EXCHANGE RATE PASS-THROUGH INTO INFLATION: THE ROLE OF ASYMMETRIES AND NONLINEARITIES

Reginaldo P. Nogueira Júnior

Miguel A. León-Ledesma

Department of Economics, University of Kent

ABSTRACT

This paper investigates the empirical evidence on exchange rate pass through (ERPT) into CPI inflation for a set of emerging and developed countries. We argue that, theoretically, ERPT may be nonlinear in contrast to standard linear estimates in the literature. We use smooth transition models to investigate several possible sources of these nonlinearities. The results suggest that, although the sources of nonlinearities vary considerably across countries, they appear to be important. We find that for four countries ERPT responds nonlinearly to inflation and for three of them it responds nonlinearly to the output gap. We also find an asymmetric response of ERPT with respect to the magnitude of exchange rate changes for only two out of six countries. Finally, for some emerging markets, ERPT seems to be affected nonlinearly by measures of macroeconomic instability.

JEL CLASSIFICATION: E31, E52, F41.

KEYWORDS: Exchange rate pass-through, Smooth transition regression models.

Acknowledgements: we would like to thank, without implicating, Jagjit Chadha, John Driffill, Mathan Satchi, and seminar participants at the 2007 Money Macro and Finance Conference for helpful comments.

Address for correspondence: Dr. Miguel León-Ledesma, Department of Economics, University of Kent, Canterbury, Kent CT2 7NP, UK. Tel: 01227 823026

EXCHANGE RATE PASS-THROUGH INTO INFLATION: THE ROLE OF ASYMMETRIES AND NONLINEARITIES

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),[1] 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.

As observed by Marazzi et al. (2005) there is little work on the issue of whether ERPT is either nonlinear or asymmetric. In a brief survey, they argued that the existing literature provides mixed evidence for the view that there are important nonlinearities in ERPT. They analysed empirically data for the United States, and were unable to find any evidence of nonlinear ERPT. Other studies that failed to find evidence on this matter include Feinberg (1989), Athukorala (1991) and Herzberg et al (2003). On the other hand, Goldberg (1995), Gil-Pareja (2000), Mahdavi (2002), Bussiere (2006) and others have found different sorts of nonlinearities, most of them associated with asymmetric behaviour to appreciation and depreciations, and to large and small exchange rate changes. Pollard and Coughlin (2004) observed that previous studies of asymmetries in ERPT have concentrated almost entirely on testing for asymmetries in the *direction* and in the *size* of exchange rate changes, thus ignoring other possible sources of nonlinearities.

Various studies have shown that ERPT has declined in recent years, especially after the adoption of Inflation Targeting (IT). The most common interpretation for this finding is that of Taylor (2000), Gagnon and Ihrig (2004), Choudhri and Hakura

---

[1] See Goldberg and Knetter (1997) for an extensive survey on this literature.

(2006) and others, that relate the decrease of ERPT to a lower inflation environment.[2] According to this view the rate of inflation affects the persistence of costs changes, which is positively correlated with ERPT. Alternative explanations for the declining ERPT include those of Mishkin and Savastano (2001) and Schimidt-Hebbel and Werner (2002) that argue that this finding is a corollary of credibility gains of monetary policy.

The objective of this paper is to provide new empirical evidence on ERPT taking into account several potential forms and sources of nonlinearity. We use smooth transition regression (STR) models,[3] and test several potentially important transition variables in order to capture possible nonlinearities. In this sense, we complement the existing literature by investigating not only nonlinearities in ERPT related to the size and the direction of exchange rate changes, but also to output gap, the inflation level, and a measure of macroeconomic stability.[4] We use data on a set of six developed and emerging market economies that adopted IT during the 1990s. The reason for focusing on countries that adopted this monetary policy framework lies on the extensive literature that discusses the importance of ERPT for countries that seek a strict inflation target[5]. As discussed by Minella et al. (2003) and others, the extent to which prices respond to *transitory* exchange rate movements determines much of the central bank's ability to attain the targets in the short-run.[6]

The results present strong evidence in favour of nonlinearities in ERPT, although the sources vary considerably across countries. We found ERPT to respond

---

[2] Campa and Goldberg (2005) challenge this conclusion, arguing that the main reason for the decline in ERPT is related to changes in the composition of the import bundle.

[3] See van Dijk, Terasvirta and Franses (2002) for an overview on STR models.

[4] We assess whether two measures of risk (EMBI+ and real interest rates differentials to the United States) provide a good fit as transition variables in the STR model. Our hypothesis is that in periods of confidence crisis, and thus when both variables are increasing, ERPT is also increasing. To some extent this hypothesis encompasses the propositions of Mishkin and Savastano (2001) and others that argued that ERPT is influenced by the credibility levels of macroeconomic policies.

[5] For a discussion on this issue, see for example Agenor (2002).

[6] Ball (2000) suggests that the inflation targets should exclude transitory exchange rate changes. The problem with this procedure in practice is the difficulty in determining which exchange rate movements are transitory and which are permanent, plus the risk of misinterpretation by the market.

nonlinearly to inflation for 4 out of the six countries analysed, namely South Africa, Mexico, United Kingdom and Canada. ERPT was found to respond nonlinearly to the size of exchange rate changes for Mexico and the United Kingdom, and to the output gap for South Africa, Mexico and Czech Republic. Finally, when considering the measures of risk (or macroeconomic *instability*) ERPT seemed to respond nonlinearly to EMBI+ spread for Brazil and Mexico, and to real interest rate differentials to Mexico and the Czech Republic. This last finding demonstrates the importance of the market's confidence in a stable macroeconomic environment to slowdown price-spirals and hence to reduce ERPT.

The remainder of the paper is structured as follows. Section 2 briefly discusses the literature on nonlinearities and asymmetries in ERPT. Section 3 presents the theoretical background and our empirical methodology. Section 4 presents the results. Section 5 concludes.

2. LITERATURE REVIEW

As mentioned earlier, evidence on nonlinear or asymmetric behaviour with respect to ERPT is mixed. Marazzi et al. (2005) observed in a brief survey that the literature has found no clear support for this matter. In an empirical examination, they investigated nonlinearities in ERPT for the United States with respect to the size of exchange rate changes, and were unable to find conclusive results.

Herzberg et al (2003) also failed to find evidence in this direction. They tried to capture asymmetric and nonlinear responses of ERPT for the United Kingdom with respect to the direction and the size of exchange changes by using three specifications (a threshold model, a spline model and a logistic STR model), but none provided evidence in favour of the hypothesis. In studies that considered a range of industries, Feinberg (1989) and Athukorala (1991) were also unable to find evidence of asymmetry with respect to exchange rate changes respectively for the United States and South Korea. Olivei (2002) checked whether ERPT of 34 industry groups in the

United States were symmetric for appreciations and depreciations, and found only weak evidence of asymmetric behaviour.[7]

On the other hand, Mahdavi (2002) looked at a range of exporting industries in the United States, and presented evidence of asymmetric response to exchange rate changes in seven out of 12 industries. Bussiere (2006) studied the G7 countries, and focused on nonlinearities in the reaction of ERPT to the size of exchange rate changes. His results also suggest that nonlinearities cannot be neglected, although their magnitude varies considerably across countries. Webber (2000) found strong support for asymmetric ERPT for five out of seven Asian countries, and observed that ERPT is higher under depreciations than appreciations of the domestic currency. Kadiyali (1997) and Goldberg (1995) focused on single industries, and also found a higher ERPT following depreciations[8].

In much of the literature the direction of the asymmetry was found to vary considerably across industries and countries, therefore not providing strong support to the view that "*prices rise faster than they fall*" (Peltzman, 2000). Gil-Pareja (2000), for example, examined the differences in ERPT in a range of industries in a sample of European economies, and found that the degree and direction of asymmetry varied across industries and countries. Similar results were found by Pollard and Coughlin (2004) that looked at 30 manufacturing industries in the United States, and found ERPT to be asymmetric in half of them.

Ohno (1989) found some evidence that changes in Japanese export prices were more frequent with large exchange rate changes than with small ones. Pollard and Coughlin (2004) have also looked at this issue, and found that for 19 of the 30 industries analysed, ERPT was statistically greater following large exchange rate changes[9].

---

[7] Only nine out of 34 industries included in Olivei's (2002) study exhibited some degree of asymmetry.

[8] Kadiyali (1997) looked at United States' import prices of photographic film from Japan, whereas Goldberg (1995) examined United Sates´ import prices of automobiles from Japan and Germany.

[9] Pollard and Coughlin (2004) used dummy variables to distinguish between large and small exchange rate changes; the latter defined as a change smaller than 3%.

These studies have focused exclusively on asymmetries with respect to the direction or magnitude of exchange rate shocks. Although the evidence is mixed with this respect, we cannot rule-out other sources of nonlinearities. Asymmetric adjustment may also be dependent on other factors emphasized by economic theory, such as the inflation environment, demand pressures, and credibility of monetary policy. All these factors may affect the degree to which exchange rate shocks are passed on to prices, constituting a potential source of nonlinearity.

3. THE NONLINEAR APPROACH TO ERPT

3.1 THEORETICAL BACKGROUND

A simple theoretical model helps illustrating the potential reasons for a nonlinear ERPT that depends on the macroeconomic environment. The model we present here is very parsimonious but it suffices to illustrate the argument.

Let us consider a foreign firm that exports its product to the domestic country. Allowing time variation, under perfect competition, a profit maximizing exporter with prices set in importing country currency will set its price at time *t* equal to the marginal cost:

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

Where P is the local currency price, C^* is the exporter's marginal cost expressed in its own currency, and E is the domestic exchange rate. Relaxing the assumption of perfect competition, the exporter's profit maximization condition includes also a mark-up θ over marginal cost:

$$P_t = \theta_t E_t C_t^* \tag{2}$$

Note that firm's marginal cost and mark-up may change independently of the exchange rate. For example, a change in the cost of a locally provided input (in the foreign country) can shift marginal cost. Also, demand shocks in the importing country can alter the exporter's mark-up. In this sense, following Bailliu and Fujii (2004), Goldberg and Campa (2005) and others, we assume the mark-up to respond to demand pressures in the importing country. In addition to demand pressures, we also assume the mark-up to depend nonlinearly on the importing country's general macroeconomic stability, so that when inflation is high, or when the economy faces a confidence crisis, ERPT is higher. We view this macroeconomic stability dependence as a firm's strategic decision on how much to pass-through exchange rate changes into prices given different macroeconomic scenarios in the importing country. Hence, the mark-up has the following functional form:

$$\theta_t = \theta\left(\rho, E^{\omega(Z)}\right) \tag{3}$$

Where ρ accounts for the demand pressures in the importing country (proxied by the aggregate output, y). As mentioned above, we also allow the mark-up to depend nonlinearly on the importing country macroeconomic stability, which is represented by a component Z. We model Z in such a way that high values imply either high inflation levels or low credibility levels and hence a bad macroeconomic environment. In other words, Z would be actually a measure of *macroeconomic instability.*[10] In this sense, the function $\omega(Z)$ can be seen as a mark-up multiplier in the following way: as firms set prices for several periods in advance, mark-ups respond more to exchange rate changes if inflation is high, and if the market's

---

[10] As general macroeconomic stability we basically refer to a low inflation environment and high credibility levels. In our empirical analysis of the model we test the inflation level and some measures of market's confidence in the economy (basically EMBI+ spreads and real interest rates differentials to the United States) as proxies for Z. The hypothesis we want to test is if when inflation is high, or when the economy faces a confidence crisis, ERPT is higher.

confidence in the economy is low. Therefore, a high inflation environment, or the advent of a confidence crisis, would tend to increase ERPT.

Consequently, a simple reduced form equation in logarithms would be[11]:

$$p_t = \alpha e_t + \omega(Z) e_t + \beta c_t^* + \kappa y_t \tag{4}$$

Where lower cases denote logarithms. Equation (4) states that there are two channels of ERPT. The first channel, which we call *direct* ERPT, is given by α and is expected to be bounded between 0 and 1. If $\alpha = 1$, the direct ERPT is complete. If $\alpha = 0$, the direct ERPT is zero. Regarding the function $\omega(Z)$, it shows an indirect channel of ERPT, which depends on the macroeconomic environment. We will further assume that there is some threshold Z^* which divides the extreme cases of good (low inflation and high credibility, translated into low values of Z) and bad (high inflation and low credibility, translated into high values of Z) macroeconomic environments.

$$\omega(Z) = \begin{cases} 0; Z \leq Z^* \\ \psi > 0; Z > Z^* \end{cases} \tag{5}$$

For these two extreme cases we find two different ERPT. If the importing country has a low inflation regime and/or a credible set of macroeconomic policy, then ERPT is equal to α. If the importing country has a high inflation regime and/or lacks credibility, then ERPT is equal to $\alpha + \psi$. We can see that ERPT is higher in the second case, as $\alpha + \psi > \alpha$. Intuitively, with low inflation and a stable macroeconomic environment, firms face competition in importing markets and cannot pass-through all exchange rate changes into prices. Hence, our model implies that higher inflation and low credibility would raise ERPT in a nonlinear way.

---

[11] Equation (4) is similar to the one developed by Goldberg and Knetter (1997), Bailliu and Fujii (2004) and Goldberg and Campa (2005), with the addition of the term $\omega(Z)e_t$.

Rewriting (4) in difference form, we have:

$$\Delta p_t = \beta \Delta c_t^* + \kappa \Delta y_t + [\alpha + \omega(Z)]\Delta e_t \tag{6}$$

The above threshold model may be plausible for one firm, but for the aggregate of firms' case we should smooth the nonlinear function (Korhonen, 2005). A potential source of smoothness may be due to interaction of heterogeneous agents at the microeconomic level. There is probably a great diversity across firms when forming opinions of the macroeconomic environment. As *Z* grows above the threshold a greater number of firms take this as an evidence of changes in the importing country macroeconomic stability, and change their pricing behaviour. Following this, we will make use of smooth transition models instead of threshold models in our empirical application.

