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Inflation Targeting and Monetary Analysis in Chile and Mexico

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Abstract

This paper studies the usefulness of real economic activity, monetary, and exchange rate indicators in predicting inflationary developments in Chile and Mexico. In so doing the investigation employs a nested P-star/Phillips curve model and alternative Phillips curve specifications. The analysis shows that in both economies real money and real output gaps are statistically significant in predicting inflation's deviations from target. In contrast, exchange rate indicators are consistently significant in predicting inflation only in Mexico. The findings are robust to using alternative money, output, and exchange rate indicators.

JEL Classifications: E50, F41.

Keywords: Inflation targeting, monetary policy, P-star, Phillips curve, forecasting.

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1 Introduction

Inflation targeting is a monetary policy strategy that became popular during the 1990s and many advanced and emerging market economies have formally adopted this policy. There is a significant literature explaining various aspects of inflation targeting; inter alia, Bernanke, Laubach, Mishkin, and Posen (1999), Amato and Gerlach (2002), Truman (2003), Bernanke and Woodford (2004), Martin and Milas (2004), Carare and Stone (2006), and Sánchez-Fung (2008). Why have countries been formally adopting inflation targeting? The reasons are diverse, including pitfalls controlling monetary aggregates, complications arising from experimentation with various exchange rate arrangements, and a resolved commitment to low and stable inflation.

In that context, this paper addresses the following question: Is there a role for monetary and open economy indicators in conducting monetary policy in these economies? Several papers tackle closely related issues, inter alia, Baltensperger, Jordan, and Savioz (2001) on Switzerland, and Rudebusch and Svensson (2002), Trecroci and Vega (2002), and Gerlach and Svensson (2003) on the Euro area. Interestingly, on money's role in economies implementing inflation targeting King (2002, page 162) notes that "…there is a paradox in the role of money in economic policy. It is this: that as price stability has become recognised as the central objective of central banks, the attention actually paid by central banks to money has declined".

To be sure, monetary policy making can benefit from using money as an information and/or predictive variable. The first role critically depends on the money demand function's stability - a topic widely studied in the literature (e.g. Goldfeld and Sichel, 1990). Basically, monetary aggregates may provide useful information if they are expected to help in predicting imperfectly observed variables of interest to the monetary authorities (e.g. aggregate output). In contrast, money's role as a predictor becomes relevant if the central bank's strategy (e.g. inflation targeting) implies using expectations of future variables that it can help to predict (see Svensson, 1997).

This paper investigates money's potential usefulness in predicting inflationary developments. Particularly, the paper inquires into monetary indicators' predictive content alongside other key macroeconomic fundamentals in two inflation targeting Latin American economies: Chile and Mexico.[1] In so doing the investigation employs an augmented P-star model and alternative Phillips curve specifications. Hallman, Porter, and Small (1991) advanced the P-star model with the aim of investigating the M2 monetary aggregate's value for inflation forecasting purposes in the USA. The paper considers an augmented version of the model, similar to that analysed by Trecroci and Vega (2002) and Gerlach and Svensson (2003), but also incorporating open economy elements. This exercise seems potentially fruitful in the light of, for instance, Calvo and Reinhart's (2002) and Edwards's (2006) arguments on the exchange rate's pervasive role in developing economies, and Svensson's (2000) insights about open economy inflation targeting.

The remaining sections proceed as follows. Section 2 recounts key developments in monetary policy and inflation targeting in Chile and Mexico. Section 3 outlines a framework nesting the Phillips curve and P-star models of inflation. Section 4 explains the data and its univariate characteristics, and estimates several gap measures. Section 5 analyses the Phillips curve/P-star model of inflation and money, and alternative Phillips curve specifications, using Granger causality tests. This part of the paper also runs checks using alternative money, output, and exchange rate indicators. Section 6 concludes.

