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Modelling Real Exchange Rate Effects on Output Performance in Latin America

By

Pablo Mejía-Reyes[†*], Denise R. Osborn[†] and Marianne Sensier[†]

[†]Centre for Growth and Business Cycle Research, School of Economic Studies, University of Manchester, Manchester, M13 9PL, UK

[*]El Colegio Mexiquense, A.C., Zinacantepec, México

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Pablo Mejía-Reyes†*
Denise R. Osborn†
Marianne Sensier†

† Centre for Growth and Business Cycles Research
School of Economic Studies
The University of Manchester
Manchester, United Kingdom

* El Colegio Mexiquense, A.C.
Zinacantepec, México

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Abstract
This paper empirically analyses real per capita GDP growth for six Latin American countries (Argentina, Brazil, Chile, Columbia, Mexico, Venezuela) in terms of real exchange rate depreciations, inflation and US interest rates, focussing on the role of the real exchange rate. We find evidence of nonlinearity in this relationship, 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 may have negative effects for growth.

JEL classification: C5, C32.

Keywords: business cycle regimes, non-linear models, smooth transition models, Latin America, real exchange rate.

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Correspondence address: Denise Osborn, Centre for Growth and Business Cycle Research, School of Economic Studies, University of Manchester, Manchester M13 9PL, UK.
Phone: + (0) 161 275 4861. Fax: + (0) 275 4812. E-mail: denise.osborn@man.ac.uk.

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 as an important tool in their stabilisation programmes[1]. 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 the depreciation and firms may choose to reduce production[2].

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 Liviatan, 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). Moreover, using different sets of countries

[1] In fact devaluation was considered as a measure to compensate the deflationary effect of other measures in the stabilization programmes. See Buira (1983) and Edwards (1989a) for extensive analyses on this issue.

[2] The actual context may have additional complications that can lead to stronger negative effects of depreciations on output growth. Some authors have highlighted the reduction in the net wealth of countries and firms because of the increased value of foreign currency debt, the loss of access to international markets due to the decline in investor confidence that follows the abrupt abandonment of peg exchange rate regime, and the inflationary effects of depreciation that can cause additional damage to growth (see van Wijnbergen, 1986; De Gregorio, 1992a, 1992b; Ahmed et al., 2000). Lizondo and Montiel (1989) present a wide review of the literature.

and a panel approach, Edwards (1986, 1989b) and Ahmed (2003) find real exchange rate depreciations to have negative effects on growth, while, by using calibration techniques, Gylfason and Schmid (1983) and Solimano (1986) also find negative effects for a set of countries. All these papers have made important contributions to our understand of the nature of depreciation effects on output in developing economies, but they can hide the experience of particular countries or rely on imputed parameters obtained from frameworks set for other purposes (see Edwards, 1986). It is interesting to observe that there have been few attempts to distinguish the conditions under which depreciations can be contractionary, with exceptions being Agénor (1991) and Ahmad et al. (2002), who underline the difference between expected and unexpected changes in the exchange rate and between developed and developing economies, respectively.

Although somewhat mixed, most existing evidence points to depreciation having contractionary effects. Yet, as mentioned above, international financial institutions still consider that depreciation can be expansionary. In this context, this paper aims to shed light on the circumstances under which depreciations can have positive or negative effects on growth. In particular, it examines the effect of changes in the real exchange rate on output in six Latin America countries utilising nonlinear smooth transition regression (STR) techniques. By using time series analysis focussed on individual countries, we allow their responses to differ. Further, the STR approach allows the 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 relationship between the rate of real output growth and its explanatory variables. Although our methodology searches over a number of possible regime indicators (past output growth, inflation, real exchange rate changes, US interest rates and, where relevant, oil production), 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 the "regimes" we identify can be associated with appreciations versus depreciations in the real exchange rate, implying that changes in the real exchange rate have asymmetric effects. Mexico is the exception, where the regime indicator in our preferred model is lagged GDP growth.

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). Nevertheless, these papers have focused on measuring and modelling asymmetries, rather on than explaining output growth or the causes of recession.

The rest of this paper is structured as follows. Section 2 outlines the economic framework that we employ, while Section 3 then discusses the variables we use and presents evidence that a nonlinear model should be employed. Section 4 reviews the STR methodology we apply to capture these nonlinearities. Substantive empirical results are presented and discussed in Section 5, including the estimated models and dynamic multipliers for the effects of appreciations and depreciations on growth. Concluding remarks are offered in Section 6.

2. The Economic Framework and Modelling Approach

The economic framework for our analysis follows Kamin and Rogers (2000). First, define real output, y, through the simple identity

$$y = d + x \tag{1}$$

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 in order to focus on real output, the domestic rate of inflation (π) and the real exchange rate (e), we have

$$d = f_d(y, \pi, e, i^{US}) \tag{2}$$

The US interest rate, i^{US}, is expected to have a negative effect, due to its impact 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 (2), the effect of higher prices of foreign goods on domestic demand through e is anticipated to be negative.

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

$$x = f_x(e, y) \tag{3}$$

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[3].

Substituting (2) and (3) into the identity of (1) yields

$$y = f(e, \pi, i^{US})\,. \tag{4}$$

[3] 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.

In this expression, the sign of the real exchange rate is ambiguous, due to the different signs anticipated in (2) and (3). 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 (4), and hence that the effects of the real exchange rate on domestic demand through (2) dominate those that operate through the net exports function of (4). Nevertheless, and as we have already noted, this finding remains controversial.

Motivated by this economic framework, we study output performance (or, more specifically, output growth) in Latin America as a function of the real exchange rate, inflation and US interest rates. In addition to these variables used for all countries, we also include the growth in oil production as an explanatory variable for Mexico and Venezuela due to its importance in these economies. In practice, the empirical literature (including Edwards, 1986, 1989a; Kamin and Rogers, 2000) assumes linearity for regressions explaining the rate of growth. However, linearity is merely a convenient assumption and our modelling examines the role of a nonlinear functional form for f.

Although much analysis of growth in developing (and developed) countries uses a panel data approach, we analyse separately each of Argentina, Brazil, Chile, Columbia, Mexico and Venezuela. In a panel data analysis (such as Edwards, 1986), a number of country-specific control variables are included in a model for output growth. However, in practice the number of country-specific controls that can be included is limited and (conditional on the included control variables) the model imposes the restriction that each country has identical responses to variables such as the real exchange rate and inflation. In contrast, our analysis allows exploration of a much richer set of possibilities for the functional form without imposing constancy across countries. Nevertheless, we recognise that different responses may be due to factors we omit such as the degree of openness of the economy (Sachs, 1990). Therefore, the approach we adopt is complementary to a panel analysis.

3. Variable Definitions and Nonlinearity Tests

There are specific issues related to the definitions of some variables we use in our models, especially the real exchange rate. These issues are considered in the first subsection. This section also presents results of tests for nonlinearity in the relationship between growth and the explanatory variables, which motivates our later nonlinear analysis.