Although the model presented above was made for import prices, we want to analyse ERPT into consumer prices in our empirical investigation, as this is the most important variable for policymakers under IT regimes.[12] In this sense we extend the model, taking as starting point the composition of the consumer price index (CPI):

$$P_{CPI} = P_H^{\phi} P_T^{1-\phi}. \tag{7}$$

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 (7) the CPI inflation equation for the economy is:

$$\pi = \phi \pi_H + (1-\phi)\pi_T. \tag{8}$$

---

[12] The inflation targets are normally expressed in terms of consumer price inflation.

Following the literature on inflation and the importance given to inertial behaviour [see, for instance, McAdam and Wilman, 2004], and assuming the same (one) period lag for both tradable and non-tradable sectors, we have:

$$\pi_{(H)t} = \delta\pi_{(H)t-1} + \varphi\Delta y_t \tag{9}$$

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

Equation (9) states that home inflation depends on the output gap and past inflation. Equation (10) shows the tradable sector prices, basically following equation (6) but allowing for price inertia. Substituting (9) and (10) into (8), yields:

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

Finally, rearranging equation (11), we have:

$$\pi_t = \delta\pi_{t-1} + [(1-\phi)\kappa + \phi\varphi]\Delta y_t + (1-\phi)\beta\Delta c_t^* + (1-\phi)[\alpha + \omega(Z)]\Delta e_t \tag{12}$$

Equation (12) shows the basic model for estimating ERPT for CPI inflation, and can be described as a *nonlinear backward-looking Phillips curve*. In the next subsection we develop this model into a proper econometric specification.

In our empirical investigation we also examine the possibility that ERPT is higher for large exchange rate changes than for small ones, as in Pollard and Coughlin (2004), Marazzi et al. (2005) and Ohno (1989). Intuitively, exporting firms may have a desired price denominated in their domestic currencies. In response to small exchange rate changes, firms are willing to hold their prices in importing country's currency to maintain market share, or maybe to avoid menu costs. Nevertheless, firms do change their prices in response to large exchange rate movements, as to prevent the selling price from getting to far away from their desired price. Finally, we also check possible nonlinearities with respect to the output gap. As argued by Goldfajn and Werlang (2000) and Correa and Minella (2006), periods where the economy is overheated may lead to stronger incentives for firms to pass-through cost increases, such as those coming from the exchange rate. In other words, increasing sales firms may find it easier to raise prices than firms that face an economy recession.

3.2 EMPIRICAL SPECIFICATION

Smooth transition regression (STR) models are a general class of nonlinear time-series models that can account for deterministic changes in parameters over time, in conjunction with regime switching behaviour. An STR model can be roughly described as the weighted average of two linear models, with weights determined by the value of the transition function. The STR model takes the following general form:

$$y_t = \beta_1 x_t + \beta_2 x_t . G(s_{t-i}, \gamma, c) + \upsilon_t \tag{13}$$

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 or location parameter. 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 and the function *G* becomes a dummy-type indicator.

A popular choice for the transition function is the logistic smooth transition (LSTR) that is given by:

$$G(s_{t-i}, \gamma, c) = \left[\left(1 + \exp\left\{ -\gamma (s_{t-i} - c) \right\} \right)^{-1} \right] \tag{14}$$

An alternative specification to the transition function is the exponential smooth transition (ESTR):

$$G(s_{t-i},\gamma,c)=\left(1-\exp\left\{-\gamma(s_{t-i}-c)^2\right\}\right) \tag{15}$$

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. In this sense, 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 $\beta_1 + \beta_1 / 2$. In the case of the ESTR model, the coefficient changes depending on whether the transition variable is close or far away from the threshold, regardless of whether this difference is positive or negative. In other words, what matters is if the shocks are large or small. Following this, we have that as $(s_t - c) \rightarrow \pm\infty$ then the coefficient on inflation becomes $\beta_1 + \beta_2$; and if $s_t = c$ it becomes β_1.

The interpretation of transition functions is crucial in understanding what the estimates mean. A priori we expect the ESTR model only to work well when using exchange rate changes as transition variable. The reason for this is that the literature highlights the role of "*menu costs*" in explaining nonlinearities in ERPT with respect to exchange rate changes. The argument is that firms are more likely to adjust their prices after large shocks than after small shocks. An ESTR model would capture exactly this kind of behaviour. On the other hand, we expect that an LSTR model would be more appropriate for the other transition variables (inflation, output and credibility measures), as it would follow the same pattern as the threshold model described in the theoretical model but assuming a smooth adjustment.

We followed the modelling approach described in Lundbergh et al. (2000), van Dijk, Terasvirta and Franses (2002) and Terasvirta (1994) in estimating the STR models. The modelling procedure consisted of the following steps: first, we tested the null of linearity of a baseline model against the smooth transition alternative, if the null was not rejected, we accepted the linear model; if the null was rejected, we estimated the nonlinear model; then, we evaluated the estimated model for general modelling misspecification, if the model failed these tests, an extended model was estimated and evaluated.

As proposed by Terasvirta (1998) and others, linearity was tested by means of LM-type tests, with the null hypothesis of linearity against STR nonlinearity. Following 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. We used two tests: an LM_2 with the null of linearity against ESTR nonlinearity, and an LM_3 with the null of linearity against LSTR nonlinearity. For a technical discussion of the tests the reader is referred to van Dijk, Terasvirta and Franses (2002).

In choosing the transition variable to be included in the final nonlinear model, we have also tested for remaining nonlinearity after estimation. We have chosen the transition variables that provided the strongest rejection of both the null of linearity of the baseline linear model, and of remaining linearity after estimation of the nonlinear model. In addition to this, we have selected models with well-defined transition functions i.e. that did not seem to depend heavily on outliers and had enough observations in the different states. Finally, we also gave preference for models that provided the highest $AdjR^2$ and the lowest Akaike Information Criteria (AIC).

The decision between an LSTR and an ESTR model was made, as proposed by Terasvirta (2001), at the evaluation stage. We initially selected the transition variable following the results of linearity tests, and then estimated the model for both LSTR and ESTR. We selected the transition function that provided the best fit to the data.

Estimation of the parameters in the LSTR model is, as discussed by van Dijk, Terasvirta and Franses (2002), a straightforward application of nonlinear least squares (NLS)[13]. Under the assumption that the errors are normally distributed, NLS is equivalent to maximum likelihood. Otherwise, the NLS estimates can be interpreted as quasi maximum likelihood estimates[14].

We use a similar type of model as in Campa and Goldberg (2005), Choudhri and Hakura (2006) and Gagnon and Ihrig (2004). However, we differ from these studies in an important way, since we assume that ERPT is nonlinear. Hence, we consider either an LSTR or an ESTR model. These functional forms take into account the pricing behaviour of exporting firms at the aggregate level, as discussed before. The estimated model has the following form:

$$\pi_t = \beta_0 + \sum_{i=1}^{n} \beta_{1,i}\pi_{t-i} + \sum_{i=0}^{n} \beta_{2,i}\Delta p_{t-i}^{imp} + \sum_{i=0}^{n} \beta_{3,i}\Delta y_{t-i} + \sum_{i=0}^{n} \beta_{4,i}\Delta e_{t-i} + \left(\beta_0^* + \sum_{i=0}^{n} \beta_{4,i}^* \Delta e_{t-i} \right) G(s_t, \gamma, c) + \varepsilon_t \qquad (16)$$

Where π is the inflation rate, Δp^{imp} is the change in import prices (in foreign currency), Δy is output growth,[15] Δe is the exchange rate change, $G(s_t, \gamma, c)$ is the nonlinear function described above and ε is an error term. When testing the model empirically, a time trend was included for some countries when it proved to be

---

[13] van Dijk, Terasvirta and Franses (2002) observe that it is quite difficult to obtain an accurate estimate of the smoothness of the transition between regimes. The estimate of γ may therefore appear to be insignificant by its t-statistic. This should not be interpreted as evidence of weak nonlinearity.

[14] 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. The values for the grid search for γ were set between 0 and 100 for increments of 1, whereas c was estimated for all the ranked values of the transition variable. The values selected are those that minimised the residual sum of squares. These results are then fed into the NLS procedure as starting values. This procedure increases the precision of the estimates and ensures faster convergence of the NLS algorithm.

[15] 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 also estimate the model using an HP-Filtered and a Band-Pass filtered output gap. The results were similar, and we opted to show here just those obtained using output growth for reasons of space.

significant. This is a standard procedure in nonlinear modelling, as for example in Clifton, Leon and Wong (2001).[16]

There are two basic outcomes for ERPT. In the first one, the transition variable is far below from the threshold for the LSTR model or near the threshold for the ESTR model. In this case ERPT is given just by the linear parameters $\sum_{i=1}^{n} \beta_{4,i}$. In the second one, the transition variable is far away from the threshold for the ESTR model, or far beyond it for the LSTR model. In this case the coefficient is the sum of the linear and nonlinear parts of the model, $\sum_{i=1}^{n} \beta_{4,i} + \sum_{i=1}^{n} \beta_{4,i}^{*}$. For the LSTR specification there is a third possible outcome: when the transition variable is equal to the threshold the ERPT is given by $\sum_{i=1}^{n} \beta_{4,i} + \sum_{i=1}^{n} \beta_{4,i}^{*} / 2$.

Monthly data was collected for 6 countries that adopted IT that may be split into two groups: the first one comprises developed economies (Canada and United Kingdom), and the second one is composed of emerging market economies (Brazil, Czech Republic, Mexico and South Africa). The period of estimation corresponds to the interval that spans from 1983M1 to 2005M12 for the developed economies, and 1992M1 to 2005M12 for the emerging markets. For Brazil the period is 1995M7 to 2005M12, and for the Czech Republic is from 1994M1 to 2005M12.[17]

Data was obtained from the IMF 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.[18] Data on import prices is the change in the series of unit value of imports (in dollars).

---

[16] For some countries there is a clear downward trend in inflation, given the adoption of IT and the strong convergence of inflation to the targets. Clifton, Leon and Wong (2001) argue that when such clear trends are present they should be included in the model.

[17] The shorter period for the emerging market economies is due to data availability. In the case of Brazil we have opted to exclude hyperinflation data, thus starting our analysis in 1995M7.

[18] For South Africa the Manufacturing Production Index was used.

This data was not available for South Africa, Mexico and Czech Republic, so the International Commodities Price Index was used as a proxy of the foreign costs faced by the importers. All the data is in log-differences[19].

4. RESULTS

4.1 TRANSITION VARIABLE: INFLATION

Taylor (2000) argued that the decline in ERPT observed in the literature over the 1990s is a by-product of lower inflation, brought about by the change in monetary policy regime with the adoption of IT. In this section we investigate whether the inflation level affects the degree of ERPT nonlinearly in specification (16).

Table 1 presents the p-values of the linearity tests with up to three periods lagged inflation. As discussed before remaining nonlinearity tests were also applied to the model to help select the transition variable.[20] A low value indicates rejection of the null of linearity. As discussed in the methodology, the LM_2 tests linearity against ESTR nonlinearity and the LM_3 tests linearity against LSTR nonlinearity. In summary, we found inflation to be an appropriate transition variable for South Africa, Mexico, Canada and the United Kingdom. After careful evaluation of different specifications an LSTR model was selected for all of them. This is in line with theoretical priors that ERPT may depend on the level of inflation.

We present for each of the cited countries the baseline linear model, and the nonlinear LSTR model with the selected transition variable. In the appendix of the paper we also provide graphs of the transition variables, the transition function over time, and the estimation residuals.

---

[19] Standard unit root tests were not able to reject the null of non-stationarity for the (log) levels, but rejected the null for the differences of the variables. We applied some cointegration tests and found no evidence of cointegration. Therefore we opted to estimate the model in (log) differences. The results of the unit root and cointegration tests are not shown here for reasons of space, but are available on request.

[20] The results of the remaining nonlinearity tests are available on request. We opted to present here just the results of the test for the selected transition variables.

Regarding the results, *Sigma* is the standard error of the regression, *AIC* is the Akaike Information Criterion, *AR* is an LM test of autocorrelation, *Hetero* is a test of heteroskedasticy. *Q(i)* is a test of autocorrelation with *i* number of lags, and *RNL* is an LM test with the null of no remaining nonlinearity; this test is either a test against LSTR or ESTR, depending on the selected transition function. Note that ** indicates significance at the 5% level and * indicates significance at the 10% level.

The results for South Africa are:

Linear Model:

$$\pi_t = 0.002^{**} + 1.229^{**}\pi_{t-1} - 0.196^{*}\pi_{t-2} - 0.090\pi_{t-3} - 0.025^{*}\Delta y_t + 0.050^{**}\Delta p_{t-1}^{imp}$$
$$-0.032^{*}\Delta p_{t-2}^{imp} + 0.038^{**}\Delta e_{t-1} - 0.034^{**}\Delta e_{t-3} + 0.016^{**}\Delta e_{t-4} + \varepsilon_t$$

Radj^2= 0.974, Sigma= 0.0053, AIC= -10.439, AR= 0.985, ARCH= 0.911, Hetero= 0.924.

Nonlinear Model:

$$\pi_t = 0.001 + 1.177^{**}\pi_{t-1} - 0.185\pi_{t-2} - 0.036\pi_{t-3} - 0.015\Delta y_t + 0.046^{**}\Delta p_{t-1}^{imp} - 0.031^{*}\Delta p_{t-2}^{imp} + 0.051^{**}\Delta e_{t-1}$$
$$-0.020\Delta e_{t-3} + 0.015\Delta e_{t-4} + \left[-0.001 - 0.019\Delta e_{t-1} - 0.028\Delta e_{t-3} + 0.057^{**}\Delta e_{t-4}\right].G(\pi_{t-2},\gamma,c) + \nu_t$$

LSTR: $G(\pi_{t-2},\gamma,c) = \left(1 + \exp\left\{-99(\pi_{t-2} - 0.071^{**})\right\}^{-1}\right)$

Radj^2= 0.978, Sigma= 0.0051, AIC= -10.456, Q(2)= 0.302, Q(4)= 0.563, Q(6)=0.705, RNL= 0.579.

