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MONETARY POLICY AND INFLATION EXPECTATIONS IN LATIN AMERICA: LONG-RUN EFFECTS AND VOLATILITY SPILLOVERS

ECONOMICS DEPARTMENT WORKING PAPERS No. 518

By
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ABSTRACT/RÉSUMÉ

Monetary policy and inflation expectations in Latin America: Long-run effects and volatility spillovers

The current monetary policy framework in several Latin American countries, combining inflation targeting and a floating exchange-rate regime, has contributed to disinflation by anchoring expectations around low, stable levels. This paper uses co-integration analysis to estimate simultaneously a monetary reaction function and the determinants of expected inflation for Brazil, Chile, Colombia and Mexico in the post-1999 period. It also tests for the presence of volatility spillovers between the monetary stance and inflation expectations based on M-GARCH modelling. The results of the empirical analysis show that: *i)* there are long-term relationships between the interest rate, expected inflation and the inflation target, suggesting that monetary policy has been conducted in a forward-looking manner and helped anchor inflation expectations in the countries under examination, and *ii)* greater volatility in the monetary stance leads to higher volatility in expected inflation in Brazil, Colombia and Mexico, suggesting that interest-rate smoothing contributes to reducing inflation expectations volatility. No volatility spillover effect was detected in the case of Chile.

JEL classification: O54, E52, C22
Key words: Brazil, Chile, Colombia, Mexico, inflation targeting, multiple co-integration, volatility spillovers, M-GARCH modelling

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Politique monétaire et anticipations d'inflation en Amérique latine: Effets à long terme et spillovers de volatilité

Le cadre courant de la politique monétaire dans plusieurs pays d'Amérique latine, qui combine un ciblage d'inflation et un régime de taux de change flottant, a contribué à la désinflation par ancrage des anticipations d'inflation à un niveau bas et stable. Ce document utilise une analyse de co-intégration pour estimer simultanément une fonction de réaction da la politique monétaire et les déterminants des anticipations d'inflation pour le Brésil, le Chili, la Colombie et le Mexique depuis 1999. Des tests sont aussi présentés sur la présence d'effets de spillover de volatilité entre la politique monétaire et les anticipations d'inflation, en s'appuyant sur le modèle M-GARCH. Les résultats de l'analyse empirique montrent que : *i*) il existe des relations de long terme entre le taux d'intérêt, les anticipations d'inflation et la cible d'inflation. Cela suggère que la politique monétaire a été conduite selon une méthode prospective et a assuré l'ancrage des anticipations d'inflation; et *ii*) une plus grande volatilité de la politique monétaire entraîne une plus importante volatilité des anticipations d'inflation au Brésil, en Colombie et au Mexique, cela laisse penser qu'un lissage des taux d'intérêt contribue à une réduction de la volatilité des anticipations d'inflation. Aucun spillover de volatilité n'a été détecté dans le cas du Chili.

JEL classification : O54, E52, C22
Mots clés : Brésil, Chili, Colombie, Mexique, ciblage d'inflation, co-intégration multiple, spillovers de volatilité, M-GARCH

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Monetary policy and inflation expectations in Latin America: Long-run effects and volatility spillovers

Luiz de Mello and Diego Moccero[1]

1. Introduction

The process of disinflation that started in the 1990s was not confined to the OECD countries. Inflation also fell steadily in many emerging markets, including the Latin American economies that had hitherto been plagued by chronically high inflation. Disinflation in commodity (especially oil) prices and globalisation are important explanatory factors (Rogoff, 2003). Changes in the nature of shocks hitting the economy and the adoption of a more forward-looking monetary policy stance have also been argued to contribute to lower volatility in inflation and output, at least as far as the experience of the United States is concerned (Ahmed, Levin and Wilson, 2002; Moreno, 2004; Stock and Watson, 2002; Cecchetti, Flores-Lagunes and Krause, 2004; Boivin and Giannoni, 2005). In the case of emerging-market economies, disinflation owes much to pro-market reforms and trade liberalisation, which underpinned structural adjustment in the 1990s. The adoption of inflation targeting as the framework for monetary policymaking has been instrumental in achieving price stability, especially after the abandonment of different forms of exchange-rate targeting since 1999 in most countries (Sterne, 2002). With greater exchange-rate flexibility, the policy question of how to gain monetary credibility by using domestic targets has featured prominently in the policy debate in the region.

The objective of this paper is two-fold: *First*, it empirically tests whether the current monetary policy framework – combining inflation targeting and a floating exchange-rate regime – has contributed to anchoring inflation expectations around the pre-announced targets in a sample of four Latin American countries: Brazil, Chile, Colombia and Mexico. The determinants of expected inflation and Taylor-type monetary reaction functions have been estimated for several countries in the region within a reduced-form single-equation framework (Schmidt-Hebbel and Werner, 2002; Minella *et al.*, 2003). Instead, this paper assesses more thoroughly the time-series properties of the data and uses multiple co-integration analysis to test for the presence of long-run co-movements among the interest rate, the inflation target and inflation expectations.

Second, the paper also tests for the existence of volatility spillover effects between the monetary stance and inflation expectations.[2] This hypothesis – which to our knowledge has not been tested empirically for Latin American countries – is important, because the monetary stance has become less volatile in the countries under examination as a result of the abandonment of exchange-rate targeting. In a fixed exchange-rate regime, monetary policy is set so as to defend a pre-announced nominal target for the exchange rate. Therefore, in a flexible exchange-rate regime, where monetary policy is conducted with the

1. The authors want to thank, without implicating, Fabio Giambiagi, Peter Jarrett, Val Koromzay, Rodrigo Valdés and Annabelle Mourougane for helpful comments and discussions, Anne Legendre for research assistance, and Mee-Lan Frank for excellent technical assistance.

2. This is, the extent by which volatility in one variable is affected by volatility in another variable.

primary aim of price stability around a pre-announced target, the presence of spillover effects running from the interest rate to expected inflation would call for greater persistence in the monetary stance. This could be achieved through interest-rate smoothing, whereby the monetary authority avoids sudden deviations in the policy interest rate from its fundamental determinants in order to anchor expectations.

The main findings of the paper are as follows:

- Based on conventional co-integration analysis, there are long-run relationships among the interest rate, the inflation target and inflation expectations in Brazil, Chile and Mexico. This suggests that the conduct of monetary policy in a regime characterised by inflation targeting and floating exchange rates has been forward-looking and effectively anchored inflation expectations in the countries under examination.

- Based on M-GARCH modelling, there appears to be volatility spillover effects running from the monetary stance to expected inflation in Brazil and Colombia. In these countries, a more volatile monetary stance leads to greater volatility in inflation expectations, making it more difficult for the monetary authority to anchor expectations. No volatility spillover effect was detected in the case of Chile, and there appears to be bi-directional effects in the case of Mexico. As a result, interest-rate smoothing, whereby the monetary authority avoids sudden changes in the policy interest rate above and below its fundamental determinants, is advisable for Brazil, Colombia and Mexico as a means of anchoring volatility in inflation expectations.

