Chapter 1

Monetary policy and macroeconomic stability in Latin America: The cases of Brazil, Chile, Colombia and Mexico

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This chapter 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.

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Introduction and summary

There is a growing empirical literature, pioneered by Taylor (2000) and Clarida *et al.* (2000), among others, on how changes in a country's monetary policy regime affect macroeconomic volatility.[1] The main finding in this literature is that, at least as far as the United States is concerned, a more pro-active policy stance since the mid-1980s, whereby the monetary authority responds strongly to changes in expected inflation, has contributed to anchoring expectations at low, stable levels and reducing business-cycle fluctuations in economic activity. Greater macroeconomic stability is also due to the fact that the shocks hitting the economy have become milder over the last 20 years or so (Ahmed, Levin and Wilson, 2002; Stock and Watson, 2002; Cecchetti, Flores-Lagunes and Krause, 2004; Boivin and Giannoni, 2005). Another factor militating in favour of lower inflation volatility in the United States is a change in price-setting mechanisms, which have been found to have become more forward-looking since the 1980s (Moreno, 2004).

A complementary strand of literature focuses on how the adoption of inflation targeting in many countries, coupled with exchange rate flexibility, has affected macroeconomic volatility. The argument is that, by allowing the exchange rate to float freely the monetary authority can respond more forcefully to changes in the inflation outlook in pursuit of its inflation target, instead of defending a nominal exchange rate peg. Empirical evidence for industrial countries suggests that, where the policy regime is credible and monetary policy is conducted in a transparent, forward-looking manner, adoption of inflation targeting has delivered lower volatility in the monetary stance (Kuttner and Posen, 1999; Woodford, 1999 and 2004). However, as suggested by the empirical evidence surveyed by Mishkin (2006), the fall in macroeconomic volatility since the 1990s is a worldwide phenomenon, and, therefore, inflation targeters in the developed world have not done better than non-inflation targeters at reducing macroeconomic volatility, although they have done a better job at anchoring expectations in the sense of reducing the sensitivity of expected inflation to shocks in current inflation. With regard to emerging-market economies, de Mello and Moccero (2006) use co-integration and M-GARCH analysis to test for the presence of long-run relationships among the policy interest rate, inflation expectations and the inflation target, as well as of volatility spillovers between inflation expectations and the monetary stance in Brazil, Chile, Colombia and Mexico. The authors conclude that the monetary stance has become more persistent under inflation targeting and exchange rate flexibility, which has contributed to anchoring inflation expectations around the pre-announced targets in these countries.

Against this background, this chapter tests the hypothesis that a change in the monetary regime has reduced macroeconomic volatility in four Latin American countries (Brazil, Chile, Colombia and Mexico), where inflation targeting has been complemented by flexible exchange rate regimes since 1999.[2] To this end, a small New Keynesian structural model comprising aggregate supply and demand equations and a monetary reaction function is estimated. Impulse response functions are computed for the structural model and for an unrestricted VAR in the interest rate, inflation and the output gap. A counterfactual exercise is performed to assess the role played by changes in the policy regime and in the shocks hitting the economy in explaining changes in macroeconomic volatility across policy regimes. The counterfactual exercise allows for the estimation of the volatilities that would arise from a given combination of shocks and monetary policy parameters, thus identifying the factors that make for greater macroeconomic stability.

Modelling and data

A simple structural model

A conventional macro-structural model is estimated to highlight the main stylised facts about how macroeconomic volatility has been affected by changes in policy and shocks across monetary regimes in Brazil, Chile, Colombia and Mexico since the mid-1990s. The New Keynesian framework has become the reference point for analysing the relationship between inflation, monetary policy and the business cycle. In its simplest form, it consists of three equations:

$$\pi_t = \delta E_t \pi_{t+1} + (1-\delta)\pi_{t-1} + \lambda y_t + u_{\pi_t}, \qquad (1.1)$$

$$y_t = \mu E_t y_{t+1} + (1-\mu) y_{t-1} - \phi(r_t - E_t \pi_{t+1}) + u_{y_t}, \qquad (1.2)$$

$$r_t = \rho r_{t-1} + (1-\rho)(\beta E_t \pi_{t+1} + \gamma y_t) + \tau e_t + u_{r_t}, \qquad (1.3)$$

where π_t, y_t, r_t and e_t denote respectively inflation, the output gap, the nominal interest rate and the nominal exchange rate at time *t*; E_t is the expectations operator conditional on information available at time *t*; and u_{π_t}, u_{y_t} and u_{r_t} are the structural errors.

Equation (1.1) is a conventional Phillips curve, including Calvo-type price stickiness, Equation (1.2) is an aggregate demand function, and Equation (1.3) is an augmented Taylor-type monetary reaction function, which includes the nominal exchange rate as a pre-determined variable. There is some controversy over whether or not the exchange rate should enter the reaction function. But we have opted for including it, because there may be more complex interactions between movements in the exchange rate and macroeconomic performance in the context of emerging-market economies that are not captured in the conventional Taylor rule. A case in point is "fear of floating" in countries that have resorted to exchange rate targeting for extended periods. Exchange rate-augmented monetary reaction functions have been estimated by Ball (1999), Mishkin and Savastano (2001), Minella *et al.* (2003), and Mohanty and Klau (2005), among others.

Monetary policy regimes: A brief summary

The four countries under consideration have upgraded their institutional setting for monetary policymaking since the 1990s as a means of entrenching macroeconomic stability (Fracasso *et al.*, 2003; Carstens and Jacome, 2005; Avendano and de Mello, 2006). Institutional reform has aimed at reducing the scope for central bank financing of budget deficits and on granting *de jure* operational autonomy to the monetary authority. Brazil is an exception, however, because the central bank is not formally independent, although it is perceived as enjoying *de facto* autonomy from the executive branch of government. Inflation targeting was formally adopted in Brazil in June 1999, following the January 1999 floating of the *real*, and in January 1999 in Mexico. Chile and Colombia had pursued some looser form of inflation targeting since the early to mid-1990s, combining pre-announced targets for both headline inflation and the exchange rate. The exchange rate was nevertheless allowed to float freely in both countries in September 1999. The levels of inflation targeted differ among countries, as well as the tolerance bands around the central targets. Chile and Mexico currently target headline inflation within a 2-4% band, whereas Brazil and Colombia target a higher level of inflation, at 4.5%. The tolerance band is wider in Brazil (2.5-6.5%) than in Colombia (4-5%).

The monetary authorities rely predominantly on open-market operations, and central bank credit and deposit facilities to conduct monetary policy in these four countries (Avendano and de Mello, 2006). The use of reserve requirements as a monetary policy instrument has become less important over time. Unremunerated reserve requirements are high in Brazil for demand deposits, but have come down, as well as in Colombia, and have been used in Chile to discourage short-term capital inflows. Interest rate controls are less widespread, although the rate on short-term demand deposits is regulated in Chile, as well as selected longer-term rates in Brazil (TR and TJLP, for example).

Empirical evidence for Brazil, Chile, and Mexico (Schmidt-Hebbel and Werner, 2002; Minella *et al*., 2003; OECD, 2005; de Mello and Moccero, 2006) suggests that inflation targeting is working well in these countries. Inflationary inertia has been reduced where the monetary authority has been forward-looking and responsive to deviations of expected inflation from the targets. The exchange rate regime has played an important role in shaping inflation dynamics in these countries, and inflation has come down more rapidly in the countries that have used exchange rate anchors to break inflationary inertia, especially where inflation had been chronically high, such as in Brazil. The reduction in inflation has been more gradual in Chile and Mexico.