The results for Mexico are:

Linear Model:

$$\pi_t = 0.001 + 1.506^{**}\pi_{t-1} - 0.613^{**}\pi_{t-2} + 0.087\pi_{t-3} - 0.010\Delta y_t - 0.0004\Delta p_t^{imp}$$
$$+0.039^{**}\Delta e_t + 0.016\Delta e_{t-1} - 0.021^{**}\Delta e_{t-2} + 0.042^{**}\Delta e_{t-3} - 0.056^{**}\Delta e_{t-4} + \varepsilon_t$$

Radj^2= 0.998, Sigma= 0.0041, AIC= -10.946, AR= 0.265, ARCH= 0.736, Hetero= 0.080.

Nonlinear Model:

$$\begin{aligned}\pi_t &= 0.002^{**} + 1.296^{**}\pi_{t-1} - 0.513^{**}\pi_{t-2} + 0.135^{**}\pi_{t-3} - 0.001\Delta y_t + \\ &0.008^{**}\Delta p_t^{imp} + 0.035^{**}\Delta e_t + 0.012\Delta e_{t-1} - 0.049^{*}\Delta e_{t-2} + 0.002\Delta e_{t-3} + 0.005\Delta e_{t-4} + \\ &\left[0.007^{**} - 0.001\Delta e_t + 0.018\Delta e_{t-1} + 0.039\Delta e_{t-2} + 0.047^{*}\Delta e_{t-3} - 0.062^{**}\Delta e_{t-4}\right].G(\pi_{t-1},\gamma,c) + \nu_t\end{aligned}$$

LSTR: $G(\pi_{t-1},\gamma,c) = \left(1 + \exp\left\{-5^{**}(\pi_{t-1} - 0.098^{**})\right\}^{-1}\right)$

Radj^2= 0.999, Sigma= 0.0036, AIC= -11.136, Q(2)= 0.273, Q(4)= 0.058, Q(6)=0.156, RNL= 0.226.

The results for the United Kingdom are:

Linear Model:

$$\begin{aligned}\pi_t &= 0.001^{*} + 1.241^{**}\pi_{t-1} - 0.240^{**}\pi_{t-2} + 0.157^{**}\pi_{t-5} - 0.184^{**}\pi_{t-6} + 0.036^{**}\Delta y_t - 0.023^{*}\Delta y_{t-1} \\ &+0.047^{**}\Delta p_t^{imp} - 0.044^{**}\Delta p_{t-1}^{imp} + 0.029^{*}\Delta e_t - 0.006\Delta e_{t-1} - 0.035^{**}\Delta e_{t-2} + 0.017^{**}\Delta e_{t-3} + \varepsilon_t\end{aligned}$$

Radj^2= 0.976, Sigma= 0.0030, AIC= -11.559, AR= 0.563, ARCH= 0.575, Hetero= 0.020.

Nonlinear Model:

$$\begin{aligned}\pi_t &= 0.001^{*} + 1.213^{**}\pi_{t-1} - 0.223^{**}\pi_{t-2} + 0.165^{**}\pi_{t-5} - 0.190^{**}\pi_{t-6} + 0.033^{**}\Delta y_t - 0.018\Delta y_{t-1} \\ &+0.033^{*}\Delta p_t^{imp} - 0.030^{*}\Delta p_{t-1}^{imp} + 0.023\Delta e_t - 0.016\Delta e_{t-1} - 0.024^{*}\Delta e_{t-2} + 0.015^{*}\Delta e_{t-3} + \\ &\left[0.0004 - 0.011\Delta e_t + 0.041^{**}\Delta e_{t-1} - 0.020\Delta e_{t-2} + 0.001\Delta e_{t-3}\right].G(\pi_{t-3},\gamma,c) + \nu_t\end{aligned}$$

LSTR: $G(\pi_{t-3},\gamma,c) = \left(1 + \exp\left\{-99(\pi_{t-3} - 0.036^{**})\right\}^{-1}\right)$

Radj^2= 0.979, Sigma= 0.0029, AIC= -11.568, Q(2)= 0.960, Q(4)= 0.979, Q(6)=0.982, RNL= 0.335.

The results for Canada are:

Linear Model:

$$\begin{aligned}\pi_t &= 0.001 + 1.030^{**}\pi_{t-1} - 0.153^{*}\pi_{t-2} + 0.074\pi_{t-3} - 0.018\Delta y_{t-4} + 0.020\Delta y_{t-5} + 0.042^{**}\Delta p_t^{imp} \\ &-0.021^{**}\Delta p_{t-1}^{imp} + 0.051^{**}\Delta e_{t-1} - 0.039^{**}\Delta e_{t-2} + \varepsilon_t\end{aligned}$$

Radj^2= 0.940, Sigma= 0.0039, AIC= -11.052, AR= 0.477, ARCH= 0.907, Hetero= 0.905.

Nonlinear Model:

$$\pi_t = 0.004^{**} + 1.005^{**}\pi_{t-1} - 0.091\pi_{t-2} + 0.053\pi_{t-3} - 0.017\Delta y_{t-4} + 0.019\Delta y_{t-5} + 0.045^{**}\Delta p_t^{imp} - 0.018^{*}\Delta p_{t-6}^{imp}$$
$$+0.070^{*}\Delta e_{t-1} - 0.093^{**}\Delta e_{t-2} + \left[0.004^{**} - 0.028\Delta e_{t-1} + 0.065^{*}\Delta e_{t-2}\right].G(\pi_{t-2},\gamma,c) + \nu_t$$

LSTR: $G(\pi_{t-2},\gamma,c) = \left(1+\exp\left\{-99(\pi_{t-2}-0.011)\right\}^{-1}\right)$

Radj^2= 0.945, Sigma= 0.0038, AIC= -11.065, Q(2)= 0.997, Q(4)= 0.256, Q(6)=0.498, RNL= 0.391.

In comparison to the linear model the standard error of the regressions is smaller, the Radj^2 is higher and the AIC is lower in the nonlinear specification. The model also passes the diagnostic tests, rejecting the presence of serial correlation of the residuals, and of remaining nonlinearity in the model. These results reinforce the view that ERPT responds nonlinearly to the inflation level for these countries.

The threshold inflation level (annual rate) varies considerably across countries. For South Africa it is around 7%; for Mexico it is approximately 10%; for the United Kingdom it is 3.6%; and for Canada it is just over 1%. For all these countries there is a positive relationship between ERPT and the inflation level, observed by the fact that the sum of the nonlinear exchange rate terms is positive. In this sense, when inflation increases above the threshold, ERPT also increases.

Table 2 shows the sum of the exchange rate coefficients when the transition function (*G*) is equal to 0, ½ and 1. For Mexico, Canada and South Africa there is a wide difference between the two states, whereas for the UK the response of ERPT to the inflation is considerably weaker.

The analysis of the graphs of the transition functions over time provides some interesting insights on the evolution of ERPT in these countries. The adoption of IT was followed by lower inflation in all four economies, and hence is translated by lower ERPT in our model. In the graphs it is possible to observe that periods where the transition function is close to 1 are more common before the change in the monetary policy framework[21]. This finding sheds some light on Taylor's (2000) hypothesis that the lower inflation of the 1990s, brought about by a change in central

---

[21] The dates of adoption of IT by each country in our sample can be seen in the Appendix.

bank's approach towards inflation, was responsible for a decline in ERPT. From our results this proposition seems plausible for the four countries analysed above.

4.2 TRANSITION VARIABLE: RIDs

Mishkin and Savastano (2001), Leiderman and Bar-Or (2000), Schimidt-Hebbel and Tapias (2002) and others have argued that credibility levels of monetary policy influence the degree of ERPT. In our theoretical model we discussed the possibility that this argument could be expanded in order to refer to general market credibility towards the macroeconomic policy. 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 our empirical investigation we used two potential indicators of macroeconomic *instability*: real interest rate differentials[22] ($Rids_t$) with the United States[23] and EMBI+ spreads. We expect that both variables would provide some proxy of the risks perceived by the market with respect to the economy under consideration; when these variables increases, ERPT would thus increase as well. The main difference between them is that EMBI+ spreads are calculated from dollar-denominated bonds, thus excluding exchange rate risk. This variable, however, was not available for all countries in our sample.

Table 3 shows the linearity tests using up to three lags of Rids as transition variable. We found evidence of nonlinear response of ERPT with respect to Rids for Mexico and Czech Republic. We show the estimation results of the baseline linear model, and of the fitted nonlinear model. An LSTR transition function provided the fest fit for both countries.

---

[22] The use of $Rids_t$ 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).

[23] To construct this variable we used data on money market interest rates for each country and for the United States. CPI inflation was used to obtain the real interest rates from the nominal interest rates collected. The data came from the IMF/IFS database.

The results for Mexico are:

Linear Model:

$$\pi_t = 0.001 + 1.506^{**}\pi_{t-1} - 0.613^{**}\pi_{t-2} + 0.087\pi_{t-3} - 0.010\Delta y_t - 0.0004\Delta p_t^{imp}$$
$$+0.039^{**}\Delta e_t + 0.016\Delta e_{t-1} - 0.021^{**}\Delta e_{t-2} + 0.042^{**}\Delta e_{t-3} - 0.056^{**}\Delta e_{t-4} + \varepsilon_t$$

Radj^2= 0.998, Sigma= 0.0041, AIC= -10.946, AR= 0.265, ARCH= 0.736, Hetero= 0.080.

Nonlinear Model:

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

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

Radj^2= 0.999, Sigma= 0.0036, AIC= -11.174, Q(2)= 0.734, Q(4)= 0.503, Q(6)=0.677, RNL= 0.152.

The results for the Czech Republic are:

Linear Model:

$$\pi_t = 0.006^{**} - 0.0001^{**}T + 0.927^{**}\pi_{t-1} + 0.046^{**}\Delta y_t + 0.009^{**}\Delta p_t^{imp} + 0.019^{**}\Delta e_t + \varepsilon_t$$

Radj^2= 0.979, Sigma= 0.0052, AIC= -10.492, AR= 0.141, ARCH= 0.887, Hetero= 0.004.

Nonlinear Model:

$$\pi_t = 0.010^{**} - 0.0001^{**}T + 0.924^{**}\pi_{t-1} + 0.043^{**}\Delta y_t + 0.001\Delta p_t^{imp}$$
$$-0.004\Delta e_t + \left[-0.087 + 0.001T + 0.289^{**}\Delta e_t\right].G(rid_{t-1},\gamma,c) + v_t$$

LSTR: $G(rid_{t-1},\gamma,c) = \left(1 + \exp\left\{-1(rid_{t-1} - 8.672^{**})\right\}^{-1}\right)$

Radj^2= 0.984, Sigma= 0.0046, AIC= -10.664, Q(2)= 0.724, Q(4)= 0.437, Q(6)=0.423, RNL= 0.499.

The nonlinear model also passes all the diagnostic tests, and provides smaller standard errors of the regression, higher Radj^2 and lower AIC. As observed before, it reinforces the quality of the nonlinear specification in comparison to the linear one. As expected there is a positive relationship between ERPT and Rids. The sum of the exchange rate coefficients under different states can be seen in Table 4. In the case of the Czech Republic the effect of Rids on ERPT is very strong, suggesting an important role for credibility. For Mexico we also capture a substantial response of ERPT to Rids, although it is weaker than the one observed with respect to inflation in the previous subsection. The threshold value for both countries is quite high, being around 8.7% for the Czech Republic, and 6.9% for Mexico.

The speed of transition is rather different for each country, being much faster for Mexico. In this sense the transition between regimes is much smoother for the Czech Republic, as can be observed in the graphs provided in the Appendix. Nevertheless it is possible to observe a spike in the transition function for the Czech Republic in 1997, around the confidence crisis that preceded the adoption of IT. Regarding Mexico, the transition function is higher mostly in the period of the troubled dollar-pegged regime that collapsed in 1995.

4.3 TRANSITION VARIABLE: EMBI+

As discussed before we use the *Emerging Markets Bond Index Plus* (EMBI+) spreads as a possible measure of macroeconomic instability. The EMBI+ spreads track total returns for traded external debt instruments in the emerging markets. The index covers mainly dollar-denominated *Brady* bonds, but for some countries it also includes loans and Eurobonds. Nonetheless, this data was available only from 1995M1 on, and just for Brazil and Mexico[24]. Table 5 presents the results of the

---

[24] Data on EMBI+ spreads was obtained from the website http://www.cbonds.info. Data used in this paper refers to the last day of the month.

linearity tests. The results show that the EMBI+ spread seems to be a possible transition variable for both countries. Again an LSTR specification was selected.

The results for Brazil are:

Linear Model:

$$\pi_t = 0.0004 + 1.560^{**}\pi_{t-1} - 0.587^{**}\pi_{t-2} + 0.012\Delta y_t + 0.003\Delta p_t^{imp}$$
$$+0.041^{**}\Delta e_{t-1} - 0.043^{**}\Delta e_{t-2} + 0.012^{*}\Delta e_{t-3} + \varepsilon_t$$

Radj^2= 0.991, Sigma= 0.0042, AIC= -10.905, AR= 0.670, ARCH= 0.366, Hetero= 0.156.