The paper is organised as follows. Section 2 presents and describes the data and the univariate time-series properties of the main variables of interest. Section 3 assesses the effectiveness of the monetary policy regime in anchoring expectations using co-integration analysis. The presence of spillover effects between the interest rate and expected inflation is tested in Section 4. Section 5 concludes.

2. Data description and unit root tests

Data

The empirical analysis presented below focuses on four Latin American inflation targeters: Brazil, Chile, Colombia and Mexico. Given the short span of time since the adoption of inflation targeting in the aforementioned countries, monthly observations are used in this study. Monthly data are available from the countries' respective central banks. The variables used in the empirical analysis are the interest rate, expected inflation, the inflation target, the exchange rate, the output gap and deviations of expected inflation from the target. 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 28-day CETES rate for Mexico. Expected inflation is defined as the 12-month-ahead 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. The exchange rate is defined as the 12-month percentage change in the nominal exchange rate (in units of domestic currency per U.S. dollar). The output gap is computed as the log difference between the actual and the HP-filtered (seasonally-adjusted) industrial production index.[3] Finally, because the inflation target is set by the monetary authorities for end-year CPI inflation, an implicit target was calculated for each month by linearly interpolating the end-year targets. Then, 12-month leads of this implicit monthly target were used to compute the deviations of expected inflation from the target (one of the focus variables in the M-GARCH analysis).

3. All percentage changes were multiplied by 100 to facilitate convergence in the estimation algorithms.

Period of analysis

The sample periods considered in the empirical analysis reported below were selected based on the date at which information on private-sector inflation forecasts started to be collected: July 2001 in Brazil, September 2001 and 2003 in Chile and Colombia, respectively, and November 2000 in Mexico. Fully-fledged inflation targeting was formally implemented in July 1999 in Brazil, September 1999 in Chile and Colombia, and January 1999 in Mexico. Although Chile formally adopted inflation targeting in 1990 as part of a more comprehensive reform package that also granted the central bank *de jure* autonomy, it was not until September 1999 that exchange-rate targeting was 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. While Mexico had already scrapped a narrow exchange-rate band in 1995, it was only in 1998 that a gradual transition to explicit inflation targeting began.[4]

The policy frameworks currently in place in all four countries exhibit the key features of fully-fledged inflation targeting (Fracasso, Genberg and Wyplosz, 2003), including: a public announcement of inflation targets and institutional commitment to reaching the pre-announced targets; a flexible exchange-rate regime and an absence of nominal anchors other than inflation as the main goal of monetary policy; an independent (*de jure* or *de facto*) monetary authority; and a transparent communication framework, with the publication of monetary policy committee minutes and Inflation Report-type documentation.[5]

Unit root properties of the data and descriptive statistics

Assessing the unit root properties of the variables is not only a pre-requisite for testing for co-integration, but it is also relevant to determine the appropriate transformation of the data needed to test for the existence of volatility spillovers between the monetary stance and inflation expectations. To test for the presence of unit roots in the data, the Phillips-Perron (PP) test was applied in the case of the variables that were found to exhibit no trend, while the Schmidt-Phillips (SP) test was used for the remaining variables (Table 1).[6] These tests are particularly attractive, because they are robust to the presence of heteroscedasticity in the residuals. The number of lags was selected on the basis of the Newey-West truncation lag selection criteria.[7] In the case of Chile, the unit root test was not applied to the inflation target and the difference between expected inflation and the target, because the target was kept unchanged at 3% over the estimation period. For Colombia, no unit root test was performed for the inflation target, because expected inflation was found to be trend-stationary; as a result, it cannot co-integrate with the target. The variables that were found to contain a unit root were differenced and unit root tests were applied to the transformed variables to verify that the variables in levels were all integrated

4. See Schmidt-Hebbel and Werner (2002) for more information on Brazil, Chile and Mexico, and Vargas (2005) for Colombia.

5. See Mishkin (2000) and Svensson (2002) for more discussion on the required institutional features of fully-fledged inflation targeting.

6. The SP test allows for comparing directly the null hypothesis of a variable being difference-stationary with drift against the alternative that the series is trend-stationary. For the PP test, the variables were first centred. This allows for testing the null hypothesis of a random-walk without drift against the alternative of stationarity around zero.

7. This is based on the number of observations used in the test regression, and consists of choosing the largest integer not exceeding $4(T/100)^{(2/9)}$. The Newey-West truncation lag was set to 3 in all cases.

Table 1. **Unit root tests: Phillips-Perron (PP) and Schmidt-Phillips (SP) statistics**[1]

Country	Variable	Test	Tau
Brazil	r_t	PP	-0.37
	$E_t\pi_{t+12}$	PP	-0.81
	π^*_{t+12}	PP	-0.09
	e_t	SP	-2.28
	y_t	PP	-2.81***
	$(E_t\pi_{t+12} - \pi^*_{t+12})$	PP	-2.29**
Chile	r_t	PP	-1.38
	$E_t\pi_{t+12}$	PP	-0.72
	e_t	SP	-1.96
	y_t	PP	-3.35***
Colombia	r_t	SP	-1.66
	$E_t\pi_{t+12}$	SP	-4.23***
	e_t	SP	-1.81
	y_t	PP	-3.78***
	$(E_t\pi_{t+12} - \pi^*_{t+12})$	SP	-3.87***
Mexico	r_t	SP	-1.34
	$E_t\pi_{t+12}$	SP	-0.99
	π^*_{t+12}	SP	-0.89
	e_t	PP	-1.65*
	y_t	PP	-3.02***
	$(E_t\pi_{t+12} - \pi^*_{t+12})$	PP	-1.04

1. r_t = interest rate, $E_t\pi_{t+12}$ = 12-month-ahead expected inflation, π^*_{t+12} = 12-month-ahead inflation target, e_t = 12-month percentage change in the exchange rate, y_t = output gap, and $(E_t\pi_{t+12} - \pi^*_{t+12})$ = difference between expected inflation and the target. (*), (**) and (***) denote statistical significance at the 10%, 5% and 1% levels, respectively. The samples are: 2001:7 to 2006:1 for Brazil, 2001:9 to 2006:1 for Chile, 2003:9 to 2005:12 for Colombia and 2000:11 to 2006:1 for Mexico.

Source: Data available from the Central Banks of Brazil, Chile, Colombia and Mexico, and authors' calculations.

of order one (results not reported). As such, based on the results reported in Table 1, the interest rate was differenced in all countries, as well as expected inflation in the case of Chile and the difference between expected inflation and the target for Mexico. In the case of Colombia, this variable was trend-adjusted (using the HP filter) and kept in levels for Brazil.