Data and times-series properties

System (1.1)-(1.3) is estimated by full information maximum likelihood (FIML) using monthly data for Brazil, Chile, Colombia and Mexico over the period spanning 1996:1 through 2006:2. The system is estimated for two sub-samples: *i*) the period prior to the abandonment of formal or informal exchange-rate targeting in Brazil, Chile and Colombia, and prior to the adoption of formal inflation targeting in Mexico, and *ii*) thereafter, when inflation targeting was complemented by exchange rate flexibility. Mexico allowed the *peso* to float at end-1994 but formally adopted inflation targeting only in 1999. Conversely, Colombia and Chile adopted inflation targeting in the early to mid-1990s, but allowed their currencies to float freely only in September 1999. The cut-off dates are therefore January 1999 for Mexico and September 1999 for Chile and Colombia. In the case of Brazil, two different cut-off dates are set: January 1999, due to the floating of the *real*, and June 1999, which corresponds to the formal adoption of inflation targeting in a floating exchange-rate regime.

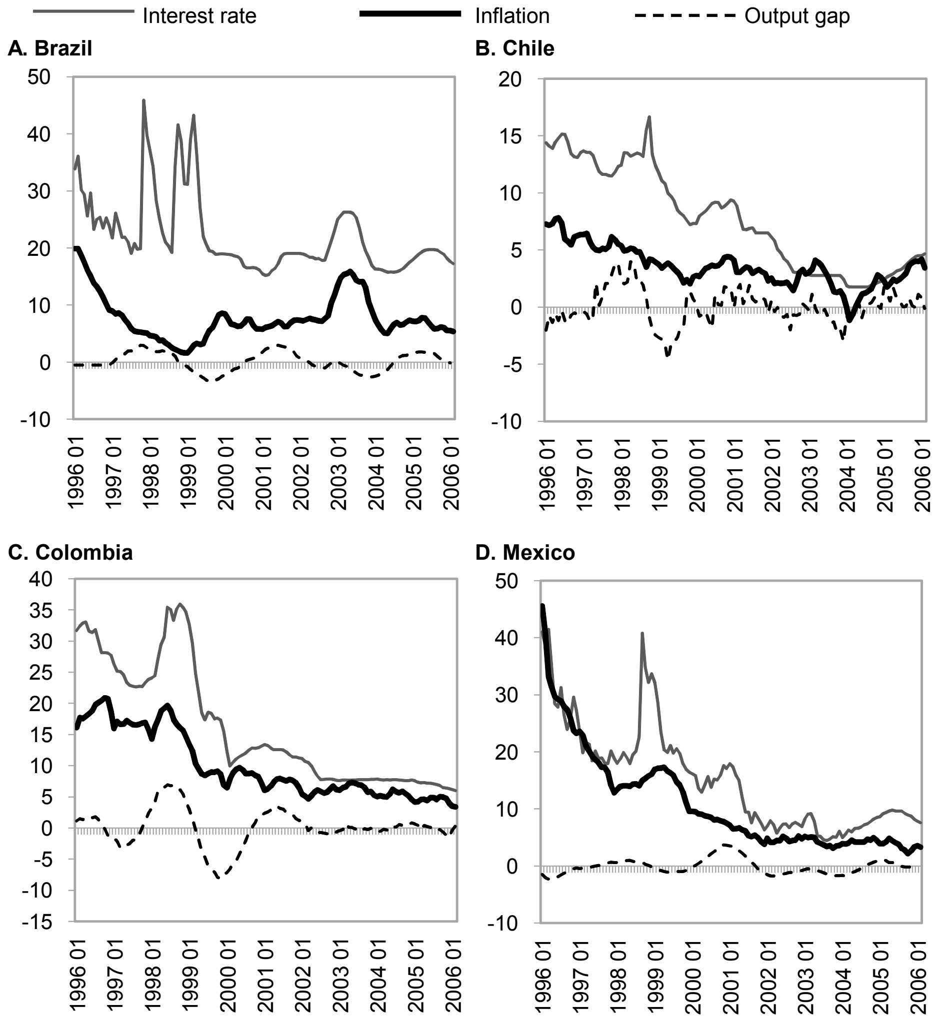
Monthly data for the four countries are available from national sources. Inflation is measured by the consumer price index (IPCA in Brazil, IPC in Chile and Colombia, and INPC in Mexico). The nominal interest rate (annualised in all countries) is the SELIC rate in Brazil, the TPM rate in Chile (inflated by the annual variation in the UF (*Unidad de Fomento*) before August 2001), the rate on 90-day deposits (CDT) in Colombia and the rate on the 28-day CETES bonds in Mexico. The output gap was computed as the percent difference between the seasonally-adjusted industrial production index and its Hodrick-Prescott (HP)-filtered trend (IMACEC index in the case of Chile). The exchange rate is the period-average rate defined as units of national currency per US dollar.

The inflation, interest rate and output gap series are depicted in Figure 1.1. Visual inspection of the data suggests that interest rates seem to have become less volatile in all countries since 1999. This is also the case of inflation for

Colombia and Mexico. The output gap does appear to have become less volatile in Colombia, where the amplitude of business-cycle fluctuations seems to have moderated somewhat since 2001 and to a lower degree in Chile, since the end of 1999.

Figure 1.1. **Trends in inflation, interest rate and output gap, 1996:1-2006:2**

In per cent



1. A 12-month moving average is reported for the output gap.

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

A battery of unit root tests was performed (results available upon request), including the augmented Dickey-Fuller (ADF), the Philips-Perron (PP) and the Zivot-Andrews tests. The latter test allows for a one-off structural change under the alternative hypothesis.[3] On the basis of these tests, the inflation, interest rate and nominal exchange rate series appear to have unit roots in all countries when the variables are defined in levels. They therefore enter the model in first differences. The output gap was found to be stationary in levels in all countries, except Chile.[4] On the basis of the Zivot-Andrews test, the interest rate series were found to have structural breaks between late 1998 and early 1999 (except in Chile), which corresponds to the selected cut-off date of end 1988 used in the empirical analysis.

Estimation of the structural model

The results of the estimation of the structural model, reported in Table 1.1, suggest a relative stability across policy regimes in the parameters of the Phillips curve and the aggregate demand equation in all countries. By contrast, the results of a similar structural model estimated by Moreno (2004) for the United States show that the Phillips curve became more forward-looking over time, suggesting an important change in price setting. The inflation and the output gap equations exhibit a fair degree of persistence, which has not changed in a discernible way across policy regimes in the Latin American countries in the sample. The output gap does not enter the Phillips curve in a statistically significant manner and the *ex ante* real interest rate does not appear to be a powerful determinant of the output gap.

Monetary policy appears to have become increasingly persistent over time in all countries, except in Mexico. This is not surprising because of the abandonment of exchange rate targeting in these countries, which allows monetary policy to pursue price stability unencumbered by the need to defend a pre-announced target for the nominal exchange rate. The monetary authority also became more forward-looking over time in Chile, as evidenced by the positively-signed and statistically significant coefficient on expected inflation (β), and in Brazil in the sample that excludes the transition period of January-June 1999. This finding is consistent with those reported by Corbo *et al*., (2002) in the case of Chile, on the basis of a one-equation monetary reaction function, and by Minella *et al*., (2003) for Brazil.

In addition, monetary policy was found to be responsive to changes in the exchange rate in Mexico (both periods) and in Brazil in the post-1999 regime, although this is not the case if the January-June transition period is excluded.