3.1 Variable Definitions

The modern definition of the real exchange rate considers it as the relative price of tradeable to nontradeable goods; see Edwards (1989b) for further details. Thus, the real exahange rate, e, is defined as

$$e = E \times \frac{P_T^*}{P_N} \qquad (5)$$

where E is the nominal exchange rate, measured as the number of units of local currency per unit of foreign currency, P_T^* is the world price of foreign tradeables in terms of foreign currency and P_N is the domestic price for nontraded goods. The real exchange rate is defined in these terms because, in the context of a model with tradeable and nontradeable goods, it is the domestic relative price of these two types of goods that is important. Thus, a higher e implies that the production of tradeables is more profitable and indicates greater international competitiveness of the domestic tradeables sector. An increase in e represents a real depreciation of the local currency relative to other currencies.

For practical purposes, the main measurement problem is the selection of the observed counterparts to P_T^* and P_N, since there are no perfect proxies available for these. In common with many other empirical studies, we use a foreign wholesale price index to measure P_T^* since it contains mainly tradeable goods, while a domestic consumer price index is used for P_N as it is more strongly influenced by nontradeables (Edwards, 1989c). In order to be able to cover a long span of data, the bilateral nominal exchange rate[4] with the US is used to measure E and for comparability, therefore, the US wholesale price index is adopted for P_T^*.

Figure 1 shows the annual percentage per capita GDP growth rate (Δy_t) for each country that we model, together with the rate of depreciation in the real exchange rate (Δe_t). The sample period used ends in 2000, while typically the sample starts during the 1950s[5]. Two features are, perhaps, particularly notable in these graphs. First, these countries sometimes experience substantial declines in per capita GDP, with all countries showing episodes with declines of greater than 5 percent in a year. Secondly, it appears that large real depreciations may be associated with these negative growth episodes. This is evidenced in Figure 1 by, for example, the experience of the 1975 depreciations for both Argentina and Chile, or 1995 for Mexico.

The basic equation that we estimate below is given by (4). However, prior to estimation, we need to decide whether the variables should be differenced. Kamin and Rogers (2000)[6] difference all variables, due to "near unit root behaviour in each series" (p.105). Whether the variables included in a regression for output growth 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

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[4] Ideally, we would prefer to use a nominal exchange rate index against a basket of currencies, but this is not available over the period from 1950.

[5] Appendix 1 contains details of the data used, including the precise sample period for each country.

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. This view is supported by Sarel (1996), who argues that the asymmetric distribution of observed inflation rates implies that large weight would be placed on a few large observations, whereas the log of inflation has a more symmetric distribution.

Our analysis treats international variables as exogenous to domestic growth. Thus, current values of US interest rates are permitted in our models. 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 this variable is also treated as predetermined. Kamin and Rogers (2000) conclude that causality runs from the real exchange rate to growth, at least in the case of Mexico[7]. As argued by Ahmed (2003), economic theory also points to the real exchange rate being causally prior to output in the contemporaneous period, since feedback effects from output to the real exchange rate through changes in domestic compared with foreign productivity is a longrun phenomenon. Nevertheless, we

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[6] It appears that Agénor (1991) may also use differenced data. Although no statement to this effect is made in the text, the explanatory variables (except for the real exchange rate) in his Table 1 are all expressed as differences.

[7] 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.

acknowledge that our treatment of the real exchange rate as exogenous is not entirely clear cut, since it partially depends on domestic inflation[8]. Only lagged values of inflation itself are used in our models to avoid endogeneity problems with this variable.

Therefore, we model the annual percentage growth in per capita GDP (Δy_t) as a function of the percentage change in the real bilateral exchange rate with the US (Δe), inflation (represented by π after taking logarithms) and the nominal US interest rate. 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 consider up to two lags of all variables, including the dependent variable. In addition, growth in oil production (Δo_t) is included for Mexico and Venezuela, but only the current value is included in order to conserve degrees of freedom.

3.2 Nonlinearity Test Results

Our primary interest focuses on the nature of the relationship given by (4), acknowledging the possibility that this relationship may be nonlinear. As a background to the detailed modelling results presented in Section 5, we here test for the presence of nonlinearity, using all variables and lags just discussed.

Results for two nonlinearity tests are presented (in the form of p-values) in Table 1, each of which considers the possibility that the nonlinearity is associated with the value of the indicated explanatory variable. The first test is based on a first-order Taylor series expansion of a nonlinear smooth transition regression (STR) model, while the second tests a simple version of a threshold model where the intercept changes in relation to an unknown threshold value of the transition variable. Although the former test is more general, it may suffer from power

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[8] We checked the extent to which the pattern in real exchange rate movements are predetermined by plotting both the contemporaneous real exchange rate depreciation variable that we use in modelling and this variable defined

problems due to the relatively large number of additional parameters (namely, 10 or 11) being subject to test in samples of only moderate size. As a single parameter test, the latter is more parsimonious. Further details of these tests can be found in Appendix 2.

With the exception of Columbia, Table 1 provides evidence of nonlinearity when (4) is estimated for each country. Although the test yielding the lowest *p*-value differs, this lowest *p*-value is obtained for Argentina, Brazil, Chile and Venezuela when the nonlinearity is a function of changes in the current real exchange rate. Except for Brazil, the statistic is significant at a level of around 1 percent or less. Thus, it appears that real exchange rate depreciations above some value for these countries may trigger different responses of the growth rate to the variables of the model. It should be noted that the evidence of the table is generally clear for these countries that nonlinearities are associated with real exchange rates, since the *p*-value for this case is markedly lower than that for other variables[9]. Of these four countries, it is only for Chile that another variable (namely, US interest rates) achieves a similar significance level.

Little evidence of nonlinearity related to real exchange rate changes is uncovered for Mexico, but strong evidence (significant at 2 percent or less) is found in relation to past growth rates, values of US interest rates and/or oil production. It is noteworthy that although nonlinearity in relation to inflation is considered as a possibility through the tests of Table 1, in no case is such a test statistic significant at the one percent level.

The results of Table 1 provide a basis on which to examine nonlinear models, with the indications being that such nonlinearity may be associated primarily with changes in the real exchange rate. However, the question of the appropriate variable generating the nonlinearity is reconsidered as part of our nonlinear modelling strategy. It is to this we now turn.

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with the current domestic CPI replaced by the CPI value lagged one year. Qualitatively, the patterns (particularly in respect of large movements in the real exchange rate) in the two series were similar.

[9] Note that the lowest *p*-value for Venezuela is for lagged, rather than current, real exchange rate depreciations.

4. Smooth Transition Regression Methodology

In this section we introduce our STR modelling methodology, with further information presented in Appendix 2. 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. This section ends with an outline of the dynamic analysis used to examine the properties of the models.

4.1 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:

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

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. Consequently, 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 Appendix 2, 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.

4.2 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 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.

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 the nonlinear dynamic responses to shocks 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.

5. Estimated Models

Since linear modelling is the standard approach, we present the results for such models before moving on to those 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 of the non-oil producing countries.

5.1 Linear Models

Table 2 presents our linear models for Argentina, Brazil, Chile, Columbia, Mexico and Venezuela. The specific explanatory variables and lags included in each case 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.

The differing responses within Latin America to real exchange rate depreciations may be associated with variables that are not included in our models, particularly the degree of openness. We anticipate that it will be easier for a more open economy to reallocate factors of production from the nontradeable to the tradeable sectors, and hence to benefit from a real exchange rate depreciation. Although he does not analyse all the countries we consider, in his discussion of the external shocks experienced by a number of Latin American countries in the late 1970s and early 1980s, Sachs (1990) argues that Brazil was a more open economy than others, which is compatible with the estimated growth-enhancing effects for that country in Table 2.

