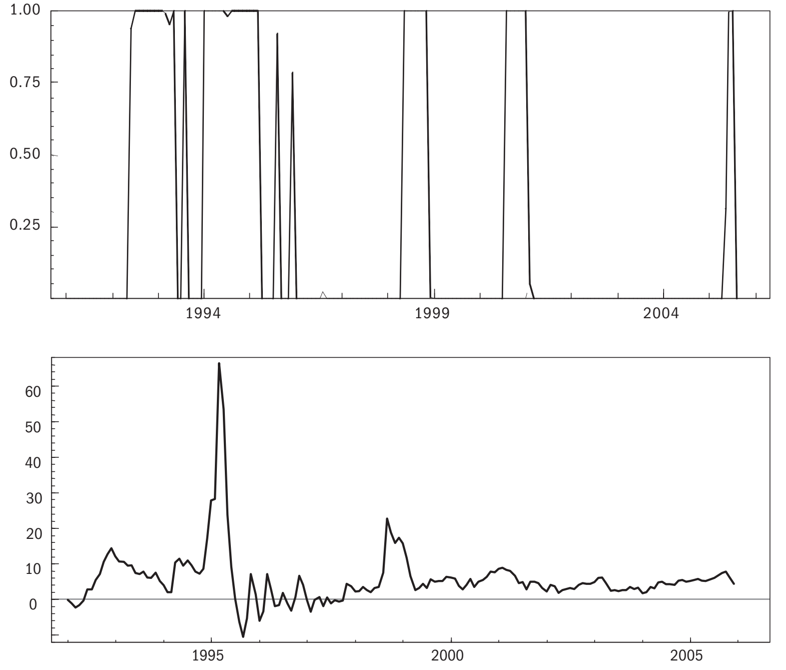
The results using EMBI+ spreads as transition variable are:

$$\begin{aligned}&\pi_t = 0.002^{**} + 1.322^{**}\pi_{t-1} - 0.428^{**}\pi_{t-2} + 0.058\pi_{t-3} + 0.018\Delta y_t - 0.009^{**}\Delta p_t^{imp} + \\&0.007\Delta e_t + 0.002\Delta e_{t-1} + 0.014\Delta e_{t-2} - 0.016\Delta e_{t-3} + 0.012\Delta e_{t-4} + \\&\left(-0.006 + 0.047^{**}\Delta e_t + 0.027\Delta e_{t-1} - 0.037\Delta e_{t-2} + 0.099^{**}\Delta e_{t-3} - 0.102^{**}\Delta_{t-4}\right)\\&.G(EMBI_{t-1}, \gamma, c) + \upsilon_t\end{aligned} \tag{16}$$

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

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

Figure 1. Transition function and transition variable (*rids*)

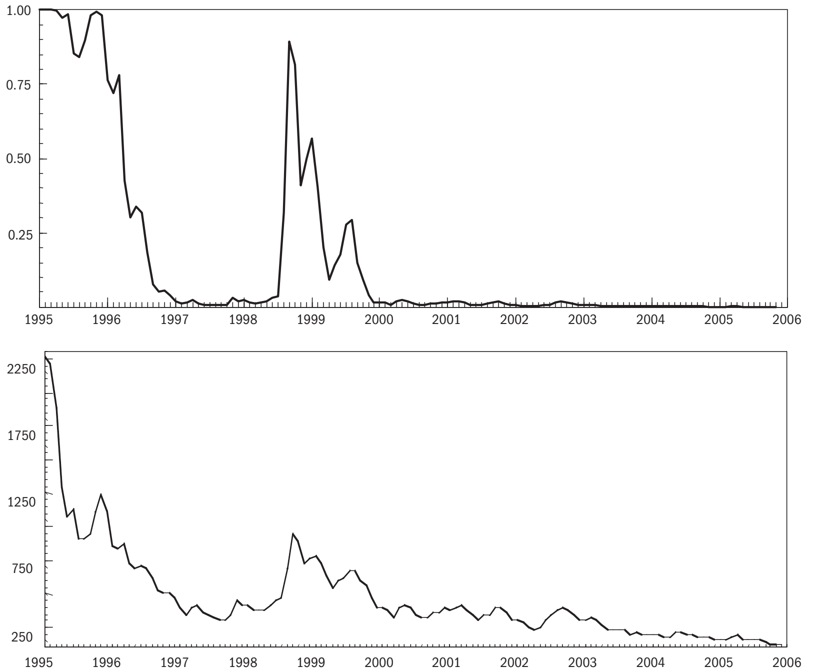


The estimated nonlinear models pass the diagnostic tests of no remaining nonlinearity and autocorrelation, and provide a good fit to the data. As expected there is a positive relationship between ERPT and our measures of macroeconomic instability, which can be verified by the fact that the sum of the nonlinear exchange rate coefficients is positive. Using these coefficients we computed the degree of ERPT over the long-run. 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 example Gagnon and Ihrig, 2004). Long-run ERPT is computed as:

$$LR = \sum_{i=0}^{n} \beta_{4,i} \Delta e_{t-i} + \left(\sum_{i=0}^{n} \beta^{*}_{4,i} \Delta e_{t-i} \right) .G(s_t;\gamma;c) \Big/ 1 - \sum_{i=1}^{n} \beta_{1,i} \pi_{t-i} \, . \tag{17}$$

Under both specifications, estimated long-run ERPT 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

Figure 2. Transition function and transition variable (EMBI + spreads)



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 (see for example Ca'Zorzi, Hahn and Sanchez 2007).