Table 1.1. **Structural model estimations: Brazil, Chile, Colombia and Mexico**[1]

	Brazil			Chile		Colombia		Mexico	
	1 1996:1 to 1998:12	2A 1999:1 to 2006:2	2B 1999:7 to 2006:2	1 1996:1 to 1999:9	2 1999:10 to 2006:2	1 1996:1 to 1999:9	2 1999:10 to 2006:2	1 1996:1 to 1998:12	2 1999:1 to 2006:2
δ	0.49**	0.50***	0.54***	0.48***	0.51***	0.60***	0.52***	0.49***	0.50***
	(0.235)	(0.054)	(0.083)	(0.167)	(0.076)	(0.084)	(0.088)	(0.066)	(0.078)
λ	0.00	0.00	0.00	-0.02	-0.12	0.00	0.00	0.00	0.00
	(0.008)	(0.003)	(0.004)	(0.142)	(0.123)	(0.001)	(0.002)	(0.009)	(0.010)
μ	0.44***	0.47***	0.46***	0.52***	0.51***	0.40**	0.53***	0.56***	0.50***
	(0.122)	(0.065)	(0.066)	(0.134)	(0.120)	(0.172)	(0.113)	(0.087)	(0.079)
ϕ	1.85	1.59	1.20	-0.48	0.00	-1.95	-10.70	1.19	-0.72
	(5.960)	(1.376)	(2.168)	(0.751)	(0.352)	(26.619)	(10.774)	(1.351)	(1.278)
ρ	0.03	0.66***	0.61***	0.30	0.63***	0.29*	0.56***	0.10	0.11
	(0.299)	(0.071)	(0.057)	(0.547)	(0.085)	(0.169)	(0.092)	(0.090)	(0.097)
β	0.54	0.14	0.19**	0.02	0.11**	0.24	-0.12	0.29	0.15
	(1.883)	(0.192)	(0.088)	(0.072)	(0.046)	(0.219)	(0.223)	(0.302)	(0.139)
γ	0.01	0.01	0.01**	-0.02	-0.02	0.01***	0.00	0.03**	0.01
	(0.031)	(0.007)	(0.003)	(0.049)	(0.051)	(0.002)	(0.004)	(0.012)	(0.006)
τ	5.08	0.47***	0.02	-0.40	-0.16	-0.24	-0.13	2.85***	3.75***
	(22.921)	(0.105)	(0.080)	(1.138)	(0.277)	(0.240)	(0.325)	(0.491)	(0.630)

1. Expected values are measured by one-period-ahead values in the relevant variables. Standard errors are reported in parentheses. (***), (**) and (*) denote, respectively, statistical significance at the 1, 5 and 10% levels.

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

When the transition period is included in the sample, the significance of the coefficient on the exchange rate (τ) is probably due to the volatility that characterised the period of overshooting following the floating of the *real* and subsequent stabilisation of the nominal exchange rate. Finally, evidence of counter-cyclicality in the monetary stance was found in Colombia and Mexico in the first period, where the coefficient of the output gap (γ) was found to be positively signed and statistically significant, and in Brazil in the current policy regime (in the sample that excludes the January-June 1999 transition period).

In sum, estimation of the structural model suggests that monetary policy has been conducted in a more gradual, forward-looking manner in Brazil and Chile since the policy regime change that occurred in 1999. The monetary stance has also become more counter-cyclical in Brazil. Instead, in the cases of Colombia and Mexico, monetary policy has become less counter-cyclical, a finding which may be associated, at least in the case of Colombia, with greater interest-rate smoothing after the policy regime change.

VAR analysis

It has become standard to assess the implications of a change in the monetary policy regime using split-sample estimates of impulse response functions derived both from unrestricted, reduced-form VARs (Boivin and Giannoni, 2002 and 2005), as well as from structural models. Therefore, impulse responses to a monetary shock, defined as a one-standard-deviation innovation to the interest rate, are computed for inflation, the output gap and the interest rate for the two sub-samples corresponding to the different monetary policy regimes. The endogenous variables enter the VAR in the following order: inflation, the output gap and the interest rate. This recursive causal ordering has become conventional (Christiano, Eichenbaum and Evans, 1998) and imposes minimum structure in the VAR in the sense that the output gap and inflation have contemporary effects on the interest rate but not the converse. The exchange rate enters the model as a predetermined variable. Lag length selection was performed on the basis of the Schwarz Information Criterion (SIC).

Stability tests conducted for the VAR representation of system (1.1)-(1.3) show that no AR root lies outside the unit circle for the full sample and for both sub-samples in all countries. Cogley and Sargent (2003) discuss the power of a host of parameter stability tests and conclude that failure to reject the hypothesis of time invariance is due to the fact that procedures are unable to detect drifting parameters. Tests performed on the individual time series (not reported) suggest the presence of parameter shifts but did not provide a consistent selection of the timing of the structural breaks for the system as a whole. Therefore, as in the case of the structural models reported above, the cut-off dates were selected on the basis of the official dates of institutional changes in the policy regimes, while making sure that the sub-samples are not too short. This procedure was also followed by Boivin and Giannoni (2002; 2005).

Impulse response functions were computed for the unrestricted VAR and for the structural model. The methodology used to compute the structural impulse response functions is reported in Annex 1.A1. In the case of the pre-1999 policy regime, there does not seem to be a stationary solution to the structural model for Brazil, Chile and Mexico; the impulse responses are therefore omitted. Failure to identify a stationary equilibrium is not surprising, given the high volatility that characterised the pre-1999 monetary periods in these countries.

