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Luiz de Mello,

Diego Moccero,

Matteo Mogliani

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ABSTRACT/RESUME

Do Latin American central bankers behave non-linearly? The experiences of Brazil, Chile, Colombia and Mexico

This papers estimates unrestricted monetary reaction functions for four Latin American countries (Brazil, Chile, Colombia and Mexico) and tests for the presence of non-linear effects in central bank behaviour. The analysis covers the post-1999 inflation-targeting period. We deal with the presence of unit roots in the data by estimating the policy rules in a co-integration setting. We test for linear and non-linear co-integration among the variables of interest. The results suggest that a non-linear specification is not rejected by the data for Brazil, Colombia and Mexico, but it is for Chile. Estimation of smooth-transition models by NLLS and EN-NLLS suggests that the central bank's response to the inflation gap (*i.e.* deviations of expected inflation from the target) is invariant across policy regimes in Colombia. It becomes stronger in Mexico as expected inflation deviates from the target. Policy responses appear to weaken in Brazil as the inflation gap widens, a finding that most probably reflects a history of adverse supply shocks and upward adjustments in targets in the early years of inflation targeting. Non-linearity is also found in the central bank's response to the exchange rate in Brazil and Colombia.

JEL codes: E52; C22; O54.
Keywords: inflation targeting; reaction function; non-linear co-integration; smooth-transition models.

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Les banques centrales d'Amérique latine se comportent-elles d'une manière non-linéaire? Les expériences du Brésil, du Chili, de la Colombie et du Mexique

Ce document estime des fonctions de réaction monétaires non-contraintes pour quatre pays d'Amérique latine (Brésil, Chili, Colombie et Mexique) et teste l'existence d'effets non-linéaires dans le comportement des banques centrales. L'analyse couvre la période post-1999 où la politique monétaire se caractérise par le ciblage d'inflation. Nous traitons la question de la présence de racines unitaires dans les données en estimant les règles de politique monétaire dans un cadre de cointégration. Nous testons l'existence d'une cointégration linéaire et non-linéaire de nos variables d'intérêt. Les résultats suggèrent que la spécification non-linéaire ne peut être rejetée pour les données brésiliennes, colombiennes et mexicaines ; elle l'est en revanche dans le cas du Chili. L'estimation de modèles de transition douce par des NLLS et EN-NLLS suggère que la réponse de la banque centrale à la différence entre l'inflation espérée et la cible d'inflation ne change pas selon le régime de politique monétaire au Chili. Elle se durcit au Mexique lorsque l'inflation espérée s'éloigne de la cible. Les réponses semblent s'assouplir au Brésil lorsque la différence entre l'inflation espérée et la cible d'inflation s'accroît ; ce résultat est certainement dû à des chocs d'offre négatifs et à des ajustements à la hausse des cibles dans les premières années de politique monétaire à cible d'inflation. Nous trouvons aussi un effet non-linéaire de la réponse de la Banque centrale au taux de change au Brésil et en Colombie.

Classification JEL : E52 ; C22 ; O54.
Mots-clés : cible d'inflation ; fonction de réaction ; cointégration non-linéaire ; modèles de transition douce.

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Table of contents

Tables

Figures

Do Latin American central bankers behave non-linearly? The experiences of Brazil, Chile, Colombia and Mexico

By
Luiz de Mello, Diego Moccero and Matteo Mogliani[1]

1. Introduction

This paper tests for the presence of non-linearity in monetary reaction functions in four Latin American inflation targeters: Brazil, Chile, Colombia and Mexico. There is a fairly large body of research on policy rules in these countries, but emphasis has so far been placed on the estimation of linear, rather than non-linear, reaction functions (Corbo, 2002; Schmidt-Hebbel and Werner, 2002; Minella *et al.*, 2003; Cerisola and Gelos, 2005; de Mello and Moccero, 2006 and 2008). A growing empirical literature nevertheless suggests the presence non-linearity, at least as far as the experience of the United States (Dolado *et al.*, 2004; Kim *et al*., 2005; Qin and Enders, 2008) and a few European countries (Bec *et al.*, 2002; Bruinshoofd and Candelon, 2004; Taylor and Davradakis, 2006) are concerned.

What are the sources of non-linearity in monetary reaction functions? In theoretical models, non-linear behaviour arises as a result of deviations from the conventional minimisation of quadratic loss functions subject to linear Phillips curves and aggregate demand schedules (Svensson, 1997; Ball, 1999). In particular, the loss function may not be quadratic because of asymmetries in the central bank's response to inflation in different points of the business cycle and/or to the size of deviations of actual inflation from the target (Bec *et al.,* 2002; Dolado *et al.,* 2004; Cukierman and Muscatelli, 2008). It is also possible that the Phillips curve may reflect more complex price-setting mechanisms than those subsumed in a linear specification (Nobay and Peel, 2003).

Non-linearity can also be tested in unrestricted monetary reaction functions. In this case, the functional specification of the reaction function does not depend on the specific parameterisation of the theoretical model describing central bank behaviour. The econometrician is agnostic about the source and type of non-linearity that may emerge in the estimation of the policy rule. Non-linearity may arise in the monetary authority's reaction to changes in inflation (or deviations of inflation expectations from the target) and/or the output gap. Regime switches, which give rise to non-linear behaviour, may be discontinuous, as in the case of threshold models (Tsay, 1998; Hansen and Seo, 2002), or continuous, as in

1. Luiz de Mello and Diego Moccero are in the OECD Economics Department and Matteo Mogliani is in the Paris School of Economics. The authors are indebted to Peter Jarrett, Elena Rusticelli, José Sanchez Fung, Lukas Vogel and the participants of the 13th Meeting of the Latin American and Caribbean Economic Association (LACEA), 20-22 November 2008, Rio de Janeiro, for helpful comments and discussions. They nevertheless remain solely responsible for any remaining errors and omissions. Special thanks go to Anne Legendre for research assistance and to Mee-Lan Frank for excellent technical preparation.

smooth-transition models (Granger and Teräsvirta, 1993; Teräsvirta, 2006). This agnosticism is convenient, because it has often been argued that parametric restrictions imposed on the functional specification of the reaction functions on the basis of structural models of central bank behaviour may generate misspecification biases.

Against this background, we estimate unrestricted non-linear reaction functions for the countries under examination over the post-1999 period, when fully-fledged inflation targeting was adopted as the institutional framework for monetary policy in these countries. We test for linear co-integration among the variables of interest, as well as for neglected non-linearity in the linear specification of the policy rules. We report the results of the estimation of linear and, where applicable, non-linear reaction functions, while controlling for possible endogeneity among the regressors. We estimate smooth-transition models, allowing for non-linearities to arise both in the inflation gap (*i.e.* deviations of expected inflation from the target) and the exchange rate. In doing so, we contribute to the empirical literature by accounting for non-linearity in central bank behaviour using integrated variables. Most of the econometric techniques available to date for testing for non-linearity require the time series to be stationary. This is a restrictive requirement for emerging-market economies, where the relevant variables often exhibit unit roots, which calls for the use of non-linear co-integration techniques that allow for integrated series.

Our main findings are as follows. *First,* we find evidence of neglected non-linearity in conventional linear reaction functions in Brazil, Colombia and Mexico, but not in Chile. *Second,* estimation of smooth-transition models for Brazil, Colombia and Mexico shows that the central bank's response to the inflation gap is invariant across policy regimes in Colombia. It becomes stronger in Mexico and weaker in Brazil, as expected inflation deviates from the target. *Third,* the central bank seems to react to the exchange rate linearly and non-linearly in Brazil and Colombia. *Finally,* the thresholds for regime changes are positive for Colombia and Mexico. But it is negative for Brazil, suggesting that the central bank may have anchored inflation at a level that has been on average somewhat lower than the target, possibly in an attempt to build credibility in the policy regime.