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 framework.

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.[9]

Consequently, if at the beginning of the 1990s Mexico's ERPT was much higher than that of most emerging economies, in the later years of our sample period the situation has changed completely. In this sense, our model suggests that the adoption of a sounder set of policies in Mexico may have had an important role in reducing ERPT, and hence in decreasing the costs of maintaining inflation stability after depreciations. In particular, the credibility gains from the adoption of Inflation Targeting may be responsible for some of the decline in ERPT. Similar results have been found for other emerging economies that adopted Inflation Targeting in the late 1990s, such as Brazil (see for example Nogueira Jr. and León-Ledesma, 2009). Although we do not want to imply that all the gain in terms of lower ERPT rates are due to better macroeconomic management, we believe that 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 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.

V. 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 the previous literature. We presented this argument in a simple mark-up model of import prices.

[9] We have also estimated a STR model (with an exponential transition function) using lagged exchange rate changes as a possible transition variable, in order to check nonlinearities with respect to the sizes of the exchange rate changes. The results found tell a very similar story about Mexico's ERPT: it is much higher than for most emerging economies, but has decreased in recent years. But the problem with this model is that most of the nonlinearity in ERPT is related solely to the exchange rate crisis in 1994-1995. The estimated threshold is 0.244 (i.e. 24% depreciation) and the speed of transition is so fast that the model almost converges to a threshold model. We believe that this exercise suggests that exchange rate volatility, at least in the case of Mexico, is a viable transition variable only in extreme cases, such as severe currency crises and massive capital flows. The results using exchange rate changes as a transition variable are available upon request from the authors.

Under bad economic conditions, firms have no incentive to absorb cost increases in their margins which thus leads to higher ERPT. From this model, we derived an empirical nonlinear model using smooth transition regressions. The model was then applied to Mexican data from 1992M1 to 2005M12.

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 United States). 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. Even though we do not believe that this is the only driver of ERPT in Mexico and other emerging countries, our results may indicate that the adoption of sounder policies in emerging markets, such as the introduction of Inflation Targeting regimes, may be an effective tool for reducing ERPT.

References

Ball, Christopher, and Javier Reyes (2004), Inflation targeting or fear of floating in disguise: The case of Mexico, *International Journal of Finance and Economics* **9:** 49-69.

Burstein, Ariel, Martin Eichenbaum, and Sergio Rebelo (2007), Modeling exchange rates pass through after large devaluations, *Journal of Monetary Economics* **54:** 346-368.

Ca'Zorzi, Michele, Elke Hahn, and Marcelo Sánchez (2007), Exchange rate pass-through in emerging markets, Working Paper 739, European Central Bank.

Choudhri, Ehsan, and Dalia Hakura (2006), Exchange rate pass-through to domestic prices: Does the inflationary environment matter?, *Journal of International Money and Finance* **25:** 614-639.

Christopoulos, Dimitris and Miguel León-Ledesma (2007), A long-run nonlinear approach to the Fisher effect, *Journal of Money, Credit and Banking* **39:** 543-559.

Clifton, Eric, Hyginus Leon, and Chorng-Huey Wong (2001), Inflation targeting and the unemployment-inflation trade-off, Working Paper 166, International Monetary Fund.

Gagnon, Joseph, and Jane Ihrig (2004), Monetary policy and exchange rate pass-through, *International Journal of Finance and Economics* **9:** 315-338.

Gil-Pareja, Salvador (2000), Exchange rates and European countries' export prices: An empirical test for asymmetries in pricing to market behaviour, *Weltwirtschatliches Archive* **136:** 1-23.

Herzberg, Valerie, George Kapetanios, and Simon Price (2003), Import prices and exchange rate pass-through: Theory and evidence from the United Kingdom, Working Paper 182, Bank of England.

Kaminsky, Graciela, Saul Lizondo, and Carmen Reinhart (1998), Leading indicators of currency crises, *IMF Staff Papers* **45**: 1-48.

Karhonen, Marko, and Juha-Pekka Juntilla (2010), Empirical evidence on the role of inflation regime in the exchange rate pass-through to import prices, Working Paper, SSRN (http://ssrn.com/abstract=1612103).

Mahdavi, Saeid (2002), The response of the US export prices to changes in the dollar's effective exchange rate: Further evidence from industrial level data, *Applied Economics* **34:** 2115-2125.

Marazzi, Mario, Nathan Sheets, Robert Vigfusson, Jon Faust, Joseph Gagnon, Jaime Marquez, Robert Martin, Trevor Reeve, and John Rogers (2005), Exchange rate pass-through to US import prices: Some new evidence, International Finance Discussion Paper 832, Board of Governors of the Federal Reserve System.

Mishkin, Frederic, and Miguel Savastano (2001), Monetary policy strategies for Latin America, *Journal of Development Economics* **66:** 415-444.

Nogueira Jr., Reginaldo, and Miguel León-Ledesma (2009), Fear of floating in Brazil: Did inflation targeting matter?, *The North American Journal of Economics and Finance* **20:** 255-266.

Taylor, John (2000), Low inflation, pass-through and the pricing power of firms, *European Economic Review* **44:** 1389-1408.

Terasvirta, Timo (2004), Smooth transition regression modelling, in H. Lutkepohl and M. Kratzig, eds., *Applied time series econometrics*, Cambridge, Cambridge University Press.

van Dijk, Dijk, Timo Terasvirta, and Philip Hans Franses (2002), Smooth transition autoregressive models – a survey of recent developments, *Econometrics Reviews* **21:** 1-47.