Nonlinear Model:

$$\pi_t = 0.001 + 1.454^{**}\pi_{t-1} - 0.488^{**}\pi_{t-2} + 0.006\Delta y_t + 0.003\Delta p_t^{IMP} + 0.031^{**}\Delta e_{t-1} - 0.024\Delta e_{t-2}$$
$$-0.003\Delta e_{t-3} + \left[-0.008 + 0.028\Delta e_{t-1} - 0.201\Delta e_{t-2} + 0.340\Delta e_{t-3}\right].G(EMBI+_{t-1},\gamma,c) + v_t$$

LSTR: $G(EMBI+_{t-1},\gamma,c) = \left(1 + \exp\left\{-1^{**}(EMBI+_{t-1} - 2005^{**})\right\}^{-1}\right)$

Radj^2= 0.993, Sigma= 0.0039, AIC= -10.974, Q(2)= 0.125, Q(4)= 0.097, Q(6)=0.225, RNL= 0.913.

The results for Mexico are:[25]

Linear Model:

$$\pi_t = 0.0004 + 1.499^{**}\pi_{t-1} - 0.590^{**}\pi_{t-2} + 0.076\pi_{t-3} - 0.013\Delta y_t + 0.002\Delta p_t^{imp}$$
$$+0.040^{**}\Delta e_t + 0.015\Delta e_{t-1} - 0.022^{*}\Delta e_{t-2} + 0.043^{**}\Delta e_{t-3} - 0.058^{**}\Delta e_{t-4} + \varepsilon_t$$

Radj^2= 0.998, Sigma= 0.0045, AIC= -10.746, AR= 0.470, ARCH= 0.902, Hetero= 0.112.

Nonlinear Model:

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

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

Radj^2= 0.999, Sigma= 0.0035, AIC= -11.174, Q(2)= 0.225, Q(4)= 0.336, Q(6)=0.230, RNL = 0.921.

---

[25] Note that due to data availability of EMBI+ spreads, the linear model for Mexico used to compare with this particular nonlinear specification was fitted for the period 1995M1 to 2005M12, and not 1992M1 to 2005M12.

The standard errors of the regressions are reduced in the nonlinear model, respectively for Brazil and Mexico, by 7% and 22%; and both AdjR^2 and AIC are improved with the LSTR specification. The threshold value for Mexico is around 760 basis points, whereas for Brazil it is around 2000 basis points. The reason for such a high threshold value for Brazil is related to the fact that EMBI+ spreads have been very high for most of the past 10 years[26].

As expected there is a positive relationship between EMBI+ spreads and ERPT, with the sum of the lagged exchange rate terms in the nonlinear part of the model being positive for both countries. Table 6 shows the sum of the exchange rate coefficients under different values for the transition function. It is easy to observe the very important role of EMBI+ spreads, and hence of credibility levels, on ERPT for both countries, but particularly for Brazil.

For both countries the estimated speed of transition is not high. In the case of Mexico there are two periods where the transition function is close to 1: just after the collapse of the Peso in 1995, and around the Russian and Brazilian crises, in late 1998 and the beginning of 1999. After the year 2000 the transition function is very close to 0, and hence ERPT is rather lower. For Brazil the transition function also shows an increase around the Russian crisis (August 1998), and the period around the breakdown of its dollar-pegged policy, in 1999. Most importantly, however, is the striking increase in the transition function, and hence in ERPT, during 2002, when a huge confidence crisis stormed the country during the presidential elections[27].

4.4 TRANSITION VARIABLE: EXCHANGE RATE

As discussed in the literature review, the literature on nonlinearities and asymmetries in ERPT has mainly looked at the response to the *magnitude* of the exchange rate changes. As mentioned earlier, we revisit this evidence.

---

[26] EMBI+ spreads have been much higher for Brazil than for most emerging markets, reflecting the general external and financial vulnerabilities that characterized the country in recent years.

[27] For an overview of the Mexican crisis of 1995, see Ball and Reyes (2004). For a discussion of the 2002 confidence crisis in Brazil, see Tombini and Alves (2006).

Following the results of the linearity tests presented in Table 7, we fitted a nonlinear model using exchange rate as transition variable for Mexico and the United Kingdom. The selected transition function was an ESTR. As argued by Terasvirta (2001), ESTR models are appropriate in situations in which the local dynamic behaviour of the process is similar at both large and small values of the transition variable and different in the middle. Hence, what matters here is the *size* of the exchange rate shock and not the *direction*. As discussed before, this characteristic of ESTR models makes it possible to capture possible nonlinearities in ERPT with respect to the existence of "menu costs", and thus if firms are more likely to increase prices after large than after small shocks. As in the previous models shown, the nonlinear model provides a better fit to the data than the baseline linear model.

The results for the United Kingdom are:

Linear Model:

$$\pi_t = 0.001^{*} + 1.241^{**}\pi_{t-1} - 0.240^{**}\pi_{t-2} + 0.157^{**}\pi_{t-5} - 0.184^{**}\pi_{t-6} + 0.036^{**}\Delta y_t - 0.023^{*}\Delta y_{t-1}$$
$$+0.047^{**}\Delta p_t^{imp} - 0.044^{**}\Delta p_{t-1}^{imp} + 0.029^{*}\Delta e_t - 0.006\Delta e_{t-1} - 0.035^{**}\Delta e_{t-2} + 0.017^{**}\Delta e_{t-3} + \varepsilon_t$$

Radj^2= 0.976, Sigma= 0.0030, AIC= -11.559, AR= 0.563, ARCH= 0.575, Hetero= 0.020.

Nonlinear Model:

$$\pi_t = 0.029^{**} + 1.200^{**}\pi_{t-1} - 0.200^{**}\pi_{t-2} + 0.134^{**}\pi_{t-5} - 0.147^{**}\pi_{t-6} + 0.031^{**}\Delta y_t - 0.015\Delta y_{t-1}$$
$$+0.045^{**}\Delta p_t^{imp} - 0.043^{**}\Delta p_{t-1}^{imp} + 0.019\Delta e_t + 0.086\Delta e_{t-1} + 0.019\Delta e_{t-2} - 0.029\Delta e_{t-3} +$$
$$\left[-0.069^{**} + 0.018\Delta e_t + 0.027\Delta e_{t-1} - 0.122^{**}\Delta e_{t-2} + 0.109^{**}\Delta e_{t-3}\right].G(\Delta e_{t-1}, \gamma, c) + \nu_t$$

ESTR: $G(\Delta e_{t-1}, \gamma, c) = 1 + \exp\left\{-0.698(\Delta e_{t-1} - 0.046)^2\right\}$

Radj^2= 0.980, Sigma= 0.0029, AIC= -11.617, Q(2)= 0.527, Q(4)= 0.016, Q(6)=0.053, RNL= 0.113.

The results for Mexico are:

Linear Model:

$$\pi_t = 0.001 + 1.506^{**}\pi_{t-1} - 0.613^{**}\pi_{t-2} + 0.087\pi_{t-3} - 0.010\Delta y_t - 0.0004\Delta p_t^{imp}$$
$$+0.039^{**}\Delta e_t + 0.016\Delta e_{t-1} - 0.021^{**}\Delta e_{t-2} + 0.042^{**}\Delta e_{t-3} - 0.056^{**}\Delta e_{t-4} + \varepsilon_t$$

Radj^2= 0.998, Sigma= 0.0041, AIC= -10.946, AR= 0.265, ARCH= 0.736, Hetero= 0.080.

Nonlinear Model:

$$\pi_t = 0.001 + 1.378^{**}\pi_{t-1} - 0.468^{**}\pi_{t-2} + 0.051\pi_{t-3} - 0.009\Delta y_t - 0.004\Delta p_t^{imp}$$
$$+0.044^{**}\Delta e_t - 0.020\Delta e_{t-1} + 0.006\Delta e_{t-2} + 0.005\Delta e_{t-3} - 0.002\Delta e_{t-4}$$
$$+\left[-0.003^{**} - 0.002\Delta e_t + 0.070^{**}\Delta e_{t-1} - 0.043\Delta e_{t-2} + 0.065^{**}\Delta e_{t-3} - 0.083^{**}\Delta e_{t-4}\right].G(\Delta e_{t-2}, \gamma, c) + \nu_t$$

ESTR: $G(\Delta e_{t-2}, \gamma, c) = 1 + \exp\left\{-6.993(\Delta e_{t-2} - 0.244)^2\right\}$

Radj^2= 0.999, Sigma= 0.0034, AIC= -11.287, Q(2)= 0.250, Q(4)= 0.139, Q(6)=0.277, RNL= 0.904.

Table 8 shows the sum of the exchange rate coefficients. For Mexico the response is smaller than for the previous transition variables. In the case of the United Kingdom the difference between the two states is not very substantial although it is statistically significant.

Both the threshold value and the speed of transition differ greatly for Mexico and the United Kingdom. For the former country the speed is much faster, and the threshold is approximately 24%. This result suggests a role of exchange rate crisis in increasing ERPT. As a matter of fact an analysis of the transition function over time for Mexico show that the main period in which ERPT is maximum is around the breakdown of the Peso, in 1995. In this sense this result should be viewed as complementary of the nonlinear models of Mexico with EMBI+ and Rids as transition variables. Regarding the United Kingdom, the speed of transition is rather slow, and the threshold value is 5%. Since the estimated threshold is not zero, its slightly positive value suggests the presence of menu costs or some adjustment costs of prices, where small exchange rate changes are not promptly transmitted to prices[28].

---

[28] As discussed by Correa and Minella (2006), if price changes are costly, a small change in the currency value can be accommodated within the mark-up margin. However, if exchange rate changes surpass some limit, the costs of not adjusting prices are higher, leading firms to change prices more promptly.

4.5 TRANSITION VARIABLE: OUTPUT GROWTH

Goldfajn and Werlang (2000) argued that one important determinant of ERPT is the output gap. Intuitively, firms would find it easier to pass-through cost changes when the economy is growing fast, rather then when it is in a recession and its sales are already falling. Following this argument, Garcia and Restrepo (2001) suggested that one of the reasons why ERPT was low in Chile in the 1990s was because negative output gaps were offsetting some of the inflationary implications of exchange rate depreciations.

In this section we analyse the extent to which ERPT may be nonlinear to output growth. Table 9 shows the linearity tests against STR nonlinearity. The results suggest possible nonlinearities of ERPT to output in Mexico, South Africa and the Czech Republic. An LSTR model was fitted for each country.

The results for the Czech Republic are:

Linear Model:

$$\pi_t = 0.006^{**} - 0.0001^{**}T + 0.927^{**}\pi_{t-1} + 0.046^{**}\Delta y_t + 0.009^{**}\Delta p_t^{imp} + 0.019^{**}\Delta e_t + \varepsilon_t$$

Radj^2= 0.979, Sigma= 0.0052, AIC= -10.492, AR= 0.141, ARCH= 0.887, Hetero= 0.004.

Nonlinear Model:

$$\begin{aligned}\pi_t &= 0.010^{**} - 0.0001^{**}T + 0.891^{**}\pi_{t-1} + 0.039^{**}\Delta y_t + 0.012^{**}\Delta p_t^{imp} \\ &+ 0.013^{**}\Delta e_t + \left[0.006^{*} - 0.0001T + 0.029^{**}\Delta e_t\right].G(\Delta y_{t-3}, \gamma, c) + \nu_t\end{aligned}$$

LSTR: $G(\Delta y_{t-3}, \gamma, c) = \left(1 + \exp\left\{-88(\Delta y_{t-3} - 0.071^{**})\right\}^{-1}\right)$

Radj^2= 0.983, Sigma= 0.0048, AIC= -10.590, Q(2)= 0.876, Q(4)= 0.275, Q(6)=0.258, RNL= 0.142.

The results for Mexico are:

Linear Model:

$$\begin{aligned}\pi_t &= 0.001 + 1.506^{**}\pi_{t-1} - 0.613^{**}\pi_{t-2} + 0.087\pi_{t-3} - 0.010\Delta y_t - 0.0004\Delta p_t^{imp} \\ &+ 0.039^{**}\Delta e_t + 0.016\Delta e_{t-1} - 0.021^{**}\Delta e_{t-2} + 0.042^{**}\Delta e_{t-3} - 0.056^{**}\Delta e_{t-4} + \varepsilon_t\end{aligned}$$

Radj^2= 0.998, Sigma= 0.0041, AIC= -10.946, AR= 0.265, ARCH= 0.736, Hetero= 0.080.

Nonlinear Model:

$$\pi_t = 0.001 + 1.513^{**}\pi_{t-1} - 0.660^{**}\pi_{t-2} + 0.119^{**}\pi_{t-3} - 0.011\Delta y_t$$
$$-0.002\Delta p_t^{imp} + 0.039^{**}\Delta e_t + 0.031^{**}\Delta e_{t-1} - 0.037^{**}\Delta e_{t-2} + 0.051^{**}\Delta e_{t-3} - 0.059^{**}\Delta e_{t-4} +$$
$$\left[0.0003 + 0.020\Delta e_t - 0.151^{**}\Delta e_{t-1} + 0.156^{**}\Delta e_{t-2} - 0.106^{**}\Delta e_{t-3} + 0.082^{**}\Delta e_{t-4}\right].G(\Delta y_{t-3},\gamma,c) + v_t$$

LSTR: $G(\Delta y_{t-3},\gamma,c) = \left(1 + \exp\left\{-6^{*}(\Delta y_{t-3} - 0.066^{**})\right\}^{-1}\right)$

Radj^2= 0.998, Sigma= 0.0038, AIC= -11.048, Q(2)= 0.104, Q(4)= 0.321, Q(6)=0.371, RNL= 0.398.

The results for South Africa are:

Linear Model:

$$\pi_t = 0.002^{**} + 1.229^{**}\pi_{t-1} - 0.196^{*}\pi_{t-2} - 0.090\pi_{t-3} - 0.025^{*}\Delta y_t + 0.050^{**}\Delta p_{t-1}^{IMP}$$
$$-0.032^{*}\Delta p_{t-2}^{IMP} + 0.038^{**}\Delta e_{t-1} - 0.034^{**}\Delta e_{t-3} + 0.016^{**}\Delta e_{t-4} + \varepsilon_t$$

Radj^2= 0.974, Sigma= 0.0053, AIC= -10.439, AR= 0.985, ARCH= 0.911, Hetero= 0.924.