The paper is organised as follows. Section 2 briefly reviews the empirical literature for Latin America. Section 3 describes the data and reports the results of unit root tests. Section 4 reports the results of linear and non-linear co-integration tests and discusses the co-integrating equations. Section 5 concludes.

2. The literature for Latin America: A brief summary

Linear reaction functions have been estimated for the four countries under examination. Minella *et al.* (2003) estimated a forward-looking reaction function for Brazil in levels and augmented it by changes in the nominal exchange rate. On the basis of OLS regressions, they report a greater-than-one coefficient on the inflation gap, a relatively weak response to the output gap and a strong reaction to changes in the exchange rate. Corbo (2002) reports the results of the estimation of linear reaction functions for a number of Latin American countries, including Chile and Colombia. In both countries, the estimated central bank response to changes in the inflation gap is statistically significant in the inflation-targeting period, correctly signed and greater than one on the basis of GMM estimations. Schmidt-Hebbel and Werner (2002) estimated linear exchange-rate-augmented reaction functions for Brazil, Chile and Mexico. The authors report mixed findings on the sign, significance and magnitude of the coefficients on the inflation gap. Monetary policy reaction to the exchange rate is found to be weak.

Most of the empirical literature for Latin America report near-unity coefficients on the lagged interest rate, a finding that may well be due to the presence of unit roots in the interest-rate series. The policy rules are estimated for variables in levels, and little attention is paid to the unit root properties of the data. This is essentially because motivation for the empirical analysis comes from the estimation of Taylor rule-type restricted reaction functions, whose specification is borrowed directly from the literature. An assessment of the time-series properties of data by de Mello and Moccero (2006 and 2008) nevertheless shows that inflation, expected inflation and the interest rate all have unit roots in the four countries under examination in the post-1999 inflation-targeting period. The authors therefore estimated both unrestricted reaction functions in an error-correction setting (de Mello and Moccero, 2006) and restricted functions in first differences derived from reduced-form New Keynesian models (de Mello and Moccero, 2008). Linear co-integration tests suggest that the interest rate, inflation and expected inflation move around a common stochastic trend in these countries. They interpret this finding as evidence of the existence of a stable long-run monetary reaction function.

Attempts to test for non-linearity in the policy rules for the countries under examination have been limited. Corbo (2002) experimented with augmenting the reaction functions to include the square of the inflation gap or interaction terms between the arguments of the reaction function and a dummy variable identifying periods when the output gap was negative. In neither case was there evidence of non-linearity in the reaction function, at least in the case of Chile.

In what follows, we depart from the literature in two main ways. *First*, we discuss the unit root properties of the data and test for linear co-integration among the variables included in the monetary reaction functions. We then test for the presence of neglected non-linearity in these linear models. *Second*, we estimate non-linear reaction functions by using a specific class of smooth-transition models that does not require stationary series. In doing so, we allow for non-linearity to emerge from the existence of different policy regimes, while explicitly modelling transition across policy regimes, rather than from the *ad hoc* augmentation of the reaction function with non-linear terms. By focusing on unrestricted policy rules, we are agnostic about parameter restrictions imposed on the estimating equation.

3. Data and unit-root tests

Data

Our empirical analysis focuses on four Latin American inflation targeters: Brazil, Chile, Colombia and Mexico. Our data set includes monthly observations available from these countries' central banks for the interest rate, inflation expectations, the inflation target and the exchange rate. The sample periods differ across countries, depending on the availability of information on survey-based inflation expectations. It starts in July 2001 for Brazil, September 2001 for Chile, September 2003 for Colombia and November 2000 for Mexico. The end of the sample is March 2008 for Brazil and Chile, and February 2008 for Colombia and Mexico.

The sample period coincides by and large with the adoption of fully-fledged inflation targeting in Brazil (July 1999), Chile and Colombia (September 1999) and Mexico (January 1999). Chile formally adopted inflation targeting in 1990 together with *de jure* central bank independence, but it was not until September 1999 that exchange-rate targeting was formally abandoned. Brazil and Colombia adopted inflation targeting following the floating of the *real* in January 1999 in Brazil and the abandonment of exchange-rate targeting in Colombia in September 1999. Mexico scrapped a narrow exchange-rate band in 1995, but a gradual transition to explicit inflation targeting began in earnest in 1998.

The interest rate is defined in nominal terms as the annualised SELIC rate for Brazil, the TPM rate for Chile, the 90-day deposit (CDT) rate for Colombia and the yield on 28-day CETES bonds for Mexico. These are the policy rates for Brazil and Chile, where monetary policy is conducted primarily through open-market operations. For Colombia and Mexico, we follow the literature and use a money-market rate for Colombia and the rate of return on the most liquid, shortest-term government security in Mexico, where monetary policy was conducted through quantitative bank reserve targets (*corto*) during most of the period of analysis.[2]

Expected inflation is defined as the 12-month-ahead expected consumer price inflation (measured by the IPCA index for Brazil, the IPC indices for Chile and Colombia, and the INPC index for Mexico) available from market surveys conducted by each country's central bank.[3] On the basis of these data, the inflation gap was computed as in de Mello and Moccero (2006). An implicit monthly target was calculated by linearly interpolating the end-year targets. In doing so, for example, the implicit monthly targets for January through November for any given year were defined as the fitted values of a line joining the end-year targets. Deviations of expected inflation from the target were then computed using 12-month leads of the implicit monthly targets.

The exchange rate is defined in units of domestic currency per US dollar. The output gap was computed as the log difference between the actual and the HP-filtered (seasonally-adjusted) industrial production index (the IMACEC, which is an economy-wide activity, rather than industrial production, indicator is used in the case of Chile).

Unit-root tests

Most of the empirical literature on monetary reaction functions for emerging-market economies is agnostic about the time-series properties of the data. Reaction functions are estimated without testing for the stationarity of the relevant series. We depart from this tradition and start by testing for the presence of unit roots using the Elliott, Rothenberg and Stock ADF-GLS test. This test is more efficient than others under the hypothesis of normal residuals. The optimal number of lags was selected on the basis of the Schwartz information criterion, starting with a maximum of 12 lags and testing for normality of the residuals. When the residuals were found to be non-normal for all lags, the Phillips-Perron test (PP) was used for the variables that did not exhibit a trend, and the Schmidt-Phillips test (SP) was used for the trended variables. In these cases, the optimal number of lags was chosen using the Newey-West truncation lag selection criteria.

On the basis of the results reported in Table 1, it appears that the null hypothesis of unit roots cannot be rejected for all variables, except the output gap, which was found to be stationary in levels in all countries barring Mexico. Visual inspection of the Mexican data nevertheless casts doubt over the presence of unit roots in the output gap, a finding that may be driven by the presence of breaks in the series, which are known to bias the parameter estimates required for unit-root testing. To be sure, we used instead the Lee and Strazicich LM test that allows for two breaks in the data.[4] The results (not reported) show that the hypothesis of a unit root can be rejected at the 5% level.[5] All in all, the results of the unit-root tests suggest

2. For a detailed description of monetary-policy instruments in these countries, see Figueiredo, Fachada and Goldenstein (2002) for Brazil; Cifuentes and Desormeaux (2005) and Loayza and Schimdt-Hebbel (2002) for Chile; Uribe (1999), Vargas (2005), Melo and Riascos (2004) and Clavijo (2004) for Colombia; and Ramos-Francia and Torres (2007) and Banco de México (2007) for Mexico.