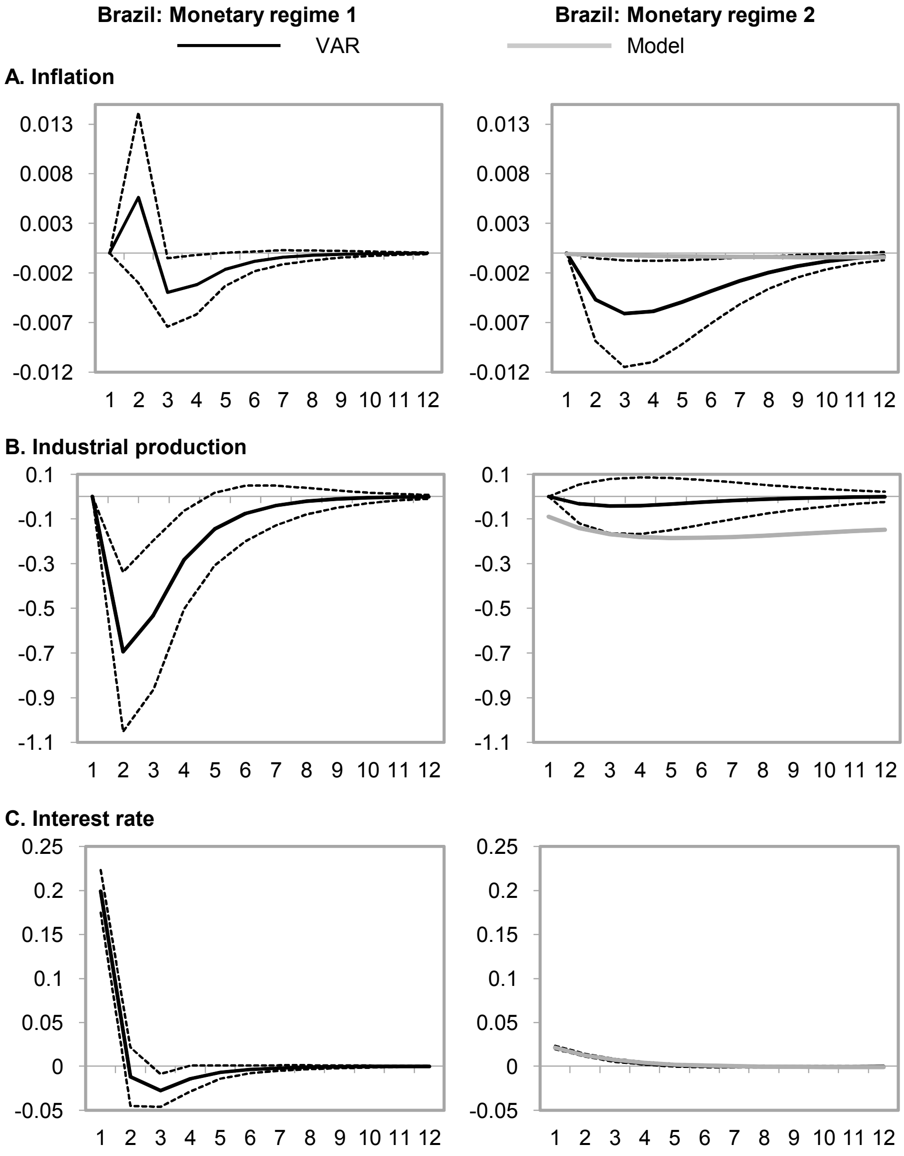
By and large, the shape of the impulse responses computed for the structural models tend to coincide with those of the VAR representation (Figure 1.2).[5] The VAR impulse responses suggest the presence of a price puzzle, whereby a positive interest rate shock is associated with an increase, rather than fall, in inflation, except for Brazil and Colombia in the current monetary regime. When a price puzzle exists, the response function is statistically significant only in the cases of Chile in the pre-1999 period and Mexico in the current policy regime. Only in the case of Colombia in the current policy regime do the impulse responses differ between the VAR and the structural model, with the latter exhibiting a price puzzle. Inclusion of a commodity price index in the VAR, as suggested in the empirical literature for the United States, does not solve the puzzle. This is consistent with the findings reported by Avendano and de Mello (2006), who estimate VAR and FAVAR models for these four countries over a similar time period and propose the inclusion of variables, such as monetary aggregates and measures of interest rate deviations from trend, to deal with the price puzzle in the impulse response functions.[6] Other studies have dealt with the price puzzle by adding considerably more structure or by including more endogenous variables in the VAR.[7] Notwithstanding these options for solving (or at least attenuating) the price puzzle, the VARs used to compute the impulse responses reported below were estimated including only inflation, the output gap and the interest rate as endogenous variables and the exchange rate as a predetermined variable, to facilitate comparison with the structural model.

The responses of the output gap to a monetary shock are negatively signed in Brazil and Chile for the unrestricted VAR in both policy regimes (even though they are statistically significant only in the first regime) and for the structural model in the post-1999 regime, which is the only period for which estimations could be performed. When a statistically significant positive reaction was computed (Colombia in the first monetary regime), it might be indicating that agents expect a reduction in future interest rates after the initial shock, which leads to a recovery in economic activity. In any case, the responses of economic activity to a monetary shock are milder in the current policy regime than in the previous one in all cases except Mexico. With regard to the interest rate response to a monetary shock, there is considerably more persistence in the current policy regime than in the pre-1999 period in Brazil and Chile and less so in Mexico.

It should also be noted that the impulse response estimates are fairly imprecise in some cases, which makes it difficult to ascertain whether they actually differ across policy regimes in a statistically significant manner. However, the results of the estimation of the structural models, showing important changes in the estimated policy reaction function, suggest that the changes in the responses across policy regimes are robust. The stronger estimated responses (except for Mexico) under managed exchange rates may be attributed to greater interest rate volatility due to the need to defend pre-announced pegs, especially against speculative attacks. Against this background, an important policy question is to identify the factors that lie behind the changes in macroeconomic volatility across policy regimes.

Figure 1.2. **Estimated impulse response to a monetary shock**

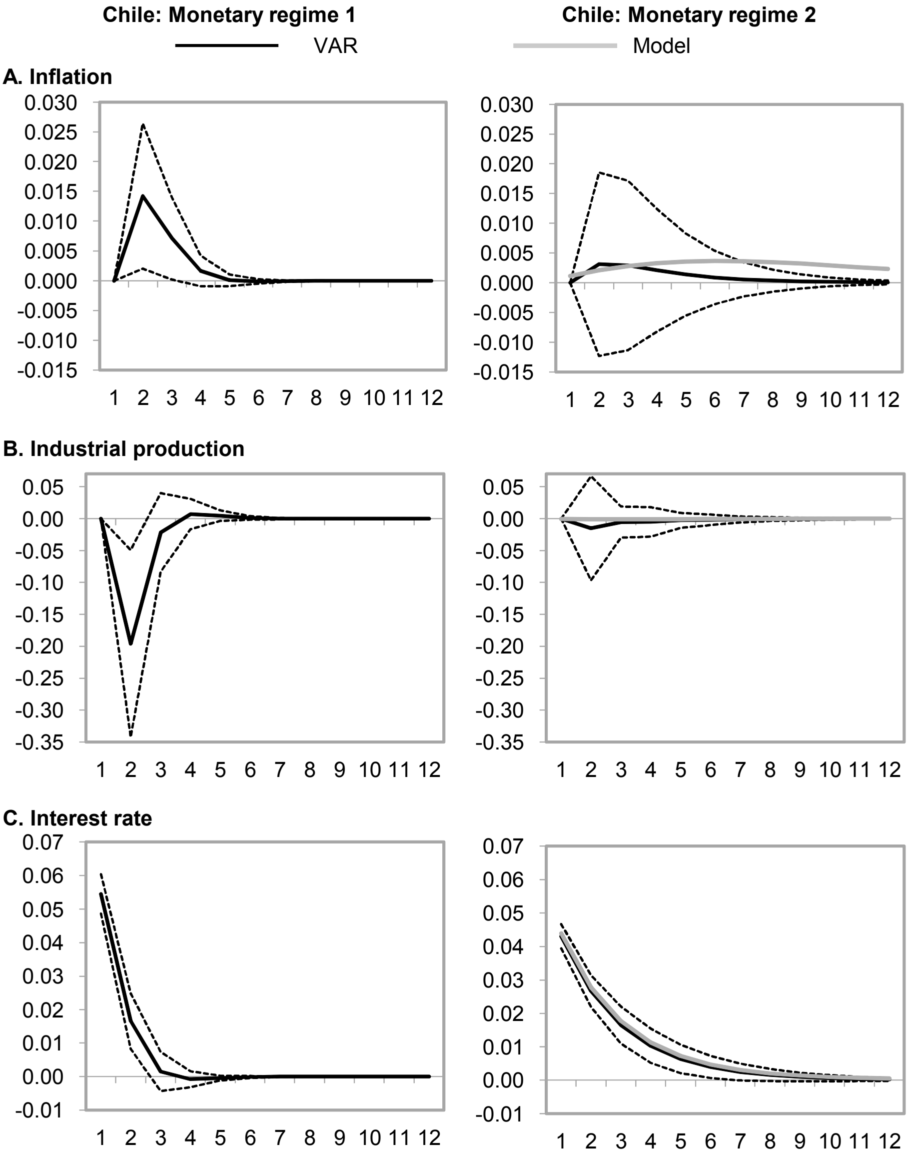
(Responses to a 1 standard-deviation innovation in the interest rate with ±2 standard-error bounds)



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

Figure 1.2. **Estimated impulse response to a monetary shock** (*cont'd*)