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 four countries (the exceptions being Chile and Columbia). In the case of Mexico, in contrast to other countries, the coefficient of the current rate is positive and strongly significant, rather than negative as anticipated from the model of Kamin and Rogers (2000) and discussed in Section 2 above. 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. It is plausible that nominal US interest rates here proxy real rates that are pro-cyclical, and hence high rates are consistent with rising Mexican exports to her main market[10]. 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. These apparently satisfactory conventional diagnostics are, however, undermined by the nonlinearity test results already discussed.

5.2 Nonlinear Models: Non-Oil Producing Countries

Prior to estimating the nonlinear STR model (6), the transition variable z_t needs to be specified. For the non-oil producing countries of Argentina, Brazil, Chile and Columbia, our search procedure over all explanatory variables selects current or lagged depreciation as the

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[10] We are grateful to an anonymous referee for this point.

transition variable z_t. More specifically, the two year lag of depreciation is selected for Argentina, but the current value is selected in the remaining three cases. For Argentina our 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). In line with the nonlinearity test results of Table 1, current depreciation is then selected as the transition variable for Argentina[11]. Therefore, all models reported in Table 2 have transition variable $z_t = \Delta e_t$, reinforcing the role found for this variable in the nonlinearity tests above.

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, indicating 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. Despite the lack of evidence of nonlinearity for Columbia in Table 1, these comments apply to this country as well as those for which we found evidence of nonlinearity associated with real exchange rate depreciations. The only exception is a partial one for Argentina, which has a smoother transition function than the other cases in Figure 2.

Table 3 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.1. The absence

[11] This selection considered lagged growth and current or lagged depreciation as the possible transition variables.

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 depreciation 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 comparisons of growth over regimes. To analyse the depreciation effects more fully, we later undertake a dynamic multiplier analysis.

Another interesting result from Table 3 is that, in contrast to the negative effects of US interest rates found for Brazil, Chile and Columbia in the linear models of Table 2, US interest rates are found to have positive contemporaneous effects on output growth in these countries during depreciation regimes[12]. In line with the interpretation made above for Mexico that US interest rate effects may here proxy the US (or international) business cycle, conditions for export-led growth in Brazil, Chile and Columbia may require both real exchange rate depreciation and rising world demand. However, the interactions are complex, since the effect

[12] We are grateful to an anonymous referee for drawing our attention to this aspect of our results.

of lagged US interest rates is also regime-dependent and these lagged effects are negative for Brazil and Chile within a depreciation regime.

Similarly to the regime-dependent coefficients for US interest rates, the roles of inflation and lagged growth sometimes differ 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.2) show that the nonlinearity evident in Table 1 is satisfactorily accounted for by our models. The other diagnostics for the nonlinear models are also generally satisfactory, although non-Normality remains in the model for Columbia.

5.3 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 satisfactory from an economic perspective[13]. 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 4 and Appendix Table A.1 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

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[13] The estimated model for Mexico replicated the finding of Table 2 that higher US interest rates lead 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.

depreciation, both of which are compatible with the results of Table 1[14]. 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 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 4, 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 4. 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.2). 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

[14] The smallest *p*-value for Venezuela in Table 1 results from current real depreciation as the transition variable,

and oil production (Table 1), while these are restricted to a constant role over regimes in Table 4.

5.4. Dynamic Analysis

As Mexico is the only case where regime changes depend on lagged growth, and hence are endogenous in our single equation context, 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 4. 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 positive shocks tend to enhance subsequent 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

although this is not substantially smaller than that for the second lag of this variable.

associated with appreciations and depreciations in the real exchange rate. As discussed in Section 4.2, 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 particularly 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. However, the effects here are relatively small and it should be recalled that our initial results (Table 1) give little support for the existence of nonlinearity in the model for Columbia. 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 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. It is clear, however, that despite the coefficients on the depreciation variables being constant over regimes in Table 4, the endogenous regime switching discussed above results in substantial asymmetries in the effects of depreciation versus appreciation[15]. The sustained real exchange rate depreciation of 10 percent per year is estimated to reduce growth on average by eight percentage points at a lag of a year, compared with an average increase of two percent for a corresponding appreciation.

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 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 3 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 5 for a continued three year depreciation are always negative, with the initial impact of a 10 percent depreciation being negative and around six 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 5 (compared with the baseline model) also reflects the differential effects of US interest rates in the two regimes.

6. Conclusions

In this paper we analyse the determination of real per capita GDP growth 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. By

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[15] In fact, by setting the initial lagged growth rate to its sample mean when computing the dynamic responses,

studying individual countries, we also do not impose the restriction that they have common responses.

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, Chile and Mexico in the short-run, with less severe negative effects for Columbia. These results are in line with the findings reported in the literature. As Ahmad et al. (2002) suggest, it may be the case that structural factors (as outlined in the Introduction above) determine the nature of the effects of the real exchange rate on growth for these Latin American countries.

Yet, some caution must be shown regarding the policy implications for this conclusion, since real depreciations may be accompanied by other constraining demand policies following periods of sustained real appreciation of domestic currencies. Thus, for policy purposes, it is important to underline that overvaluation should not be allowed given that the correction of the exchange rate would only be a matter of time[16]. On the other hand, most of the countries we study have implemented deep reforms that have transformed the structure of their economies, which may imply different effects coming form changes in the exchange rate in the future. For example, they have become more open to international trade and capital transfers, and hence their productive sectors have been exposed to greater competition. It would be expected that in the medium run the tradable sectors would represent a greater share in domestic production, so

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changes in regime can occur, because the threshold value for the regimes is fairly close to this mean.

real depreciations could have positive impacts on overall output growth. This does not imply targeting the real exchange rate at too high levels, given the inflationary effects linked to nominal depreciation that some authors have highlighted.

In the cases of Brazil and Venezuela, the estimated effects of depreciations are positive for growth, which may be link to the fact that the former country has been a more open economy than the others during most of the sample period. However, in both of these countries, the asymmetric responses imply larger negative growth consequences of real appreciation than the positive effects of depreciation. Although on different grounds to that above for the other countries, we could argue in favour of avoiding overvaluation of the real exchange rate, since for Brazil and Venezuela it seems to imply a loss of competitiveness that slows down economic growth. Yet depreciation is not suggested either, given that its benefits for growth are relatively small and short-lived, but it may trigger increasing inflation rates similar to those experienced by these countries in recent decades.

Overall, we suggest that neither real overvaluation nor real undervaluation should be pursued to encourage growth in any country. Rather, in the new context of more open economies and an increasing number of countries adopting floating exchange rate regimes, it would be more appropriate to adopt policies that maintain the real exchange rate around its equilibrium level. The definition of such a level is not an aim of this paper, but it is a matter of future research.

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

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[16] This result is evident with the recent currency crisis in Argentina in 2001/2002 when the release from the peg with the US dollar saw Argentina's peso devalue by 70%. This was accompanied by recession and unemployment reached 30 percent.

