Modelling Real Exchange Rate Effects on Growth in Latin America

by

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Abstract
This paper empirically analyses the dynamics of real per capita GDP growth for six Latin American countries (Argentina, Brazil, Chile, Columbia, Mexico, Venezuela), focusing on the role of the real exchange rate. We argue that real exchange rate effects on growth may be inherently nonlinear, which we capture through a smooth transition regression model. With the exception of Mexico, nonlinearity in economic growth is associated with changes in the real exchange rate, with depreciations leading to different relationships compared with appreciations. Regimes for Mexico are associated with the business cycle through past growth rates, with effectively symmetric effects of real exchange rate changes. Overall, our results are in accord with other recent literature that depreciations have negative effects for growth.

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1. Introduction

The effect of the real exchange rate on output growth is extremely important for developing countries and is a subject of great debate among economists. The controversy concerns the effect to the economy from depreciation in the real exchange rate. The orthodox view sees this as expansionary with the substitution of imports with home goods and increased exports putting the economy on a path of greater sustained growth (Dornbusch, 1988). This has led the World Bank and the IMF to use devaluations in their stabilisation programmes. The *New Structuralist* school (see inter alia Díaz-Alejandro, 1963; Krugman and Taylor, 1978; van Wijnbergen, 1986), on the other hand, emphasises that a real depreciation could be contractionary. They argue that in a typical semi-industrialised economy, inputs for manufacturing are largely imported and working capital from banks is subject to rationing. In this context, a sudden devaluation will sharply increase firms' input costs and the need for working capital. As additional funds required by firms may need to be obtained in the informal loan market at high interest rates, these effects may offset the positive impact of lower export prices and firms may choose to reduce production.

According to recent empirical evidence, contractions in output are frequently preceded by overvaluation of the real exchange rate, with positive growth episodes accompanied by appreciation of the real exchange rate (see Agénor, 1991; Kiguel and Leviatan, 1992; Pérez-López Elguezabal, 1995; Razin and Collins, 1997; Papazoglou, 1999). Within Latin America, it appears that abrupt devaluations of the exchange rate are associated with recessions (Edwards, 1995a, 1995b; Kamin and Rogers, 2000). Using the same Latin American countries as the present paper and a panel approach, Ahmed (2003) also finds real exchange rate depreciations to have negative effects on growth.

This paper examines the effect of changes in the real exchange rate on growth in six Latin America countries utilising nonlinear smooth transition regression (STR) techniques.

This approach allows sufficient flexibility to capture possibly asymmetric effects and hence to explore whether real depreciations versus real appreciations, or the magnitudes of these, have different effects. Through the STR methodology we seek to distinguish two "regimes" in the growth rate of the economy. In the present application, we find changes in the real exchange rate to be the indicator of regime switches for five of the six countries, namely Argentina, Brazil, Chile, Columbia and Venezuela. Indeed, in most cases we find that the "regimes" we identify can be associated with appreciations versus depreciations in the real exchange rate, implying that changes in the real exchange rate has asymmetric effects. Mexico is the exception, where the regime indicator in our preferred model is lagged GDP growth.

The rest of this paper is structured as follows. Section 2 outlines the variables used, including the economic model that underlies the choice of these, while Section 3 reviews the STR methodology we apply. Substantive empirical results are presented and discussed in Section 4, including the estimated models and dynamic multipliers for the effects of appreciations and depreciations on growth. Concluding remarks are offered in Section 5.

2. Variables and Economic Motivation

The dependent variable is the annual rate of growth of output, measured by per capita gross domestic product (GDP). In addition to the depreciation in the real exchange rate, we use lags of the growth rate and of inflation to control for the domestic economic situation and US interest rates to allow for international influences. The choice of the included variables is motivated by the recent analysis of Kamin and Rogers (2000) for Mexico.

The real exchange rate, e, is defined as

$$e = E \times \frac{P_T^{US}}{P_N} \qquad (1)$$

where E is the nominal exchange rate, measured as the number of units of local currency per US dollar, P_T^{US} is the foreign denominated price index of traded goods (measured by the US wholesale price index)[1] and P_N is the domestic price index for non-traded goods (measured by the domestic consumer price index). Issues in the definition of the real exchange rate are discussed at some length in Edwards (1989), with the definition in (1) being widely used in empirical applications. An increase in e represents a real depreciation of the local currency relative to the US dollar.

As a framework for the empirical models of the effect of the real exchange rate on output, our discussion follows Kamin and Rogers (2000). Real output, y, can be defined through the simple identity

$$y = d + x \tag{2}$$

where d and x represent domestic demand and net exports, respectively. Domestic demand depends on many factors, including financial variables such as bank credit and the rate of interest. Substituting these out to focus on real output, the domestic rate of inflation (π) and the real exchange rate, we have

$$d = f_d(y, \pi, e, i^{US}) \tag{3}$$

The US interest rate, i^{US}, enters through its negative effect on net capital inflows. These inflows have been particularly important for growth in Latin America during the 1990s (Calvo, Leiderman and Reinhart, 1996). Since the usual positive effect of devaluation through exports is excluded from (3), the effect of higher prices of foreign goods on domestic demand is anticipated to be negative.

Net exports are determined by the real exchange rate and real domestic GDP:

$$x = f_x(e, y) \tag{4}$$

[1] Since exchange rates relative to the US dollar are often available for a longer period than other exchange rate series, the use of the rate relative to the US dollar implies that a longer data series is available than when, say, a

with the anticipated sign for e being positive. Net exports may also depend on demand factors, with a positive effect anticipated for world real income (for example, Agénor, 1991). However, our models do not include such a variable as it was not statistically significant in our estimated equations[2].

Substituting (3) and (4) into the identity of (2) yields

$$y = f(e, \pi, i^{US}) . \qquad (5)$$

In this expression, the sign of the real exchange rate is ambiguous, due to the different signs anticipated in (3) and (4). The literature of the effect of the real exchange rate on output is reviewed well by Agénor (1991), Kamin and Rogers (2000) and others. Although there have been relatively few econometric studies to date, Kamin and Rogers (p.94) summarise these by: "the econometric analyses indicate almost uniformly that devaluations lead to reduced output", implying that $\partial y/\partial e < 0$ in (5), and hence that the effects of the real exchange rate on domestic demand through (3) dominate those that operate through the net exports function of (4). Nevertheless, this finding remains controversial.

Our study estimates (5) as a reduced form equation; details are given in Section 4. However, a number of important questions need to be tackled prior to estimation. Firstly, there is the nature of the functional form assumed for f. In practice, the empirical literature (including Edwards, 1989, Kamin and Rogers, 2000) measures real income and possibly the real exchange rate in logs and otherwise assumes linearity. However, linearity is merely a convenient assumption and, as argued above, real exchange rate effects on output may be inherently nonlinear. In order to allow such effects, our empirical specification permits a nonlinear functional form for f.

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basket of currencies is employed.

[2] More precisely, we included growth in US per capita income as a proxy for world income growth. However, since this variable was not significant in the linear model for growth in any of our six countries, it is not included in the models reported.

A second practical issue concerns the question of whether the variables in (5) should be differenced. Kamin and Rogers (2000)[3] difference all variables, due to "near unit root behaviour in each series" (p.105). Alternatively, in a related specification, Edwards (1986) uses levels with the addition of a time trend. In order to focus explicitly on economic growth, we use the percentage growth in per capita GDP as the dependent variable in our models. It is widely agreed in the applied macroeconomics literature that the growth rate is stationary, whereas the level of GDP is not.

Whether the remaining variables included in a growth regression should be differenced is not a simple question. Statistically, unit root tests on the real exchange rate, inflation and interest rates do not always give clear answers. For the real exchange rate, we use the rate of depreciation (or appreciation) in preference to the level in order to eliminate the apparently nonstationary trending behaviour exhibited by some Latin American countries over our sample period. Both the level and difference of US interest rates were investigated, with essentially similar results. The reported results are based on the level, so that (through the estimation of coefficients) the model can select the difference where appropriate. Finally, we use the logarithm of domestic inflation. We prefer this to the change in inflation, since recent empirical studies of growth in Latin America have almost uniformly found the level of inflation to negatively impact on growth; see, among others, De Gregorio (1992a, 1992b), Easterly, Loayza and Montiel (1997). We have found taking the logarithm improves the models, since this reduces the otherwise extreme values observed in periods of very high inflation.

Annual data are employed on all variables. The sample period used ends in 2000, while typically the sample starts during the 1950s. Appendix 1 contains details of the data used, including the precise sample period for each country. The data series for the variables of primary interest to us, namely per capita GDP growth (Δy_t) and the rate of depreciation (Δe_t),

[3] It appears that Agénor (1991) may also use differenced data. Although no statement to this effect is made in the

are shown for all six countries in Figure 1. At a superficial level from an examination of these graphs, it appears that large real depreciations may be associated with negative growth episodes; for example, in 1975 for both Argentina and Chile, or 1995 for Mexico.

3. Smooth Transition Regression Methodology

Since the 1980s there has been an increasing interest in the use of nonlinear models to capture the dynamics of growth in developed economies. Typically, the emphasis in these studies has been on business cycle recessions and expansions in the tradition of Burns and Mitchell (1946), with Hamilton (1989), Teräsvirta and Anderson (1992) and Potter (1995) being prominent examples. However, this kind of analysis has only recently begun to be undertaken for developing economies (Greenaway, Leybourne, and Sapsford, 1997, Chen and Lin, 2001). For the Latin American economies, some authors have argued that business cycles can be characterised as exhibiting nonlinear behaviour arising from the existence of asymmetric dynamics over the business cycle (Mora, 1997; Mejía-Reyes, 1999, 2000). However, these papers have focused on measuring and modelling asymmetries, rather on than explaining the dynamics of growth or the causes of recession. Sensier, Osborn and Öcal (2002) have recently applied this methodology to the UK, modelling quarterly output growth in terms of lagged changes in short-term interest rates.

In this section we introduce our STR modelling methodology, with further information on model specification, estimation and diagnostic checking presented in Appendix 2. This section ends with an outline of the dynamic analysis used in Section 4 to examine the properties of the models.

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text, the explanatory variables (except for the real exchange rate) in his Table 1 are all expressed as differences.