Nonlinear Model:

$$\pi_t = 0.003^{**} + 1.232^{**}\pi_{t-1} - 0.159\pi_{t-2} - 0.135^{**}\pi_{t-3} - 0.036^{**}\Delta y_t + 0.049^{**}\Delta p_{t-1}^{imp} - 0.033^{**}\Delta p_{t-2}^{imp} + 0.036^{**}\Delta e_{t-1}$$
$$-0.038^{**}\Delta e_{t-3} + 0.020^{**}\Delta e_{t-4} + \left[0.003^{**} + 0.059^{**}\Delta e_{t-1} - 0.037\Delta e_{t-3} - 0.001\Delta e_{t-4}\right].G(\Delta y_{t-2},\gamma,c) + v_t$$

LSTR: $G(\Delta y_{t-2},\gamma,c) = \left(1 + \exp\left\{-98(\Delta y_{t-2} - 0.062)\right\}^{-1}\right)$

Radj^2= 0.978, Sigma= 0.0051, AIC= -10.456, Q(2)= 0.851, Q(4)= 0.684, Q(6)=0.673, RNL= 0.376.

The nonlinear specification passes the basic diagnostic tests. As in the other cases analysed, the AdjR^2 was higher, and the AIC was lower for the STR model. The sum of the exchange rate coefficients is shown in Table 10. As expected the sum of the nonlinear exchange rate terms is positive, indicating that when the economy is overheated i.e. above the threshold, ERPT is higher. Mexico's response, however, is very weak. For South Africa and the Czech Republic, on the other hand, the response is rather strong. The threshold value (annual growth) is quite similar for the three countries: 6% for South Africa, 7% for the Czech Republic, and 6.5% for Mexico.

5. CONCLUSIONS

In this paper we have investigated the role of nonlinearities in exchange rate pass-through into inflation (ERPT) for a set of emerging and developed economies that adopted Inflation Targeting (IT). We presented a simple mark-up model of import prices where we posit that nonlinearities may arise due to changes in macroeconomic conditions. Previous work has focused only on asymmetries arising from the size and direction of exchange rate shocks. We analyse nonlinearities not only with respect to the exchange rate, but use a wide range of threshold variables that could potentially influence ERPT.

The results suggest that, although the sources of nonlinearities vary considerably across countries, they cannot be neglected. There is widespread evidence of nonlinear ERPT. We found a nonlinear response of ERPT with respect to the magnitude of exchange rate changes for Mexico and the United Kingdom only. For South Africa, Czech Republic and Mexico, ERPT seems to respond nonlinearly to the output gap. We found that ERPT responds nonlinearly to inflation in South Africa, Mexico, Canada, and the United Kingdom. Finally, for Mexico, Brazil and Czech Republic, ERPT seems to be affected nonlinearly by some measure of macroeconomic instability.

Our results provide some evidence with respect to the recent decline in ERPT observed in the literature. As proposed by Taylor (2000), we found that higher inflation leads to higher ERPT, and thus the lower ERPT of the 1990s may be related to the fall in the rates of inflation. Regarding some emerging markets, however, we observed the important role of macroeconomic measures of risk for ERPT. Confidence crisis, brought about by poor macroeconomic policies, seem to increase ERPT. In this sense the adoption of better economic policies, leading to lower inflation and higher credibility levels, appear to be effective tools to reduce ERPT.

APPENDIX I

Date of adoption of the Inflation Targeting regime

Canada	February 1991
United Kingdom	October 1992
Brazil	June 1999
Czech Republic	January 1998
Mexico	January 1999
South Africa	February 2000

REFERENCES

Agenor, P. (2000). "Monetary policy under flexible exchange rates: an introduction to Inflation Targeting". The World Bank. Mimeo.

Athukorala, P. (1991). "Exchange rate pass-through: the case of Korean exports of manufactures". Economics Letters, 35. pp. 79-84.

Ball, C and Reyes, J (2004). "Inflation Targeting or Fear of Floating in disguise: the case of Mexico". International Journal of Finance and Economics. 9. pp 49-69.

Bailliu, J. and Fujii, E. (2004). "Exchange rate pass-through and the inflation environment in industrialized countries: an empirical investigation". Bank of Canada, Working Paper nº 2004/21.

Bussiere, M. (2006). "Exchange rate pass-through in the G7 economies: the role of nonlinearities and asymmetries". European Central Bank. Mimeo.

Campa, J. and Goldberg, L. (2005). "Exchange rate pass-through into import prices". Review of Economics and Statistics, 87. pp. 679-690.

Choudhri, E. and Hakura, D. (2006). "Exchange rate pass-through to domestic prices: does the inflationary environment matter?" Journal of International Money and Finance, 25, pp. 614-639.

Christopoulos, D.K., León-Ledesma, M.A. (2007). "A long-run nonlinear approach to the Fischer effect". Journal of Money, Credit and Banking, 39, pp. 543-559.

Clifton, E., Leon. H. and Wong, C. (2001). "Inflation Targeting and the unemployment-inflation trade-off". IMF, Working Paper nº 166.

Correa, A. and Minella, A. (2006). "Nonlinear mechanisms of exchange rate pass-through: a Phillips curve model with threshold for Brazil". Central Bank of Brazil, Working Paper nº 122.

Feinberg, R. (1989). "Symmetric pass-through of tariffs and exchange rates under imperfect competition: an empirical test". Review of Economics and Statistics, 71. pp. 505-511.

Gagnon, J. and Ihrig, J. (2004). "Monetary policy and exchange rate pass-through". International Journal of Finance and Economics, 9. pp. 315-338.

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

Gil-Pareja, S. (2003). "Pricing to market behaviour in European car markets". European Economic Review, 47. pp. 945-962.

Goldberg, K. (1995). "Product differentiation and oligopoly in international markets: the case of the US automobile industry". Econometrica, 63. pp. 891-951.

Goldberg, K. and Knetter, M. (1997). "Goods prices and exchange rates: what have we learned?" Journal of Economic Literature, 35. pp. 1243-1272.

Goldfajn, I. and Werlang, S. (2000). "The pass-through from depreciation to inflation: a panel study". Central Bank of Brazil, Working Paper nº 5.

Herzberg, V., Kapetanios, G. and Price, S. (2003). "Import prices and exchange rate pass-through: theory and evidence from the United Kingdom". Bank of England, Working Paper nº 182.

Kadiyali, V. (1997). "Exchange rate pass-through for strategic pricing and advertising: an empirical analysis of the photographic film industry". Journal of International Economics, 43. pp. 437-461.

Kaminsky, Graciela, Lizondo, Saul and Reinhart, Carmen (1998). "Leading indicators of currency crises". IMF Staff Paper, 45.

Karhonen, M. (2005). "Nonlinearities in exchange rate: evidence from smooth transition regression model. University of Oulo. Mimeo.

Leiderman, L. and Bar-Or, H. "Monetary Policy Rules and Transmission Mechanisms under Inflation Targeting in Israel." Working paper, Bank of Israel.

Lundbergh, S., Terasvirta, T. and van Dijk, D. (2000). "Time-varying smooth transition autoregressive models." Working Paper Series in Economics and Finance nº 376, Stockholm School of Economics.

Mahdavi, S. (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. pp. 2115-2125.

Mann, C. (1986). “Prices, profit margins, and exchange rates”. Federal Reserve Bulleting, 72. pp. 366-379.

Marazzi, M. et al. (2005). “Exchange rate pass-through to US import prices: some new evidence”. Board of Governors of the Federal Reserve System, International Finance Discussion Paper, nº 832.

McAdam, P. and Willman, A. (2004). “Supply, Factor Shares and Inflation Persistence: Re-examining Euro-area New-Keynesian Phillips Curves.” Oxford Bulletin of Economics & Statistics, 66, pp. 637-670

Minella, A. de Freitas, P.S., Goldfajn, I. and Muinhos, M.K. (2003). “Inflation targeting in Brazil: constructing credibility under exchange rate volatility”. Journal of International Money and Finance. 22. pp. 1015-1040.

Mishkin, F.S. and Savastano, M. (2001). “Monetary policy strategies for Latin America”. Journal of Development Economics. 66. pp. 415-444.

Ohno, K. (1989). “Export pricing behaviour of manufacturing: a US-Japan comparison”. IMF Staff Papers, 36. pp. 550-579.

Olivei, G. (2002). “Exchange rates and the prices of manufacturing products imported into the United States”. New England Economic Review, 1. pp. 3-18.

Pelzman, S. (2000). “Prices rise faster than they fall”. Journal of Political Economy, 108. pp. 466-502.

Pollard, P. and Coughlin, C. (2004). “Size matters: asymmetric exchange rate pass-through at the industrial level”. University of Nottingham, Research Paper Series nº 2004/13.

Schmidt-Hebbel, K. and Tapia, M. (2002). “Inflation targeting in Chile”. The North American Journal of Economics and Finance. 13. pp. 125-146

Steel, D. and King, A. (2004). “Exchange rate pass-through: the role of regime changes”. International Review of Applied Economics, 18. pp. 301-322.

Taylor, J. (2000). “Low inflation, pass-through and the pricing power of firms”. European Economic Review, 44. pp. 1389-1408.

Terasvirta, T. (1998). “Modelling economic relationships with smooth transition regressions”. In Handbook of Applied Economics Statistics; Marcel Dekker (ed.): New York. pp. 507-552.

Terasvirta, T. (2004). “Smooth transition regression modelling”. In Applied Time Series Econometrics; Helmut Lutkepohl and Markus Kratzig (eds.): Cambridge University Press. pp. 222-242.

Tombini, A.A. and Lago Alves, S.A. (2006). “The recent Brazilian disinflation process and costs”. Central Bank of Brazil. Working Paper nº 109.

Weber, A. (2000). “Newton’s gravity law and import prices in the Asia Pacific”. Japan and the World Economy, 12. pp. 71-87.

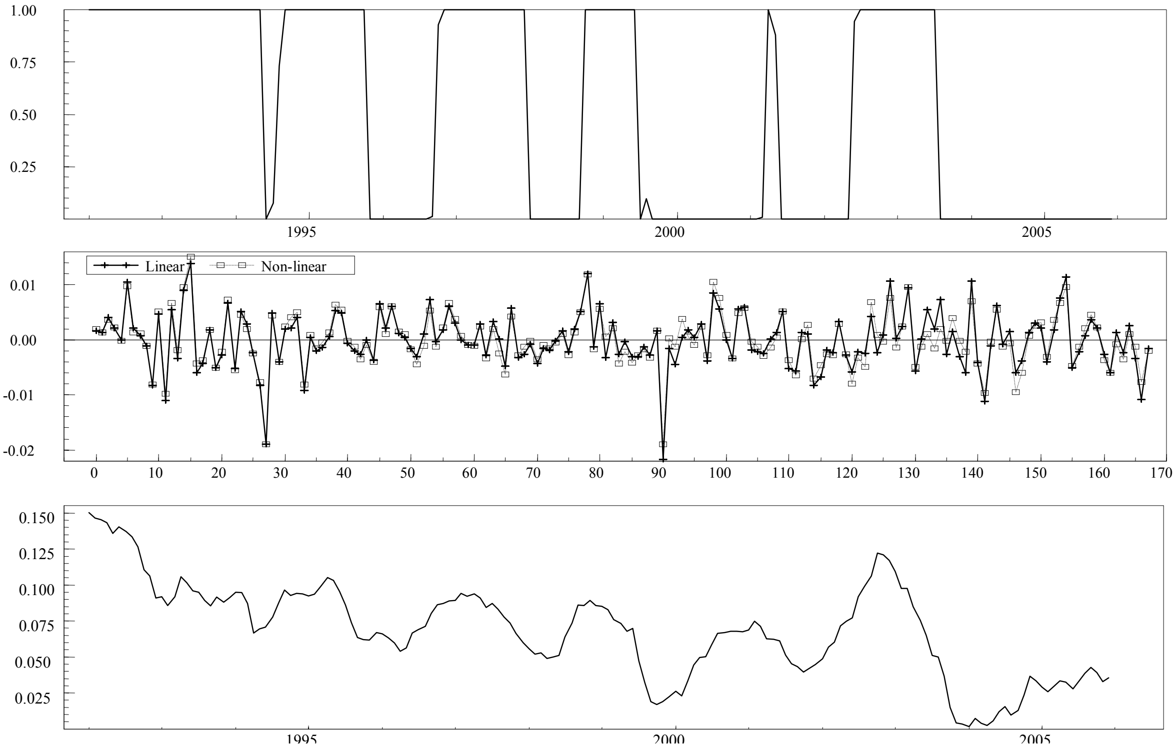
van Dijk, D. Terasvirta, T. and Franses, P. H. (2002). “Smooth transition autoregressive models – a survey of recent developments”. Econometrics Reviews, 21. pp. 1-47.

FIGURES

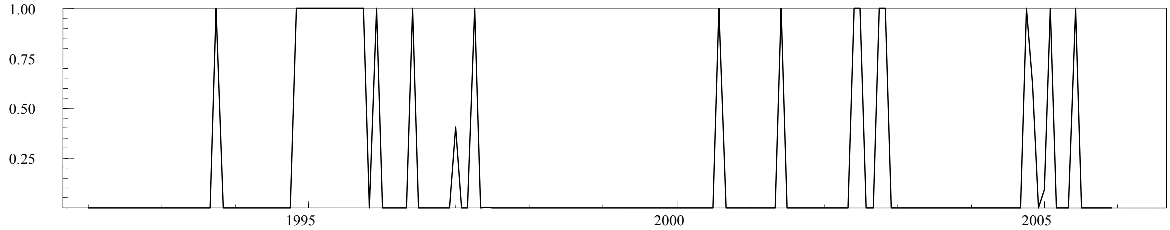
Note: The graphs show the transition function over time, the estimation residuals of both the linear and nonlinear models, and the transition variable over time.