Table 2 presents selected descriptive statistics for the two focus variables used in the M-GARCH estimations: the interest rate, measuring the monetary stance, and the difference between expected inflation and the target.[8] Accordingly, interest-rate volatility has been higher in Mexico and Brazil than in Chile and Colombia. In almost all cases, there appears to be excess kurtosis, indicating that the tails of the distribution of the relevant variables are thicker than those of a normal distribution. Also, gauged by the first-order autocorrelation coefficients and the Ljung-Box (LB) tests, the relevant variables appear to exhibit considerable autocorrelation in Brazil and Chile and less so in the cases of Mexico and Colombia. Autocorrelation is also present in the squared variables for Brazil, Chile and Colombia, which suggests the existence of non-linearities in the data when the variables are defined in levels, possibly due to a changing second-order conditional moment.

Excess kurtosis and non-linearity are pre-conditions for the existence of volatility spillovers. The last two lines in Table 2 report the estimated impact of one lagged variable on the present value of the other, both in squared form. Spillovers are easily seen for Brazil and, to a lesser extent, for Mexico. For Chile and Colombia, these basic statistics are not powerful enough to capture such effects. All in all, these results justify modelling together the first- and second-order conditional moments of the data, allowing for the existence of volatility spillovers effects.

3. Is monetary policy anchoring expectations?

Co-integration analysis

Visual inspection of the interest rate, the inflation target and the inflation expectations series suggests that these variables have tended to move together in the countries under examination, following the adoption of inflation targeting and the liberalisation of the exchange-rate regimes (Figures 1-4). These co-movements provide *prima facie* evidence that monetary policy has been successful in anchoring inflation expectations. But, to be sure, a more formal test is required, consisting of estimating long-run relationships among these variables using co-integration analysis. This is the appropriate technique to use to estimate an expected inflation equation and a monetary reaction function for the countries under examination, because the variables of interest were found to be integrated of order one.

8. For Chile, as the inflation target was constant over the reference period, attention is focused on expected inflation.

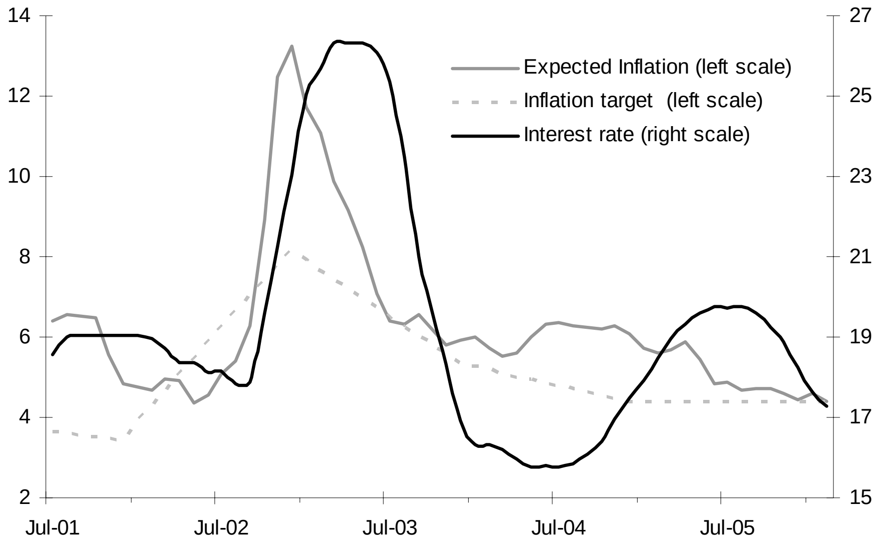
Table 2. **Descriptive statistics**[1]

	Brazil		Chile		Colombia		Mexico	
	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$	r_t	$E_t\pi_{t+12}$	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$
No. of obs.	55	55	53	53	28	28	63	63
Mean	0.01	1.10	-0.04	-0.01	-0.05	0.00	-0.13	-0.01
Std. Dev.	0.81	1.36	0.22	0.17	0.10	0.12	0.94	0.17
Excess Kurtosis	2.15	0.98	1.78	0.44	3.42	-0.33	1.35	3.53
Skewness	-0.17	0.78	-1.31	-0.66	-1.42	0.51	-0.85	-1.09
Normality test (KS)	0.15***	0.12**	0.30***	0.17***	0.14***	0.10***	0.10**	0.14***
ρ(1)	0.81	0.86	0.72	0.31	0.23	0.19	0.16	-0.14
LB(6)	75.94***	83.08***	82.06***	11.47**	2.71	4.48	4.69	19.85***
ρ2(1)	0.64	0.78	0.25	0.15	0.05	-0.13	0.31	0.04
LB2(6)	36.07***	51.07***	11.59**	4.40	0.33	2.42	14.92**	27.51***
$Corr\left[r_t^2, (E_t\pi_{t+12} - \pi^*_{t+12})^2_{t-1}\right]$	0.28**		0.07		-0.20		0.15	
$Corr\left[r_{t-1}^2, (E_t\pi_{t+12} - \pi^*_{t+12})^2_{t}\right]$	0.44***		-0.09		0.22		0.37***	

1. KS is the Kolmogorov-Smirnov test of the null hypothesis of normality. *ρ(1)* and *ρ2(1)* denote the autocorrelations of order one for the variables in levels and squared, respectively. LB(6) and LB2(6) are the Ljung-Box statistics for autocorrelation of order 6 calculated for the variables in levels and squared, respectively. (*), (**) and (***) denote statistical significance at the 10%, 5% and 1% levels, respectively. The samples are: 2001:7 to 2006:1 for Brazil, 2001:9 to 2006:1 for Chile, 2003:9 to 2005:12 for Colombia and 2000:11 to 2006:1 for Mexico.

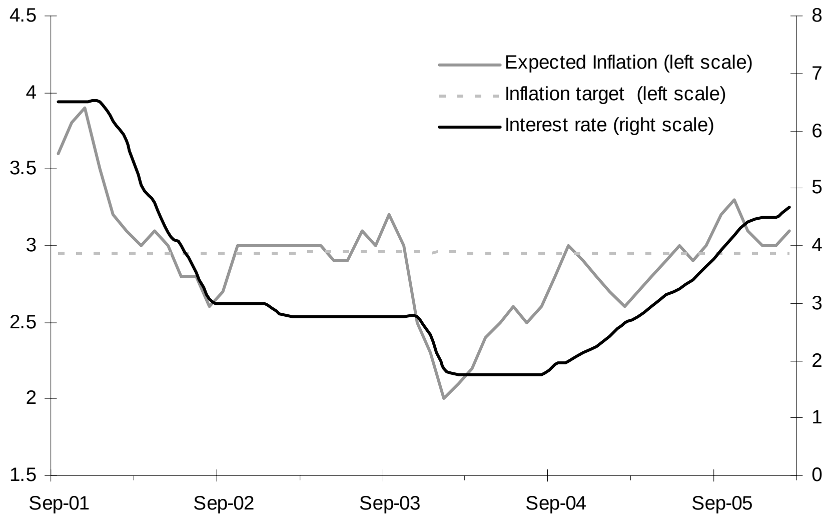
Source: Data available from the Central Banks of Brazil, Chile, Colombia and Mexico, and authors' calculations.