3.1 STR Modelling

The STR model we utilise can be defined as follows:

$$\Delta y_t = \alpha_0 + \sum_{i=1}^{n} \beta_i x_{ti} + F(z_t)\left(\alpha_1 + \sum_{i=1}^{n} \eta_i x_{ti}\right) + \varepsilon_t \qquad (6)$$

where the dependent variable (Δy_t) is the annual growth rate in per capita real GDP, x_{ti} are the observations on n explanatory variables (i = 1, …, n), ε_t is an independent and identically distributed disturbance, with mean zero and variance σ^2, while $F(z_t)$ is a transition function between regimes. Nonlinearity is captured through $F(z_t)$, which is defined as a function of an explanatory variable, conveniently denoted by z_t. This function F is bounded, $0 \leq F \leq 1$, with the extremes of $F = 0$ and $F = 1$ corresponding to distinct "regimes". Within each regime, a different linear relationship applies between y_t and the explanatory variables. For example, when $F = 0$, the variable x_{ti} has coefficient β_i, whereas when $F = 1$, the coefficient of x_{ti} becomes $\beta_i + \eta_i$. Similarly, the intercept in (6) is α_0 when $F = 0$ and $\alpha_0 + \alpha_1$ when $F = 1$. Intermediate values of F define situations where the model is a mixture of the linear models corresponding to these two regimes. These models are now widely in a univariate context, for which van Dijk, Teräsvirta and Franses (2002) provide a review. Teräsvirta (1998) discusses the regression counterpart we employ in (6).

We do not specify *a priori* the transition variable z_t that determines the regimes in (6). Nevertheless, for plausible transition variables such as real depreciation or lagged growth, we anticipate regimes associated with high versus low values of z_t; this could, for example, reflect regimes associated with the business cycle, as in Sensier *et al.* (2002). Therefore, following the usual approach in this STR literature, we define F through the logistic function:

$$F(z_t) = \frac{1}{1 + \exp\{-\gamma(z_t - c)\}}, \qquad \gamma > 0 \qquad (7)$$

where γ is the slope of the transition function, and c is the threshold parameter that indicates its location in relation to observations on z_t. One attraction of the logistic function is that it is a monotonically increasing function of z_t. Therefore, the value of F increases with z_t, so that (depending on the values of γ and c), small z_t yield F close to zero and large z_t in F close to one. At the location parameter value, $z_t = c$, then $F = 0.5$.

A crucial issue in applying (6) and (7) in practice is the selection of the transition variable, z_t. As explained in more detail in the Appendix, our procedure for selecting z_t utilises a search over all explanatory variables (including lags), in order to find the one that yields the lowest residual sum of squares in (6). Subsequent to this selection, and in order to produce a reasonably parsimonious model, we drop redundant explanatory variables using the Akaike Information Criterion (AIC). Estimation of the final model consisting of (6) and (7) is undertaken by nonlinear least squares.

Various statistics relating to the estimated models are presented. These include the Akaike information criterion (AIC) and the Schwarz information criterion (SIC), as well as the residual standard error and the conventional R^2 for model comparison purposes. Other diagnostic statistics are discussed in Appendix 2 and presented in the detailed Appendix tables.

3.2 Dynamic Analysis

In order to analyse the dynamic implications for growth implied by our models, we compute dynamic multipliers for the effects of depreciation and appreciation in the real exchange rate. Since our models are reduced form equations, the use of dynamic multipliers aids their interpretation. However, because the STR model is nonlinear, the responses to (say) a 10 percent depreciation and a 10 percent appreciation are not necessarily symmetric. Further, the effects of these changes are (in general) state-dependent, so that the values taken by other explanatory variables can play a role in the estimated responses to changes in the real exchange rate. When the transition variable is endogenous, the computation of dynamic multipliers requires the use of simulation techniques, developed for the computation of impulse response functions in nonlinear models by Koop, Pesaran and Potter (1996) and Potter (1998). However, in our context, this applies only when a lagged value of the dependent variable is the transition variable. This is the case only for Mexico, and calculation of the dynamic multipliers for a real depreciation/appreciation in the two regimes for Mexico are discussed in Appendix 2. With this exception, z_t in our models is exogenous. Since the regime is then also exogenous, we do not have to resort to simulations for other countries.

In order to concentrate on the impact of changes in the real exchange rate on growth, when calculating dynamic multipliers we generally assume that the control variables (the inflation rate and the US interest rate) take values equal to their mean over the sample period used for model estimation. Similarly, the lagged growth rates needed to initialise the model are set at their sample mean. The exception again is Mexico, where the lagged growth rate generates the initial regime.

To examine the effects of a depreciation or appreciation, we compute three sets of dynamic forecasts for the per capita GDP growth rate. The first “baseline” set specifies changes to the real exchange rate as zero. The two further sets assume a given rate of depreciation or appreciation (as appropriate). By subtracting the baseline forecasts from these values, we obtain the estimated dynamic multiplier effects.

In practice, our forecasts for comparison with the baseline case assume a ten percent depreciation or appreciation for each of the first three periods, with no further changes in the real exchange rate thereafter. While we could examine the effect of a one-off change, with a nonzero depreciation/appreciation only in the first period, this would not capture all the dynamic interactions between the transition function and lagged changes in real exchange rates. To reflect these interactions, we prefer to examine the dynamic effects on growth of a sustained depreciation that takes place for three years, rather than a one-off depreciation/appreciation. Dynamic multipliers are reported for the period of the initial change and ten subsequent periods.

As the model for Mexico implies endogenous regime switching, in this case we also explore these nonlinear dynamics through the computation of generalised impulse response functions, as in Koop *et al.* (1996) and Potter (1998). It should be noted that although estimated effects are dynamic, all are based on single equation models and consequently do not allow for feedback between, for example, inflation and changes in the real exchange rate.

4. Empirical Results

We model the relationship between annual percentage growth in per capita GDP (Δy_t) and the percentage change in the real bilateral exchange rate with the US (Δe), with inflation (represented by π after taking logarithms) and the US interest rate as control variables, as explained in Section 2. As Δe_t measures real exchange rate depreciation, a negative value represents an appreciation. Because it is important to capture the dynamics of growth, we use up to two lags of all variables, including the dependent variable. In addition to these variables used for all countries, we also include the growth in oil production (Δo_t) as an explanatory variable for Mexico and Venezuela due to its importance in these economies, but only the current value is included in order to conserve degrees of freedom.

International variables are treated as exogenous to domestic growth, so that current values of US interest rates and the change in the real exchange rate are permitted. For this purpose, oil production is also treated as exogenous on the grounds that it depends primarily on international demand. Although the real exchange rate is influenced by domestic as well as international factors, we take the view that past (rather than current) performance of the domestic economy plays the main role in determining real exchange rate changes, so that these can be treated as predetermined. Further, Kamin and Rogers (2000) conclude that causality runs from the real exchange rate to growth, at least in the case of Mexico[4]. Only lagged values of inflation are used to avoid endogeneity problems with this variable.

We consider linear models, before moving on to our main results employing nonlinear STR models. For the latter results, it is convenient to separate the discussion for the oil producing countries of Mexico and Venezuela from those for the non-oil producing countries.

4.1 Linear Models

Linear model results for each country (Argentina, Brazil, Chile, Columbia, Mexico and Venezuela) are reported in Table 1, where the specific explanatory variables included have been selected according to AIC. Values shown in parentheses are *t*-ratios.

The implications of the linear models for the contemporaneous effects of a depreciation are clear. Where current depreciation enters the model (that is, for Argentina, Chile, Columbia and Mexico), the corresponding coefficient is negative, implying that depreciation causes a reduction of growth in the current year. However, when lags are considered, the picture is less clear. Although depreciation has negative lagged coefficients for Argentina and Mexico, the coefficients for Brazil and Venezuela are positive. In the case of Columbia, the initial negative impact is partially redressed by a positive coefficient for lagged depreciation. Overall, therefore, there appears to be a set of countries where depreciation has a negative effect on growth, with this being evident within the current year. Where this is not the case (Brazil and Venezuela), the delayed effect of depreciation is to enhance growth.

Turning briefly to the remaining explanatory variables, inflation enters for Brazil, Mexico and Venezuela, with a negative effect in each case. US interest rates are included for

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[4] We also computed bivariate causality tests for growth and changes in the real exchange rate for the six Latin American countries. Using a ten percent significance level, only Chile provided evidence of causality running from growth to depreciation. As usual, these tests were computed using lagged values only.

four countries (the exceptions being Chile and Columbia). However, in the case of Mexico the coefficient of the current rate is positive (and strongly significant) rather than negative as anticipated from the model of Kamin and Rogers (2000). Although they do not comment on this feature, it is notable that the linear impulse response functions plotted by Kamin and Rogers (Figure 3) also imply a positive dynamic effect of US interest rates on growth in Mexico. Ahmed (2003) also finds an important effect on growth in these countries from the US real interest rate. Dynamics of growth appear through the lagged dependent variable in four models, with the effect generally being negative and at a lag of two years. When oil production is included for Mexico and Venezuela, its impact on growth is illustrated by the significant positive coefficient.

With the exception of Mexico, the explanatory power of these linear models, as measured by R^2, is relatively modest. This is especially so in the cases of Brazil, Chile and Columbia. There are some indications of possible nonlinearity or outliers in the significant non-Normality of the residuals for Chile and Columbia. In general, however, the conventional diagnostics for autocorrelation and heteroscedasticity are satisfactory.

As explained in Appendix 2, we also explicitly test for possible STR nonlinearity, considering each explanatory variable in the initial model as the possible explanatory variable, with the results shown in Appendix Table A.1. At a conventional 5 percent significance level, four of the six countries show evidence of nonlinearity (the exceptions being Brazil and Columbia), with the nonlinearity for Chile, Mexico and Venezuela being particularly marked and associated with more than one of the variables of the model. Although we interpret these results as indicative of possible nonlinearity, we believe that the test may lack power due to the relatively small sample sizes available, and hence we investigate nonlinear models for all six countries.

4.2 Nonlinear Models: Non-Oil Producing Countries

Prior to estimating the nonlinear STR model (6), the transition variable z_t needs to be specified. As explained in Appendix 2, our procedure is to search over all explanatory variables (including lags) considered in the general linear models, and examine each in turn as the possible z_t. For the non-oil producing countries of Argentina, Brazil, Chile and Columbia, current or lagged depreciation is selected as the transition variable. More specifically, the two year lag of depreciation is selected for Argentina, but the current one in the remaining three cases. However, in the case of Argentina, our modelling procedure leads to a highly parameterised model with 19 estimated coefficients, which we consider unreliable given the relatively small sample size available. Therefore, for this case, the regime-dependent coefficients are restricted to apply only to the variables of central interest, namely lagged growth and (current or lagged) depreciation, with other $\eta_i = 0$ in (6). Current depreciation is selected as the transition variable for Argentina, with this selection taken over lagged growth and current or lagged depreciation. Therefore, all models reported in Table 2 have transition variable $z_t = \Delta e_t$.