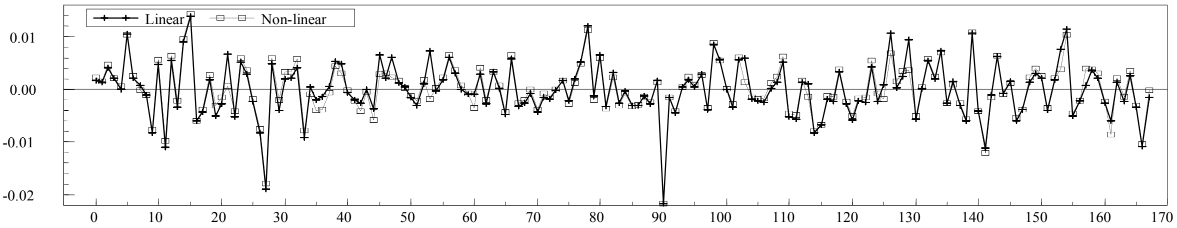
South Africa

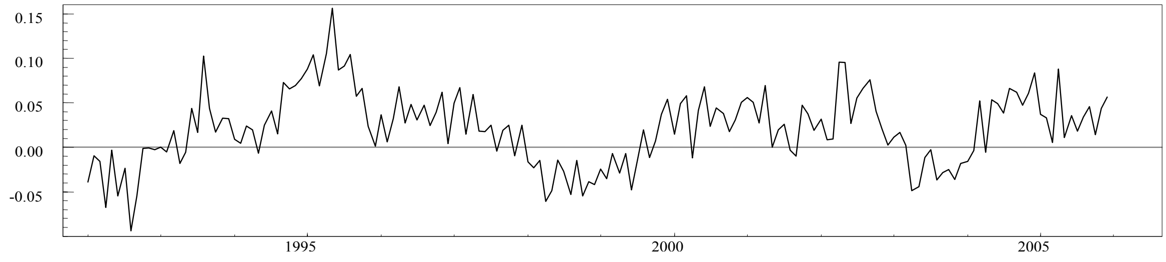
1. Transition Variable: Inflation



2. Transition Variable: Output Growth

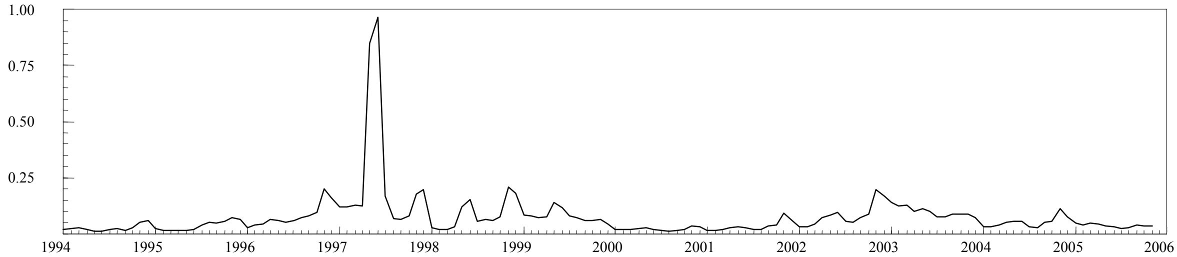


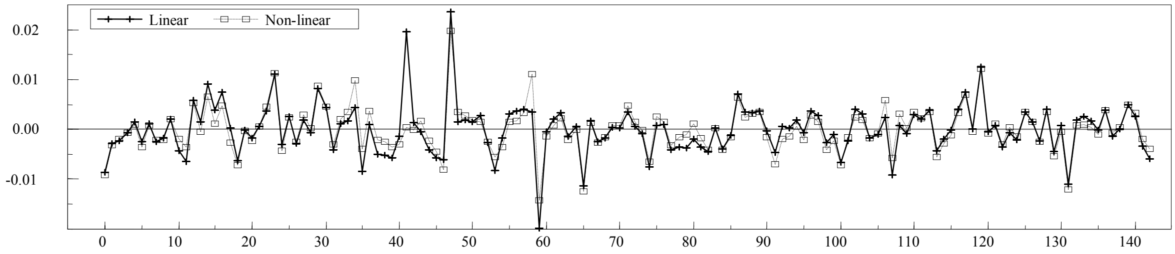


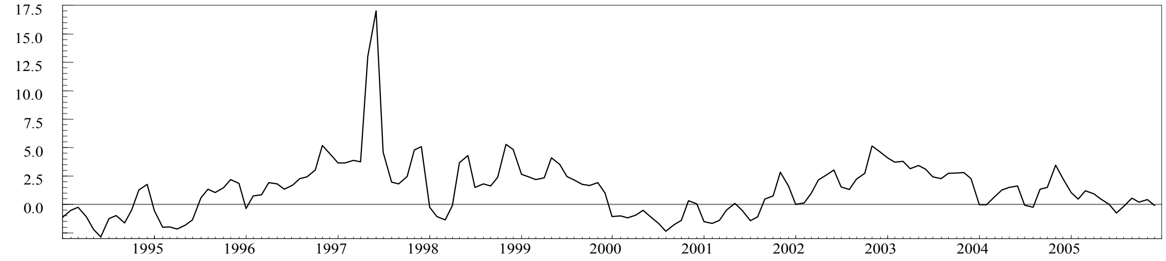


Czech Republic

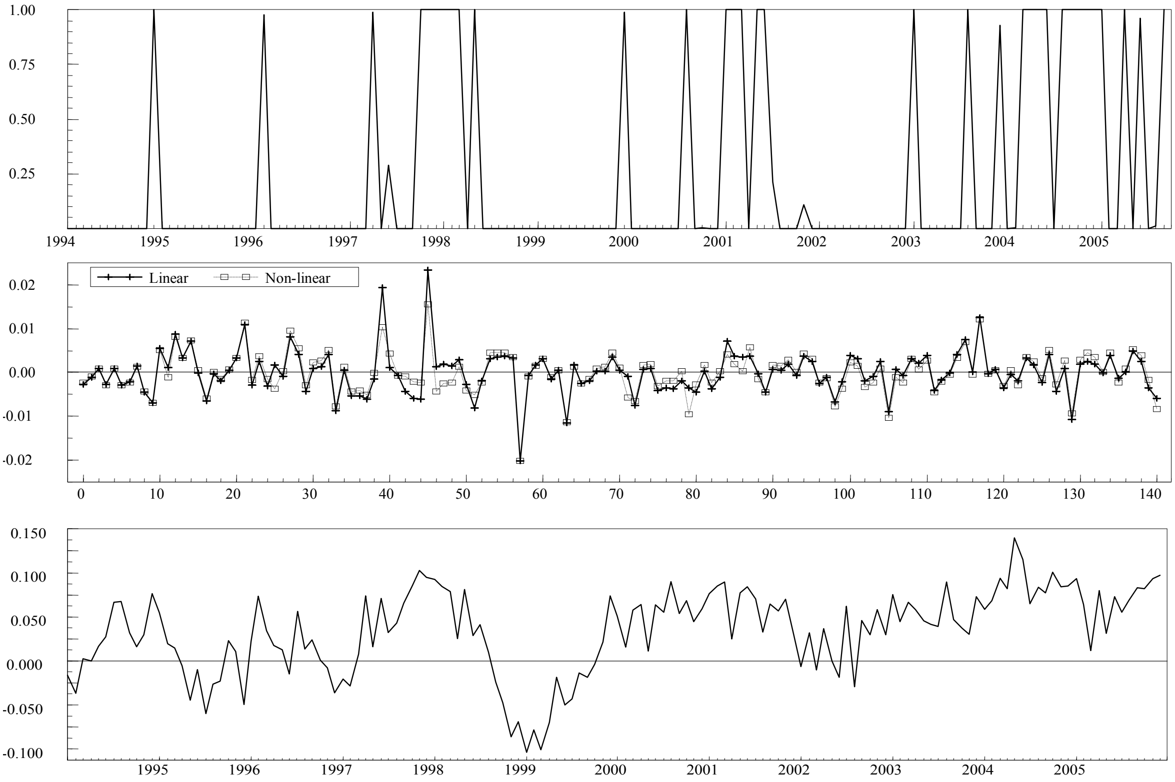
1. Transition Variable: Rids





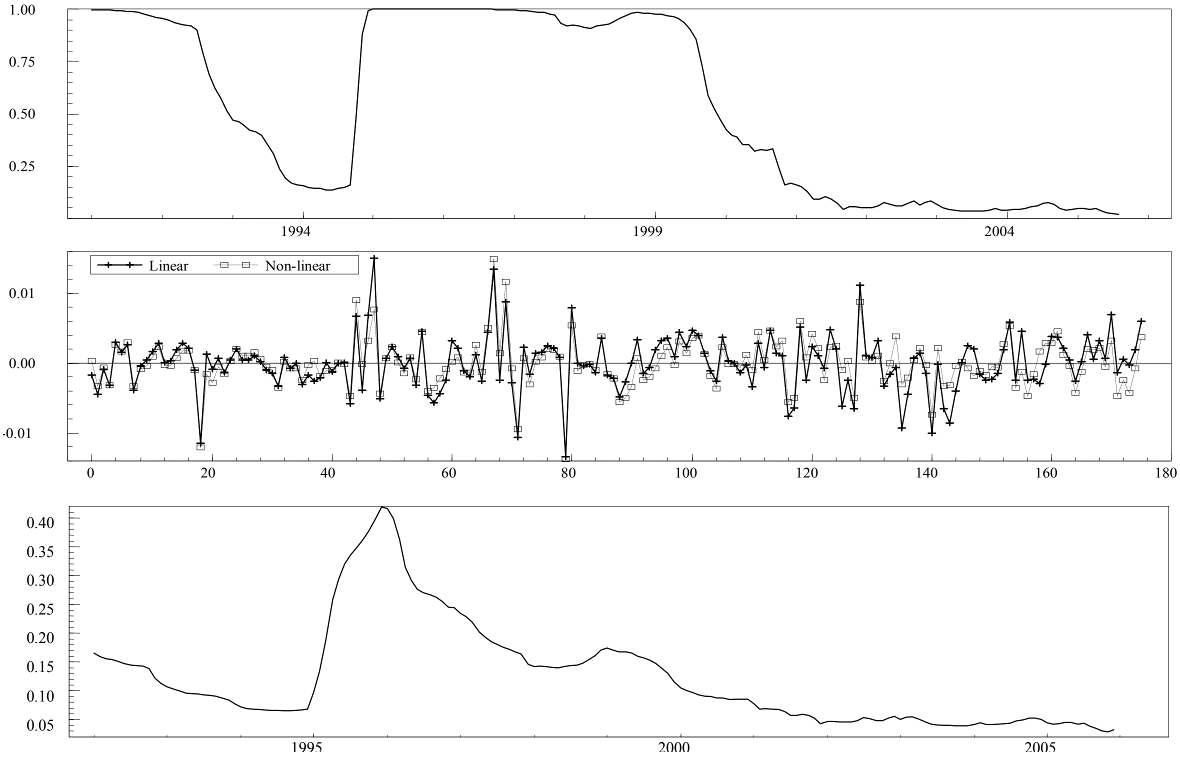


2. Transition Variable: Output Growth

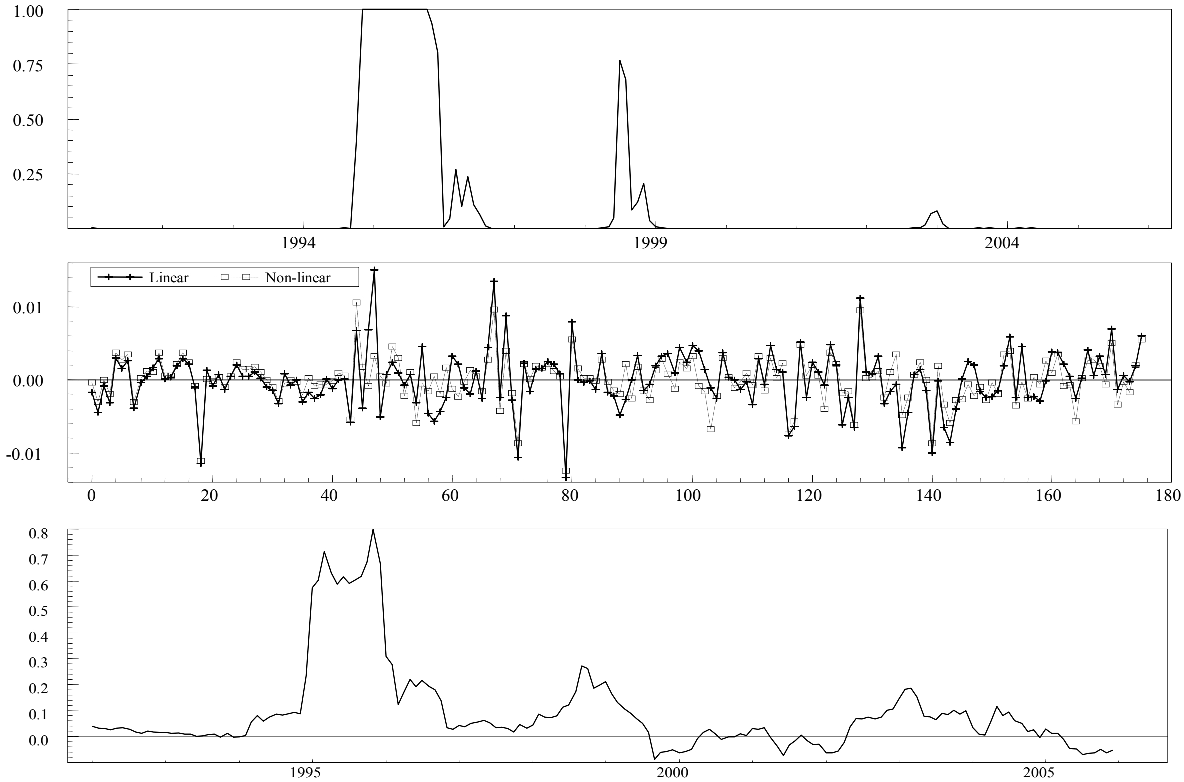


Mexico

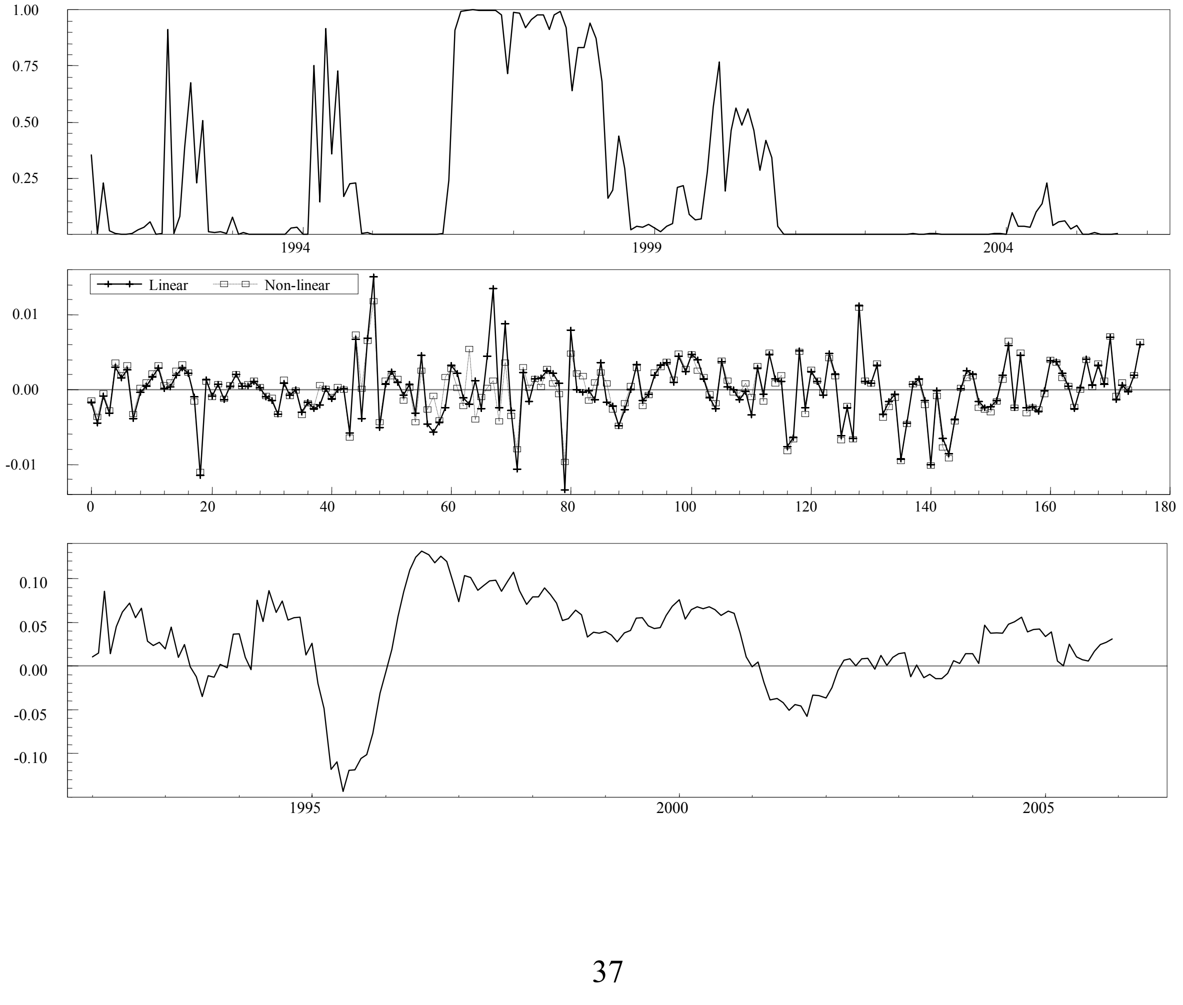
1. Transition Variable: Inflation



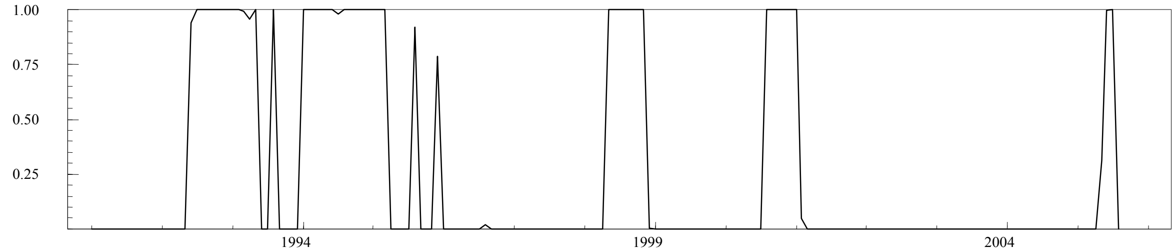
2. Transition Variable: Exchange rate

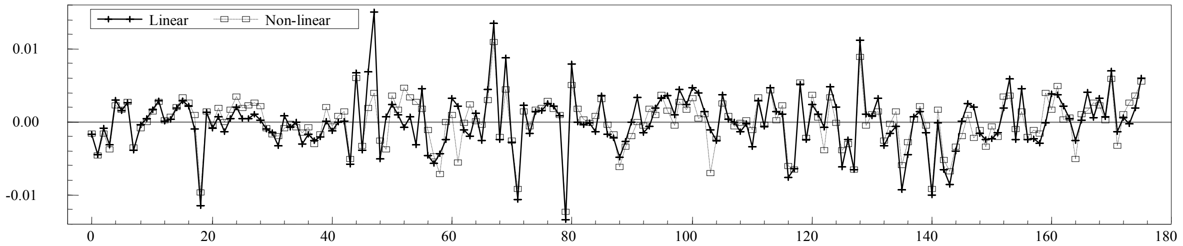


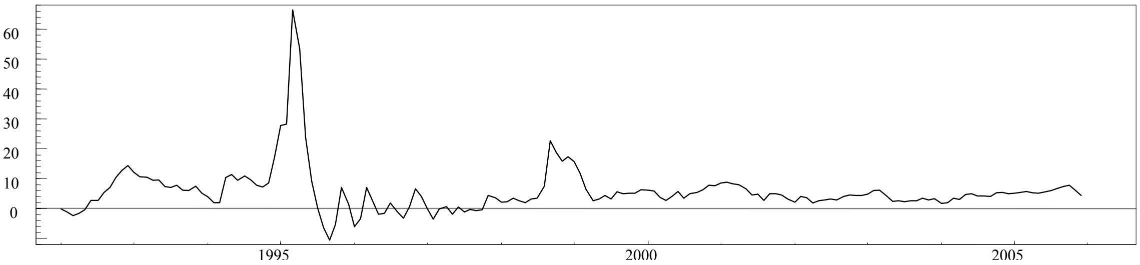
3. Transition Variable: Output Growth



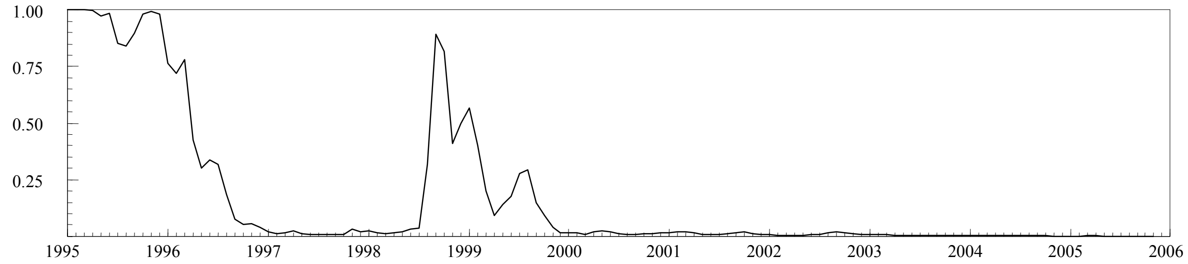
4. Transition Variable: Rids

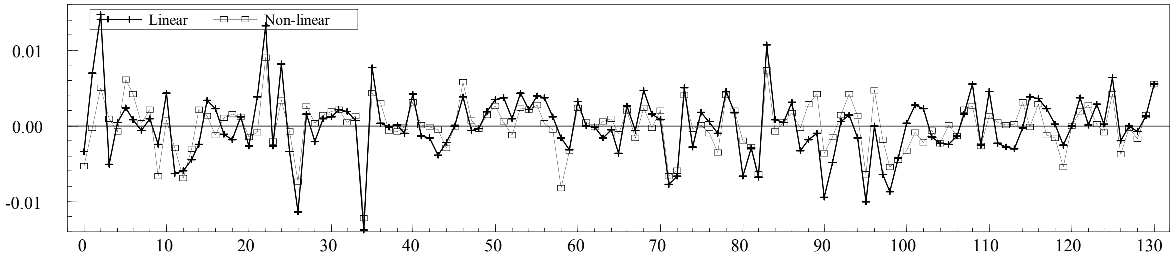


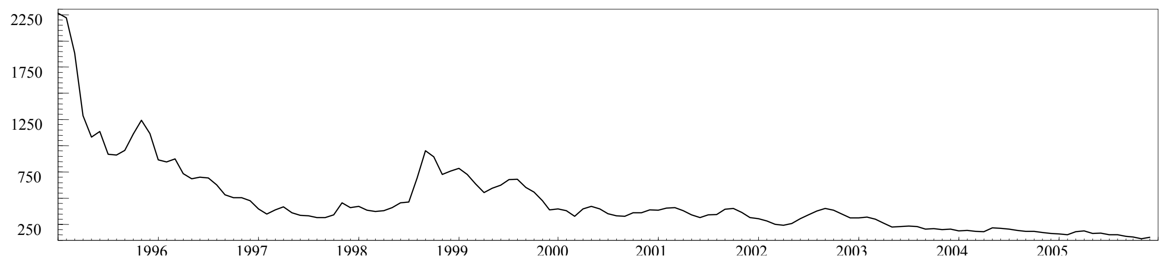




5. Transition Variable: EMBI+

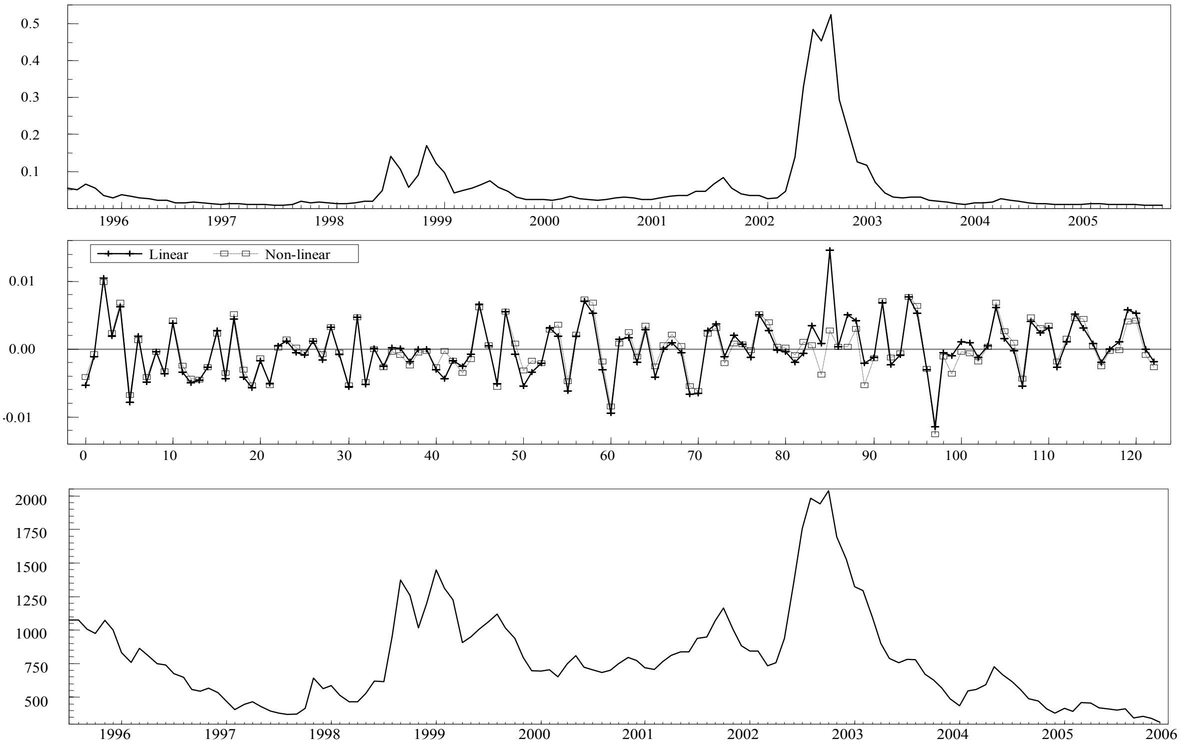






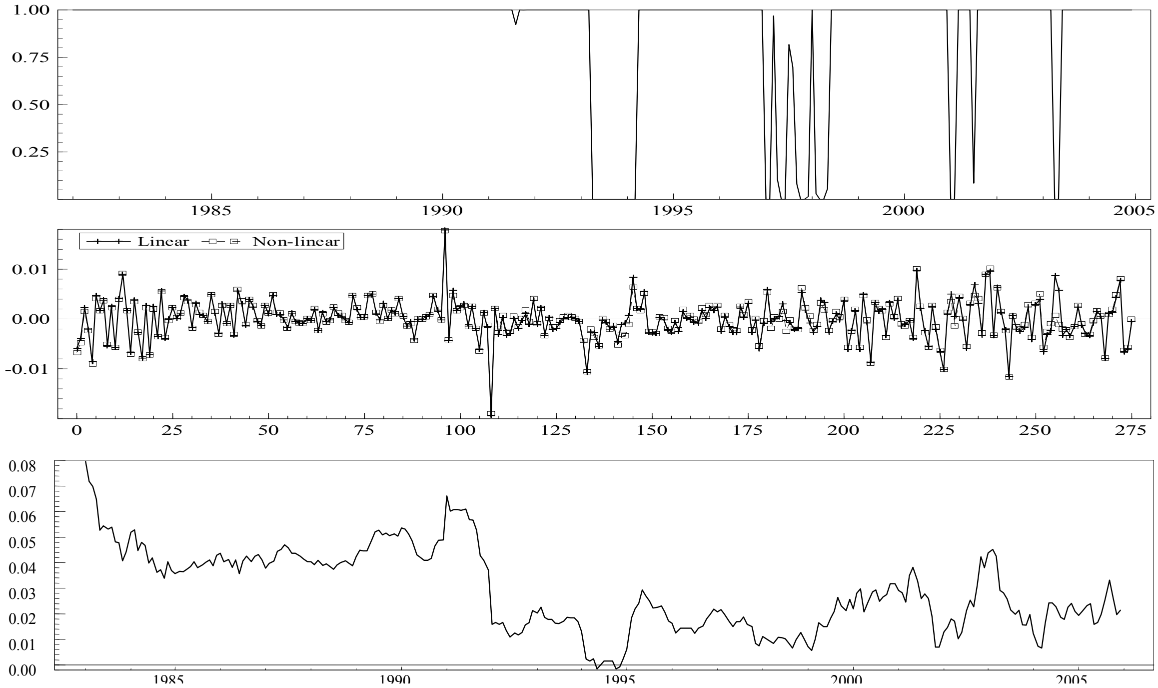
Brazil

1. Transition Variable: EMBI+



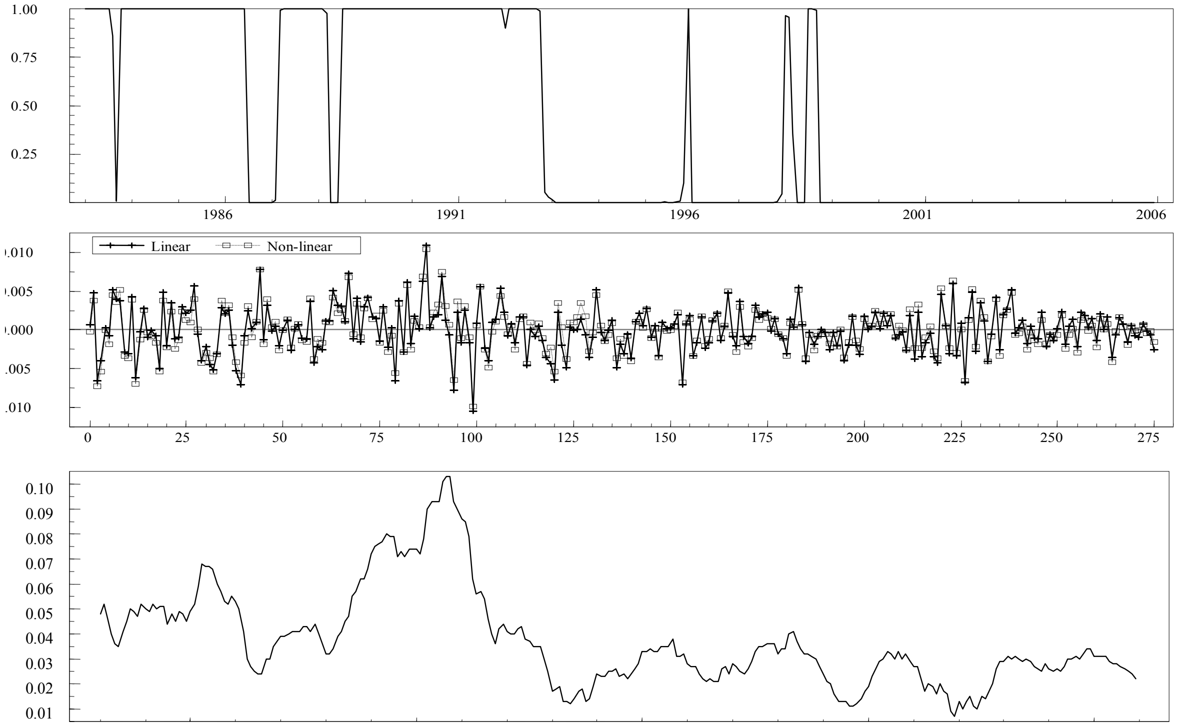
Canada

1. Transition Variable: Inflation

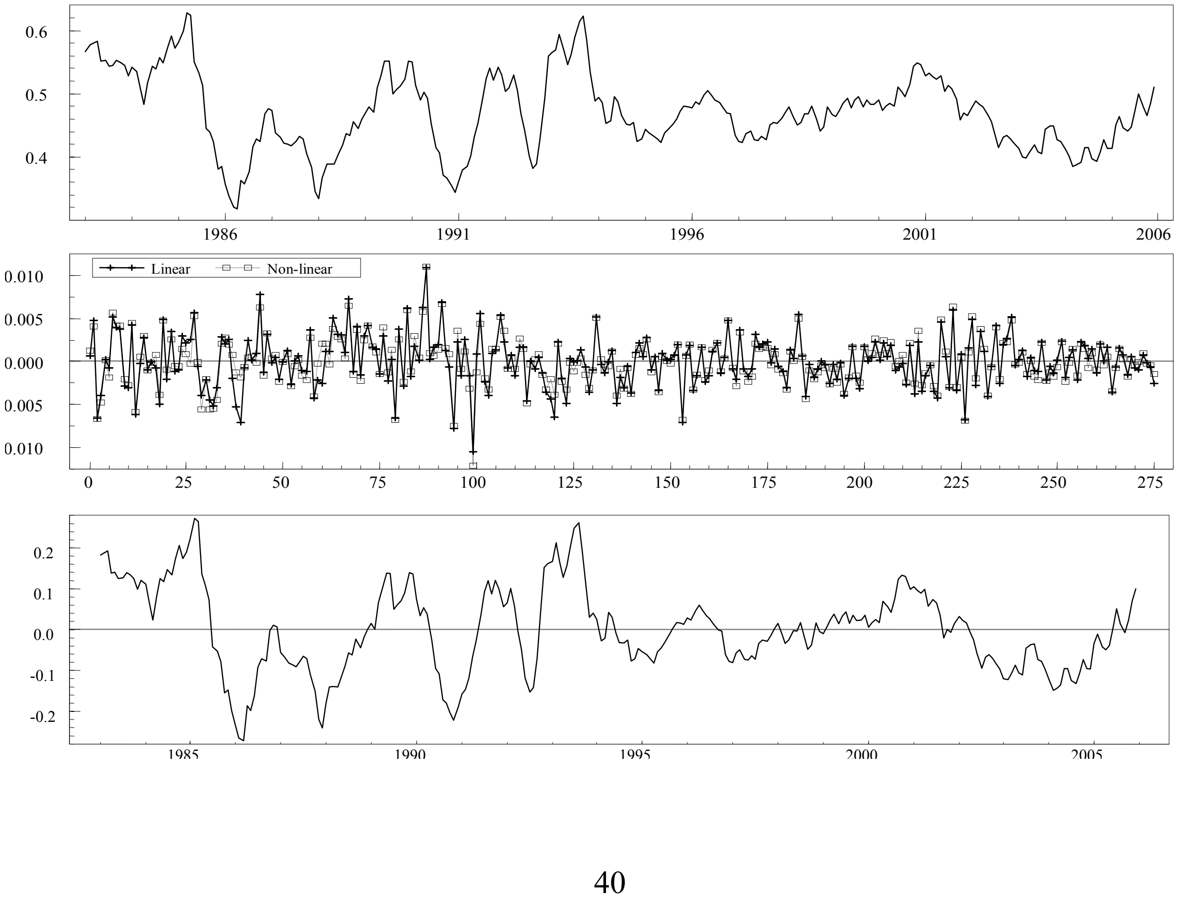


United Kingdom

1. Transition Variable: Inflation



2. Transition Variable: Exchange rate



TABLES

Table 1: *Linearity tests – Transition variable: Inflation*

	Brazil			Czech Republic			Mexico		
	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-1}	π_{t-2}	π_{t-3}
LM_2	0.133	0.176	0.168	0.856	0.938	0.880	0.037	0.143	0.206
LM_3	0.677	0.868	0.717	0.713	0.126	0.923	0.001	0.114	0.174
	South Africa			Canada			United Kingdom		
	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-1}	π_{t-2}	π_{t-3}
LM_2	0.194	0.209	0.256	0.796	0.659	0.547	0.043	0.032	0.034
LM_3	0.031	0.036	0.045	0.011	0.032	0.154	0.079	0.055	0.065