3. Market surveys are conducted among key financial institutions and consulting firms in the four countries. The surveys include about 90 respondents in Brazil, 35 in Chile, and 30-40 in both Colombia and Mexico (depending on survey wave).

4. This test has the advantage of allowing for breaks both under the null (unit root with breaks) and the alternative hypotheses (stationarity around breaks). See Lee and Strazicich (2003) for more information.

5. The tests for the null hypothesis of unit root with one or two breaks (Lee and Strazicich, 2003 and 2004) were also applied to the other series. The results (available upon request) are consistent with those reported in Table 1.

that variables included in the policy rules should be tested for co-integration. These findings also cast doubt over the appropriateness of including the output gap in monetary reaction functions.

Table 1. **Unit root tests**[1]

Country	Variable	Test	Test statistics
Brazil	r_t	SP	-1.89
	$E_t\pi_{t+12} - \overline{\pi}_{t+12}$	ADF-GLS	-0.06
	e_t	SP	-1.79
	y_t	ADF-GLS	-2.61***
Chile	r_t	ADF-GLS	-1.39
	$E_t\pi_{t+12} - \overline{\pi}_{t+12}$	ADF-GLS	-1.10
	e_t	ADF-GLS	-2.59
	y_t	ADF-GLS	-1.95**
Colombia	r_t	ADF-GLS	-1.05
	$E_t\pi_{t+12} - \overline{\pi}_{t+12}$	ADF-GLS	-0.76
	e_t	ADF-GLS	-1.43
	y_t	ADF-GLS	-2.44**
Mexico	r_t	SP	-1.39
	$E_t\pi_{t+12} - \overline{\pi}_{t+12}$	ADF-GLS	-1.25
	e_t	ADF-GLS	-1.29
	y_t	ADF-GLS	-1.48

1. r_t = interest rate, $E_t\pi_{t+12}$ = 12-month-ahead expected inflation, $\overline{\pi}_{t+12}$ = 12-month-ahead inflation target, e_t = nominal exchange rate, and y_t = output gap. (*), (**) and (***) denote statistical significance at the 10, 5 and 1% levels, respectively.

Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

4. Monetary reaction functions: The co-integrating equations

Linear co-integration

Because the variables of interest were found to have unit roots, we tested for linear co-integration among the interest rate, the inflation gap and the nominal exchange rate in the four countries under examination and estimated a co-integrating equation defined as follows:

$$r_t = \alpha + \beta_1(E_t\pi_{t+12} - \overline{\pi}_{t+12}) + \beta_2 e_t + u_t \, , \tag{1}$$

where r_t is the interest rate, E_t is the expectation operator conditional on the information set available at time *t*, π_t denotes inflation, $\overline{\pi}_{t+12}$ is the 12-month ahead inflation target, e_t is the nominal exchange rate, and u_t is an error term.

Table 2. **Linear co-integration test and estimates[1]**

Coefficients	Brazil	Chile	Colombia	Mexico
α	4.39***	11.66***	4.92***	6.36***
	[1.36]	[1.68]	[1.71]	[2.51]
β_1	0.47**	5.32***	3.62***	3.91*
	[0.21]	[0.56]	[0.69]	[2.13]
β_2	5.01***	-0.013***	0.001	…
	[0.55]	[0.003]	[0.001]	
Trend				-0.03
				[0.03]
No. obs.	81	79	54	88
Co-integration test				
λ_y	-0.05	-0.09	-0.12	-0.18
t-statistic	-3.30	-3.98	-3.04	-4.97
p-value	0.08	0.02	0.14	0.00

1. Standard errors are reported in brackets. The Phillips and Hansen (1990) fully-modified estimator is used to estimate the co-integrating equation. A quadratic-spectral kernel and an automatic bandwidth selection are also implemented. (*), (**) and (***) denote significance at the 10, 5 and 1% levels, respectively. The Banerjee, Dolado and Mestre (1998) procedure is used to test for co-integration. The number of lags for the first-differenced terms in the ECM equation was chosen on the basis of the Schwartz information criterion. Exact *p*-values are provided by Ericsson and MacKinnon (2002).

Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

Equation (1) was estimated by fully-modified OLS (Phillips and Hansen, 1990). The results are reported in Table 2. Parameter estimates are in general signed as expected and of reasonable magnitudes, except for the coefficient on the inflation gap, which is lower than unity in Brazil. In the cases of Mexico and Colombia, parameter estimates are somewhat poorly determined. These results may be indicative of some model misspecification. We therefore shed additional light on this issue by proceeding to test for neglected non-linearity in the linear co-integrating equations.

The Banerjee, Dolado and Mestre (1998) procedure, which involves the estimation of an error-correction equation, was used to test for co-integration.[6] The hypothesis of co-integration is tested on the basis of the *t*-ratio associated with the estimated coefficient on the lagged level of the dependent variable (λ_y). Ericsson and MacKinnon (2002) tabulated the unbiased exact *p*-values for the test statistic. The co-integration tests, reported in the lower panel of Table 2, show that the null hypothesis of absence of co-integration cannot be rejected for Colombia and Brazil (albeit only at the 10% level in this latter case), but this is not the case for Chile. As for Mexico, a model that excludes the exchange rate and includes a

6. The main advantages of this methodology over residuals-based univariate cointegration tests are twofold (Ericsson and MacKinnon, 2002). *First*, the long-run coefficients (on which the hypothesis of cointegration is tested) are not biased, due to the inclusion of the short-run dynamics of the model into the test equation. *Second*, no restrictions are imposed on the long- and the short-run coefficients, unlike the two-step cointegration tests, given that the equilibrium and the dynamic relationships described by the model are estimated simultaneously.

linear trend in the co-integrating equation was found to perform better than the specification used for the other countries. In this case, the null hypothesis of absence of linear co-integration can be rejected.

Non-linear co-integration

Testing for non-linearity against linearity in co-integrating equations

We applied the Breitung (2001) rank procedure to test for the presence of neglected non-linearity in linear co-integration equations.[7] Based on the reaction functions defined by Equation (1), non-linearity may arise as follows:

$$r_t = \alpha + \beta_1 (E_t \pi_{t+12} - \overline{\pi}_{t+12}) + \beta_2 e_t + f^* (E_t \pi_{t+12} - \overline{\pi}_{t+12}, e_t) + u_t \,, \tag{2}$$

where the variables are defined as in Equation (1), $f^*(.,.)$ describes the non-linear part of the model, and u_t is an error term.

The test procedure was implemented as follows. Although the actual specification of f^* is unknown, it can be approximated by Fourier series and neural networks (Granger, 1995). To some extent, this approximation is related to the rank transformation of $(E_t \pi_{t+12} - \overline{\pi}_{t+12}, e_t)$. Therefore, Breitung (2001) proposes the use of a multiple of the rank transformation $\theta \cdot R_T (E_t \pi_{t+12} - \overline{\pi}_{t+12}, e_t)$ instead of $f^*(E_t \pi_{t+12} - \overline{\pi}_{t+12}, e_t)$ itself, where θ is a parameter and R_T is a rank function. The null hypothesis of linear against non-linear co-integration was then tested in three steps. *First*, following Breitung (2001), a linear co-integrating equation was estimated by DOLS (Stock and Watson, 1993), using the specifications in Table 2. *Second*, the residuals from the linear co-integrating equations were regressed on the same regressors and the rank transformation $R_T (E_t \pi_{t+12} - \overline{\pi}_{t+12}, e_t)$, therefore embedding the alternative hypothesis of neglected non-linearity. *Finally*, the score statistic $T \cdot R^2$ was computed to test for the significance of $R_T (E_t \pi_{t+12} - \overline{\pi}_{t+12}, e_t)$; it is distributed asymptotically as χ^2 with *k* degrees of freedom (where k is the number of integrated regressors) under the null hypothesis of linearity.