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. Nonlinearity Test Results

Transition variable	Argentina	Brazil	Chile	Colombia	Mexico	Venezuela
Δy_{t-1}	0.878 0.379	0.186 0.158	0.177 0.878	0.754 0.350	0.079 0.910	0.559 0.696
Δy_{t-2}	0.524 0.658	0.752 0.852	0.118 0.609	0.964 0.456	0.474 0.017	0.163 0.934
Δe_t	0.267 0.004	0.243 0.024	0.003 0.892	0.521 0.800	0.452 0.966	0.013 0.889
Δe_{t-1}	0.371 0.503	0.825 0.952	0.322 0.810	0.235 0.233	0.182 0.354	0.152 0.107
Δe_{t-2}	0.548 0.247	0.434 0.432	0.343 0.078	0.576 0.160	0.406 0.962	0.017 0.069
π_{t-1}	0.316 0.018	0.680 0.098	0.069 0.763	0.411 0.479	0.134 0.058	0.104 0.525
π_{t-2}	0.016 0.042	0.594 0.196	0.510 0.399	0.693 0.949	0.046 0.244	0.291 0.243
i_t^{US}	0.287 0.586	0.237 0.308	0.009 0.889	0.274 0.296	0.006 0.353	0.049 0.118
i_{t-1}^{US}	0.169 0.741	0.242 0.774	0.006 0.076	0.652 0.532	0.037 0.437	0.344 0.428
i_{t-2}^{US}	0.325 0.206	0.117 0.371	0.004 0.410	0.661 0.869	0.174 0.139	0.401 0.546
Δo_t	N/A N/A	N/A N/A	N/A N/A	N/A N/A	0.117 0.020	0.172 0.581

Note: Results are presented as p-values. For each potential transition variable, the results of two tests are presented; the first is obtained as a significance test for the additional interaction variables in equation (A.1), while the second is a test of linearity against the threshold model (with unknown threshold value) of (A.2); see Appendix A.2. N/A indicates not available because Δo_t does not appear in the regression model for that country.

Table 2. 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 3. Estimated Nonlinear Models for Per Capita GDP Growth in Non-Oil Producing Countries

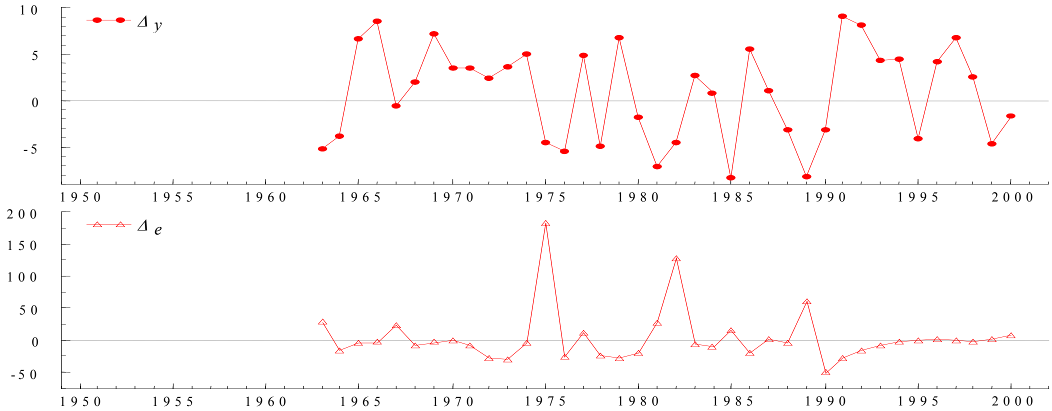
Coefficient	Argentina		Brazil		Chile		Columbia	
	Large appreciation ($F = 0$)	**Small appreciation/ Depreciation ($F = 1$)**	**Appreciation ($F = 0$)**	**Depreciation ($F = 1$)**	**Appreciation ($F = 0$)**	**Depreciation ($F = 1$)**	**Appreciation ($F = 0$)**	**Depreciation ($F = 1$)**
Intercept	13.87	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 4. Estimated Nonlinear Models for Per Capita GDP Growth in Oil Producing Countries

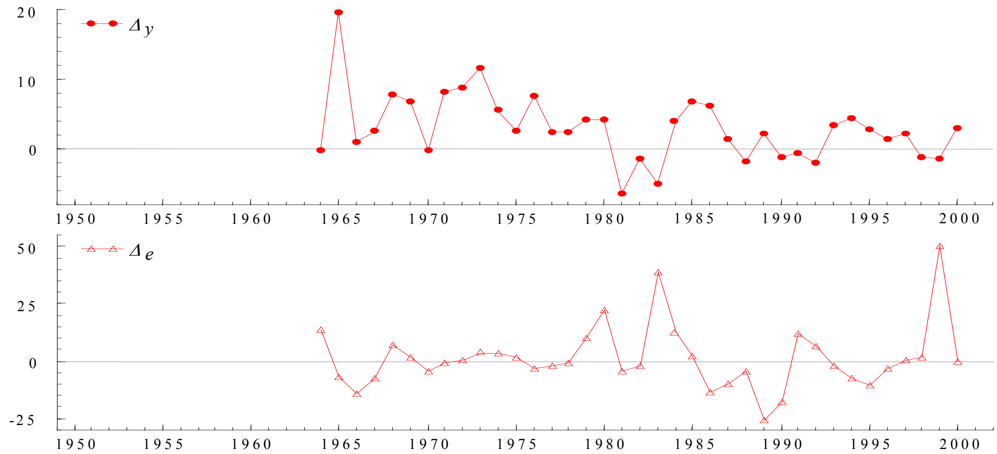
Coefficient	México		Venezuela	
	Moderate/ negative growth ($F = 0$)	**High growth ($F = 1$)**	**Large appreciation ($F = 0$)**	**Small appreciation/ Depreciation ($F = 1$)**
Intercept	3.157	-4.540	3.158	6.291
Δy_{t-1}	0.0972	0.0972		-0.205
Δy_{t-2}		1.020	-0.174	-0.174
Δe_t	-0.165	-0.165		
Δe_{t-1}	-0.036	-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}	
Goodness of fit measures				
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