Note: the numbers are p-values of F variants of the LM-type tests of linearity against STR nonlinearity.

Table 2: *Sum of the exchange rate coefficients - Inflation*

	South Africa	Mexico	United Kingdom	Canada
G = 0	0.046	0.005	-0.002	-0.023
G = ½	0.051	0.026	0.004	-0.005
G = 1	0.056	0.046	0.009	0.014

Note: the numbers show the sum of the exchange rate coefficients when the transition function is equal to zero (G=0), ½ (G=½), and when it is equal to one (G=1).

Table 3: *Linearity tests - Rids*

	Brazil			Czech Republic			Mexico		
	$Rids_{t-1}$	$Rids_{t-2}$	$Rids_{t-3}$	$Rids_{t-1}$	$Rids_{t-2}$	$Rids_{t-3}$	$Rids_{t-1}$	$Rids_{t-2}$	$Rids_{t-3}$
LM_2	0.592	0.618	0.810	0.000	0.000	0.568	0.000	0.008	0.017
LM_3	0.271	0.207	0.314	0.000	0.000	0.561	0.002	0.000	0.000
	South Africa			Canada			United Kingdom		
	$Rids_{t-1}$	$Rids_{t-2}$	$Rids_{t-3}$	$Rids_{t-1}$	$Rids_{t-2}$	$Rids_{t-3}$	$Rids_{t-1}$	$Rids_{t-2}$	$Rids_{t-3}$
LM_2	0.113	0.280	0.183	0.570	0.102	0.132	0.560	0.644	0.361
LM_3	0.377	0.831	0.827	0.472	0.850	0.567	0.369	0.095	0.289

Note: the numbers are p-values of F variants of the LM-type tests of linearity against STR nonlinearity.

Table 4: *Sum of the exchange rate coefficients - Rids*

	Czech Republic	Mexico
G = 0	-0.004	0.021
G = ½	0.141	0.028
G = 1	0.285	0.034

Note: the numbers show the sum of the exchange rate coefficients when the transition function is equal to zero (G=0), ½ (G=½), and when it is equal to one (G=1).

Table 5: *Linearity tests – EMBI+*

	Brazil			Mexico		
	$EMBI_{t-1}$	$EMBI_{t-2}$	$EMBI_{t-3}$	$EMBI_{t-1}$	$EMBI_{t-2}$	$EMBI_{t-3}$
LM_2	0.002	0.005	0.014	0.000	0.000	0.000
LM_3	0.035	0.052	0.153	0.001	0.000	0.000

Note: the numbers are p-values of F variants of the LM-type tests of linearity against STR nonlinearity.

Table 6: *Sum of the exchange rate coefficients – EMBI+*

	Brazil	Mexico
G = 0	0.004	0.019
G=½	0.088	0.036
G = 1	0.171	0.053