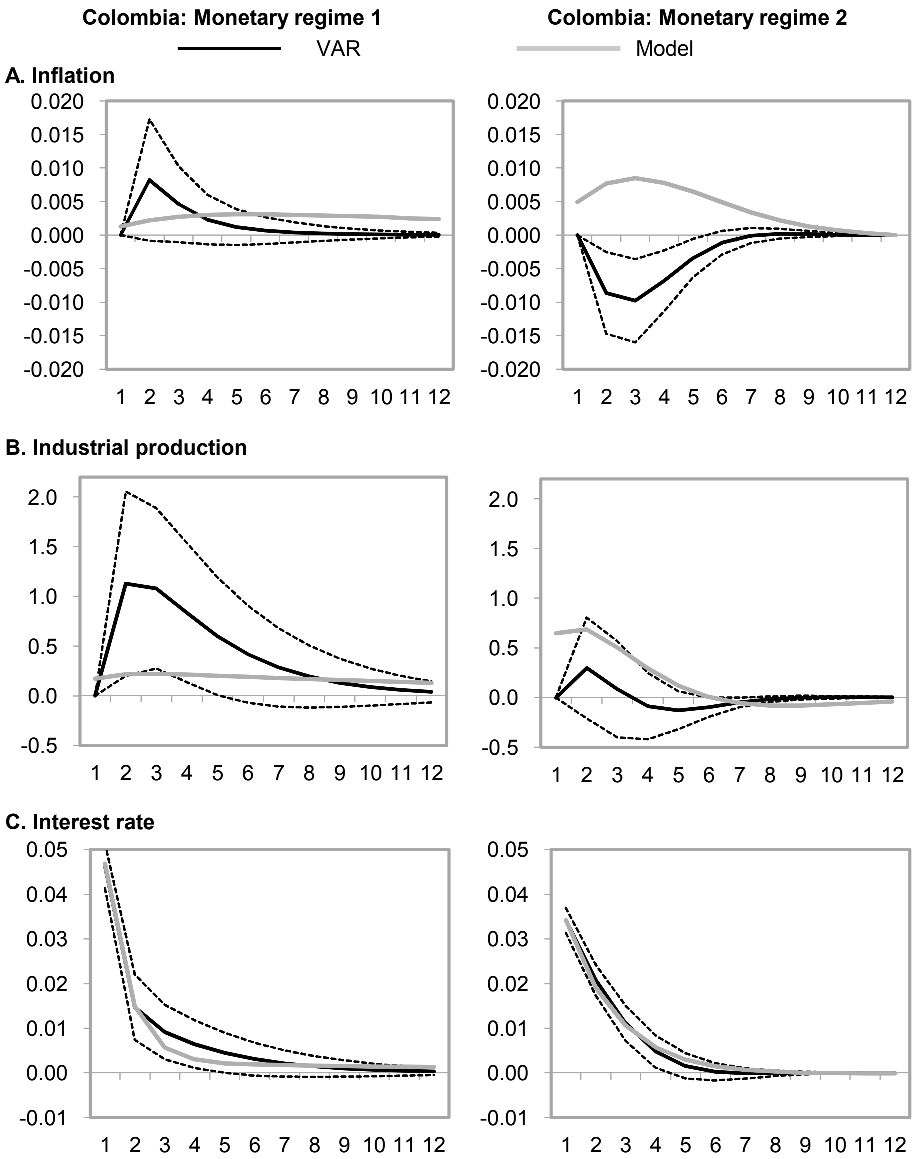
(Responses to a 1 standard-deviation innovation in the interest rate with ±2 standard-error bounds)



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

Figure 1.2. **Estimated impulse response to a monetary shock** (*cont'd*)

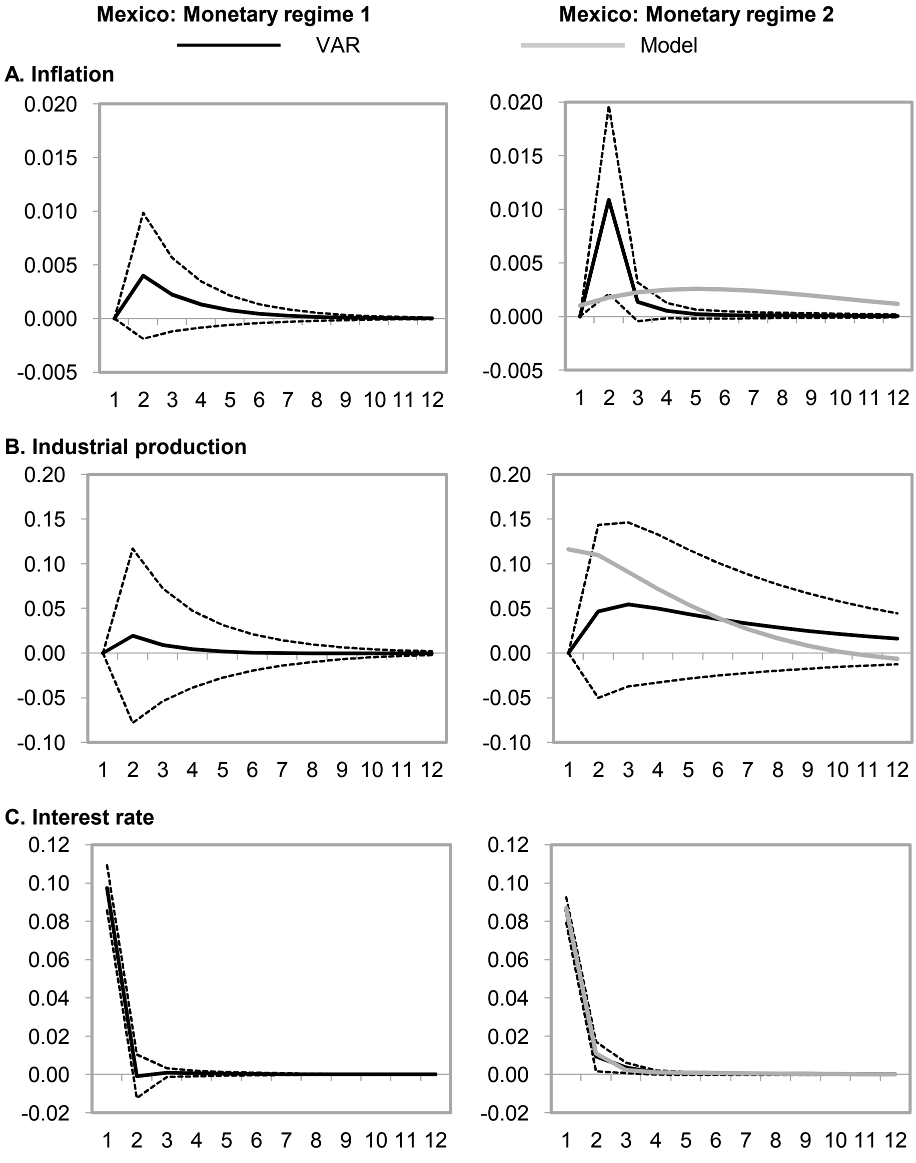
(Responses to a 1 standard-deviation innovation in the interest rate with ±2 standard-error bounds)



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

Figure 1.2. **Estimated impulse response to a monetary shock** (*cont'd)*

(Responses to a 1 standard-deviation innovation in the interest rate with ±2 standard-error bounds)