2 Monetary Policy and Inflation Targeting in Chile and Mexico

Several Latin American countries have adopted inflation targeting (IT) as their monetary policy strategy.[2] Amongst this group, Chile and Mexico provide interesting cases, not least because similar circumstances and corresponding policy responses have in the past been observed in these economies (e.g. Edwards, 1998). It is worth noting that a monetary policy framework consists of various inputs. Amongst these is an inflation target, alongside accountability, transparency, and the analytical ability to operate the regime. In Chile and Mexico these requisites evolved during the 1990s. Particularly, the Central Bank of Chile adopted inflation targeting with the aim of achieving disinflation. In contrast, the Bank of

[1] It is worth pointing out that these countries have in the past provided a fertile ground for comparative economic analyses. See, for example, Arrau and De Gregorio (1993), Edwards (1998), and Schmidt-Hebbel and Werner (2002).
[2] Mishkin and Savastano (2000), and Corbo and Schmidt-Hebbel (2001) provide further insights on IT in Chile and Mexico and in other Latin American economies.

Mexico did so in the context of a relatively low and stable inflation. What follows briefly overviews key developments in these economies' monetary policy making.

2.1 Chile

Chile formally granted independence to its central bank in 1989. Alongside this step came a mandate to adopt inflation as monetary policy's main objective. In 1990 a more formal monetary policy stance was signalled by announcing an inflation target range for 1991. It is worth noting that from the mid-1980s until August 1999 the Central Bank of Chile had an exchange rate band regime in place (see Morandé and Tapia, 2002). Nevertheless, following the formal introduction of inflation targeting the authorities have always made clear that their primary objective is achieving the pre-announced inflation target. And as can be seen in Table 1, and graphically in Figure 3, Chile's inflation targeting regime has helped in bringing down inflation from double to single digits. Importantly, a sound fiscal stance and a solid financial system regulation underpin Chile's success in operating an inflation targeting strategy.[3]

2.2 Mexico

In December 1987 Mexico introduced the 'Pacto' stabilisation programme. A byproduct of this effort was that between 1987 and 1991 monetary policy was based on a pegged exchange rate regime. The strategy also led to reducing government spending, and this prudent fiscal management helped in taming inflation. Moreover, Cecchetti, Flores-Lagunes and Krause (2000) show that monetary policy in Mexico became more efficient, as measured by the lower variability of inflation and output, from 1991 onwards. This was at least in part due to the policies undertaken during the late 1980s.

Accordingly, and with the aim of establishing a more flexible exchange rate framework, the Bank of Mexico adopted exchange rate bands in 1991. However, the regime was put under pressure from December 1994 in the midst of the well-known Mexican crisis. The economic and political developments leading to that meltdown implied a substantial widening of the prevalent exchange rate band and a corresponding fall in the peso.

So at the beginning of 1995 the Mexican authorities aborted the pegged exchange rate regime, switching to a policy in which the monetary base played a major role (see Khamis and Leone, 2001). The Bank of Mexico continued to announce its targets for money and inflation in the next two years, but without much success. Still, following the formal implementation of an inflation targeting strategy in 1999 (as dated by Mishkin and Schmidt-Hebbel, 2001, Table 2), Mexico's monetary policy has been consolidating.

Table 1: Chile and Mexico
Inflation targeting: Adoption dates and targets width

Country	Date introduced	Target price index	Target width	
Chile	January 1991	Headline CPI	1991 1992 1993 1994 1995 1996 1997 1998 1999 2000 2001	15-20% 13-16% 10-12% 9-11% ±8% ±6.5% ±5.5% ±4.5% ±4.3% ±3.5% 2-4%
Mexico	January 1999	Headline CPI	1999 2000 2001	13% <10% 6.5%

Source: Mishkin and Schmidt-Hebbel (2001), Table 2.
Notes: CPI = consumer price index.