Although the STR models are specified and estimated separately for each country, there is a remarkable similarity in the estimated transition functions, shown in Figure 2 in terms of observed values of Δe_t. Specifically, the logistic transition functions for all four countries are centred very close to zero with a steep slope. This indicates that the regimes detected by the nonlinear model relate to depreciations, with $F(z_t) = 1$, versus appreciations, $F(z_t) = 0$. In other words, there appear to be asymmetric responses of growth to positive and negative Δe, with one linear model applying in periods of depreciation and a different model applying in periods of appreciation. Our procedure does not impose this or even impose depreciation as the transition variable; rather this outcome is selected as the "best fit" nonlinear model for these countries. The only exception is a partial one for Argentina, for which the transition function is smoother than in other cases and is centred below zero, at the relatively small appreciation of 3.2 percent.

Table 2 presents the estimated models for these countries separately for the two regimes of $F(z_t) = 0$ and $F(z_t) = 1$, with detailed estimation results in Appendix Table A.2. The absence of any estimated coefficient for Δe in the cases of Brazil and Columbia within the appreciation regime implies that the magnitude of an appreciation has no effect on growth in these countries. However, appreciation has a negative contemporaneous effect in both Argentina and Chile, which is partly off-set in the former case by a lagged value of the opposite sign. (Note that since the explanatory variable Δe_t is the rate of depreciation, an appreciation corresponds to a negative value of this variable.) Turning to the depreciation regime, only for Chile and (to a very small extent, Argentina) is the amount of depreciation found to have a negative impact on growth in the current year. With the exception of Argentina, and still within the depreciation regime $F = 1$, the extent of a previous depreciation has a positive effect on growth (as indicated by the positive coefficients for Δe_{t-1} or Δe_{t-2}). Thus, the negative effect of depreciation may be short-lived.

Nevertheless, such a simple analysis based on looking at the coefficients of the depreciation variable is fraught with difficulty, because these coefficients do not take full account of the dynamics. Further, the estimated intercept is regime-dependent for Argentina and Columbia, and this also affects growth rate comparisons over regimes. To analyse the depreciation effects more fully, we later undertake a dynamic multiplier analysis.

It is notable that international conditions, captured here by US interest rates, play an important role in these models. Further, the coefficients can change between the depreciation and appreciation regimes. Similar comments apply to the role of inflation and lagged growth generally differing over regimes. This indicates that the impact of a depreciation (or appreciation) on growth will depend on the national and international environment applying at the time, although for the practical reasons explained above, we do not allow this flexibility in the model for Argentina.

The diagnostic tests for the estimated nonlinear models (see Appendix Table A.3) show that the earlier evidence of nonlinearity for Argentina and Chile is satisfactorily accounted for by our models. For Chile, in particular, this is notable, since there are strong indications of nonlinearity in Appendix Table A.1. The other diagnostics for the nonlinear models are also generally satisfactory, despite strong evidence of non-Normality and some indication of change over time in the intercept of the model for Columbia.

4.3 Nonlinear Models: Oil Producing Countries

Applying the modelling methodology of Appendix 2 leads to some difficulties in the cases of Mexico and Venezuela. When all parameters are allowed to change over regimes, the first lag of the US interest rate is selected as the transition variable z_t for both countries. However, the models were not economically plausible[5]. We attribute these difficulties to the relative small sample sizes available in order to estimate these models, and hence restrict the models in the same manner as adopted above for Argentina. Thus, the results presented in Table 3 and Appendix Table A.2 are based on models that allow regime-dependent coefficients only for growth and depreciation, with the transition variable also chosen from these variables. In the case of Mexico, the selected transition variable is the second lag of growth in per capita GDP, while for Venezuela it is the second lag of depreciation. In this later case, the transition function is very steep and centred at a value close to zero, again implying different responses to appreciations and depreciations, although now after a lag of two years; see the first panel of

[5] The estimated model for Mexico strengthened the apparently perverse finding (see Table 1) of higher US interest rates leading to higher growth, while that for Venezuela implied that high US interest rates (above 6.1%) lead to a regime with unstable oscillations in growth, due to a lagged dependent variable coefficient of –1.1.

Figure 3. Transition between regimes for Mexico is also steep, with past per capita growth rates above and below 3.4 percent implying different relationships.

Turning to the estimated coefficients of Table 3, the effects of depreciation are not regime-dependent in the case of Mexico, with both current and past depreciation having negative effects on growth. For Venezuela, although the coefficient is again constant over regimes, depreciation has a positive effect on growth after one year. Indeed, the growth effects of a depreciation in Venezuela appear to be enhanced by the estimated intercept being larger in the depreciation regime.

Inflation has negative effects on growth after a lag of one year in the case of Mexico, but no role of inflation is found for Venezuela. Not surprisingly, the nonlinear models continue to show that increases in oil production are good for per capita GDP growth. The positive coefficient of US interest rates for Mexico, noted above for the linear model, remains in Table 3. However, US interest rates have negative effects in the case of Venezuela. It should also be noted that within the high growth regime for Mexico, the coefficient on the two year lag of growth is marginally greater than unity and the dynamics of growth within this regime are consequently unstable. However, this does not necessarily imply instability for growth once endogenous changes in regime are considered, as in the dynamic analysis below.

Finally, the nonlinear models for both Mexico and Venezuela are satisfactory, despite some indication of non-constancy over time for the intercept for Mexico (Appendix Table A.3). The satisfactory diagnostics and lack of evidence of nonlinearity is notable especially for Mexico, since the linear model shows significant nonlinearity in relation to US interest rates (Appendix Table A.1), while these are restricted to have a constant role over regimes in Table 3.

4.4. Dynamic Analysis

The model for Mexico is the only case where regime changes depend on lagged growth, and hence are endogenous in our single equation context, and we begin our dynamic analysis with an investigation of these implied nonlinear dynamics. Figure 4 shows the generalised impulse responses of this model to a disturbance shock of ±1 and ±2 residual standard deviations, with moderate/negative current growth being considered separately from the case of a high growth rate. When the initial regime is the "normal" one of moderate or negative growth, the impulse response is close to being linear. However, at a horizon of two years there is evidence of asymmetry to positive and negative shocks, which arises because the shock at t may be sufficiently large and positive to cause a switch to the high growth regime ($F = 1$) for period $t + 2$. When such a switch occurs, the lower intercept in that regime takes effect, resulting in larger reductions in growth at the horizon of 2 in comparison with the positive responses at this horizon to negative shocks in period 0.

For the high growth regime, note first that the dynamic responses are stable with endogenous regime switching, despite the apparently unstable coefficient on lagged growth in Table 3. The second point to note is the clear asymmetry between large positive and negative disturbance shocks of ±2 standard deviations, with both implying positive growth after two periods. Overall, the model implies that shocks tend to enhance growth in the high growth regime, but not in the moderate/negative growth one.

The dynamic multipliers for the estimated nonlinear models for all countries are shown in Figure 5. As noted above, with the single exception of Mexico, the regimes can be associated with appreciations and depreciations in the real exchange rate. As discussed in Section 3.2, in order to capture the dynamic responses of growth to changes in the real exchange rate, we consider the effect of a depreciation (appreciation) that continues for three years at the rate of 10 percent per year.

The results imply that the effects of a depreciation are severe for some countries. For instance, in the case of Argentina, the impact of a 10 percent depreciation is seen immediately in a reduction in growth of more than three percentage points, which increases to a four percent reduction after a year. If the depreciation is sustained over three years, growth does not rise above the baseline level (associated with no change in the real exchange rate) until three years after the initial depreciation. Notice that although the model for Argentina is asymmetric, this asymmetry is not very marked in practice despite Δe_t being the transition variable. In addition to Argentina, substantial negative effects of depreciation are seen in Figure 5 for Chile and Mexico, with lesser negative effects for Columbia.

The case of Columbia is interesting, because although the pattern of response to a sustained depreciation is similar to (though less marked than) Argentina, there is effectively no response of growth to an appreciation. It is only in the case of Brazil that the dynamic multipliers suggest that a depreciation unambiguously increases growth, while an appreciation decreases it. Although as in Venezuela there is an initial increase in growth after a depreciation, this is later partially reversed. Real appreciations in both Brazil and Venezuela are associated with lower growth of around four percentage points after two years.

As already discussed, Mexico is the only case where regimes are not associated with changes in the real exchange rate. Indeed, the pattern for Mexico in Figure 5 is relatively straightforward, and effectively symmetric[6]. A real exchange rate depreciation is estimated to reduce growth, even at a horizon of ten years after the initial depreciation. Conversely, a real appreciation leads to higher growth of a similar magnitude.

It needs to be emphasised that the dynamic multipliers presented in Figure 5 are only indications of the responses implied by the nonlinear models to positive and negative changes

[6] The results shown in Figure 5 for Mexico relate to the case where the economy is initially at negative or moderate growth. The responses are, however, very similar to those shown when the initial growth corresponds to the case of “high growth”.

in the real exchange rate. Indeed, the reason for a particular outcome can lie in the apparently innocuous assumptions made. One such point is the implied growth response in Chile to a depreciation. It appears from the coefficient estimates of Table 2 that, while the initial effect of a depreciation in Chile will be negative, after a lag of two years positive effects will feed through. However, the effects in Figure 4 for a continued three year depreciation are always negative, with the initial impact of a 10 percent depreciation being negative and around 6 percentage points. In terms of the model, this large effect is due primarily to the US interest rate, which has an estimated overall negative effect in the depreciation regime, but an overall positive one in the appreciation regime. Since $F = 0$ in the baseline simulation with $\Delta e = 0$, then the effect of a depreciation shown in Figure 4 (compared with the baseline model) also reflects the differential effects of US interest rates in the two regimes.

5. Conclusions

In this paper we analyse the dynamics of the growth rate of real GDP per capita for six important Latin American countries using nonlinear smooth transition regression models focusing on the effects of changes in the real exchange rate on growth. Through this analysis we particularly hope to contribute to the literature by allowing the possibility that different growth effects may apply depending on the sign and magnitude of the real depreciation.