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

Counterfactual analysis

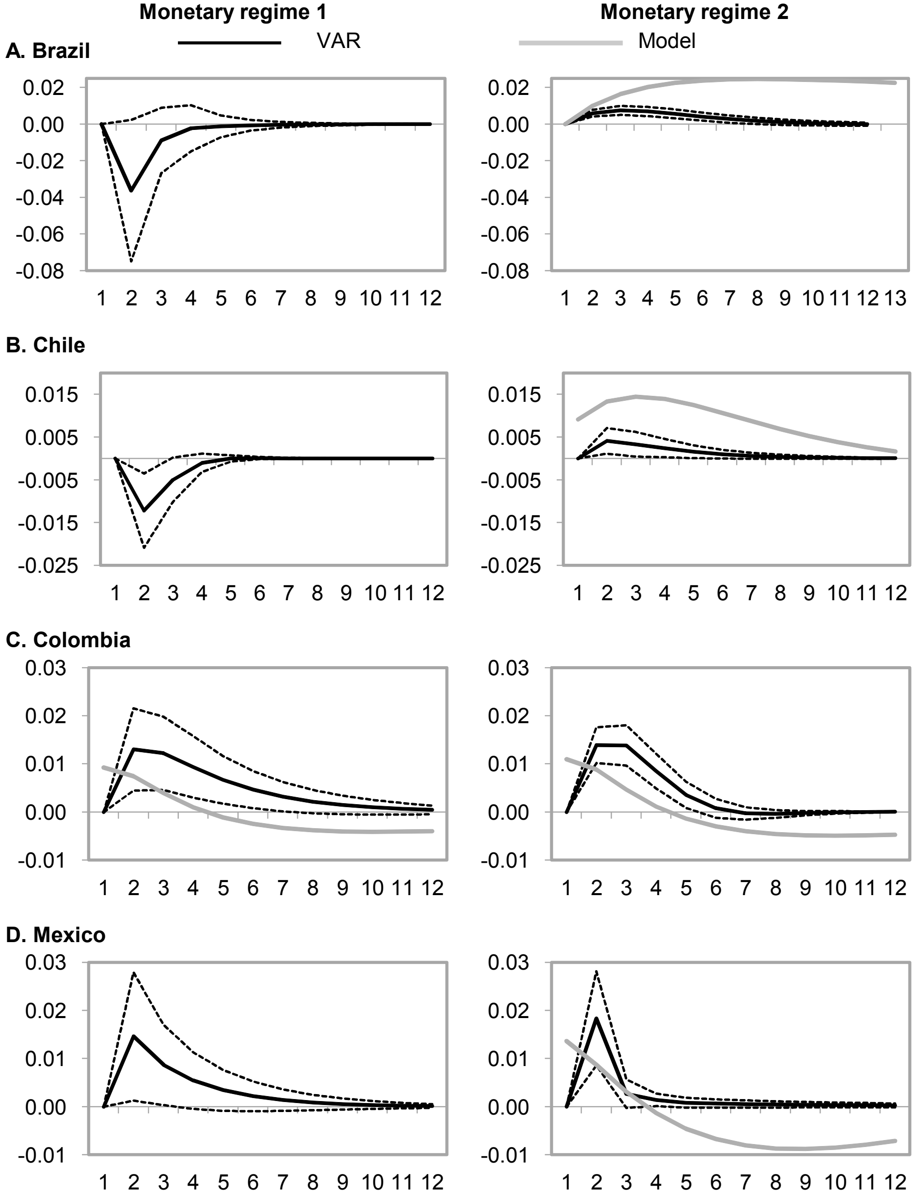
Policy versus shocks across policy regimes

Changes in output, inflation and interest rate volatility may be due to a shift in the policy regime but also to changes in the nature of the shocks hitting the economy. Different methodologies have been used to decompose the effect of policy and shocks on the variability of inflation and output over time. Cecchetti, Lopes-Lagunes and Krause (2004) construct an output-inflation variability frontier derived from a structural model and conclude that policy accounted for the bulk of the improvement in economic performance from the 1980s to the 1990s in a sample of industrial and emerging-market economies. An alternative approach consists of comparing volatility outcomes across time periods for different combinations of shocks and policy parameters. Using structural VAR analysis, Boivin and Giannoni (2002 and 2005) show that monetary policy has become more stabilising in the United States since the 1980s, because it has become more responsive to changes in inflation expectations. In turn, this stabilisation effect on both inflation and output is stronger than that of shocks, which have become milder. Instead, Ahmed, Levin and Wilson (2002) also estimate a structural VAR for the United States and show that milder shocks have accounted for most of the decline in output volatility since the 1980s. Likewise, the evidence reported by Moreno (2004), also for the United States, shows that more forward-looking price setting, due to greater flexibility in indexation mechanisms for wages and financial contracts, rather than the conduct of monetary policy *per se*, was the most important contributor to the fall in inflation variability in the 1990s relative to the 1980s.

When conducted under a managed exchange-rate regime, the focus of monetary policy is to defend a nominal peg, rather than to respond to changes in inflation expectations. Impulse responses can be computed for the unrestricted VAR and the structural model to assess the reaction of monetary policy to innovations in inflation across policy regimes (Figure 1.3). The unrestricted VAR estimations suggest that the central bank has reacted to inflationary shocks more strongly in the current policy regime both in Brazil and Chile, where monetary responses have also been more persistent.[8] Rather, monetary responses do not appear to have become stronger nor more persistent in the current policy regime in Colombia and Mexico. The impulse responses computed for the structural model are stronger than those computed for the unrestricted VAR in Brazil and Chile, and more persistent in all countries.

Figure 1.3. **Estimated monetary response to an inflationary shock**

(Responses of the interest rate to a 1 standard-deviation innovation in inflation with ±2 standard-error bounds)



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

To shed further light on the role of policy and shocks, a counterfactual exercise in the spirit of Ahmed, Levin and Wilson (2002), Stock and Watson (2002) and Moreno (2004) was conducted as follows. The volatility of variable *X* in vector *V*, for $V = (\pi, y, r)$, given the monetary policy regime of period *j* (P_j) and the structure of shocks of period *i* (S_i), for $j = (1,2)$ and $i = (1,2)$, is denoted by $\sigma_X^2(P_j, S_i)$. Based on this notation, accepting the null hypothesis $H_0^1 : \sigma_X^2(P_1, S_1) / \sigma_X^2(P_2, S_2) > 1$ implies that volatility fell in the current policy regime (period 2) relative to the period-1 policy regime, based on each period's own shocks and policy parameters. This can be due to changes in the nature of shocks and/or the policy regime. The fall in volatility due to shocks can be assessed by testing the following two hypotheses: $H_0^2 : \sigma_X^2(P_1, S_1) / \sigma_X^2(P_1, S_2) > 1$, where volatility falls in period 2 relative to period 1 by holding the policy setting in period 2 unchanged relative to period 1 (*i.e.* the parameters of the monetary reaction function do not change across monetary regimes), and $H_0^3 : \sigma_X^2(P_2, S_1) / \sigma_X^2(P_2, S_2) > 1$, where volatility falls in period 2 relative to period 1 holding the policy setting in period 1 constant as in period 2 (*i.e.* the parameters of the monetary reaction function estimated for period 2 are set to hold in period 1).

A fall in volatility due to policy can be assessed by testing the following two hypotheses: $H_0^4 : \sigma_X^2(P_1, S_2) / \sigma_X^2(P_2, S_2) > 1$, where volatility falls in period 2 under the period-2 policy setting, holding shocks constant as in period 1 (*i.e.* the parameters of the monetary reaction function differ across policy regimes, but the shocks estimated in period 2 are applied to period 1); and $H_0^5 : \sigma_X^2(P_1, S_1) / \sigma_X^2(P_2, S_1) > 1$, where volatility falls in period 1 under the period-2 policy setting, holding shocks constant as in period 1 (*i.e.* the parameters of the monetary reaction function differ across policy regimes, but the shocks estimated in period 1 are applied to period 2). The counterfactual exercise therefore allows the estimation of the volatilities that would arise from a given combination of shocks and monetary policy parameters, keeping the remaining structural parameters in the model unchanged across monetary regimes. The factors that militate in favour of greater macroeconomic stability can therefore be uncovered from the data.