Figure 1. **Brazil: Monetary stance and expected inflation**



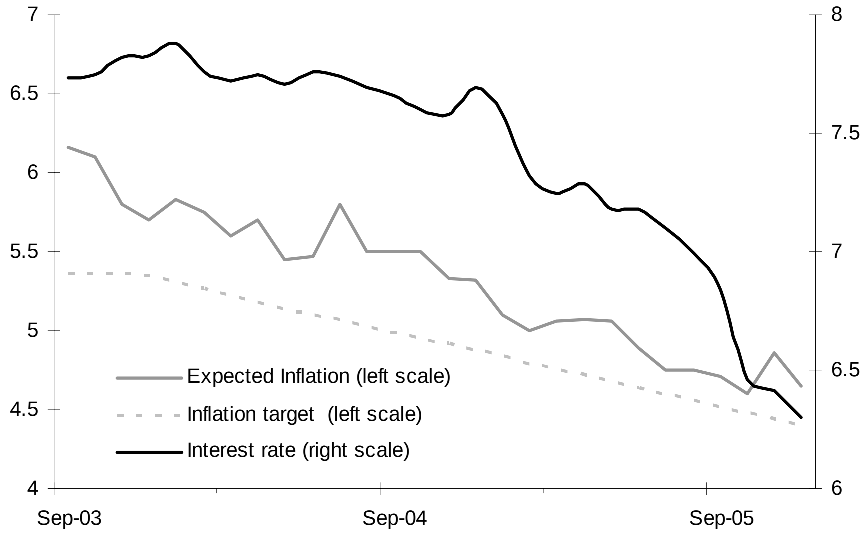
Source: Central Bank of Brazil and authors' calculations.

Figure 2. **Chile: Monetary stance and expected inflation**



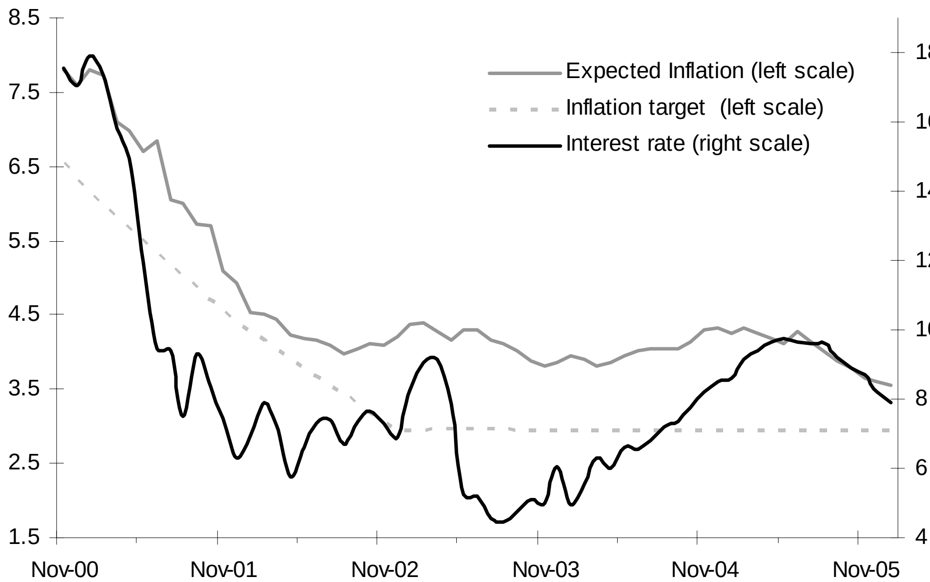
Source: Central Bank of Chile and authors' calculations.

Figure 3. **Colombia: Monetary stance and expected inflation**



Source: Central Bank of Colombia and authors' calculations.

Figure 4. **Mexico: Monetary stance and expected inflation**



Source: Central Bank of Mexico and authors' calculations.

The co-integration tests will be performed using the Johansen-Juselius methodology, because it allows for the existence of multiple co-integrating vectors. This is important in the current setting, since more than two variables are included in some of the estimations. Therefore, two such long-run relationships may potentially emerge: with *n* variables, there may exist at most *n–1* co-integrating relationships. In this case, the first co-integrating vector would define the process whereby inflation expectations are formed. If the monetary authorities are successful in anchoring expectations, expected inflation should respond negatively to changes in the interest rate and positively to the inflation target. The second co-integrating relationship that could potentially emerge would define the monetary authority's (reduced-form) reaction function. If the central bank conducts monetary policy in a forward-looking manner, the policy interest rate should respond positively to changes in expected inflation.[9] Also, for a given level of expected inflation, the interest rate should respond positively to changes in the target, because the inflation target and the interest rate are expected to have opposite affects on inflation expectations (based on the first co-integrating vector). In other words, a given level of expected inflation can be reached by positive combinations of the interest rate and the inflation target.

The co-integrating vectors

Based on the results of the unit root tests presented above, the three variables (interest rate, inflation expectations and the inflation target) will be tested for co-integration in the cases of Brazil and Mexico, where all of the variables were found to be integrated of order one. No co-integration test was performed for Colombia, because expected inflation was found to be trend-stationary and, as such, it cannot co-integrate with the other variables. Finally, in the case of Chile, co-integration will be tested between the interest rate and expected inflation, because the inflation target was constant during the period since inflation expectations started to be collected. The results of the co-integration tests are presented in Annex A1, while the estimated co-integrating vectors are reported in Table 3.

Two long-run relationships were found for Brazil and Mexico and one for Chile. In all cases, the interest rate reacts to changes in expected inflation, suggesting that monetary policy is conducted in a forward-looking manner. The estimated coefficient on expected inflation is particularly sizeable in the cases of Mexico and Chile. The estimated parameters suggest that a one-percentage-point increase in expected inflation leads to a 0.5 percentage-point increase in the interest rate in Brazil, 1.9 in Mexico and 5.5 in Chile. Also, the coefficient on the inflation target in the interest-rate equation is positive, as expected. Moreover, expected inflation appears to respond to both the inflation target and the interest rate over the long run, suggesting that the central bank is conducting monetary policy in a credible manner. The sensitivity of expected inflation to the target is very strong: a one-percentage-point reduction in the target leads to 1.5 and 3.0 percentage-point reductions in expected inflation in Brazil and Mexico, respectively.[10] The magnitude of this estimated sensitivity suggests that the conduct of monetary policy has managed to de-link private-sector inflation forecasts from realised inflation outcomes.[11]

9. The monetary authority may react to other variables, in addition to expected inflation, when setting monetary policy, such as the output gap and the exchange rate. These variables are taken into account in the next section.