The results of the rank tests are reported in Table 3. The hypothesis of linear co-integration is strongly rejected for Brazil, Colombia and Mexico (only when two leads and lags are used in the auxiliary DOLS regression for the latter country). On the other hand, the hypothesis of linearity cannot be rejected for Chile at classical levels of significance. On the basis of these findings, the linear co-integration framework appears to be more suitable for the estimation of monetary reaction functions in the case of Chile, but not for the other countries in the sample.

7. We thank Jörg Breitung for sharing his GAUSS code and for useful discussions.

Table 3. **Rank test for neglected non-linear co-integration**[1]

Country	Leads and lags in DOLS equation	Score statistic	*p*-value
Brazil	2	7.011	0.030
	3	11.320	0.003
Chile	2	2.743	0.254
	3	1.152	0.562
Colombia	2	9.265	0.010
	3	10.025	0.007
Mexico	2	7.388	0.006
	3	0.409	0.522

1. The Breitung (2001) test for neglected non-linearity in co-integrating vectors is used. The null hypothesis is that of linear co-integration. The score statistic is distributed as χ^2 χ^2 with 2 degrees of freedom.

Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

Estimating the non-linear co-integrating equations

Non-linear co-integrating equations were estimated for Brazil, Colombia and Mexico as smooth-transition regression (STR) processes, defined as follows:[8]

$$r_t = \alpha_1 + \beta_{11}(E_t\pi_{t+12} - \bar{\pi}_{t+12}) + \beta_{12}e_t + (\alpha_2 + \beta_{21}(E_t\pi_{t+12} - \bar{\pi}_{t+12}) + \beta_{22}e_t)G(\gamma, c, E_t\pi_{t+12} - \bar{\pi}_{t+12}) + u_t \quad (3)$$

where $G(\gamma, c, E_t\pi_{t+12} - \bar{\pi}_{t+12}) = \left(1 - \exp\left\{-\left(\dfrac{\gamma}{\sigma^2_{E\pi_t - \bar{\pi}}}\right) \times \left[\left(E_t\pi_{t+12} - \bar{\pi}_{t+12}\right) - c\right]^2\right\}\right)$.

The inflation gap is the variable describing transition across policy regimes. Selection of this variable was motivated by previous empirical evidence for the countries in the sample that the monetary authority responds more forcefully to deviations of expected inflation from the target than to changes in the output gap. The slope parameter ($\gamma > 0$), which defines the degree of smoothness of the transition function across regimes, $G(\gamma, c, E_t\pi_{t+12} - \bar{\pi}_{t+12})$, is normalised by the variance of the transition variable (inflation gap), as suggested by Granger and Teräsvirta (1993).[9] Parameter c is the threshold location parameter, implying that there are two possible policy regimes. It can easily be seen that, if $\gamma \to 0$, then $G(\bullet) \to 0$ and the reaction function becomes $r_t = \alpha_1 + \beta_{11}(E_t\pi_{t+12} - \bar{\pi}_{t+12}) + \beta_{12}e_t$. By contrast, when $\gamma \to +\infty$, the reaction

8. An advantage of smooth-transition regression models is that they are locally linear, which facilitates interpretation. Also, from a theoretical point of view, the assumption of just a few discrete states can be too restrictive compared with the continuum of states implied by STR models.

9. This normalisation is to avoid numerical problems in the estimation of Equation (3) arising from the fact that γ depends on the scale of the transition variable.

function tends to a threshold model.[10] For intermediate values of γ, monetary responses move smoothly from one regime to another. Moreover, the transition function is a bounded (between 0 and 1), continuous function of the transition variable.

We focus on the exponential specification of the transition function. Unlike the standard framework for stationary series, no test is currently available for selecting among competing non-linear functional forms, which are usually exponential or logistic. However, to the extent that the monetary authority behaves symmetrically for positive and negative deviations of expected inflation from the target, the exponential function would provide a more accurate representation of central bank behaviour than its logistic counterpart, which would imply asymmetric responses. In any case, and from a purely statistical point of view, if the true transition function were in fact logistic, the estimated threshold location parameter (reported below) would tend to be low in comparison with the range of values of the transition variable. Only in such a case would it be advisable to change the specification of the transition function.

There is a large literature on the asymptotic properties and estimating procedures for broad classes of non-linear models, such as those described by Equation (3).[11] In particular, we are interested in a narrower class of non-linear regressions that include integrated series. An asymptotic theory for such models was developed by Park and Phillips (2001). In this seminal paper, the authors analyzed the consistency of the non-linear least squares (NLLS) estimator for two types of functions: integrable functions (I-regular) and asymptotically homogeneous functions (H-regular). The authors show that when the regularity conditions describing these two classes of functions hold, an asymptotic theory applies and the NLLS estimator is consistent.

Nevertheless, the asymptotic properties of the NLLS estimator are valid so long as all integrated regressors are exogenous.[12] Chang, Park and Phillips (2001) extended the theory developed in Park and Phillips (2001) and proposed a simple methodology, closely related to the fully-modified regression model (Phillips and Hansen, 1990) and the canonical co-integration regression procedure (Park, 1992), to deal with endogeneity in a non-linear co-integrating regression.[13] This is important in the present case, because expected inflation and the exchange rate can also react to interest-rate changes. We followed the Chang, Park and Phillips procedure and computed an efficient non-stationary non-linear least squares estimator (EN-NLLS), which corrects for the correlation between the residuals of the long-run regression and innovations in the integrated regressors.[14]

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10. When $\gamma \to +\infty$, the reaction function becomes $r_t = \alpha_1 + \beta_{11}(E_t\pi_{t+12} - \bar{\pi}_{t+12}) + \beta_{12}e_t$, if $E_t\pi_{t+12} - \bar{\pi}_{t+12} = c$, and $r_t = (\alpha_1 + \alpha_2) + (\beta_{11} + \beta_{21})(E_t\pi_{t+12} - \bar{\pi}_{t+12}) + (\beta_{12} + \beta_{22})e_t + u_t$, if $E_t\pi_{t+12} - \bar{\pi}_{t+12} \neq c$.

11. For recent surveys see Granger and Teräsvirta (1993) and Teräsvirta (2006).

12. The authors show that least-square regressions are consistent even when the model is non-linear, but the rates of convergence can differ from the case of regressions with stationary data. Also, with many explanatory variables, the asymptotic distribution of the NLLS estimator is in general non-Gaussian, which implies that standard hypothesis testing is invalid. Only in the special case where the integrated regressors are strictly exogenous is the asymptotic distribution of the NLLS estimator mixed-normal.

13. Chang, Park and Phillips (2001) consider a non-linear regression model that includes stationary and non-stationary regressors, as well as deterministic trends. We applied their asymptotic theory and the correction procedure described in their paper, while ignoring the deterministic trends and stationary components. We thank Peter Phillips for this suggestion.

14. Another theoretically efficient estimator consists of including leads and lags of the first-differenced regressors, as suggested by Choi and Saikkonen (2004). We experimented with the leads-and-lags estimator, using one lead and lag for each non-stationary regressor, and the results (not reported) are

The NLLS estimator was applied using the Marquardt-Levenberg optimisation algorithm.[15] Since many parameters have to be estimated, we proceeded with a simple identification strategy. Because the parameters of the transition function are the most difficult to estimate, we first concentrated on the linear and non-linear parameters, fixing γ γ equal to 1 and c equal to 0 to obtain a set of initial values for these parameters.[16] Then, in a second round of estimations, the restriction on the constant c was relaxed, and a grid search for the optimal initial value of c was implemented using an interval delimited by the minimum and maximum values of the transition variable. These estimated parameters were in turn used as initial values in the third step of the estimations, where the restriction on the slope parameter γ was relaxed, and a grid search was performed again to obtain the optimal initial values for γ within an interval from 1 to 10.[17] Finally, we computed initial values for all relevant parameters and implemented the estimation procedure for the last time on the basis of these new initial values. Insignificant parameters were discarded, and a new estimation loop was run until a final model could be obtained.