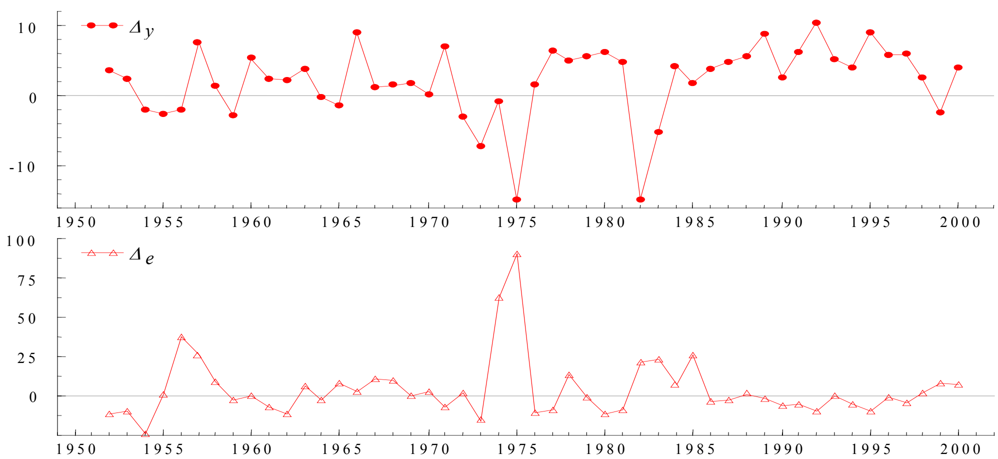
Argentina



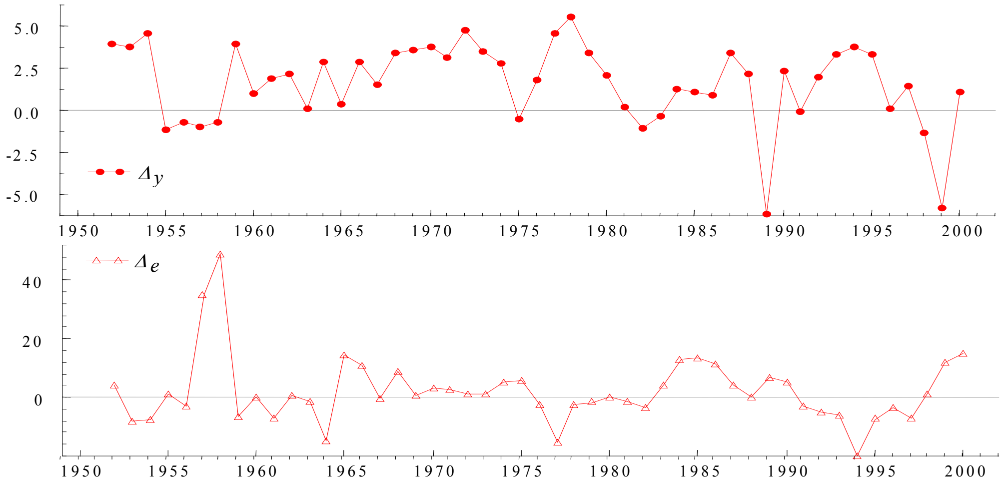
Brazil



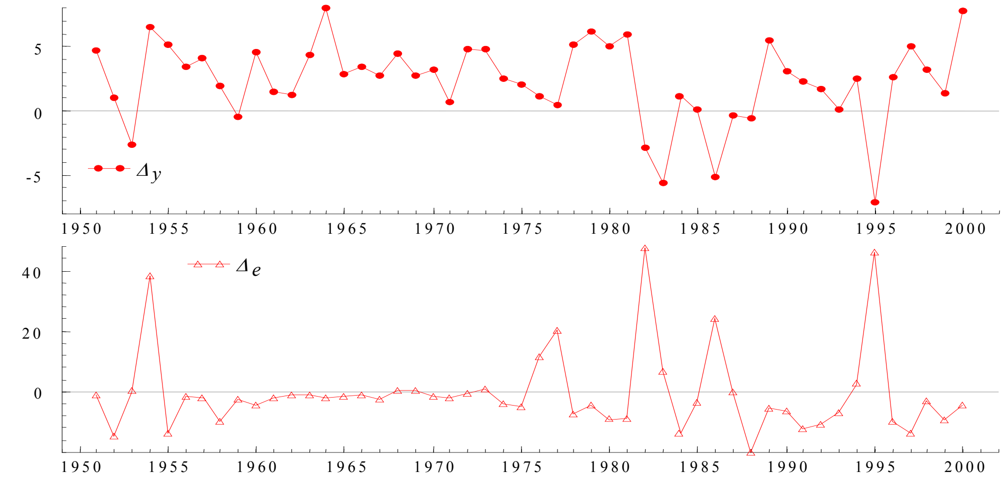
Chile



Colombia



Mexico



Venezuela

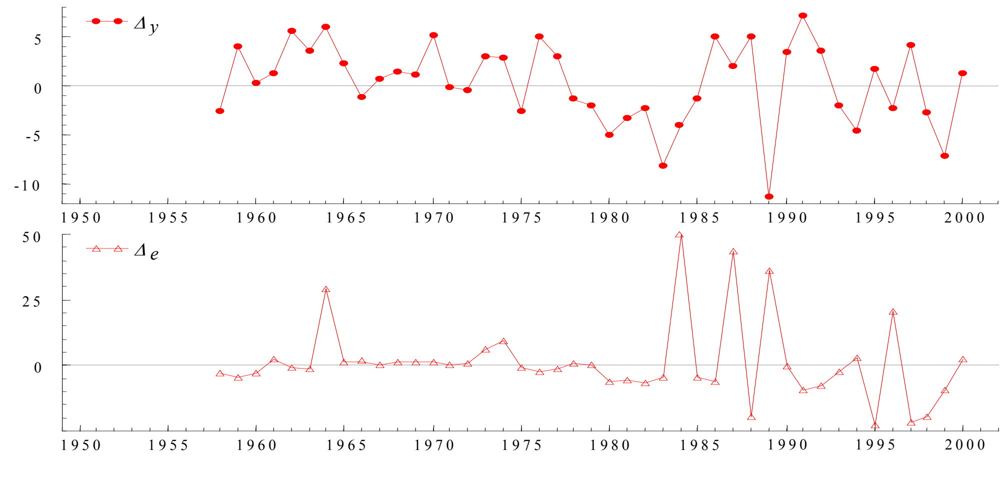
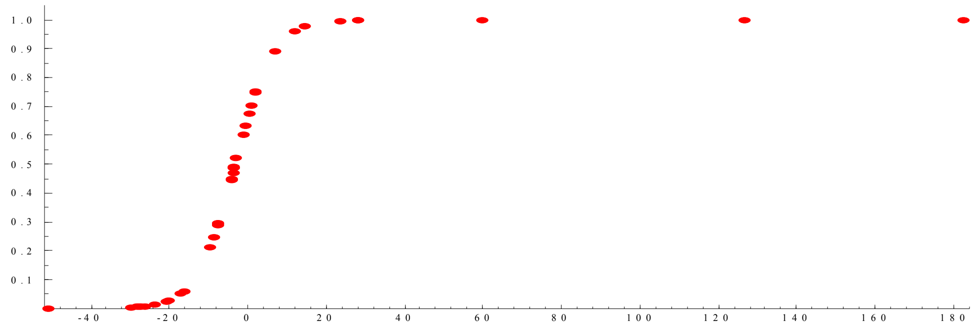
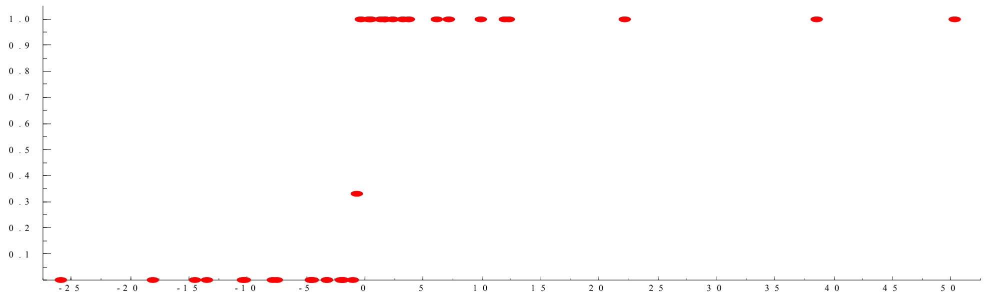