3 A Model of Inflation and Money

Gerlach and Svensson (2003), and Trecroci and Vega (2002) consider a nested version of the price gap and Phillips curve models for empirical purposes. The model comprises the following elements

$$\pi_t = \pi^e_{\langle t|t-1\rangle} + \eta(y_{t-1} - y^*_{t-1}) + \varphi(\tilde{m}_{t-1} - \tilde{m}^*_{t-1}) + \xi_t , \tag{1}$$

$$\pi^e_{\langle t|t-1\rangle} = \hat{\pi}_t , \tag{2}$$

[3] For further insights on inflation targeting's evolution in Chile see Schmidt-Hebbel and Tapia (2002).

$$\tilde{m}_t \equiv (m_t - p_t) = \beta_0 + \beta_y y_t + \beta_R R_t + \nu_t\,, \tag{3}$$

$$p_t^* \equiv m_t - \beta_0 - \beta_y y_t^* - \beta_R R_t^*\,, \tag{4}$$

$$\tilde{m}_t - \tilde{m}_t^* \equiv (m_t - p_t) - (m_t - p_t^*) \equiv -(p_t - p_t^*) = (m_t - p_t) - \beta_0 - \beta_y y_t^* - \beta_R R_t^*\,. \tag{5}$$

Rewriting equation (5) to highlight the indicators underlying it yields

$$p_t^* - p_t \equiv \beta_y\left(y_t - y_t^*\right) + \beta_R\left(R_t - R_t^*\right) + \nu_t \equiv \tilde{m}_t - \tilde{m}_t^*\,. \tag{5a}$$

- Equation (1) is a mixture of a Phillips curve and a price gap model of inflation, where π, π^e, y, y^*, $\tilde{m}$, and $\tilde{m}^*$ are actual and expected inflation rates, output and potential output, and actual and long-run real money balances, respectively. η and φ are parameters to be estimated empirically. ξ is expected to be a well-behaved disturbance term.
- Equation (2) specifies how inflation expectations (π^e) are formed, and $\hat{\pi}$ is the central bank's inflation objective. Note that this implies that, in an inflation targeting regime, the model focuses on explaining deviations of inflation from its target.
- Equation (3) is a long-run money demand function, which assumes a standard specification. y is real output, p is price, and R measures the opportunity cost of holding money. β_0, β_y, and β_R are parameters to be estimated, and ν is expected to be a well-behaved disturbance term.
- Equation (4) generates the equilibrium price level (p^*), for a level of the money stock assuming that the other variables in the model are at their equilibrium levels, by inverting the long run money demand equation (3).
- Equation (5) defines the real money gap as the negative of the price gap. Additionally, equation (5a) underlines the model's accounting for the role of information from the good's market through the output gap, the stance of monetary policy via the interest rate gap, and the money market in the form of the monetary overhang ν_t.

A further important element in empirically assessing this model is the role of the exchange rate. Svensson (2000) advances three key reasons why this is particularly important for inflation targeting open economies. Firstly, the exchange rate explicitly allows for an additional channel via which monetary policy can be transmitted. Secondly, the exchange rate is a forward-looking variable, and therefore can provide valuable information in designing and implementing monetary policy. Thirdly, foreign shocks mainly propagate through the exchange rate.

In addition to the above factors, Calvo and Reinhart (2002, page 394) note that "…central bankers in emerging market economies appear to be extremely mindful of external factors in general and the foreign exchange value of their currency, in particular". Some reasons Calvo and Reinhart highlight in rationalising the exchange rate's significant role in monetary policy making, and which give rise to what they label *fear of floating*, even under an inflation targeting regime, are

- liability dollarisation,
- output costs associated with exchange rate fluctuations,
- inelastic supply of funds during crises, and
- lack of credibility and fear of loss of access to the international capital markets.

Given these factors the study augments equation (1) as follows:

$$\pi_t = \pi^e_{\langle t|t-1\rangle} + \eta(y_{t-1} - y^*_{t-1}) + \varphi(\tilde{m}_{t-1} - \tilde{m}^*_{t-1}) + \alpha(e_{t-1} - e^*_{t-1}) + \xi_t \,. \tag{6}$$

Equation (6) incorporates departures of actual (e) from equilibrium (e^*) exchange rates, where α is a parameter to be estimated.