The results of the counterfactual exercise, reported in Table 1.2, suggest that inflation became more volatile across policy regimes in Brazil and Colombia, essentially as a result of shocks. This may be associated with an increased volatility in agricultural commodity prices after 1999, whereas the shocks that affected these economies in the second half of the 1990s had operated predominantly through the exchange rate, rather than directly through inflation. On the other hand, the interest rate became less volatile in the current policy regime in all countries, except Chile, which is due to a change in the nature of shocks that have hit these economies since the change in these countries' exchange-rate regimes. Colombia is the only country where the change in policy regime is associated with a fall in output volatility, a finding that is essentially due to the change in the nature of the shocks that have hit the economy over the period of analysis.

The counterfactual analysis also shows that, controlling for differences in shocks, the policy regime shift appears to have increased interest-rate volatility in Brazil and Chile. Applying the period 1 policy parameters under the current shock scenario would have reduced volatility interest-rate volatility in these countries. This is most likely because shocks have become milder: a less persistent, less forward-looking monetary response, as in the previous regime, would have allowed the interest rate to revert to its pre-shock level more swiftly, thus reducing volatility. But less persistent, less forward-looking monetary responses might not have stabilised, and subsequently anchored, expectations during the initial phase of confidence-building after the regime transition. The counterfactual exercise also shows that applying the current monetary responses to the previous policy regime would have increased interest-rate volatility, although not by an amount that would make these volatilities statistically distinguishable. Given the magnitude of the shocks that hit these economies in the second half of the 1990s, the greater persistence in the monetary stance pursued in the current regime would have been inconsistent with exchange rate targeting and, therefore, exacerbated volatility.

These findings are related to a growing literature on the effect of gradualism in the conduct of monetary policy on inflation expectations. If the policy setting is credible and implemented in a transparent, forward-looking manner, monetary responses to inflationary shocks are mild (Woodford, 1999 and 2004), and the market's ability to forecast monetary policy is enhanced (Lange *et al.*, 2001). The empirical evidence reported by de Mello and Moccero (2006) for the same sample of Latin American countries under the current policy regime shows that greater variability in the monetary stance leads to higher volatility in expected inflation in Brazil, Colombia and Mexico. This suggests that interest-rate smoothing contributes to reducing volatility in inflation expectations, which makes it easier for the monetary authority to anchor expectations around the targeted level.

Table 1.2. **Counterfactual analysis: Brazil, Chile, Colombia and Mexico**[1]

	π		y		r	
			Brazil			
$H_0^1 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_2)$	0.29	***	1.39		33.38	***
$H_0^2 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_1,S_2)$	0.29	***	1.39		85.85	***
$H_0^3 : \sigma_X^2(P_2,S_1)/\sigma_X^2(P_2,S_2)$	0.28	***	1.38		43.52	***
$H_0^4 : \sigma_X^2(P_1,S_2)/\sigma_X^2(P_2,S_2)$	1.00		1.00		0.39	***
$H_0^5 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_1)$	1.01		1.01		0.77	
			Chile			
$H_0^1 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_2)$	0.93		1.07		0.80	
$H_0^2 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_1,S_2)$	0.88		1.07		1.49	
$H_0^3 : \sigma_X^2(P_2,S_1)/\sigma_X^2(P_2,S_2)$	0.92		1.07		1.04	
$H_0^4 : \sigma_X^2(P_1,S_2)/\sigma_X^2(P_2,S_2)$	1.05		1.00		0.53	**
$H_0^5 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_1)$	1.00		1.01		0.76	
			Colombia			
$H_0^1 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_2)$	0.60	**	1.65	**	1.72	**
$H_0^2 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_1,S_2)$	0.60	**	1.59	*	2.36	***
$H_0^3 : \sigma_X^2(P_2,S_1)/\sigma_X^2(P_2,S_2)$	0.60	**	1.93	***	1.62	*
$H_0^4 : \sigma_X^2(P_1,S_2)/\sigma_X^2(P_2,S_2)$	1.00		1.04		0.73	
$H_0^5 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_1)$	0.99		0.86		1.06	
			Mexico			
$H_0^1 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_2)$	1.09		0.74		2.31	***
$H_0^2 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_1,S_2)$	1.10		0.75		2.08	***
$H_0^3 : \sigma_X^2(P_2,S_1)/\sigma_X^2(P_2,S_2)$	1.10		0.75		2.95	***
$H_0^4 : \sigma_X^2(P_1,S_2)/\sigma_X^2(P_2,S_2)$	1.00		0.99		1.11	
$H_0^5 : \sigma_X^2(P_1,S_1)/\sigma_X^2(P_2,S_1)$	1.00		1.00		0.78	

1. The numbers reported are the ratios between the two standard deviations. (***), (**) and (*) denote statistical significance at the 1%, 5% and 10% levels, respectively, on the basis of an *F* test.

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

Conclusions

This chapter estimated a simple New Keynesian structural model for four Latin American countries that adopted a monetary regime characterised by inflation targeting and floating exchange rates in 1999: Brazil, Chile, Colombia and Mexico. The chapter's main findings are as follows:

- The post-1999 regime, characterised by inflation targeting and exchange rate flexibility, has been associated with stronger, more persistent responses by the monetary authority to changes in expected inflation in Brazil and Chile. The monetary stance has become less counter-cyclical in Colombia and Mexico than in the previous policy regime, but more counter-cyclical in Brazil. Mexico is the only country in the sample where changes in the nominal exchange rate were found to be statistically significant in the central bank's reaction function.

- The impulse reaction functions computed for unrestricted VARs on the interest rate, inflation and the output gap, as well as for the structural model, suggest that the responsiveness of monetary policy to inflationary shocks became stronger and more persistent in Brazil and Chile in the current regime. This is consistent with the formal abandonment of exchange rate targeting in these countries and the adoption of a more pro-active policy stance underpinned by the shift to inflation targeting.

- Lower interest-rate volatility in the post-1999 period owes much to a more benign economic environment. The change in monetary regime has not yet resulted in a reduction in output volatility, a finding that may be attributed to the relatively short time span of analysis. Colombia is nevertheless an exception, where greater output stability was found to be due essentially to milder shocks in the current policy regime, rather than the regime change itself. Inflation volatility was found to have increased in Brazil and Colombia, again due to a change in the nature of shocks. Given that the current policy regime has been characterised by milder shocks, the remaining volatility in the interest rate is due to greater monetary policy responsiveness of expected inflation, despite interest-rate smoothing, at least in the cases of Brazil and Chile.

Notes

1. See Cecchetti and Debelle (2006) for evidence and a survey of the recent literature on univariate analysis, Cecchetti, Flores-Lagunes and Krause (2004) for cross-country evidence based on structural models, and Boivin and Giannoni (2005) for evidence based on VAR modelling.

2. Peru is another inflation targeter in Latin America, but it was not included in the sample. This is because inflation targeting was formally adopted in 2002, which would have severely reduced degrees of freedom in the post-regime change sample. The Peruvian economy is also highly dollarised, which makes the monetary transmission mechanism somewhat different from those in the countries under examination (Leiderman *et al.*, 2006).

3. The timing of this structural break does not need to be known *a priori*. The date of the break is estimated from the data as the observation that maximises the absolute value of the unit root statistics (Zivot and Andrews, 1992).