10. The coefficient for Brazil is higher here than in previous long-run estimations (Cerisola and Gelos, 2005).

11. This finding is consistent with the evidence for industrial countries reported by Levin *et al.* (2004) that inflation is much less persistent in countries that have explicit targets for inflation, where there is no correlation between private-sector inflation forecasts and lagged inflation (Levin *et al.*, 2004).

Table 3. **Estimated co-integrating vectors**[1]

	Equations	
	$E_t\pi_{t+12}$	r_t
	Brazil	
Intercept	-0.74	15.25
r_t	-0.04	…
π^*_{t+12}	1.50	0.10
$E_t\pi_{t+12}$	…	0.49
	Chile	
Intercept	…	-12.69
r_t	…	…
π^*_{t+12}	…	…
$E_t\pi_{t+12}$	…	5.53
	Mexico	
Time trend	0.01	0.18
r_t	-0.04	…
π^*_{t+12}	2.96	0.88
$E_t\pi_{t+12}$	…	1.87

1. r_t = interest rate, $E_t\pi_{t+12}$ = 12-month-ahead expected inflation, π^*_{t+12} = 12-month-ahead inflation target and *t* = linear trend. The samples are: 2001:7 to 2006:1 for Brazil, 2001:9 to 2006:1 for Chile, 2003:9 to 2005:12 for Colombia and 2000:11 to 2006:1 for Mexico.

Source: Data available from the Central Bank of Brazil, Chile, Colombia and Mexico, and authors' calculations.

4. Monetary policy and inflation expectations: Are there volatility spillovers?

The link between the conduct of monetary policy and the process whereby inflation expectations are formed is crucial under inflation targeting. Based on the empirical evidence reported above, monetary policy has been conducted in a forward-looking manner, and the existence of a stable long-run relationship between the interest rate, the inflation target and inflation expectations suggests that the current policy regime has contributed to anchoring expectations. The empirical link between monetary policy and inflation expectations has been the focus of part of the empirical literature on monetary policy in Latin America (Schmidt-Hebbel and Werner, 2002; Minella *et al.*, 2003; Cerisola and Gelos, 2005; Leiderman *et al.*, 2006), but the existence of volatility spillovers between the monetary stance and inflation expectations has so far not been tested empirically for the countries under examination. In addition, the estimation of monetary reaction functions has conventionally been conducted within a reduced-form single-equation econometric framework, thereby ruling out the possibility that the shocks affecting the relevant variables may be correlated.

To our knowledge, the only related literature that focuses on (conditional) interest-rate volatility under inflation targeting is that of Chadha and Nolan (2001) and Connolly and

Kohler (2004). Chadha and Nolan assess the impact on interest-rate volatility of information flows based on announcements of interest rate decisions by the monetary authority, as well as the publication of monetary policy committee meetings and inflation reports. They focus on the three-month Sterling LIBOR rate during the period of narrow-band inflation-targeting in the United Kingdom (1997:5 – 1999:5). Connolly and Kohler perform a similar exercise but focus on interest-rate expectations (instead of effective interest rates) for a panel including Australia, Canada, the Euro-area countries, New Zealand, the United Kingdom and the United States for the period spanning 1997:1 through 2004:6. While Chadha and Nolan find no impact of information flows on interest-rate volatility in the United Kingdom, Connolly and Kohler report evidence of such spillover effects in their sample.

Against this background, the presence of volatility spillovers between the interest rate and expected inflation will be tested below for Brazil, Chile, Colombia and Mexico.[12] The main caveat that should be acknowledged when interpreting the empirical findings is that, because the experience of the countries under examination with inflation targeting is relatively short, it is difficult to take full account in the empirical analysis of the structural and behavioural changes that were brought about by policy regime switches. This may be particularly important for the Colombian case, for which the relevant data starts only in late 2003.

The econometric model

The presence of volatility spillovers can be tested within a system defined as follows:[13]

$$r_t = a_{10} + a_{11} r_{t-1} + a_{12} (E_t \pi_{t+12} - \pi^*_{t+12})_{t-1} + a_{13} e_{t-1} + a_{14} y_{t-1} + \varepsilon_{1t}\,, \tag{1}$$

$$(E_t \pi_{t+12} - \pi^*_{t+12})_t = a_{20} + a_{21} r_{t-1} + a_{22} (E_t \pi_{t+12} - \pi^*_{t+12})_{t-1} + a_{23} e_{t-1} + a_{24} y_{t-1} + \varepsilon_{2t}\,, \tag{2}$$

where r_t is the interest rate, ($E_t\pi_{t+12}$- π^*_{t+12}) denotes deviations of 12-month-ahead expected inflation from the implicit target, e_t is the exchange rate, y_t is the output gap, and ε_{it} ($i=1,2$) are error terms.[14]

The auto-regressive coefficients (a_{11} and a_{22}) capture persistence in the monetary stance and in expected inflation, and are hypothesised to be positively signed. The cross-variable coefficients (a_{12} and a_{21}) measure the importance of one variable as a determinant of the other.

The exchange rate and the output gap are included in Equation (1) because inflation targeters may respond to information other than inflation expectations when setting monetary policy (coefficients a_{13} and a_{14}). This is especially the case in Latin America, where external shocks have been sizeable, leading to considerable exchange-rate volatility. Policymakers may then try to damp exchange-rate movements, not only via direct interventions in the foreign exchange market, but also through the level of the interest rate. Including the exchange rate as an explanatory variable in the

12. For more information on tests of volatility spillovers see the financial econometric literature linking co-movements in asset returns and volatilities over time and across assets and markets (Koutmos and Booth, 1995; Ng, 2000; So, 2001; Miyakoshi, 2003; Kim *et al.*, 2005; Savva *et al.*, 2005, among others).

13. In the case of Chile, the inclusion of two dummy variables in Equations (1) and (2) was found to lead to substantial improvements in the model's diagnostic tests, while not affecting the rest of the results. These dummies take the value of one in March 2002 and September 2003, respectively, and zero otherwise.

14. The inflation target was not included as an additional variable in the system, because a three-equation system would have been difficult to estimate. Also, the exchange rate and the output gap enter the system in lagged form to partially account for potential endogeneity problems.

Taylor rule in Equation (1) therefore captures the objective of the monetary authority to avoid excessive exchange-rate volatility, over and above its impact on inflation expectations.[15] The goal of short-term demand management by the central bank justifies inclusion of the output gap as another explanatory variable in the Taylor rule (Svensson, 2005). Both variables are also included in Equation (2) as they are expected to affect the rate of inflation in the future, via a pass-through effect and through labor demand and wage setting.