Figure 2. Transition Functions for Non-Oil Producing Countries

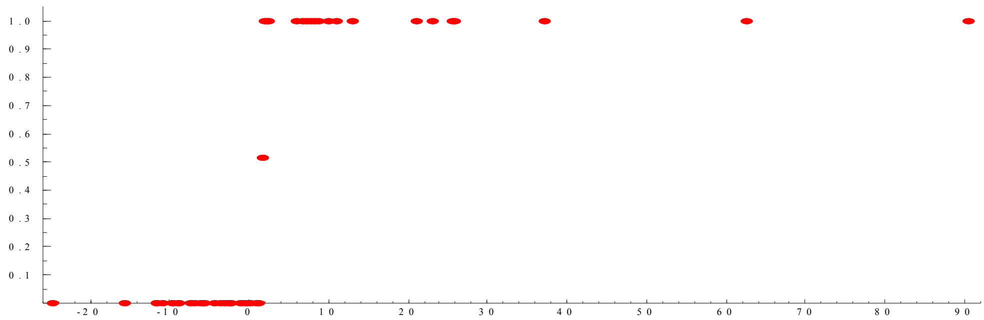
Argentina: F versus Δe_t



Brazil: F versus Δe_t



Chile: F versus Δe_t



Colombia: F versus Δe_t

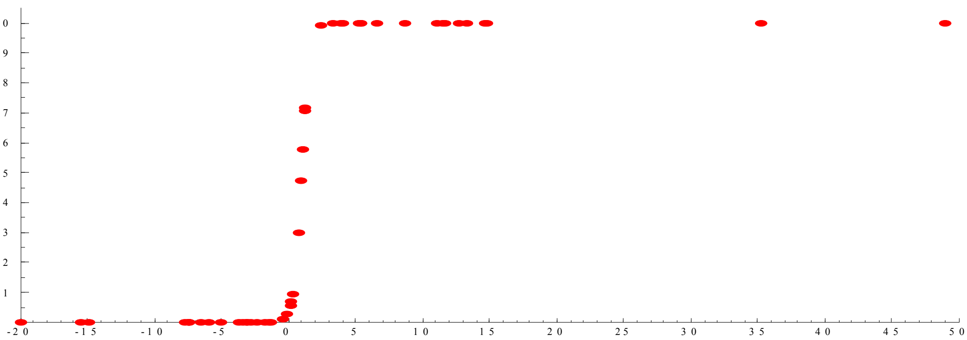
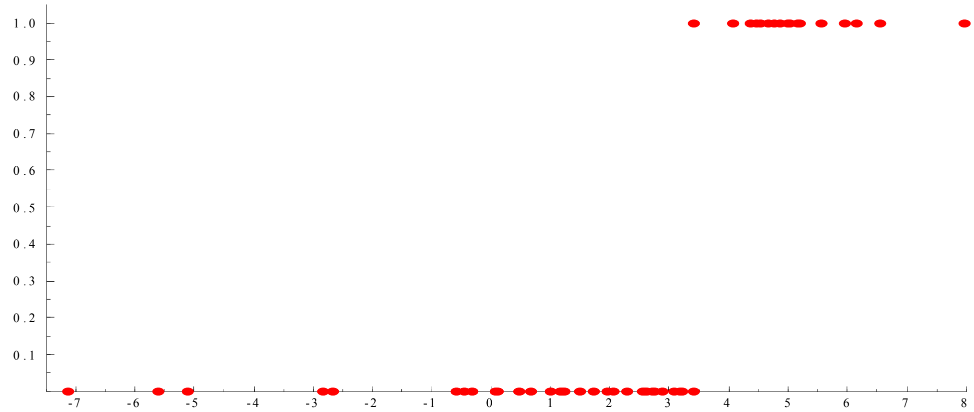


Figure 3. Transition Functions for Oil Producing Countries

Mexico: *F* versus Δy_{t-2}



Venezuela: *F* versus Δe_{t-2}

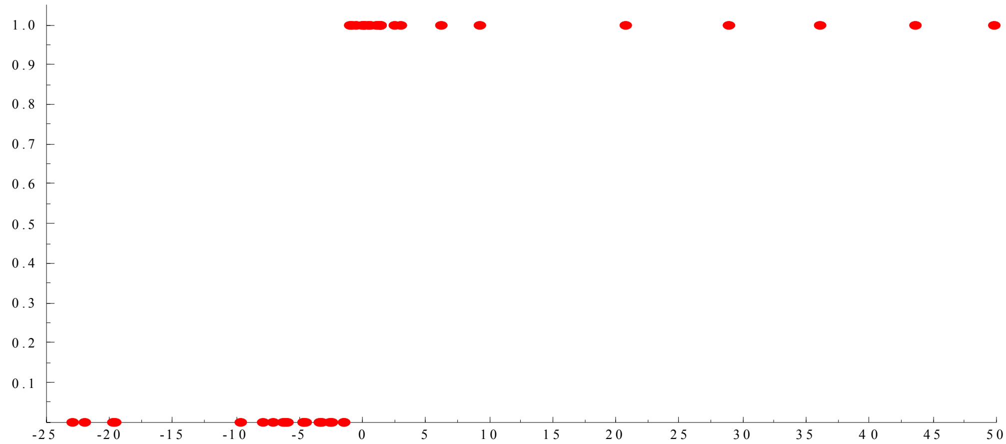
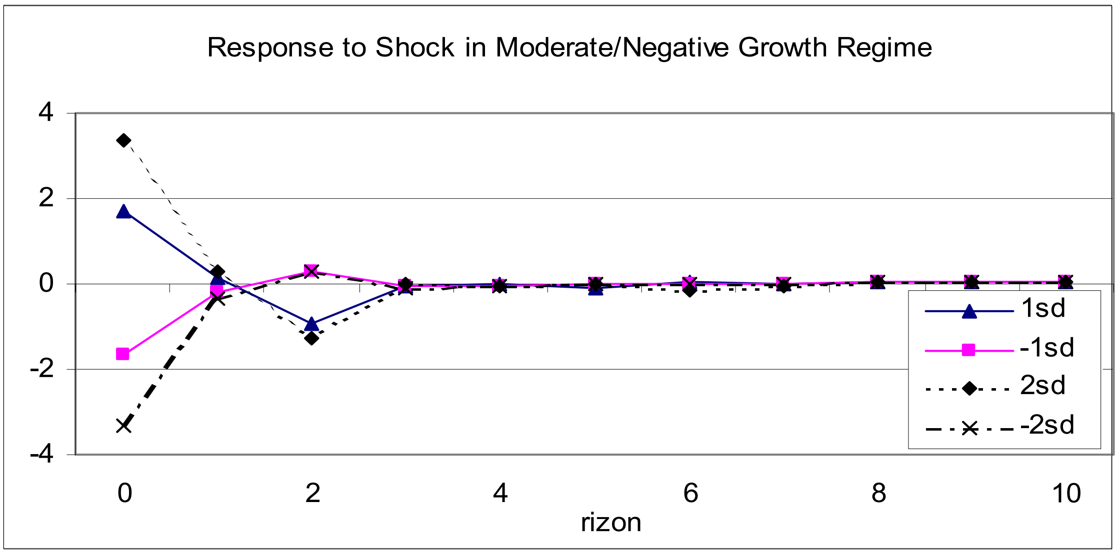