A few steps were necessary to compute the EN-NLLS estimator. We first computed the fitted residuals $\hat{u}_t$ from the estimated baseline non-linear model (Equation 3). We then ran a first-order autoregressive regression for the innovations of the integrated variables, $v_t = \Delta(E\pi_t - \bar{\pi}, e_t)$ $(\boldsymbol{v_t} = \Delta \boldsymbol{x_t})$ and kept the fitted residuals, $\hat{\eta}_t$. We subsequently retrieved the one-period-ahead fitted residuals ($\hat{\eta}_{t+1}$) to compute $\hat{\sigma}_{u\eta} = \frac{1}{n}\sum_{t=1}^{n} \hat{u}_t \hat{\eta}'_{t+1}$ and $\hat{\Sigma}_{\eta\eta} = \frac{1}{n}\sum_{t=1}^{n} \hat{\eta}_t \hat{\eta}'_t$. Finally, we defined a modified dependent variable as $r_t^* = r_t - \hat{\sigma}_{u\eta} \hat{\Sigma}_{\eta\eta}^{-1} \hat{\eta}'_{t+1}$ $\boldsymbol{y_t^*} = \boldsymbol{y_t} - \hat{\boldsymbol{\sigma}}_{\boldsymbol{u\varepsilon}} \hat{\boldsymbol{\Sigma}}_{\boldsymbol{\varepsilon\varepsilon}}^{-1} \hat{\boldsymbol{\varepsilon}}'_{\boldsymbol{t+1}}$ and re-estimated the non-linear models as:

$$r_t^* = \tilde{\alpha}_1 + \tilde{\beta}_{11}(E\pi_t - \bar{\pi}) + \tilde{\beta}_{12} e_t + (\tilde{\alpha}_2 + \tilde{\beta}_{21}(E\pi_t - \bar{\pi}) + \tilde{\beta}_{22} e_t) G(\tilde{\gamma}, \tilde{c}, E\pi_t - \bar{\pi}) + u_t^*, \qquad (4)$$

following the identification strategy described in the previous paragraph.

The parameter estimates

The results of the non-linear co-integration models are reported in Table 4. The parameters of interest are signed as expected and of reasonable magnitudes. The NLLS and EN-NLLS estimators produce comparable estimates, suggesting that endogeneity may not be a major problem in these estimations. In the case of Mexico, it was not possible to compute the EN-NLLS estimator, because the algorithm failed to converge.

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consistent with those obtained on the basis of the Chang-Park-Phillips estimator. However, it was difficult to estimate the regressions with more than one lead and lag because of the limited number of observations and the fall in the degrees of freedom resulting from the inclusion of additional parameters.

15. We also experimented with the Gauss-Newton algorithm, but the results did not differ significantly from those obtained with the Marquardt-Levenberg algorithm.

16. As mentioned before, while γ is a free parameter, other than the restriction that it be greater than zero, c must lie between the minimum and the maximum values of the transition variable to be economically interpretable. We therefore fixed γ equal to 1 and c equal to 0, which lies between the maximum and the minimum values of the inflation gap for Brazil and is very close to the minimum value for Colombia and Mexico, as initial values for the NLLS estimates.

17. Even if 10 is very far from infinity, a value for γ at which the smooth-transition model collapses into a threshold model, this is a reasonable upper-band value to keep the grid search computationally feasible.

Table 4. **Non-linear co-integration equations, Brazil, Colombia and Mexico**

Coefficients	Brazil (NLLS)	Brazil (EN-NLLS)	Colombia (NLLS)	Colombia (EN-NLLS)	Mexico (NLLS)
α_1	-10.36**	-10.89**	-	-	11.88***
	[5.02]	[5.35]			[1.09]
β_{11}	6.84**	6.30**	2.64***	2.62***	-
	[3.18]	[3.19]	[0.35]	[0.38]	
β_{12}	12.80***	12.98***	0.003***	0.003***	…
	[2.25]	[2.35]	[0.0001]	[0.0001]	
α_2	21.88***	22.15***	6.31***	6.32***	-5.46***
	[6.33]	[6.79]	[0.19]	[0.18]	[1.53]
β_{21}	-5.99*	-5.59*	-	-	4.36***
	[3.24]	[3.27]			[1.18]
β_{22}	-10.59***	-10.63***	-0.003***	-0.003***	…
	[2.61]	[2.76]	[0.0001]	[0.0001]	
c	-0.39***	-0.40***	0.26***	0.27***	0.63***
	[0.07]	[0.07]	[0.03]	[0.03]	[0.04]
γ	7.01	6.79	3.91	4.03	3.53
Trend					-0.06***
					[0.01]
No. obs.	81	78	54	51	88
Adj-R^2	0.65	0.64	0.56	0.52	0.31
Wald test			3.24	3.34	
p-value			0.07	0.07	

1. Approximated standard errors are reported in brackets. "–" indicates that the coefficient was insignificant and then excluded from the final model. "…" means that variable was not included in the model. The Wald test was performed to test the restriction that $(\beta_{12}+\beta_{22}=0)$ $(\beta_{12}+\beta_{22}=0)$. (*), (**) and (***) denote significance at the 10, 5 and 1% levels, respectively.

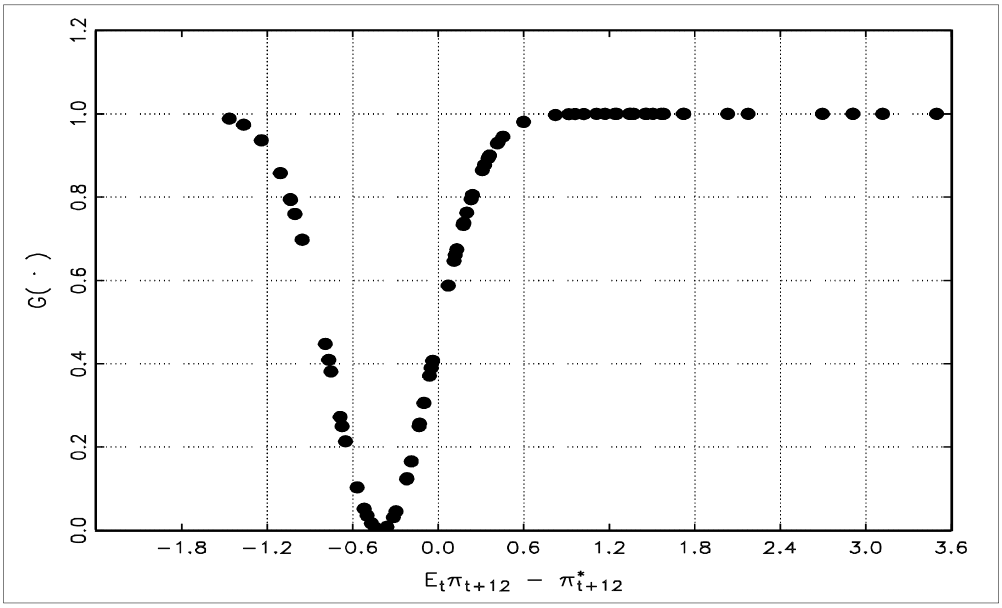
Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