Note: the numbers show the sum of the exchange rate coefficients when the transition function is equal to zero (G=0), ½ (G=½), and when it is equal to one (G=1).

Table 7: *Linearity tests – Exchange rate*

	Brazil			Czech Republic			Mexico		
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}
LM_2	0.753	0.885	0.737	0.330	0.516	0.315	0.000	0.000	0.000
LM_3	0.347	0.515	0.237	0.098	0.259	0.384	0.000	0.000	0.000
	South Africa			Canada			United Kingdom		
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}
LM_2	0.897	0.953	0.998	0.770	0.886	0.964	0.057	0.087	0.131
LM_3	0.176	0.377	0.235	0.726	0.796	0.713	0.005	0.005	0.004

Note: the numbers are p-values of F variants of the LM-type tests of linearity against STR nonlinearity.

Table 8: *Sum of the exchange rate coefficients – Ex. Rate*

	United Kingdom	Mexico
G = 0	0.095	0.033
G = 1	0.137	0.040

Note: the numbers show the sum of the exchange rate coefficients when the transition function is equal to zero (G=0) and when it is equal to one (G=1).

Table 9: *Linearity tests – Output Growth*

	Brazil			Czech Republic			Mexico		
	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}
LM_2	0.310	0.402	0.843	0.348	0.209	0.029	0.000	0.000	0.005
LM_3	0.347	0.663	0.260	0.271	0.003	0.029	0.240	0.042	0.032

	South Africa			Canada			United Kingdom		
	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}
LM_2	0.815	0.995	0.682	0.229	0.279	0.273	0.843	0.832	0.569
LM_3	0.247	0.003	0.996	0.602	0.283	0.267	0.486	0.255	0.383

Note: the numbers are p-values of F variants of the LM-type tests of linearity against STR nonlinearity.

Table 10: *Sum of the exchange rate coefficients - Output*

	South Africa	Mexico	Czech Republic
G = 0	0.018	0.025	0.013
G=½	0.029	0.026	0.028
G = 1	0.039	0.026	0.042

Note: the numbers show the sum of the exchange rate coefficients when the transition function is equal to zero (G=0), ½ (G=½), and when it is equal to one (G=1).