One strong result is that changes in the real exchange rate generally acts as the transition variable in our growth rate regressions. Indeed, without any imposition in the form of this function, the overall implication is that appreciations and depreciations act as different “regimes”, between which the nature of the economic determination of growth differs. The main exception is Mexico where past growth acts as the transition variable.

Our results imply that real depreciations have particularly severe contractionary consequences for Argentina and Chile in the short-run, while the effect for Mexico is not only negative but it persists even after a lag of ten years. Overall, we conclude that real depreciations have adverse consequences for growth in four countries (Argentina, Chile, Columbia and Mexico). In the cases of Brazil and Venezuela, the estimated effects of depreciations are positive for growth. However, in these countries, the asymmetric responses imply larger negative growth consequences of real appreciation than the positive effects of depreciation.

There are many directions in which the present analysis can be extended, in order to verify the central role found for real exchange rate depreciations for the growth rate regime and to uncover the causes for the apparently different directions of responses over countries to real depreciations. Our analysis has been restricted in terms of sample size and one obvious direction is to explore the use of higher frequency (quarterly or monthly) data. Another important extension is to model the joint determination of the key variables of growth and depreciation, together with the inflation rate. Meanwhile, we believe that our study has emphasised the importance of real exchange rate depreciations and that a linear framework may be too restrictive to satisfactorily capture their effects on growth.

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Table 1. Estimated Linear Models for Per Capita GDP Growth

Coefficient	Argentina	Brazil	Chile	Colombia	México	Venezuela
Intercept	7.070 (3.500)	9.864 (4.130)	2.730 (3.911)	1.224 (2.910)	3.121 (4.611)	4.843 (3.371)
Δy_{t-1}				0.246 (1.660)		
Δy_{t-2}	-0.263 (-1.936)				-0.122 (1.308)	-0.219 (1.631)
Δe_t	-0.037 (-2.257)		-0.135 (-3.836)	-0.084 (-2.714)	-0.147 (-6.339)	
Δe_{t-1}	-0.037 (-2.292)			0.051 (1.592)	-0.032 (-1.596)	0.088 (2.606)
Δe_{t-2}		0.079 (1.388)				
π_{t-1}					-1.152 (-5.333)	-0.452 (-1.714)
π_{t-2}		-0.520 (-1.399)				
i_t^{US}	-0.843 (-2.882)				0.670 (3.049)	-0.615 (-2.844)
i_{t-1}^{US}		-0.763 (-2.990)			-0.440 (-1.978)	
i_{t-2}^{US}						
Δo_t	N/A	N/A	N/A	N/A	0.132 (2.758)	0.437 (2.654)
Goodness-of-fit measures						
s	3.809	3.352	4.557	2.095	1.740	2.923
R^2	0.44	0.27	0.25	0.22	0.70	0.48
AIC	2.924	2.648	3.118	1.649	1.441	2.438
BIC	3.144	2.826	3.197	1.807	1.753	2.689
Diagnostic tests (p-values)						
Normality	0.484	0.556	0.004	0.000	0.701	0.602
Autocorrelation	0.969	0.729	0.549	0.947	0.857	0.099
Heteroscedasticity	0.343	0.278	0.871	0.834	0.514	0.814

Note: Numbers in parentheses are t-statistics.

Table 2. Estimated Nonlinear Models for Per Capita GDP Growth in Non-Oil Producing Countries

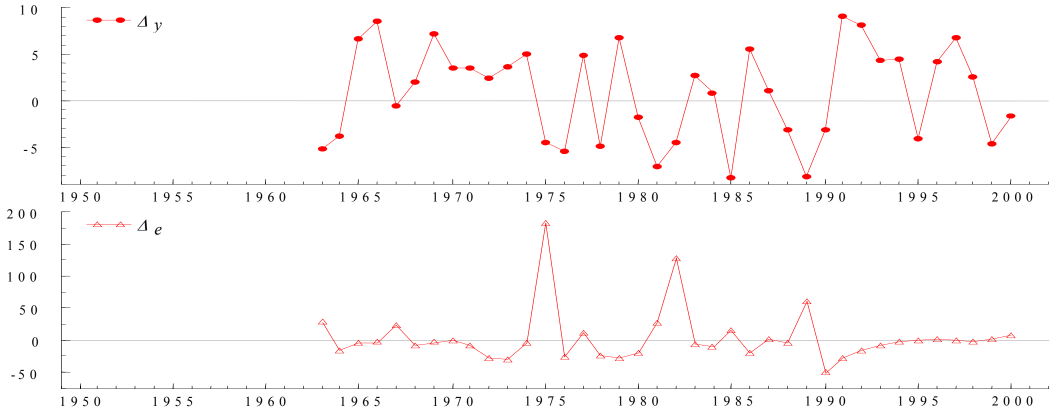
Coefficient	Argentina		Brazil		Chile		Columbia	
	Large appreciation ($F = 0$)	Small appreciation/ Depreciation ($F = 1$)	Appreciation ($F = 0$)	Depreciation ($F = 1$)	Appreciation ($F = 0$)	Depreciation ($F = 1$)	Appreciation ($F = 0$)	Depreciation ($F = 1$)
Intercept	13.87	2.020	6.011	6.011	4.246	4.246	1.645	7.504
Δy_{t-1}	-0.477	0.060		0.500		-0.290	0.514	-0.085
Δy_{t-2}		-0.571				-0.313		
Δe_t	0.260	-0.005			0.174	-0.064		
Δe_{t-1}	-0.057	-0.057						
Δe_{t-2}				0.158		0.139		0.209
π_{t-1}				3.205	-3.752	-3.752		-4.527
π_{t-2}				-0.202	3.021	3.021	-0.280	0.457
i_t^{US}	-0.632	-0.632	-1.367	1.599	-0.828	3.166	0.409	0.409
i_{t-1}^{US}				-3.473	1.451	-2.264	-1.302	0.522
i_{t-2}^{US}			0.706	0.706		-1.054	0.941	-0.276
z_t	Δe_t		Δe_t		Δe_t		Δe_t	
Goodness of fit measures								
s	3.516		2.065		2.863		1.817	
R^2	0.659		0.81		0.80		0.62	
AIC	2.761		1.702		2.358		1.448	
SIC	3.245		2.191		2.949		2.039	

Table 3. Estimated Nonlinear Models for Per Capita GDP Growth in Oil Producing Countries

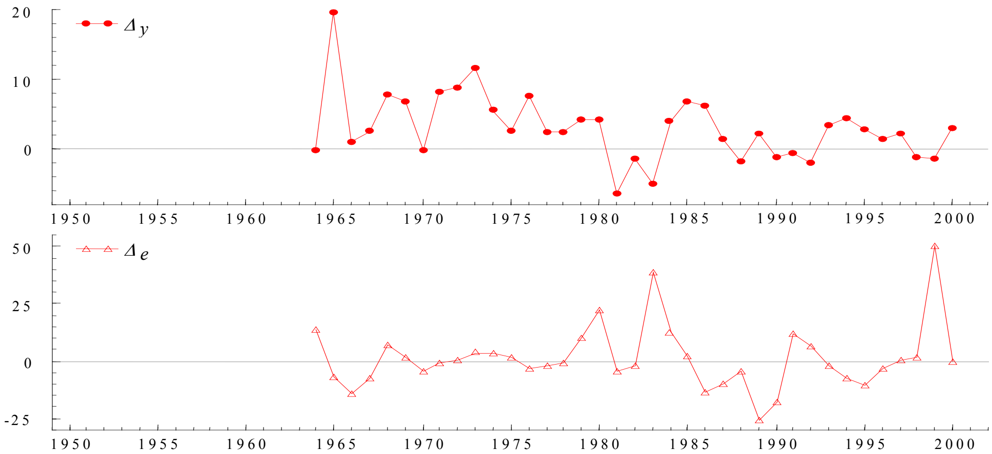
Coefficient	México		Venezuela	
	Moderate/ negative growth ($F = 0$)	**High growth ($F = 1$)**	**Large appreciation ($F = 0$)**	**Small appreciation/ Depreciation ($F = 1$)**
Intercept	3.157	-4.540	3.158	6.291
Δy_{t-1}	0.0972	0.0972		-0.205
Δy_{t-2}		1.020	-0.174	-0.174
Δe_t	-0.165	-0.165		
Δe_{t-1}	-0.036	-0.036	0.123	0.123
Δe_{t-2}				
π_{t-1}	-1.097	-1.097		
π_{t-2}				
i_t^{US}	0.202	0.202	-0.736	-0.736
i_{t-1}^{US}				
i_{t-2}^{US}				
Δo_t	0.201	0.201	0.448	0.448
z_t	Δy_{t-2}		Δe_{t-2}	
<u>Goodness of fit measures</u>				
s	1.667		3.075	
R^2	0.786		0.552	
AIC	1.217		2.438	
SIC	1.641		2.814	