In the case of Brazil, the estimated coefficients on the inflation gap are statistically significant, positively signed in the linear part of the model and negatively signed in the non-linear part. These coefficients suggest that, for an inflation gap close to the estimated threshold for regime change (-0.4), the transition function tends to 0, and the linear part of the model dominates. However, non-linear responses emerge when the inflation gap deviates from the estimated location threshold, and the transition function tends to 1. The net effect, calculated as the sum of the coefficients estimated for both regimes (0.71, on the basis of the EN-NLLS estimates), is nevertheless positive. In other words, the central bank reacts to an increase (decrease) in the inflation gap by tightening (loosening) monetary policy, but this policy response appears to weaken as the inflation gap widens. Central bank responses to the exchange rate also appear to vary across policy regimes. The net effect is positive (2.35, for the EN-NLLS estimates), suggesting that the central bank reacts by hiking the interest rate when the exchange rate depreciates. But this response loses vigour as the inflation gap widens. In other words, the central bank's responses (to the inflation gap and the exchange rate) seem to be stronger when the inflation gap is close to the threshold.

We attribute these findings to the sequence of adverse supply shocks that hit the Brazilian economy in the early phase of inflation targeting, rather than a lack of resolve on the part of the monetary authorities to act decisively when confronted with large inflation surprises. These shocks, such as a severe energy shortage in 2001 and a confidence crisis in the run-up to the presidential election of October 2002, when commitment by the front-running candidate to macroeconomic austerity was in doubt, resulted in a sizeable exchange-rate depreciation and an attendant impact on inflation and inflation expectations. Cognizant that monetary action under such circumstances would be overly destabilising, the monetary authority opted for pursuing adjusted targets while committing to tackling the second-round effects of exchange-rate devaluations on inflation.[18]

Two final observations are noteworthy in the case of Brazil. *First*, the estimated threshold is negative, as noted above, albeit small in magnitude (-0.4). We interpret this finding as reflecting the fact that the central bank may have attempted to consolidate credibility in inflation targeting by seeking to anchor inflation expectations somewhat below the target. *Second*, the slope parameter, which reflects the smoothness of the transition function, is high in magnitude (6.8, on the basis of the EN-NLLS estimates).[19] This implies that regime shifts are fairly swift, once the transition variable has reached the estimated threshold. This is illustrated by the transition function and regime changes depicted in Figures 1 and 2.

Figure 1. **Brazil: Estimated transition function**

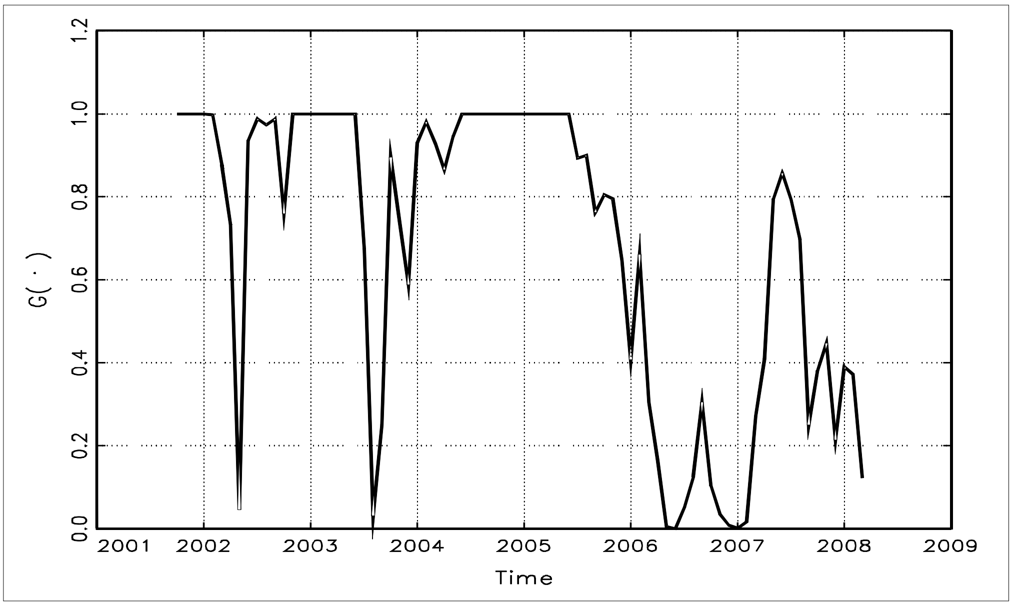


Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

18. See OECD (2005) and Bevilaqua *et al*. (2008) for descriptive accounts of these episodes.

19. The significance of this parameter is not reported, because, as Saikkonen and Choi (2004) pointed out, conventional hypothesis testing cannot be used for the parameters inducing non-linearity due to identification issues.

Figure 2. **Brazil: Estimated regime changes**



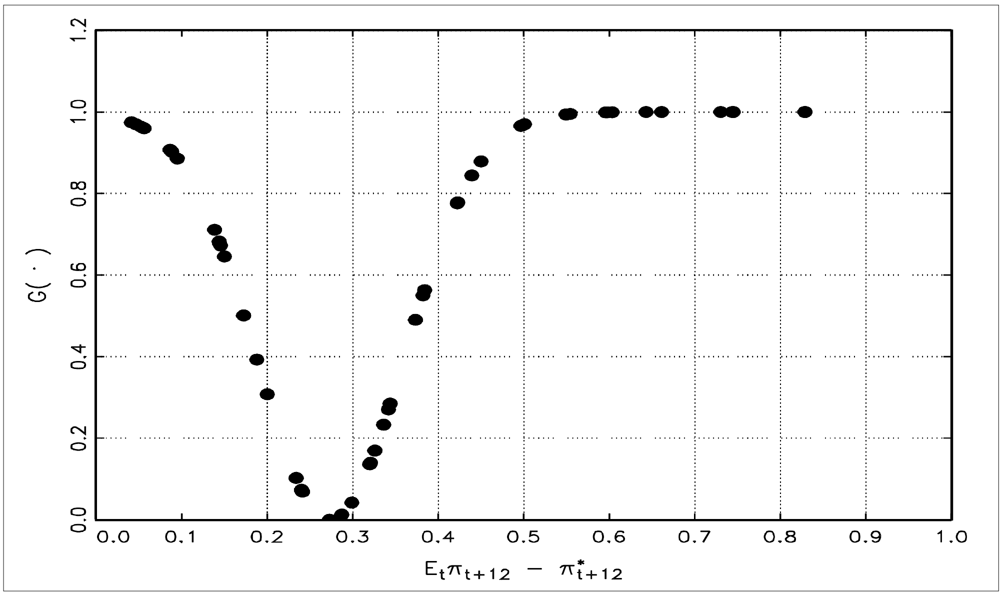
Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

The estimated transition function can be used to assess the behaviour of the central bank as the inflation gap approaches the edges of the inflation-target band (Figure 1). In Brazil, the central bank operates in the "lower-response" regime, where the value of the transition function is 1, when expected inflation exceeds the central target by about 0.5 percentage points or falls short of it by about 1.4 percentage points. This suggests a policy regime shift that takes place within a range of values for the inflation gap that is much narrower than the formal width of the tolerance band around the central target, which is currently plus or minus 2 percentage points. Also, central bank responses have strengthened since March 2005, because the transition function has never reached the value of 1, when the "low-response" regime dominates, since then, and its values have been lower on average (Figure 2).