Figure 4. Generalised Impulse Response Functions for Nonlinear Model for Mexico



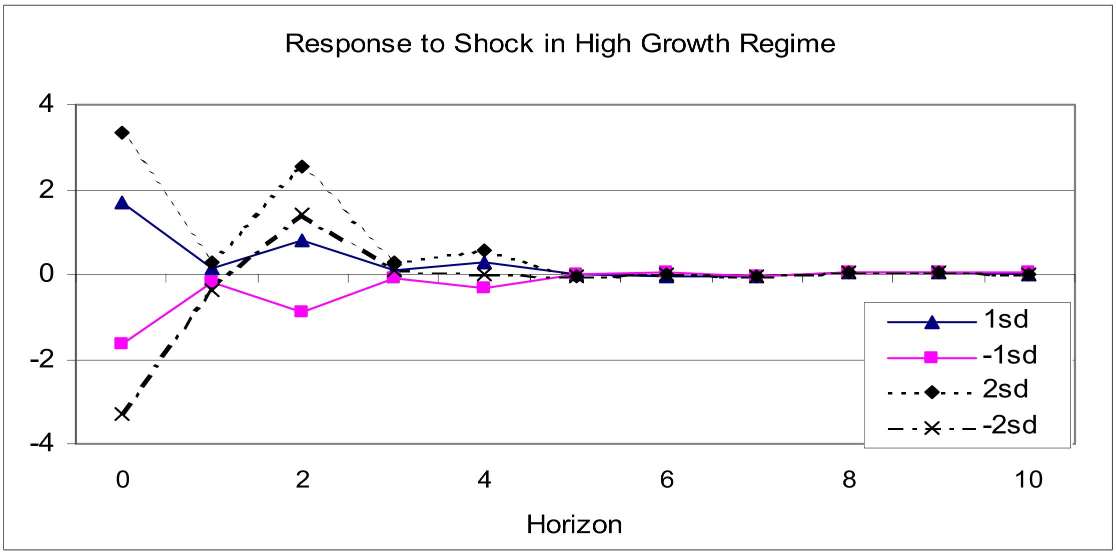
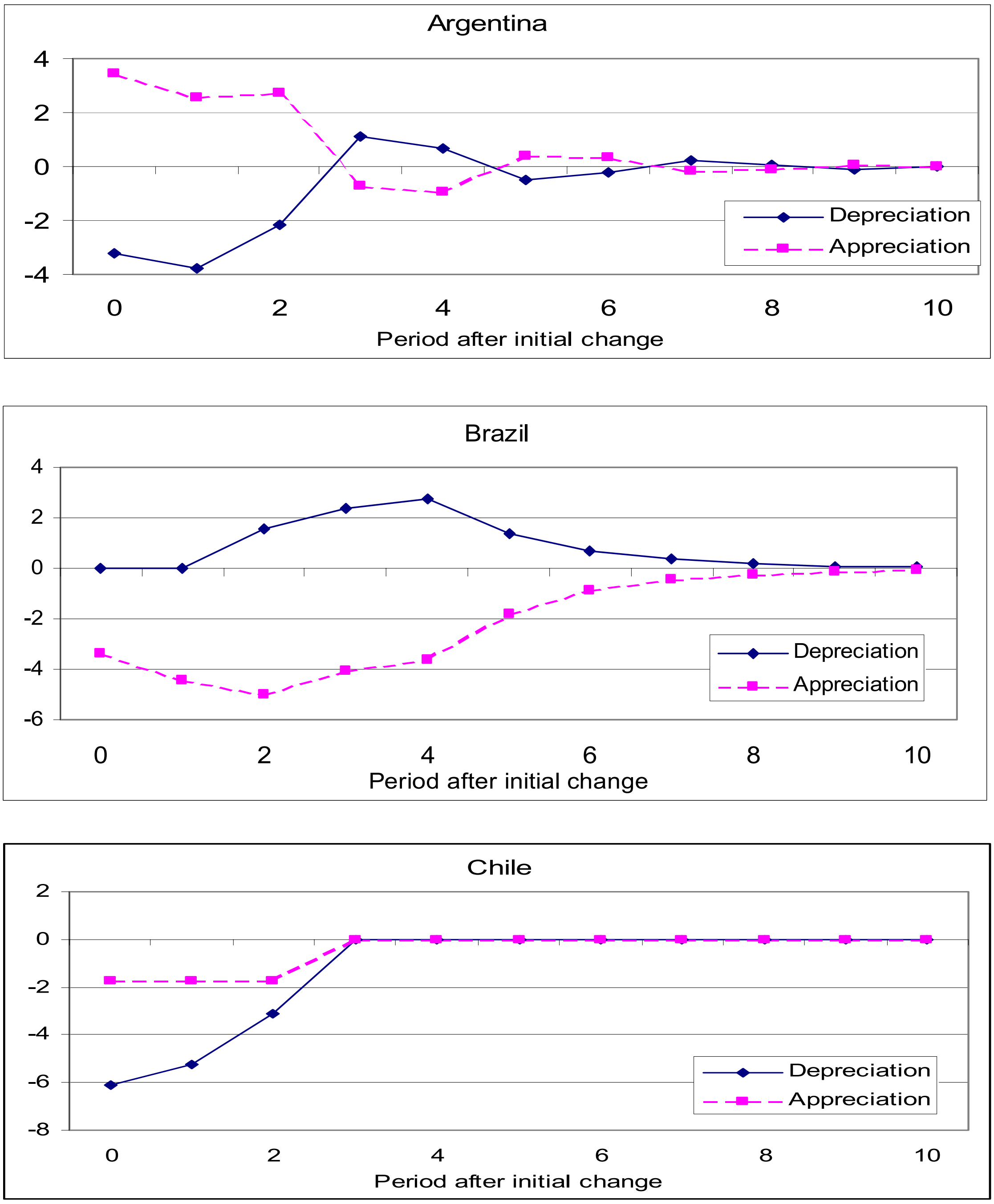
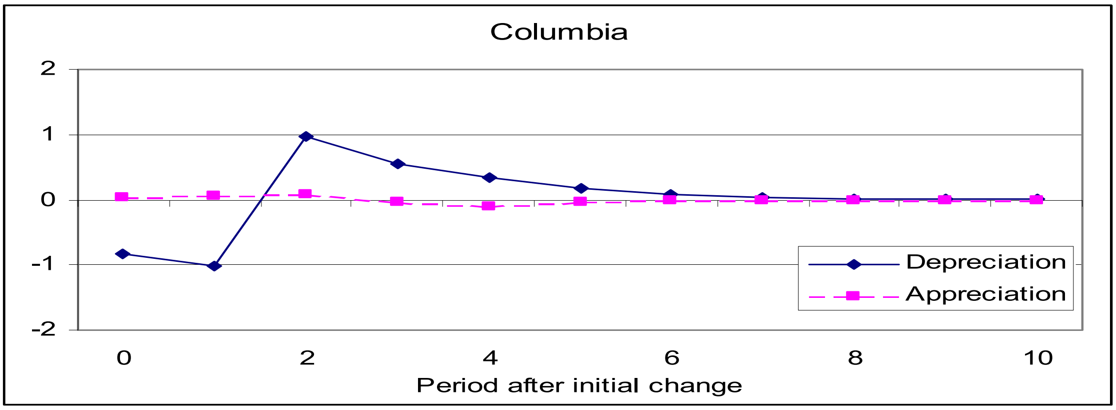
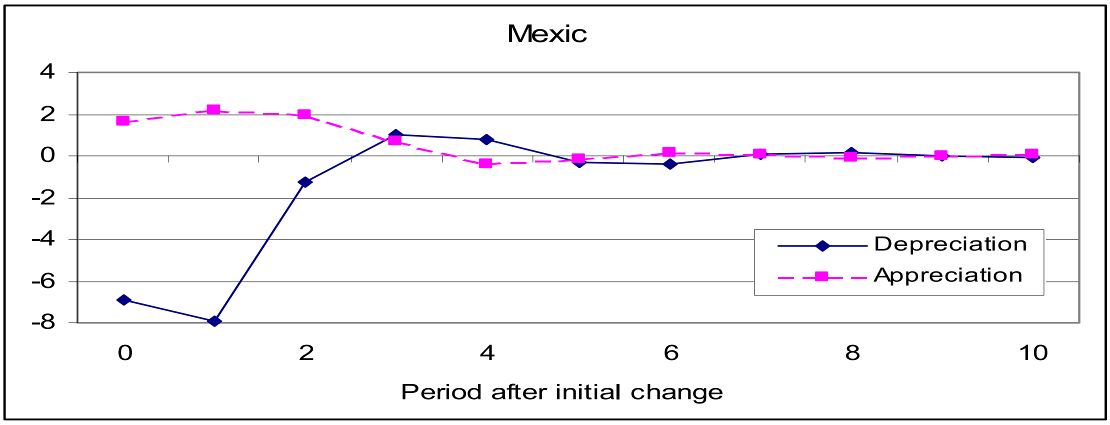
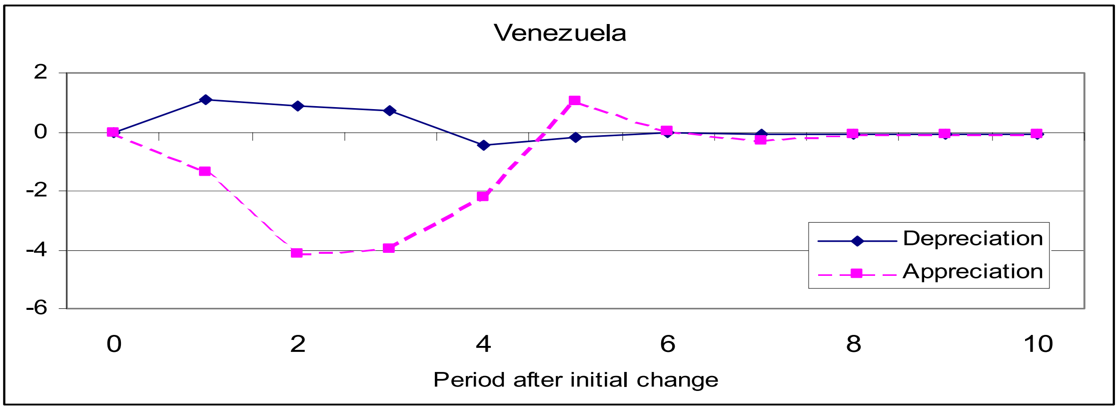