4 Data

The econometric modelling employs data on money, real income, prices, interest rates, exchange rates, and inflation objectives. The data are monthly, and span from 1990.01 to 2001.06 for Chile, and from 1986.01 to 2001.06 for Mexico.[4] In what follows all the time series are expressed in logs, excepting the interest rates - which are expressed in percentage points. Table A1 in Appendix 1 contains data definitions and sources.

In determining the data's order of integration the study applies the augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1979).[5] Panel A in Table 2 displays the results from applying this test. All the series seem to contain a unit root in their levels, but become stationary after being differenced once. The exceptions are e and the producer price (p^D) for

[4] Note that for Chile the data covers only the inflation targeting regime. For Mexico the statistical information encompasses several periods: a pegged exchange rate policy (1987-1991), an exchange rate band (1991-1994), monetary targeting (1995-1998), and inflation targeting (1999-present).

[5] Sekioua and Karanasos (2006), and Sekioua and Zakane (2007), inter alia, employ alternative unit root tests.

Chile, which appear to be $I(2)$. But p^D is transformed before entering the purchasing power test (PPP) test below. And, in fact, $\left(p^D - p^*\right)$ is $I(1)$.

Table 2: Chile and Mexico
Unit roots and cointegration tests

A. ADF unit root tests				
	Levels		Annual differences ($\Delta 12$)	
Variables	Chile (1991.02-2001.06)	Mexico (1987.02-2001.06)	Chile (1992.02-2001.06)	Mexico (1988.02-2001.06)
$\tilde{m}$	-2.36	-0.63	-2.93*	-5.19**
y	-2.42	-0.72	-3.01*	-4.36**
R	-2.71	-2.51	-3.48*	-3.93**
e	0.33	-2.48	-1.67	-3.82**
p	-3.27	-2.04	-2.88*	-6.78**
$p^D - p^*$	-0.78	-1.94	-3.55**	-7.23**
B. Long run solutions to ADL equations				
	Chile (1991.01-2001.07)		Mexico (1988.01-2001.06)	
Variables	Money demand Equation (1)	PPP Equation (2)	Money demand Equation (3)	PPP Equation (4)
Cons	-1.21(0.44)**	6.24 (0.02)**	9.03 (0.69)**	1.53 (0.0)**
β_y	1.72 (0.07)**		1.38 (0.14)**	
β_R	-0.02 (0.01)*		-0.003 (0.001)**	
λ		0.92 (0.09)**		0.88 (0.087)**
ADF test	-8.05 (-4.40)	-7.90 (-3.98)	-9.46 (-4.37)	-8.90 (-3.96)
$WALD-\chi^2$	493.74 (2)**	97.84 (1)**	262.72 (2)**	101.36 (1)**

Notes: In Panel A, ** and * denote rejection of the unit root hypothesis at the 1% and 5% level, respectively. In Panel B, coefficients' standard errors are inside parentheses. ADL = autoregressive distributed lag. Critical values (1%) for the ADF test applied to the residuals of the cointegrating relations are from MacKinnon (1991), and are shown in parentheses next to the corresponding ADF statistic. A significant test means rejection of the hypothesis of nonstationarity, i.e. a cointegrating relationship exists between the variables under analysis. ** and * denote a coefficient/test is significant at the 1% and 5% levels, respectively. The ADL equations allow for twelve lags of each of the variables considered. Further details on these procedures can be obtained from the author upon request. $WALD-\chi^2$ is a test of the null that all long-run coefficients are zero, with $\chi^2(\cdot)$ distribution.

The only remaining concern relates to the annual change $(\Delta 12)$ in Chile's exchange rate. Even so, an alternative transformation shows that in fact the monthly change in the exchange rate (Δe) is also stationary according to an ADF test statistic (-6.77) significant at the 1% level. Therefore nonstationarity does not seem to be an inherent feature of the change in the exchange rate in Chile, with the results for the annual change likely to be driven by abnormal values in that particular transformation of the series. The time series properties of the data can be further examined by inspecting Figures 1 and 2 for Chile and Mexico, respectively.