The relationship between the interest rate and deviations of inflation expectations from the target is defined in *levels* in Equations (1)-(2) (mean equations). But, as hypothesised, volatility spillovers (the second conditional moment of the data) may also be present. As a result, not only may changes in the monetary stance affect expected inflation, but also an increase in interest-rate volatility may render inflation expectations more volatile, and vice versa. Calling $\varepsilon_t = (\varepsilon_{1t}\ ,\ \varepsilon_{2t})$ the vector of random errors, the volatility spillover hypothesis can be tested using the BEKK representation of the conditional variances and co-variances of vector ε_t:[16]

$$\varepsilon_t / F(t-1) \sim \mathrm{N}(0, \mathrm{H}_t),$$

where $F(t\text{-}1)$ is the information set up to period t-1, and H_t denotes the conditional covariance matrix associated with ε_t. This matrix is defined as $H_t = c_0^{/} c_0 + B^{/} \varepsilon_{t-1} \varepsilon_{t-1}^{/} B$, where C_0 and B are parameter matrices of the form: $c_0 = \begin{bmatrix} C_{11} & C_{12} \\ 0 & C_{22} \end{bmatrix}$ and $B = \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix}$. The elements of H_t can be represented in univariate form as follows, where spillover effects exist if b_{21} and/or b_{12} are estimated to be different from zero:[17]

$$h_{11,t} = c_1 + b_{11}^2 \varepsilon_{1,t-1}^2 + 2 b_{11} b_{21} \varepsilon_{1,t-1} \varepsilon_{2,t-1} + b_{21}^2 \varepsilon_{2,t-1}^2 \tag{3}$$

$$h_{12,t} = h_{21,t} = c_2 + b_{11} b_{21} \varepsilon_{1,t-1}^2 + (b_{12} b_{21} + b_{11} b_{22}) \varepsilon_{1,t-1} \varepsilon_{2,t-1} + b_{12} b_{22} \varepsilon_{2,t-1}^2 \tag{4}$$

$$h_{22,t} = c_3 + b_{12}^2 \varepsilon_{1,t-1}^2 + 2 b_{12} b_{22} \varepsilon_{1,t-1} \varepsilon_{2,t-1} + b_{22}^2 \varepsilon_{2,t-1}^2 \tag{5}$$

For example, if only b_{21} is statistically significant, then there are unidirectional spillovers running from expected inflation to the monetary stance. If both coefficients are statistically different from zero, then there are bi-directional spillovers between the monetary stance and inflation expectations. Equations (1) and (2), joint with the BEKK specification for the error terms, are estimated by Maximum Likelihood.

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15. Exchange rate-augmented monetary reaction functions were estimated by Mishkin and Savastano (2001) and Mohanty and Klau (2005). Based on a sample of developed and developing inflation targeters, Ho and McCauley (2003) concluded that the monetary authorities care about the exchange rate and its variability because of their impact on inflation (via relative prices and expectations, for instance), on the performance of the external sector, investment and growth (via trade competitiveness), on financial and public debt sustainability (via balance sheet effects) and on the development of foreign-exchange and capital markets.

16. The BEKK model is a generalisation of univariate GARCH models. See Bauwens *et al.* (2006) for more information.

17. Note that c_i (i=1,2,3) are combinations of the elements in c_0. Also note that only the ARCH specification is retained in order to save degrees of freedom.

The results

Based on the unit root tests reported above, the variables enter the system as follows. The interest rate and the exchange rate have unit roots in all countries (except in the case of Mexico for the last variable) and, therefore, they enter the system in first differences. By contrast, the output gap is always stationary and will therefore be included in levels. The deviation of expected inflation from the target is stationary in Brazil and will therefore be included in levels, while in the case of Colombia the series is trend-stationary and will therefore be HP-filtered. For Chile and Mexico, the variable was found to have a unit root and will be included in first differences. For Chile, the focus is on expected inflation, rather than the deviation of expected inflation from the target, as noted above.[18]

In the cases of Brazil and Colombia, as the interest rate has a unit root while the deviation of expected inflation from the target does not, there is no need to test for co-integration between these variables. For Chile, as both variables have unit roots and are co-integrated, the previously estimated error-correction term will be included as a predetermined variable in the system.[19] For Mexico, as both the interest rate and the deviation of expected inflation from the target have unit roots, a co-integration test was performed based on the Johansen-Juselius methodology (results not reported). However, convergence was not achieved when the error-correction term was included in the system; it was therefore excluded from the estimation.[20]

The results of the estimation of Equations (1)-(2) are presented in the upper panel of Table 4. Both the interest rate and the deviation of expected inflation from the target exhibit a positive, relatively high and significant degree of persistence in the cases of Brazil and Chile. Over the period of analysis, the interest rate was estimated to respond positively to lagged changes in expected inflation in Chile and to lagged changes in the deviation of expected inflation from the target in Mexico. The variable is not significant in Brazil and has a counterintuitive negative, statistically significant sign in Colombia. The empirical findings also suggest that monetary policy is conducted in a counter-cyclical manner in all countries, with the interest rate reacting positively to the output gap, and that the monetary authority responds to changes in the exchange rate only in the case of Brazil. For Chile and Colombia, the exchange rate was not significant in either equation and was therefore excluded from the system.[21] Evidence of high persistence and greater responsiveness of monetary

18. The presence of unit roots in the data also affects the deterministic components of the system. In Brazil, given that the deviation of expected inflation from the target is stationary (not necessarily around a zero mean), a constant term should be included in Equation (2). However, due to a lack of convergence in the algorithm used to estimate the model, the constants were eliminated. This is also the case for Mexico and Colombia, where a constant should in principle be included in Equation (1), because the interest rate was found to follow a random walk with drift. For Chile, the inclusion of intercepts is not required, because the variables of interest are difference-stationary without drift.

19. This is defined as: $ECT_t = r_t - 5.53 E_t \pi_{t+12} + 12.69$.

20. The algorithm converged when the exchange rate was excluded and the ECT was included in the system, with the ECT having the expected signs in both equations. The hypothesis of volatility spillovers continued not to be rejected (results not reported).

21. The fact that the exchange rate is not significant in the case of Colombia may not come as a surprise, since the central bank reacted more strongly to the appreciation of the *peso* by conducting discretionary interventions than by reducing the interest rate during 2004 (only two reductions of 25bp took place). The other event in which the central bank reacted to exchange rate developments was the depreciation of the *peso* from mid-2002 to early 2003. During this period the central bank followed a more aggressive strategy, increasing the interest rate by 100bps on two occasions. The sample used in the estimations excludes this episode and contributes then to weakening the finding of a statistical

policy to deviations of inflation from the target than to the output gap among Latin American inflation targeters has been reported by Corbo and Schmidt-Hebbel (2001), among others.