Figure 5. Estimated Dynamic Response to Change in the Real Exchange Rate









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...[17]	
Chile	1952-2000	22899BVPZF...[18]	22899Z..ZF...	228..RF.ZF...	22864...ZF...	
Colombia	1952-2000	23399BVPZF...[19]	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...).

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

[18] 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.

[19] 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 procedures we adopt for testing for nonlinearity, model specification and evaluation, and computation of the dynamic multipliers for the effect of a real depreciation in Mexico.

Testing for Nonlinearity

As an indication of possible nonlinearity in our model, we test the null hypothesis of linearity in two ways. Firstly, we test against a nonlinear STR specification using the specification:

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

where the terms in z_t derive from a first-order Taylor series approximation to $F(z_t)$ defined in (7) 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)[20]. Each explanatory variable x_{ti} is considered in turn as the possible transition variable z_t, with results shown in the first row of Table 1 for each transition variable. 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.

The second form of nonlinearity test we employ is based on a threshold specification for the nonlinear model, which assumes that the transition between regimes is

[20] We prefer to use the general linear model rather than the specific one of Table 2 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).

abrupt. The general form of the threshold model is identical to that of the STR model of (6), except that

$$F(z_t) = \begin{cases} 0 & z_t \le c \\ 1 & z_t > c \end{cases} \tag{A.2}$$

Hansen (1997) discusses inference for threshold models of this type, where c is estimated as the value of z_t that yields minimum residual sum of squares when the model is estimated. However, with 50 or fewer observations available for estimation and 10 or 11 explanatory variables in our models, combined with the need to "trim" the observed distribution of z_t to avoid regimes associated with extreme values, testing in the context of the general threshold model is impractical in our context. For this reason, we apply the approach to the specification

$$\Delta y_t = \alpha_0 + \sum_{i=1}^{n} \beta_i x_{ti} + F(z_t)\alpha_1 + \varepsilon_t \tag{A.3}$$

where only the intercept changes with the regime defined by (A.2) and hence α_1 is analogous to a dummy variable coefficient. We search for c over the observed values of z_t, after trimming by removal of the eight largest and eight smallest values (out of 40 to 50 observations). The p-values reported for this test in Table 1 are obtained by simulation, using the method described by Hansen (1997) applied in the context of (A.2) and based on 10,000 replications.

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 for all variables equal to two (with the exception of oil production for Mexico and Venezuela). 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 2.

A grid search of the general STR 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. 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. This standardisation is useful for comparing the properties of the estimated transition functions across different countries, and also aids estimation. All 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.

Conditional on this initial 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, 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 STR specification, including γ and c, is estimated by nonlinear least squares. 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 Tables 3 and 4 of the text, with details in Appendix Table A.2.

Diagnostic tests are applied to the estimated STR 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 latter 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.2 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 $\hat{\gamma}$ is large (so that the STR model effectively becomes a threshold model) and, as pointed out by Eitrheim and Teräsvirta (1996), the moment matrix of the regressors in the auxiliary models used in computing the test statistics approach singularity. In such cases (namely, when computing the results in Appendix Table A.2 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 reported estimation results have been obtained using GAUSS 3.2 and the programs used for the STR computations are originally due to Timo Teräsvirta.

Dynamic Analysis

Since, with the exception of Mexico, the effects of a given depreciation/appreciation on the rate of growth are linear, we generally use deterministic forecasts (with zero disturbances) in the computation of the dynamic responses.

The dynamic multiplier response to depreciation for Mexico is investigated through techniques 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[21]. 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.

[21] 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.

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 ω_{t-1}, the dynamic multiplier for a depreciation of δ is defined as

$$E[y_{t+h} | \Delta e_t = \Delta e_{t+1} = \Delta e_{t+2} = \delta, \omega_{t-1}] - E[y_{t+h} | \Delta e_t = \Delta e_{t+1} = \Delta e_{t+2} = 0, \omega_{t-1}] \qquad \text{(A.4)}$$

for horizons h = 0, 1, 2, …. The computation of the multiplier is based on stochastic simulations of the estimated equation (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 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.

We explore the endogenous dynamics for Mexico through computation of the generalized impulse response function (GIRF), with shocks of ±1 and ±2 standard deviations applied at t and random shocks thereafter. Regarding the history ω_{t-1} in (A.4), we simulate the responses in three different episodes corresponding to each regime[22], and then average to obtain the regime-dependent multiplier. This allows us to investigate the extent to which the endogenous dynamics of the growth rate are regime dependent, as has been reported in the literature (see Potter, 1995 and Öcal and Osborn, 2000, for example).

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[22] 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.

Table A.1. Detailed 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.2. 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").