With regard to Colombia, the inflation surprise coefficient is only significant in the linear part of the model. In this case, as opposed to Brazil, the central bank's response to the inflation gap does not change across regimes. Nevertheless, this coefficient is positive and higher than one, as expected. As for the exchange rate, policy responses are positive in the linear part of the model and negative in the non-linear part. We nevertheless cannot reject the null hypothesis of equality of these coefficients (tested through a Wald test); as a result, only the constrained coefficients are reported.[20] This implies that the net effect is nil, suggesting that the monetary authorities react to exchange-rate depreciations by increasing interest rates when the inflation gap is around the threshold level, but responses weaken as the inflation gap widens. Moreover, the transition parameter is positive (0.3, on the basis of the EN-NLLS estimates) and the slope parameter is fairly high (4.0, on the basis of the EN-NLLS estimates), suggesting that regime shifts are fairly swift, as in the case of Brazil. The estimated transition function and regime changes are depicted in Figures 3 and 4, respectively. In particular, the increase in the value of the transition function suggests that authorities may be increasingly reluctant to use monetary policy as an exchange-rate management tool (Figure 4).

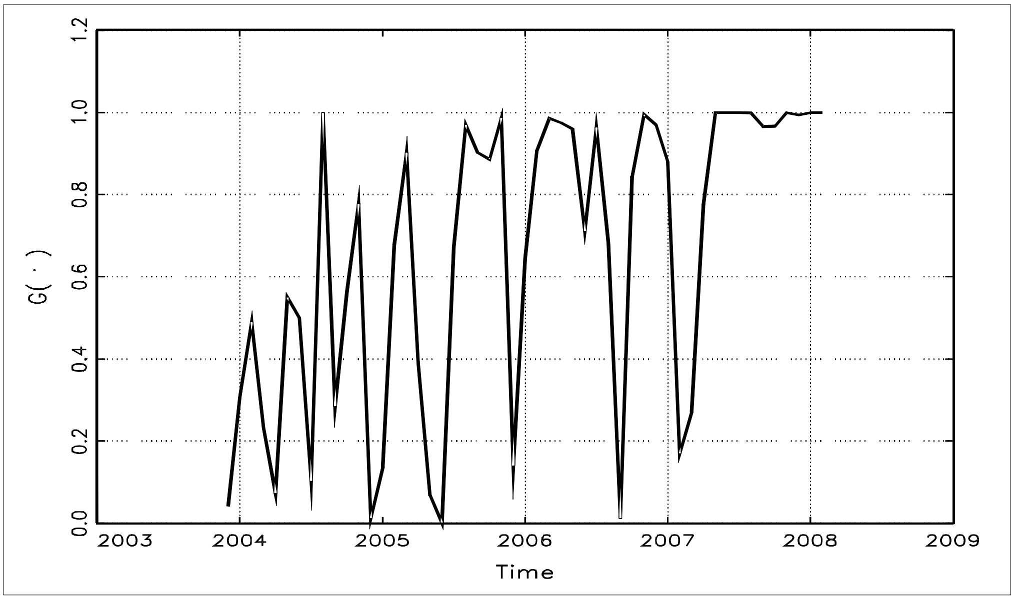
20. It follows from Theorems 4.1 and 5.1 in Park and Phillips (2001) and from Chang, Park and Phillips (2001) that standard hypothesis testing, such as Wald tests, is valid for both linear and non-linear parameters.

Figure 3. **Colombia: Estimated transition function**



Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

Figure 4. **Colombia: Estimated regime changes**

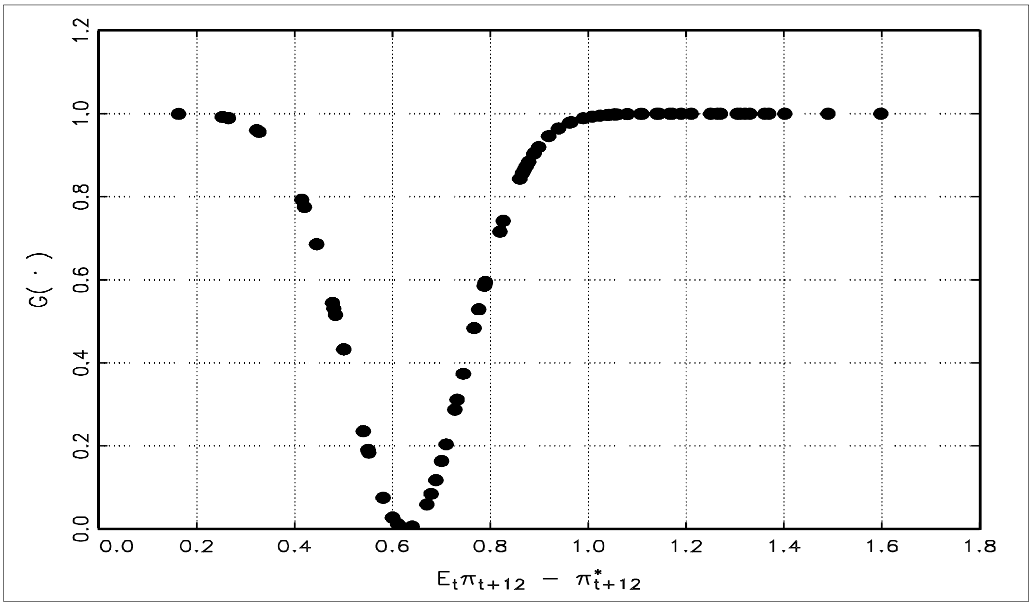


Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

As motivated by the linear co-integration analysis reported above, in the case of Mexico, a different specification was used for the reaction function, which excludes the exchange rate and includes a time trend. The NLLS estimation results show that the inflation gap is statistically different from zero only in the non-linear part of the model, where it is positively signed and greater than one in magnitude, as expected. Unlike the findings for Brazil and Colombia, the Mexican monetary authority does not seem to react to the inflation gap when it is close to the threshold of 0.6 percentage point. However, sizeable inflation gaps prompt a strong response by the central bank. Finally, the estimated transition threshold is positive, as in the case of Colombia, and the smoothness parameter is fairly high, as in the other countries.

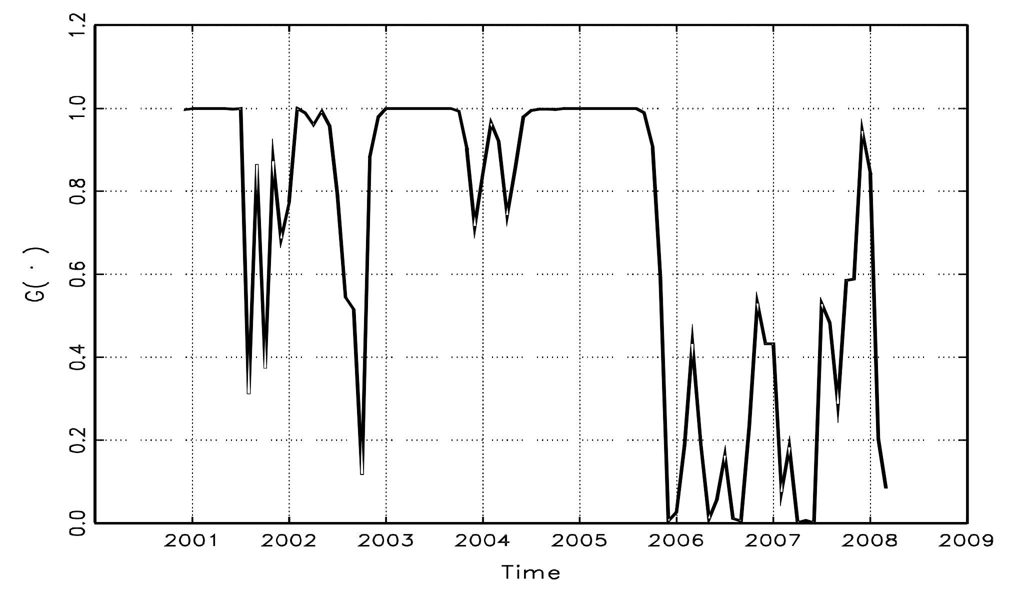
The estimated transition function and regime changes are depicted in Figures 5 and 6, respectively. According to the estimated transition function, a shift to the "stronger response" regime takes place when expected inflation exceeds the central target by around 1 percentage point, which coincides with the ceiling of the tolerance band around the central target. There is also evidence of non-linear behaviour as expected inflation exceeds the central target by about 0.2 percentage points. This asymmetry may reflect the central bank's evaluation of risks associated with deviations of inflation expectations from the central target. The central bank may evaluate the deflationary risks related to small negative inflation gaps, which call for monetary loosening, to be higher than the inflationary risks associated with positive inflation gaps, which call for a tightening of the policy stance. Moreover, the fall over time in the value of the transition function suggests that the monetary authorities reacted particularly aggressively to the inflation gap until mid-2005 (Figure 6).