4. The ADF test, unlike the other unit root tests, suggested the output gap exhibits a unit root only in Mexico.

5. The second monetary regime for Brazil excludes the transition period January-June 1999.

6. The authors suggest that introduction of an indicator of interest rate misalignment, defined as the interest rate deviations from a HP-filtered series, as an endogenous variable in the VAR instead of the policy interest rate solves the price puzzle for Chile and reduces it considerably for Mexico.

7. See Avendano and de Mello (2006) for a review of the empirical literature on monetary VARs for Latin American countries.

8. In Chile, a positive inflationary shock induces a negative reaction by the interest rate in the first period.

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Annex 1.A1

Solving the rational expectations model

This Annex describes the rational-expectations solution to the structural model defined by System (1.1)-(1.3) and computes the associated impulse response functions (IRFs). The system can be re-written in matrix form as:

$$\begin{bmatrix} 1 & -\lambda & 0 \\ 0 & 1 & \phi \\ 0 & -(1-\rho)\gamma & 1 \end{bmatrix} \begin{bmatrix} \pi_t \\ y_t \\ r_t \end{bmatrix} = \begin{bmatrix} \delta & 0 & 0 \\ \phi & \mu & 0 \\ (1-\rho)\beta & 0 & 0 \end{bmatrix} E_t \begin{bmatrix} \pi_{t+1} \\ y_{t+1} \\ r_{t+1} \end{bmatrix} + \begin{bmatrix} (1-\rho) & 0 & 0 \\ 0 & (1-\mu) & 0 \\ 0 & 0 & \rho \end{bmatrix} \begin{bmatrix} \pi_{t-1} \\ y_{t-1} \\ r_{t-1} \end{bmatrix} +$$

$$+ \begin{bmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & \tau \end{bmatrix} \begin{bmatrix} 0 \\ 0 \\ e_t \end{bmatrix} + \begin{bmatrix} u_{\pi_t} \\ u_{y_t} \\ u_{r_t} \end{bmatrix}$$

Or, in compact notation:

$$BX_t = AE_t X_{t+1} + CX_{t-1} + TH_t + u_t, \qquad (1.A1.1)$$

where $u_t \sim iid(0, D)$ is the vector of structural errors, D is the variance-covariance matrix, 0 denotes a 3x1 vector of zeros, $X_t = (\pi_t \, y_t \, r_t)'$, $H_t = (0\,0\,e_t)'$, and *B*, *A*, *C* and *T* are matrices of structural parameters.

When a unique rational-expectations solution to System (1.A1.1) exists, it can be written in reduced form as (Anderson, 2006):

$$X_t = \Omega X_{t-1} + \sum_{s=0}^{+\infty} F^s \Phi u_{t+s} + \sum_{s=0}^{+\infty} F^s \Phi T H_{t+s}, \qquad (1.A1.2)$$

where $\Phi = (B - A\Omega)^{-1}$ and $F = \Phi A \Omega$.

Given that $E_t(u_{t+s}) = 0, \forall s \geq 1$, and assuming a random-walk process for the (log) exchange rate, Equation (A1.2) simplifies to:[1]

$$X_t = \Omega X_{t-1} + \Phi u_t + \Phi TH_t \,. \qquad (1.A1.3)$$

Using an undetermined coefficient approach to solve (1.A1.1), it follows that Ω must satisfy the following matrix polynomial equation: $\Omega = (B - A\Omega)^{-1}C$. For Ω to be an admissible solution to the system, it must by real-valued and exhibit stationary dynamics. Also, because Ω is a non-linear function of the structural parameters, there can be potentially no solution to the system or, instead, multiple stationary equilibria. Uhlig (1997) proposes an algorithm to solve for Ω and characterises the uniqueness and stationarity conditions for the solution.[2] However, when multiple equilibria exist, the algorithm says nothing about how to choose a particular solution among them. In such a case, the Blanchard and Kahn (1980) criterion is used.

The IRFs can be computed as the derivative of X_t with respect to u_{t-s}, using the following infinite-vector moving-average representation (VMA) derived from Equation (1.A1.3):

$$X_t = \sum_{s=0}^{+\infty} \Omega^s \Phi u_{t-s} + \sum_{s=0}^{+\infty} \Omega^s \Phi TH_{t-s} \,.$$

Notes

1. Remember that *et* is the first difference of the (log) exchange rate. Also, note that the implied reduced-form of the structural model defined by System (1.1)-(1.3) is simply a VAR(1) model with highly non-linear parameter restrictions (Moreno, 2004).

2. Anderson (2006) compares Uhlig (1997)'s algorithm with other methods to solve rational expectations models.

List of acronyms

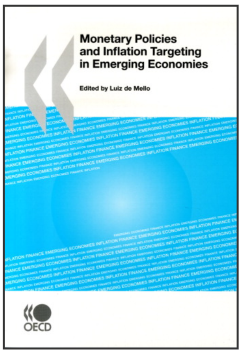
ADF	Augmented Dickey-Fuller Test
AR	Autoregressive Model
BCB	Central Bank of Brazil (*Banco Central do Brasil*)
BER	Bureau for Economic Research
BI	Bank Indonesia
BIS	Bank for International Settlements
BRSA	Banking Regulatory and Supervisory Agency
CBC	Central Bank of Chile
CBT	Central Bank of Turkey
CDT	Colombian Interest Rate (*Certificado de Depósito a Término*)
CETES	Mexican Interest Rate (*Certificados de la Tesorería de la Federación*)
CMN	Brazilian National Monetary Council (*Conselho Monetário Nacional*)
CNB	Czech National Bank
COPOM	Brazilian Monetary Policy Committee (*Comitê de Política Monetária*)
CPI	Consumer Price Index
CPIX	Consumer Price Index (excluding mortgage interest costs)
DSGE	Dynamic Stochastic General Equilibrium
EMBI	Emerging Market Bond Index
FAVAR	Factor-Augmented Vector Autoregressive Model
FDI	Foreign Direct Investment
FIML	Full Information Maximum Likelihood
FPAS	Forecasting and Policy Analysis System
FTO	Fine-Tuning Operations
GDP	Gross Domestic Product
GNP	Gross National Product
IGP-DI	Brazilian General Price Index (*Índice Geral de Preços - Disponibilidade Interna*)

IMACEC Chilean Monthly Economic Activity Index (*Indicador Mensual de Actividad Económica*)
IMF International Monetary Fund
IMT Inflation Management Team
INPC Mexican Consumer Price Index (*Índice Nacional de Precios al Consumidor*)
IPC Chilean and Colombian Consumer Price Indices (*Índice de Precios al Consumidor*)
IPCA Brazilian Consumer Price Index (*Índice Nacional de Preços ao Consumidor*)
IRF Impulse Response Functions
IT Inflation Targeting
M-GARCH Multivariate Generalised Autoregressive Conditional Heteroskedasticity Model
MPC Monetary Policy Committee
MPR Monetary Policy Report
OMO Open-Market Operations
PP Phillips-Perron Test
QPM Quarterly Projection Model
SADC Southern African Development Community
SARB South African Reserve Bank
SBI Bank Indonesia Certificates (*Sertifikat Bank Indonesia*)
SDIF Savings Deposit Insurance Fund
SEE Survey of Economic Expectations
SELIC Brazilian Policy Interest Rate
SIC Schwarz Information Criterion
TJLP Brazilian Long-Term Interest Rate (*Taxa de Juros de Longo Prazo*)
TPM Chilean Policy Interest Rate (*Tasa de Política Monetaria*)
TR Brazilian Reference Interest Rate (*Taxa Referencial de Juros*)
UF Chilean *Unidad de Fomento*
VAR Vector Autoregressive Model
VMA Vector Moving Average

Table of contents

Tables

Figures



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