Figure 1. Per Capita GDP Growth and the Rate of Depreciation

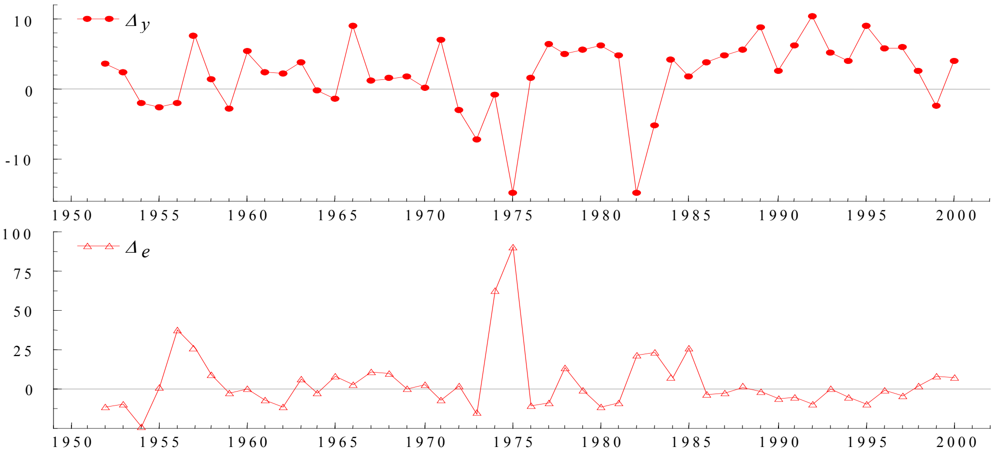
Argentina



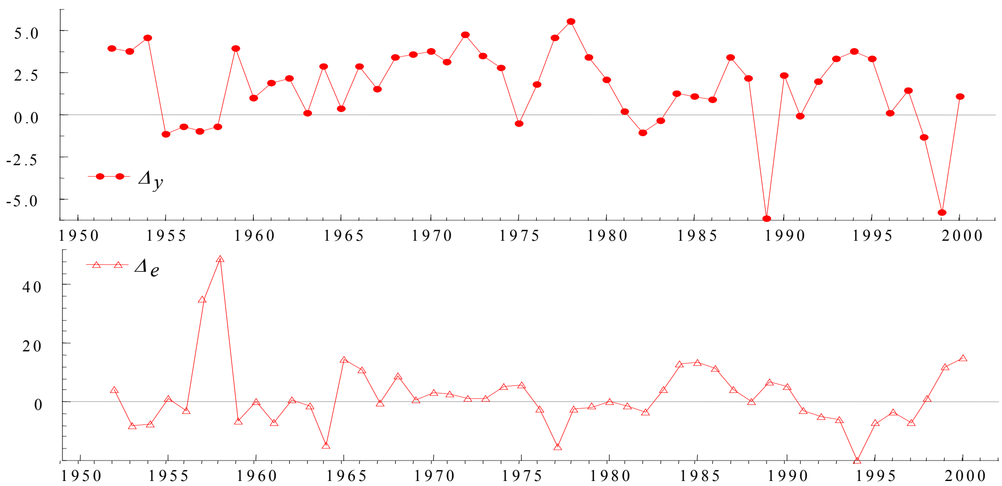
Brazil



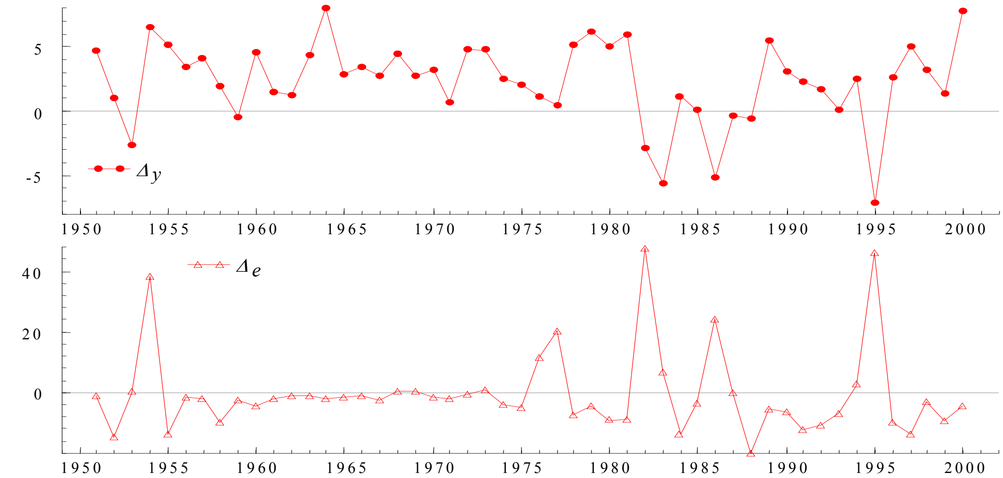
Chile



Colombia



Mexico



Venezuela

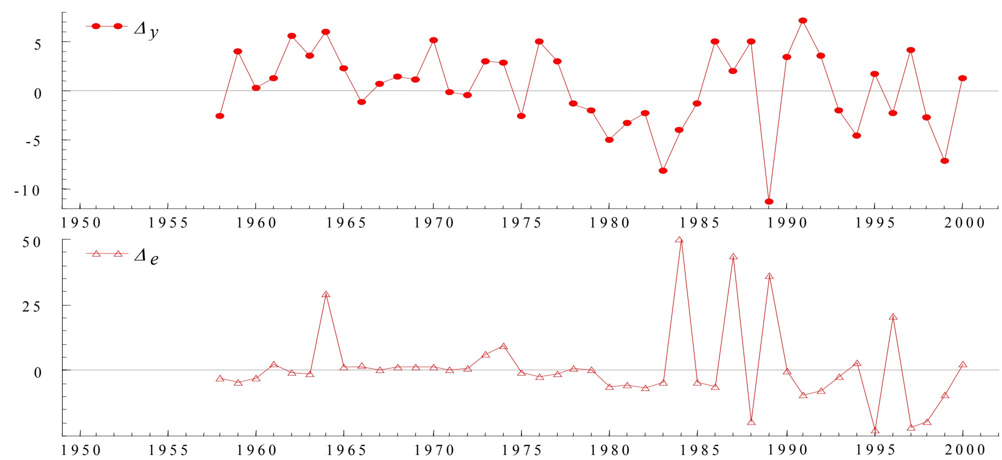
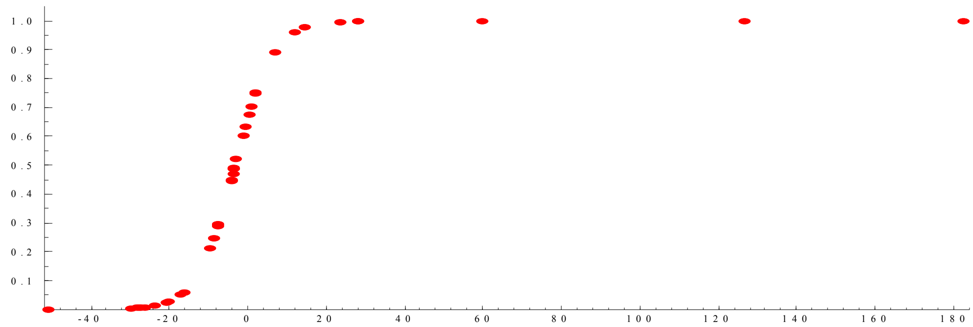
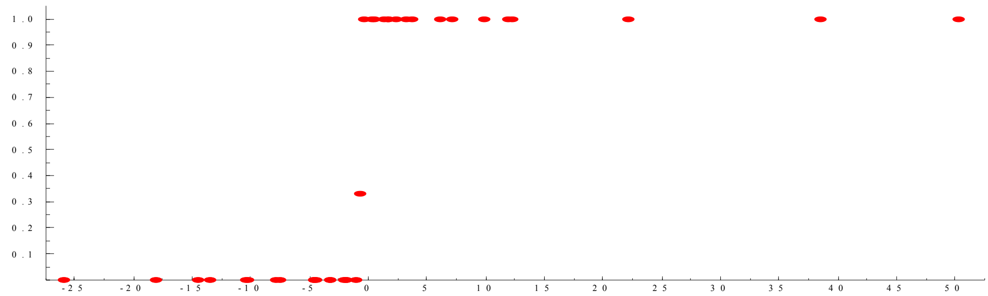


Figure 2. Transition Functions for Non-Oil Producing Countries

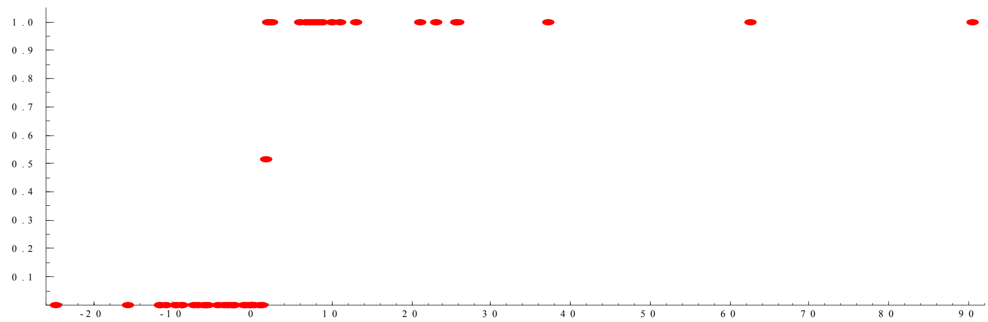
Argentina: *F* versus Δe_t



Brazil: *F* versus Δe_t



Chile: *F* versus Δe_t



Colombia: *F* versus Δe_t

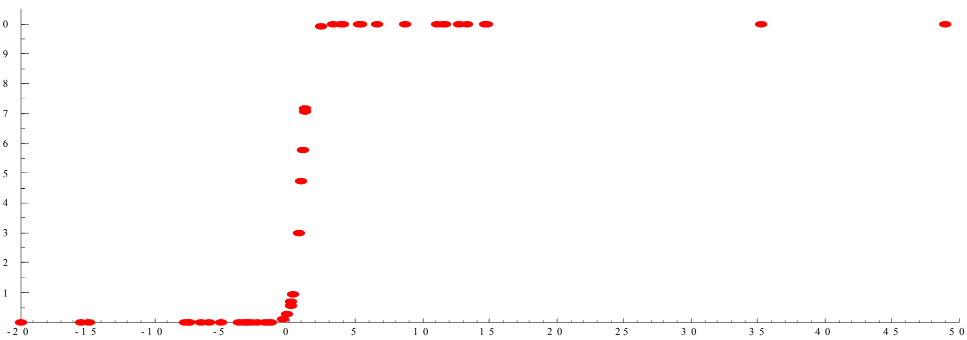
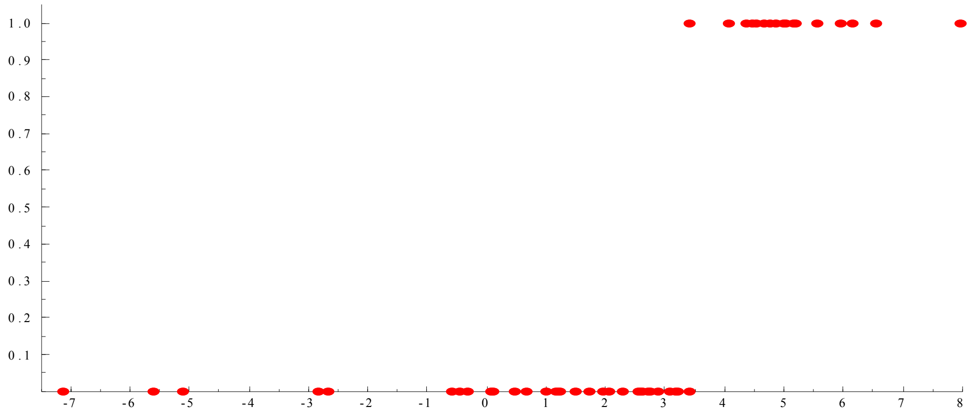