With regard to the expected-inflation equation, the interest rate was found to have the expected negative sign in the cases of Brazil and Colombia.[22] After all, the objective of the monetary authorities is in part to prevent increases in inflation by means of anchoring inflation expectations. The exchange rate is positively signed, as expected, in Brazil and Mexico, although the estimated coefficient is rather small in the case of Mexico. In addition, the output gap is positively signed in all countries. In Chile, adjustment towards long-run equilibrium is achieved through changes in expected inflation. This is because the error-correction term is positively signed in expected-inflation equation, while it is not significant in the Taylor rule.[23]

The estimated coefficients of Equations (3)-(5), allowing for the analysis of volatility spillover effects, are reported in the middle panel of Table 4.[24] ARCH effects are present for the deviation of expected inflation from the target in Brazil and Colombia, and for the interest rate in Chile and Colombia. Regarding volatility spillover effects, the estimation results for Brazil and Colombia suggest that they are unidirectional, running from the interest rate to expected inflation. This implies that volatility in the monetary stance leads to volatility in inflation expectations. Spillover effects could not be detected in the Chilean data, while for Mexico there are bi-directional spillovers between the interest rate and expected inflation. With regard to the magnitude of these estimated effects, the greatest spillover from interest rate volatility to expected inflation seems to be present in Mexico. Assuming that there is no innovation to expected inflation, the impact of a one-standard-deviation shock to the interest rate on expected inflation amounts to almost 90% of the mean of its conditional volatility. To put this magnitude in perspective, a similar shock would have an impact on expected inflation of 61% of its mean conditional volatility in Colombia and 52% in Brazil.[25]

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relationship between the interest rate and the exchange rate in the Colombian case. See Vargas (2005) for a detailed assessment of these policy episodes.

22. In other studies for Brazil, this variable was found to be statistically insignificant for the inflation targeting period (Minella *et al.*, 2003; Cerisola and Gelos, 2005).

23. Celasun *et al.* (2004) and Cerisola and Gelos (2005) suggest that improvements in the fiscal balance may help to reduce expected inflation by allaying concerns regarding fiscal dominance. This variable was not found to be significant and was therefore excluded from the estimations.

24. For the BEKK representation retained in this study, the condition for second-order stationarity is that the moduli of the eigenvalues of the matrix $(B \otimes B)$ are less than one (Engle and Kroner, 1993). This condition is fulfilled by all of the countries except Colombia.

25. These impacts are calculated as: $imp = \left. \left(\frac{\partial h_{22}}{\partial \varepsilon_{1,t-1}^2} \right) (\varepsilon_1)^2 \middle/ \overline{h_{22}} \right.$, where *imp* is the spillover effect mentioned in the text, $(\varepsilon_1)^2$ is the square of a one-standard-deviation shock to the interest rate, and $\overline{h_{22}}$ is the sample mean of the conditional volatility of expected inflation.

Table 4. **Monetary reaction function and inflation expectations: M-GARCH analysis of volatility spillover effects**[1]

	Brazil		Chile		Colombia		Mexico	
	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$	r_t	$(E_t\pi_{t+12} - \pi^*_{t+12})$
r_{t-1}	0.79	-0.43	0.62	0.06	-0.11	-0.58	-0.03	0.04
	(24.12)***	(-6.19)***	(9.64)***	(0.75)	(-0.84)	(-2.89)***	(-0.32)	-1.35
$(E_t\pi_{t+12} - \pi^*_{t+12})_{t-1}$	-0.03	0.86	0.19	0.42	-0.19	0.16	1.68	-0.12
	(-1.10)	(33.60)***	(2.36)**	(3.96)***	(-2.91)***	(0.87)	(4.06)***	(-1.14)
e_{t-1}	0.01	0.03					-0.01	0.00
	(2.81)***	(6.21)***					(-0.58)	(2.05)**
y_{t-1}	0.07	0.16	0.04	0.09	0.01	0.01	0.13	0.04
	(4.04)***	(6.15)***	(2.15)**	(3.46)***	(6.63)***	(2.58)**	(2.03)**	(3.00)***
ECT (t-1)			-0.00	0.07				
			(-0.18)	(4.95)***				
Variance equations								
C$_{11}$	0.41		0.07		0.02		0.84	
	(4.15)***		(0.48)		(0.06)		(8.87)***	
C$_{12}$	0.48		0.00		-0.06		0.01	
	(9.37)***		(0.01)		(-2.08)**		(0.52)	
C$_{22}$	0.10		0.11		0.07		0.07	
	(0.35)		(1.04)		(0.21)		(0.25)	
B$_{11}$	0.48		0.89		1.32		0.09	
	(1.43)		(3.63)***		(6.2)***		(0.04)	
B$_{21}$	0.07		0.21		0.03		2.11	
	(0.54)		(1.24)		(0.25)		(2.04)**	
B$_{12}$	-1.28		0.27		1.08		-0.18	
	(-5.33)***		(1.15)		(2.66)**		(-4.47)***	
B$_{22}$	-0.59		0.13		-0.70		0.15	
	(-2.32)***		(0.73)		(-1.71)*		(1.04)	
Diagnostic tests								
Log-likelihood	-42.83		-177.10		-111.33		-78.13	
AIC	-3.04		-8.97		-7.86		-4.96	
SBC	-2.74		-8.51		-7.58		-4.68	
Normality test (KS)	0.11	0.11*	0.11*	0.06	0.09	0.14	0.07	0.07
LB(6)	3.59	6.24	6.12	1.90	0.84	3.14	8.44	4.30
ARCH(6)	1.42	1.02	4.41	8.70	5.83	1.16	5.42	7.55

1. The numbers in parentheses are *t* statistics. AIC and SBC are the Akaike and Schwarz Bayesian information criteria, respectively. KS is the Kolmogorov-Smirnov test of the null hypothesis of normality. LB (6) and ARCH(6) are the Ljung-Box test for autocorrelation and the Lagrange Multiplier (LM) test for heteroscedasticity up to 6 lags, respectively, performed on the univariate standardised residuals. In the case of Chile, $(E_t\pi_{t+12} - \pi^*_{t+12})$ refers to expected inflation. ECT = error-correction term. (*), (**) and (***) denote statistical significance at the 10%, 5% and 1% levels, respectively. The samples are: 2001:7 to 2006:1 for Brazil, 2001:9 to 2006:1 for Chile, 2003:9 to 2005:12 for Colombia and 2000:11 to 2006:1 for Mexico.

Source: Data available from the Central Banks of Brazil, Chile, Colombia and Mexico, and authors' calculations.

Finally, the lower panel of Table 4 reports conventional diagnostic test statistics. On the basis of these tests, the M-GARCH model adequately describes the interactions between the monetary stance and expected inflation**.** Except for the multivariate AIC and SBC criteria, the tests are performed on the univariate standardised residuals.[26] The null hypothesis of normality is rejected, albeit only at the 10% level, for the deviation of expected inflation from the target in Brazil and for the interest rate in Chile. The absence of autocorrelation indicates that the mean equations are well specified. Also, the BEKK representation seems to be capturing well the volatility dynamics present in the data, since no conditional heteroscedasticity is left in the standardised residuals.