Figure 5. **Mexico: Estimated transition function**



Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

Figure 6. **Mexico: Estimated regime changes**



Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

Testing for non-linear co-integration

We tested for stationarity in the residuals of the non-linear reaction functions estimated above to make sure that the variables of interest do co-integrate. Many procedures are available in the literature to test for linear co-integration. In a recent paper, Choi and Saikkonen (2004) extended this analysis to the non-linear case.[21] The authors proposed a test statistic based on the residuals of the non-linear equations, where non-linear co-integration is the null hypothesis against an alternative of no co-integration.[22] However, tabulations of this statistic are impractical due to non-linearity, a problem that cannot be solved through bootstrapping or sub-sampling methods. The authors therefore proposed a new test statistic using sub-residuals and the Bonferroni procedure, defined as:

$$C^{b,i}_{EN-NLLS} = b^{-2}\hat{\omega}^{-2}_{u^*}\sum_{t=i}^{i+b-1}\left(\sum_{j=i}^{t}\hat{u}^*_j\right)^2 , \qquad (5)$$

where $\hat{\omega}^2_{u^*}$ is a consistent estimator of the long-run variance ($\omega^2_{u^*}$) of the residuals from Equation (4) ($\hat{u}^*_t$), b is the block size of sub-residuals, and i denotes the starting point of the sub-residuals.

21. We thank In Choi for sharing his GAUSS code.

22. The procedure is akin to the Shin (1994) test for the null of cointegration, following the tradition of the Kwiatkowski, Phillips, Schmidt and Shin (1992) univariate test (henceforth KPSS) for the null of stationarity.

The test statistics is computed as $C^{b,i}_{EN-NLLS} = \max(C^{b,i_1}_{EN-NLLS},...,C^{b,i_M}_{EN-NLLS})$, whose α α-level critical values are taken from the distribution of $\int_0^1 W^2(s)ds$ $\int_0^1 W^2(s)ds$, where W is a standard Brownian motion, using level α/M $\frac{\alpha}{M}$. For a given block size b, the authors propose a simple rule to choose the optimal number of starting points of sub-residuals $(i_1,...,i_M)$ $(i_1,...,i_M)$ and the number of sub-residuals-based tests (M) used in the Bonferroni procedure. In addition, the selection of the block size b can be carried out by using a fixed rule, such as by fixing $b = \lfloor T^{\delta} \rfloor$, with $(0 \leq \delta \leq 1)$ $(0 \leq \delta \leq 1)$, or a minimum-volatility rule. In the latter case, b should be chosen so as to minimise the standard deviation of the test statistics for each value of b from $b_i = b_{small}$ $b_i = b_{small}$ to $b_i = b_{big}$ $b_i = b_{big}$ (Romano and Wolf, 2001), with $b_i = (small + m,...,big - m)$ $i = (small + m, ..., big - m)$.[23]

The results of the non-linear co-integration tests are reported in Table 5. The null hypothesis of non-linear co-integration cannot be rejected at the adjusted 5% level for all countries, regardless of whether the fixed or the minimum-volatility rules are used. These findings strongly suggest the existence of a stationary relationship in our non-linear monetary reaction functions for Brazil, Colombia and Mexico.

Table 5. **Choi and Saikkonen test for non-linear co-integration**

Country	Rule	Block size	M	*p*-value	Adjusted 5% level (α/M)
Brazil	FX	51	2	0.10	0.025
	MV	43	2	0.51	0.025
Colombia	FX	35	2	0.36	0.025
	MV	25	3	0.07	0.017
Mexico	FX	57	2	0.07	0.025
	MV	48	2	0.31	0.025

1. The null hypothesis is that of non-linear co-integration (stationary residuals). The criterion for non-rejection of the null at the 5% level of significance is *p*-value > Adjusted 5% level (α/M). FX and MV denote the fixed and minimum-volatility rules, respectively. The fitted residuals from EN-NLLS estimates ($\hat{u}^*_t$) are used for Brazil and Colombia, while the residuals from NLLS estimates ($\hat{u}_t$) are used for Mexico.

Source: Data available from the central banks of Brazil, Chile, Colombia and Mexico, and authors' estimations.

5. Conclusions

This paper estimated unrestricted monetary reaction functions for Brazil, Colombia, Chile and Mexico using monthly data for the post-1999 inflation-targeting period. We tested for co-integration among the inflation gap, the interest rate and the exchange rate due to the presence of unit roots in these series. Strong evidence was found of neglected non-linearity in the linear co-integrating equations for Brazil, Colombia and Mexico on the basis of the Breitung test, as well as of non-linear co-integration in the data for these three countries using the Choi-Saikkonen test. Smooth-transition models with exponential transition functions, where the transition variable is the inflation gap, were estimated using the non-linear least

23. As in Romano and Wolf (2001), the value of m is set = 2 by Choi and Saikkonen (2004).

squares (NLLS) and the efficient non-stationary non-linear least squares (EN-NLLS) methodologies developed by Park and Phillips (2001) and Chang, Park and Phillips (2001). In the case of Chile, only linear co-integration was tested for, given that no evidence of neglected non-linearity was found.

The results reported above suggest that central banks react to increases in the inflation gap by tightening monetary policy. Central bank behaviour is linear in Colombia, because the estimated policy response is statistically insignificant in the non-linear part of the model, and in Chile, where statistical tests rejected non-linearity. In both cases the coefficients are greater than one, as expected. Evidence of non-linear monetary responses was found for Brazil and Mexico. In Brazil, a negative coefficient on the inflation gap in the non-linear part of the model suggests that policy responses weaken when the inflation gap widens. Nevertheless, the net effect across policy regimes is still positive. In addition, central bank responses have strengthened since March 2005, because the transition function has never reached the value of 1, when the “low-response” regime dominates, since then, and its values have been lower on average. By contrast, the central bank of Mexico does not seem to respond to the inflation gap when it is close to the estimated threshold parameter, while policy responses become stronger as the inflation gap widens.

As for the exchange rate, there also appears to be different policy regimes in Brazil, with a still positive net effect across regimes. In the case of Colombia, however, the estimated coefficients were found to be of the same magnitude and opposite signs in both regimes, suggesting that the overall response is muted, once non-linear behaviour is taken into account. All in all, in both countries a nominal exchange-rate depreciation seems to trigger a tightening of monetary policy. On the basis of the linear co-integration analysis, the central bank’s response to the exchange rate is very small in magnitude in Chile.

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