Figure 3. Transition Functions for Oil Producing Countries

Mexico: *F* versus Δy_{t-2}



Venezuela: *F* versus Δe_{t-2}

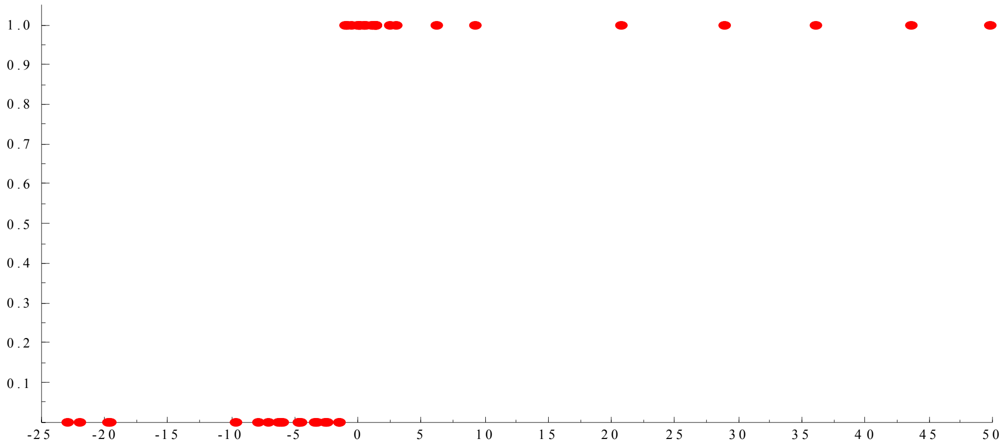
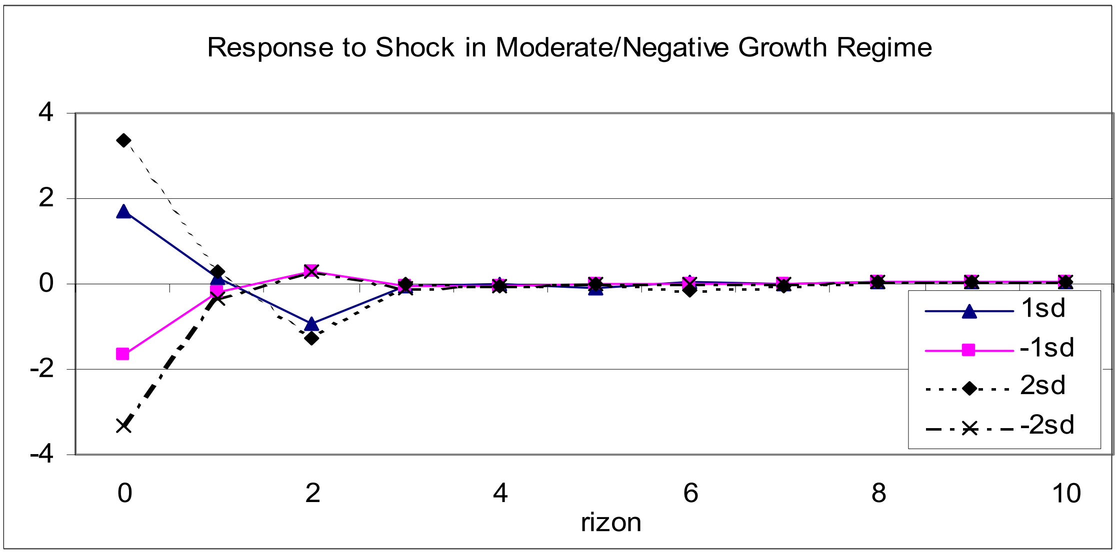


Figure 4. Generalised Impulse Response Functions for Nonlinear Model for Mexico



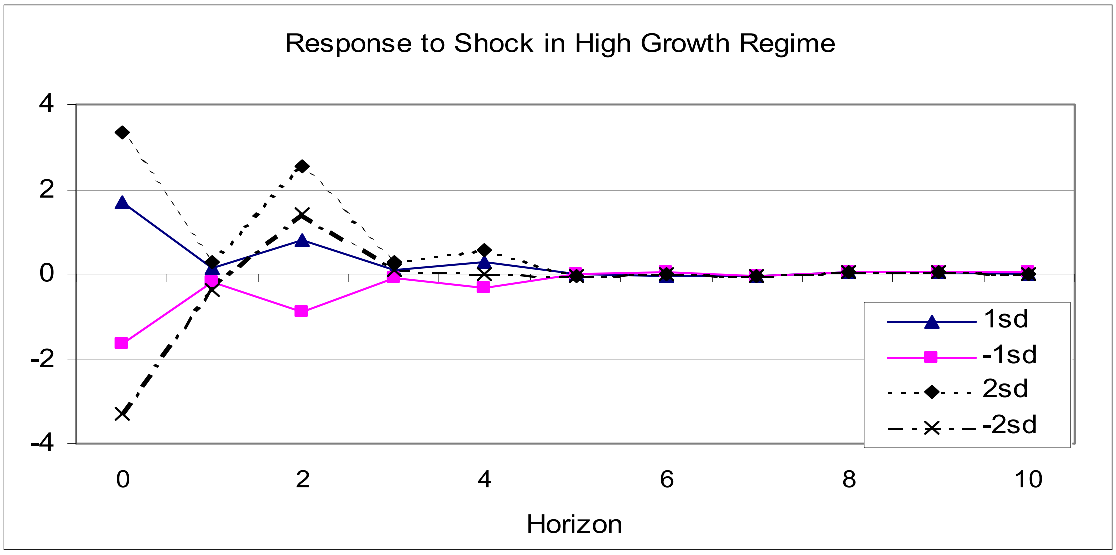
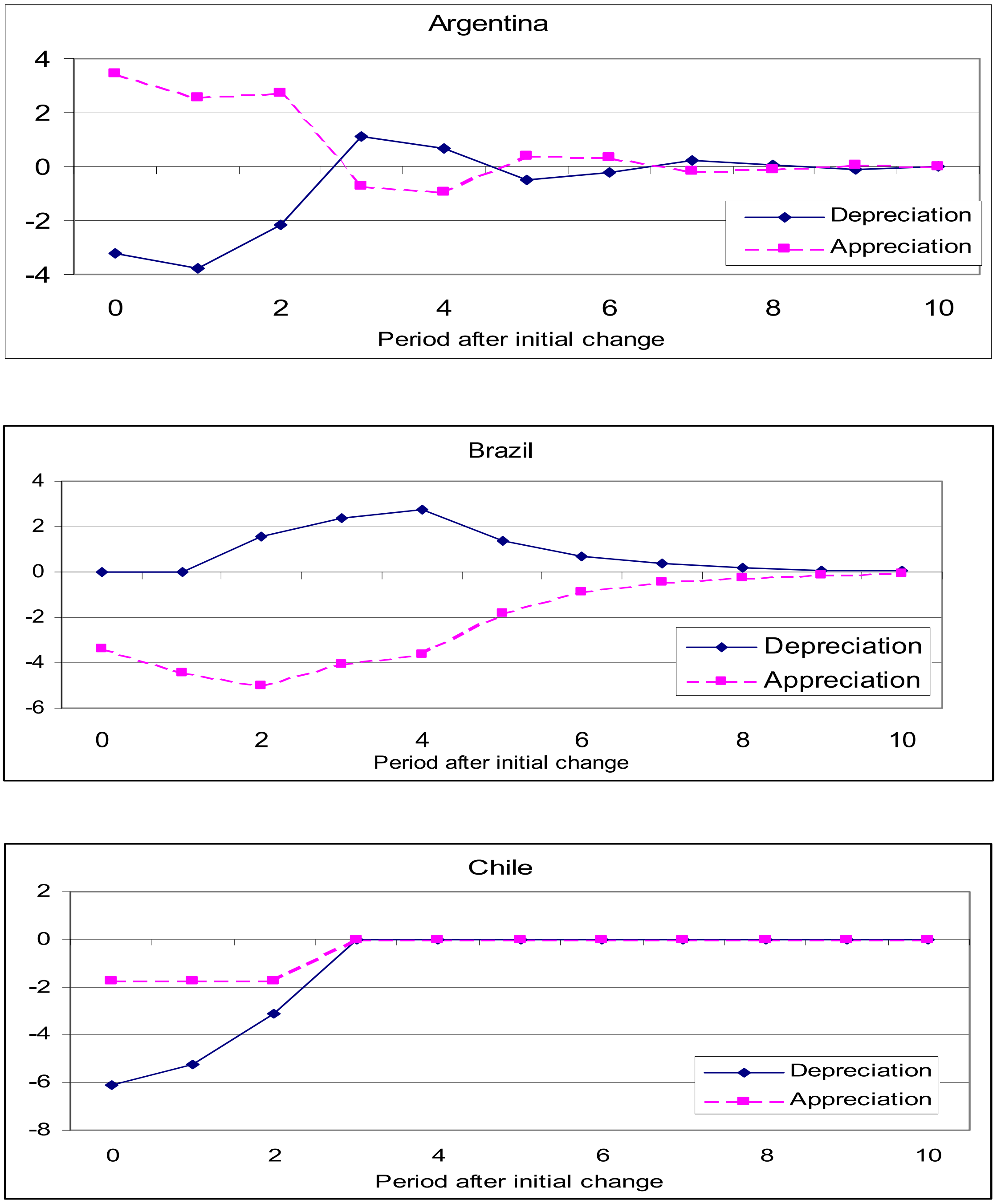
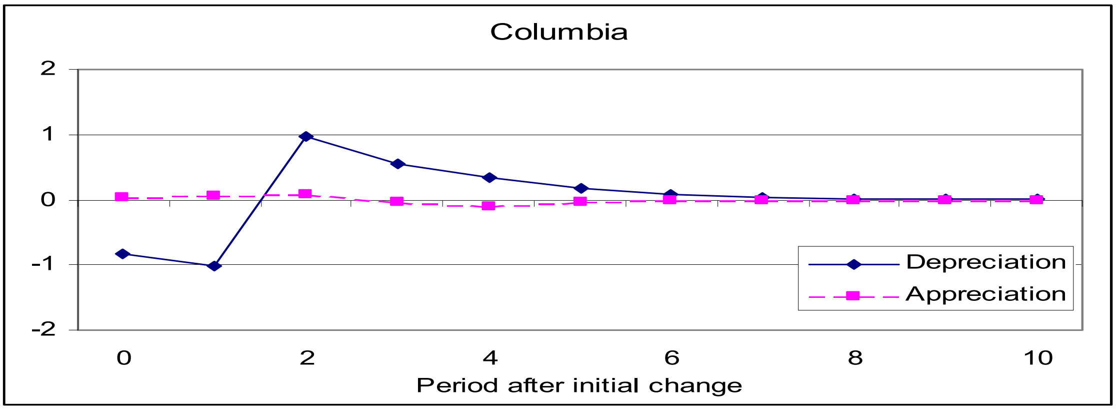
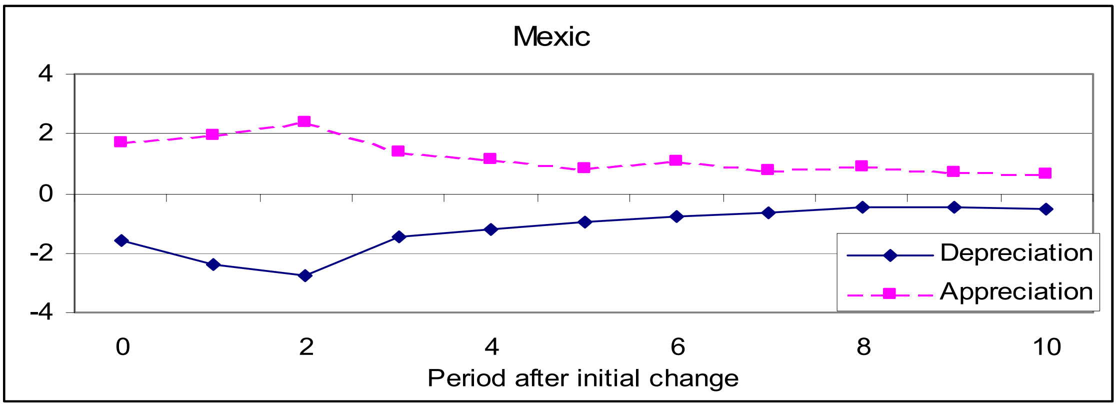
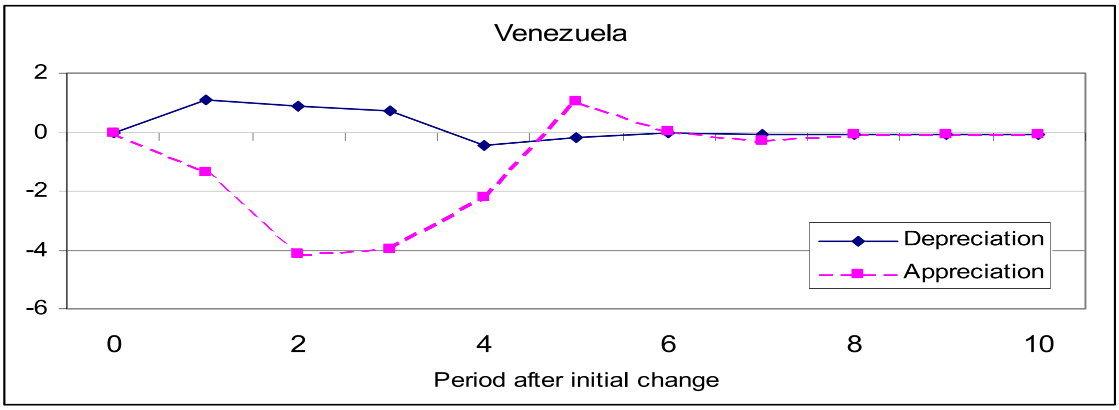