5. Conclusion

Two well recognised goals of inflation targeting are to reduce inflation and to anchor expectations around the pre-announced targets. The latter objective implies that it is important not only to have a better understanding of the empirical link between the conduct of monetary policy and the process whereby inflation expectations are formed, but also to test for the presence of volatility spillovers between the monetary stance and expected inflation.

The results of the empirical analysis reported above show that the interest rate, the inflation target and inflation expectations tend to move together in the countries under consideration. This suggests that the conduct of monetary policy in the current regime has been effective in anchoring inflation expectations. The hypothesis that greater volatility in the monetary stance may lead to greater volatility in expected inflation (and *vice versa*) has also been tested using an M-GARCH modelling technique. This is important because reducing volatility in inflation expectations is the ultimate test of whether the monetary regime can anchor expectations around the pre-announced targets or not. The main findings of the empirical analysis show that volatility in the monetary stance is conducive to volatility in inflation expectations at least in the cases of Brazil and Colombia. No volatility spillover effect was detected in the case of Chile, and there appears to be bi-directional effects in the case of Mexico. On the basis of these findings, interest-rate smoothing, whereby the monetary authority avoids sudden changes in the policy interest rate above and below its fundamental determinants, is advisable at least in the cases of Brazil, Colombia and Mexico as a means of reducing volatility in inflation expectations.

26. These are defined as: $\eta_{it} = \frac{\varepsilon_{it}}{\left(h_{ii,t}\right)^{1/2}}$, where ε_{it} (*i = 1, 2*) are the residuals from the mean equations, and $h_{ii,t}$ is the conditional variance for variable *i*.

Annex A1

Co-integration test results

This Annex reports the results of the Johansen-Juselius co-integration tests. The methodology consists of testing for the rank of matrix Π in the following vector error-correction model (VECM):

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Phi_i y_{t-i} + AD_t + \varepsilon_t$$

In this model, y_t is the vector of relevant variables to be tested for co-integration, Δ is the difference operator, Φ and *A* are square matrices of parameters, D_t denotes deterministic terms, and ε_t is a multivariate white-noise process. If matrix Π has full-rank, all the components of y_t are I(0) (*i.e.* the variables are stationary in levels); if rank is 0, then, y_t is stationary in first differences, but the variables are not co-integrated. Finally, the rank of Π equals the number of co-integrating vectors present in the data (the number of long-run relationships) when it lies between zero and the number of variables included in the model.

When testing for co-integration using this methodology the results are sensitive to the specification for the deterministic components (D_t) and the number of lags included in the model ($p-1$). Regarding the deterministic component, the specification in which the VECM contains a constant restricted in the co-integrating vector was used for Brazil. The choice is based on the evidence that the variables behave as random walks without drift, and, as such, no constant should be included in the model for the differenced variables. In this case, the co-integration test is that of a linear combination of the variables being stationary around a constant. The same is true for the case of Chile. In the case of Mexico, the variables were found to follow random walks with drift (they show a downward trend); as a result, the VECM included a trend restricted to lie in the co-integrating vector. This accounts for the possibility that the variables may not contain the same linear trend.[1]

The number of lags (*p-1*) included in the VECM was selected by applying two different multivariate lag selection criteria: the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC). With monthly data, the maximum number of lags was originally set at twelve in all cases. Then, it was reduced to seven for Brazil, eleven for Chile and nine for Mexico, since for the highest lags the results indicated that all variables should be stationary in levels, which is inconsistent with the results of the unit root tests.

For Brazil, the optimal lag structure was found to include three lags on the basis of SBC and seven lags based on AIC. In both cases, two co-integrating relationships were detected among the variables, as hypothesised.[2] The results reported in Table A1.1 are based on an optimal structure of seven

1. That the deterministic terms can be included in the co-integrating vectors was not rejected by the data, once the number of co-integrating vectors was determined. See next.

2. Giving the short samples under consideration, a 10% significance level was used in order to compensate for the lack of power of tests.

lags, because with three lags the two co-integrating vectors (not reported) were found to have implausible values for the long-run coefficients. In the case of Chile, an optimal lag length of two was selected on the basis of SBC and seven on the basis of AIC. The results reported in Table A1.2 are based on the optimal structure selected according to SBC, since no co-integrating vector was found with seven lags. Finally, in the case of Mexico, an optimal structure of four lags was selected on the basis of SBC and nine lags on the basis of AIC. Table A1.3 reports the results based on four lags, since with nine lags the estimated monetary reaction function exhibited a coefficient for the inflation target that was at odds with economic priors.[3]

It should be noted that, with multiple co-integrating vectors, any linear combination of those vectors will be another co-integrating vector. As such, if there are *r* co-integrating vectors in an *n*-variable system, then there is also a co-integrating vector for each subset of ($n - r + 1$) variables. With three variables, there is a co-integrating vector for each pair of variables. When these are computed for Brazil and Mexico, very similar values for the coefficients presented in Table A1.1 and Table A1.3 are obtained.

Table A1.1. **Brazil: Co-integration test results**

	MAX test			**Trace test**		
H_0	r = 0	r = 1	r = 2	r = 0	r ≤ 1	r ≤ 2
H_1	r = 1	r = 2	r = 3	r ≥ 1	r ≥ 2	r ≥ 3
Statistics	45.21	16.63	3.81	65.65	20.44	3.81
Critical value (at 10% confidence level)	19.77	13.75	7.52	31.88	17.79	7.50

Source: Data available from the Central Bank of Brazil and OECD calculations.

Table A1.2. **Chile: Co-integration test results**

	MAX test		**Trace test**	
H_0	r = 0	r = 1	r = 0	r ≤ 1
H_1	r = 1	r = 2	r ≥ 1	r ≥ 2
Statistics	16.05	5.72	21.77	5.72
Critical value (at 10% confidence level)	13.75	7.52	17.79	7.50

Source: Data available from the Central Bank of Chile and OECD calculations.

Table A1.3. **Mexico: Co-integration test results**

	MAX test			**Trace test**		
H_0	r = 0	r = 1	r = 2	r = 0	r ≤ 1	r ≤ 2
H_1	r = 1	r = 2	r = 3	r ≥ 1	r ≥ 2	r ≥ 3
Statistics	46.82	22.37	9.39	78.58	31.76	9.39
Critical value (at 10% confidence level)	23.11	16.85	10.49	39.08	22.95	10.56

Source: Data available from the Central Bank of Mexico and OECD calculations.

3. In the cases of Brazil and Chile, a likelihood-ratio test does not reject the null hypothesis that the constant may be included in the co-integrating vector (the *p*-values obtained were 0.69 and 0.74, respectively). For Mexico, the *p*-value (0.27) does not reject the null that the linear trend can be included in the co-integrating vector.

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