Figure 5. Estimated Dynamic Response to Change in the Real Exchange Rate









Note: In all cases it is assumed that a 10 percent real depreciation or appreciation takes place in each of three successive years.

APPENDIX 1

Data

The source for all data is International Financial Statistics, published by the IMF

Country	Sample	GDP volume (1995=100)	Population	Exchange rate to US dollar	Consumer Prices	Petroleum production
Argentina	1963-2000	21399BVPZF...	21399Z..ZF...	213..RF.ZF...	21364...ZF...	
Brazil	1964-2000	22399BVPZF...	22399Z..ZF...	223..RF.ZF...	22364...ZF...[7]	
Chile	1952-2000	22899BVPZF...[8]	22899Z..ZF...	228..RF.ZF...	22864...ZF...	
Colombia	1952-2000	23399BVPZF...[9]	23399Z..ZF...	233..RF.ZF...	23364...ZF...	
Mexico	1951-2000	27399BVRZF...	27399Z..ZF...	273..WF.ZF...	27364...ZF...	27366AA.ZF...
Venezuela	1958-2000	29999BVPZF...	29999Z..ZF...	299..RF.ZF...	29964...ZF...	29966AA.ZF...

Also used is the US Treasury Bill Rate (code: 11160C..ZF...).

[7] This series was available electronically from the IFS from 1980, prior to this the data were taken from the IFS Yearbook 1981.

[8] This series was available electronically from the IFS from 1960, prior to this the data were taken from the IFS Yearbook 1981 in 1975 prices.

[9] This series was available electronically from the IFS from 1968, prior to this the data were taken from the IFS Yearbook 1981 in 1975 prices.

APPENDIX 2

Modelling Methodology

This Appendix discusses details of the model specification and evaluation, together with computation of the dynamic multipliers for the effect of a real depreciation in Mexico.

Model Specification and Evaluation

Our modelling starts with a linear specification. A general-to-specific strategy is followed to get a parsimonious model, with minimum *AIC* being the criterion for model choice. We start with the maximum number of lags of y_t and x_t equal to two. Current values of the depreciation in the real exchange rate and US interest rates are included, but only lags of inflation. Therefore, counting each lag as a separate variable, a total of ten explanatory variables are initially included (eleven for Mexico and Venezuela, with the inclusion of oil production growth). After estimation by ordinary least squares (OLS) and calculation of *AIC*, the variable with the smallest *t*-statistic is deleted. Proceeding in this way, variables are dropped one by one, thereby allowing the possibility of "holes" in the lag structure. The procedure stops when deletion of a variable leads to an increase in *AIC*. The linear model with minimum *AIC* for each country is presented in Table 1.

As an indication of possible nonlinearity in our model, we test the null hypothesis of linearity against a nonlinear STR specification using the following equation:

$$y_t = \alpha + \sum_{i=1}^{n} \beta_i x_{ti} + \sum_{i=1}^{n} \theta_i \, x_{it} z_t + v_t \tag{A.1}$$

where the terms in z_t derive from a first order Taylor series approximation to $F(z_t)$. This test is a simplified version of the test in Teräsvirta (1994), with the simplification being the use of a first order Taylor expansion due to the relatively short sample sizes available to us. A test of the joint null hypothesis $\theta_i = 0$, $i = 1, \ldots, n$, in (A.1) is a test for linearity against STR nonlinearity with the known transition variable z_t. This test is computed as an F-test using the full initial linear model, with two lags of all variables (plus current values of US interest rates and depreciation)[10]. Each explanatory variable x_{ti} is considered in turn as the possible transition variable z_t, with results shown in Appendix Table A.1 as p-values. Although Teräsvirta (1994) recommends that comparison of these p-values can be used to determine the transition variable, we prefer to directly compare the fit of possible nonlinear models through the grid search procedure below.

A grid search of the general LSTR model of equations (6) and (7) is undertaken, using each of the ten explanatory variables x_{ti} in turn as the potential transition variable. This search, using OLS regression, is conducted over a range of γ and c values that define the logistic function (7). More specifically, the grid search examines 150 values of γ and 20 values of c within the observed range of the potential transition variable z_t. All ten explanatory variables x_{ti} are included in the general model. By choosing the combination that minimises the residual sum of squares, an initial estimation of the transition function $F(z_t)$ is obtained.

Based on this initial estimated transition function from the grid search, a general-to-specific approach is taken to obtain the explanatory variables included in our nonlinear model. In this case, the set of explanatory variables considered are the intercept, the explanatory variables x_{ti} (i = 1, …, 10), the transition function, F, and the interaction terms between the transition function and the explanatory variables, Fx_{ti} (i = 1, …, 10). Estimation is again by OLS, with AIC used to select the explanatory variables included,

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[10] We prefer to use the general linear model rather than the specific one of Table 1 to ensure that we do not miss possible nonlinear effects. In other words, we believe that overspecification is generally preferred to underspecification at this stage; see also Teräsvirta (1994). Another related issue is that neglected heteroscedasticity may lead to spurious rejection of the linearity null hypothesis. Consequently, some authors have suggested robustifying the linearity test. However, since this robustification may remove most of the power of the linearity test, robustification cannot be recommended, when the aim is to find and model nonlinearity in the conditional mean (see van Dijk, Teräsvirta and Franses, 2002, and references therein).

using the same procedure as for the linear model. Conditioning on the transition function simplifies estimation, since OLS can be employed rather than nonlinear least squares.

Having determined the variables to be included in the nonlinear model, the full LSTR specification, including γ and c, is estimated by nonlinear least squares. As suggested by Teräsvirta and Anderson (1992) and Teräsvirta (1994), each term in the exponent of F is scaled by dividing by the sample standard error of the transition variable, to assist in estimation. This standardisation is also useful for comparing the properties of the estimated transition functions across different countries. Further coefficients may be dropped at this stage, again based on the smallest t-statistic and the model is re-estimated by nonlinear least squares after each deletion. Values from the previous model are used as the initial values in estimation of the next model. Selection of the final model is based on minimum *AIC*, with the resulting models shown in Table 2 of the text with further details in Appendix Table A.2.

Diagnostic tests are applied to the estimated LSTR models. In particular, we present the ARCH test of Engle (1982) and the Jarque-Bera normality test (Jarque and Bera, 1980). In addition, we present tests designed by Eitrheim and Teräsvirta (1996) and Teräsvirta (1998) specifically for STAR and STR models. These tests are all computed using F-statistics for the significance of additional terms in the linearised version of the model. These include a test for second order autocorrelation in the residuals and a parameter constancy test which tests against the possibility that the intercept changes monotonically or non-monotonically with time. In addition, we test for the possibility of additional nonlinearity through a second transition function, which is represented by a first-order Taylor series approximation. Each explanatory variable in turn is considered as the second transition variable. Results of all diagnostic tests are reported in Appendix Table A.3 as p-values under the null hypothesis that the model is correctly specified.

One problem that can arise in the computation of some diagnostic tests is that $\widehat{\gamma}$ is large (so that the LSTR model becomes a threshold model) and, as pointed out by Eitrheim and Teräsvirta (1996), the moment matrix of the regressors in the auxiliary regressions used in computing the test statistics effectively becomes singular. In such cases (namely, when computing the results in Appendix Table A.3 for Brazil and Chile), we follow the recommendation of Eitrheim and Teräsvirta by omitting the terms corresponding to γ and c from the moment matrix for the computation of these test statistics.

All results reported in this paper have been obtained using GAUSS 3.2 and the programs used for the STR computations are originally due to Timo Teräsvirta.

Dynamic Analysis

The dynamics of nonlinear models can be investigated through the use of a technique analogous to the generalised impulse response functions, proposed by Koop, Pesaran and Potter (1996) and Potter (1998) as generalisation of the concept of impulse response functions used in linear models[11]. In doing so we follow current practice in the analysis of nonlinear model dynamics (see, for example Potter, 1995; Skalin and Terävirta, 1999; Öcal and Osborn, 2000; van Dijk, Teräsvirta and Franses, 2002). As already explained, these are required for the case of Mexico, since the transition variable is the lagged dependent variable in the growth rate regression.

In our context, we analyse the effect a depreciation (or appreciation) of 10 percent for each of three consecutive years, that is the effect of $\Delta e_t = \Delta e_{t+1} = \Delta e_{t+2} = \delta$, where $\delta = \pm 10$. For a history $\omega_{t-1,}$ the dynamic multiplier for a depreciation of δ is defined as

[11] A traditional impulse response function has well-known properties when the underlying model is linear. It has a *history independent property*, which implies that it is independent on the particular history ω_{t-1} (for example, changes occurring in a contraction have the same effect as those in an expansion). Also, it has a *symmetry property* in the sense that a change of $-\delta$ in any variable, or a disturbance shock of this size, has exactly the opposite effect of a change of $+\delta$. Finally, it has a *shock linearity property*, as the impulse response is proportional to the size of the shock.

$$E[y_{t+h} | \Delta e_t = \Delta e_{t+1} = \Delta e_{t+2} = \delta, \omega_{t-1}] - E[y_{t+h} | \Delta e_t = \Delta e_{t+1} = \Delta e_{t+2} = 0, \omega_{t-1}] \quad \text{(A.2)}$$

for horizons h = 0, 1, 2, …. The computation of the multiplier is based on stochastic simulations of (6) for h = 0, 1, 2, … using random disturbances generated from a normal distribution with zero mean and variance equal to the corresponding residual variance of the estimated STR model. In practice, the effect of the depreciation by δ in (A.2) is obtained by comparing the average path with $\Delta e_t = \Delta e_{t+1} = \Delta e_{t+2} = \delta$ (and subsequent Δe_{t+k} zero) to the average path when all $\Delta e_{t+k} = 0$ (k = 0, 1, 2, …). We compute each simulation using 10,000 replications and different draws for each path.

Regarding the history ω_{t-1} in (A.2), we simulate the responses of the model for Mexico in three different episodes corresponding to each regime[12], and then average to obtain the regime-dependent multiplier. This allows us to investigate whether the depreciation response is dependent on the growth regime.

Since the transition variable for Mexico is endogenous (namely, a lagged dependent variable), it is important to understand the nature of the endogenous dynamics, since these influence the estimated responses to a depreciation. This is effected through computation of the generalized impulse response function, with shocks of ±1 and ±2 standard deviations applied at t and random shock thereafter. Computations are otherwise the same as those of the dynamic multipliers (but with all $\Delta e_{t+k} = 0$), including the episodes over which averaging takes place. This allows us to investigate whether dynamics are regime dependent, as has been reported in the literature (see Potter, 1995 and Öcal and Osborn, 2000, for example).

[12] In the computation of the GIRF we consider the following particular years corresponding to the high growth regime: 1975, 1981 and 1990. For the moderate/negative growth regime, the corresponding years are 1959, 1983 and 1977. The time horizon is restricted in this case to eleven years to have enough observations to compute the GIRF for recent years.

Appendix Table A.1. Linearity Tests Against LSTR Model

Linearity tests	Argentina	Brazil	Chile	Colombia	Mexico	Venezuela
Δy_{t-1}	0.878	0.186	0.177	0.754	0.079	0.559
Δy_{t-2}	0.524	0.752	0.118	0.964	0.474	0.163
Δe_{t}	0.267	0.243	0.003	0.521	0.452	0.013
Δe_{t-1}	0.371	0.825	0.322	0.235	0.182	0.152
Δe_{t-2}	0.548	0.434	0.343	0.576	0.406	0.017
π_{t-1}	0.316	0.680	0.069	0.411	0.134	0.104
π_{t-2}	0.016	0.594	0.510	0.693	0.046	0.291
i_{t}^{US}	0.287	0.237	0.009	0.274	0.006	0.049
i_{t-1}^{US}	0.169	0.242	0.006	0.652	0.037	0.344
i_{t-2}^{US}	0.325	0.117	0.004	0.661	0.174	0.401
Δo_{t}	N/A	N/A	N/A	N/A	0.117	0.172

Note: Results are presented as *p*-values.

Table A.2. Full Estimation Results for Nonlinear Models

Coefficient	Argentina	Brazil	Chile	Colombia	Mexico	Venezuela
Intercept	13.87 (3.157)	6.011 (4.895)	4.246 (2.333)	1.645 (1.799)	3.157 (6.641)	3.158 (1.947)
Δy_{t-1}	-0.477 (-2.046)			0.514 (2.355)	0.0972 (1.117)	
Δy_{t-2}						-0.174 (-1.343)
Δe_t	0.260 (1.875)		0.174 (1.816)		-0.165 (-8.759)	
Δe_{t-1}	-0.057 (-3.585)				-0.036 (-1.859)	0.123 (3.640)
Δe_{t-2}						
π_{t-1}			-3.752 (-4.062)		-1.097 (-6.494)	
π_{t-2}			3.021 (3.521)	-0.280 (-1.285)		
i_t^{US}	-0.632 (-2.271)	-1.367 (-6.241)	-0.828 (-2.171)	0.409 (1.912)	0.202 (1.798)	-0.736 (-3.185)
i_{t-1}^{US}			1.451 (3.265)	-1.302 (-3.431)		
i_{t-2}^{US}		0.706 (3.123)		0.941 (2.650)		
Δo_t		N/A	N/A	N/A	0.201 (4.540)	0.448 (2.804)
$F(z_t)$	-11.85 (-2.392)			5.859 (2.643)	-7.697 (-3.472)	3.133 (2.847)
$F(z_t) \times \Delta y_{t-1}$	0.537 (1.495)	0.500 (3.937)	-0.290 (-1.903)	-0.599 (-2.129)		-0.205 (-0.841)
$F(z_t) \times \Delta y_{t-2}$	-0.571 (-2.297)		-0.313 (-1.877)		1.020 (2.506)	
$F(z_t) \times \Delta e_t$	-0.265 (-1.936)		-0.238 (-2.063)			
$F(z_t) \times \Delta e_{t-1}$						
$F(z_t) \times \Delta e_{t-2}$		0.158 (3.207)	0.139 (2.255)	0.209 (3.034)		
$F(z_t) \times \pi_{t-1}$		3.205 (3.922)		-4.527 (-4.870)		
$F(z_t) \times \pi_{t-2}$		-2.020 (-2.764)		0.737 (2.192)		
$F(z_t) \times i_t^{US}$		2.966 (7.131)	3.994 (5.726)			
$F(z_t) \times i_{t-1}^{US}$		-3.473 (-6.858)	-3.715 (-3.754)	1.824 (3.370)		
$F(z_t) \times i_{t-2}^{US}$			-1.054 (-1.840)	-1.217 (-2.540)		
$F(z_t) \times \Delta o_t$		N/A	N/A	N/A		
γ	8.873 (1.716)	999.50 (3617.0)	6711.00 (0.00002)	37.60 (0.932)	16760 (0.001)	896.1 (3601)
c	-3.179 (-0.833)	-0.719 (-46.38)	1.850 (457.1)	1.003 (3.581)	3.416 (4.988)	-1.166 (-4.676)

Note: Values in parentheses are t-ratios.

Appendix Table A.3. Diagnostic Tests for Nonlinear Models.

Test	Argentina	Brazil	Chile	Colombia	México	Venezuela
Normality	0.255	0.325	0.375	0.000	0.156	0.230
ARCH (2)	0.713	0.509	0.295	0.903	0.562	0.134
Autocorr. (2)	0.405	0.160	0.039	0.984	0.708	0.210
Intercept Constancy	0.498	0.644	0.563	0.027	0.033	0.313
Additional nonlinearity tests (*p*-value for each possible transition variable)						
y_{t-1}	0.657	0.137	0.754	0.768	0.700	0.251
y_{t-2}	0. 830	0.790	0.490	0.563	0.682	0.814
e_t	0.587	0.811	0.429	0.172	0.949	0.046
e_{t-1}	0.701	0.589	0.407	0.342	0.450	0.120
e_{t-2}	0.591	0.717	0.214	0.507	0.941	0.030
π_{t-1}	0.733	0.222	0.702	0.904	0.429	0.366
π_{t-2}	0.555	0.299	0.848	0.998	0.412	0.332
i_t^{US}	0.447	0.988	0.343	0.966	0.219	0.843
i_{t-1}^{US}	0.381	0.996	0.311	0.963	0.415	0.850
i_{t-2}^{US}	0.495	0.991	0.582	0.956	0.706	0.500
o_t	N/A	N/A	N/A	N/A	0.828	0.426

Notes: All results are presented as *p*-values. ARCH (2) is the Lagrange multiplier test of second order of Engle (1982); Normality is the test of Jarque and Bera (1980). Tests of no autocorrelation to order 2 and Intercept constancy are those of Eitrheim and Teräsvirta (1996). The additional nonlinearity test (Eitrheim and Teräsvirta, 1996) is the test of no missing linear terms, and no additional nonlinearity (not ignoring the “holes”).