Figure 1

Chile, 1990-2001.

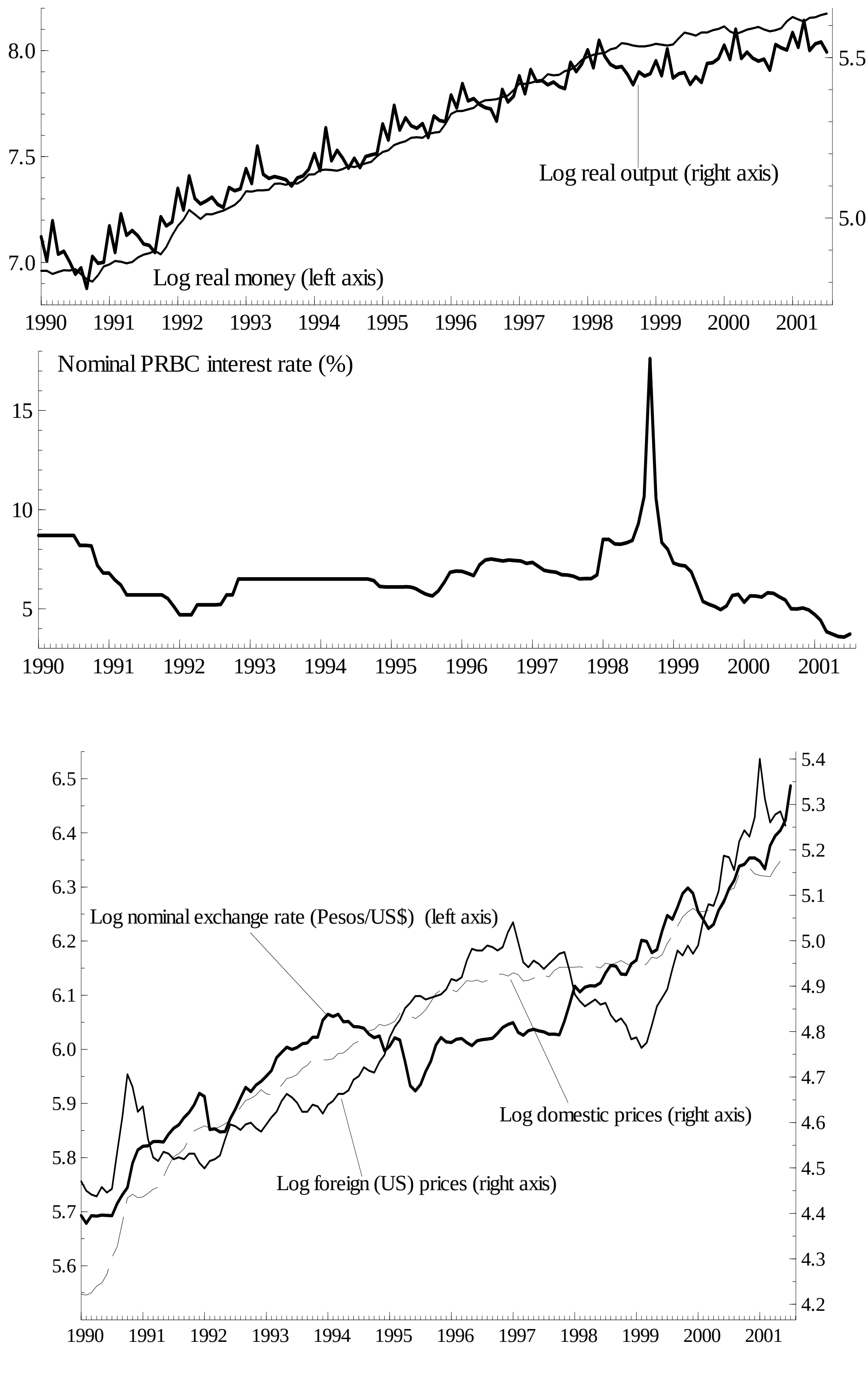
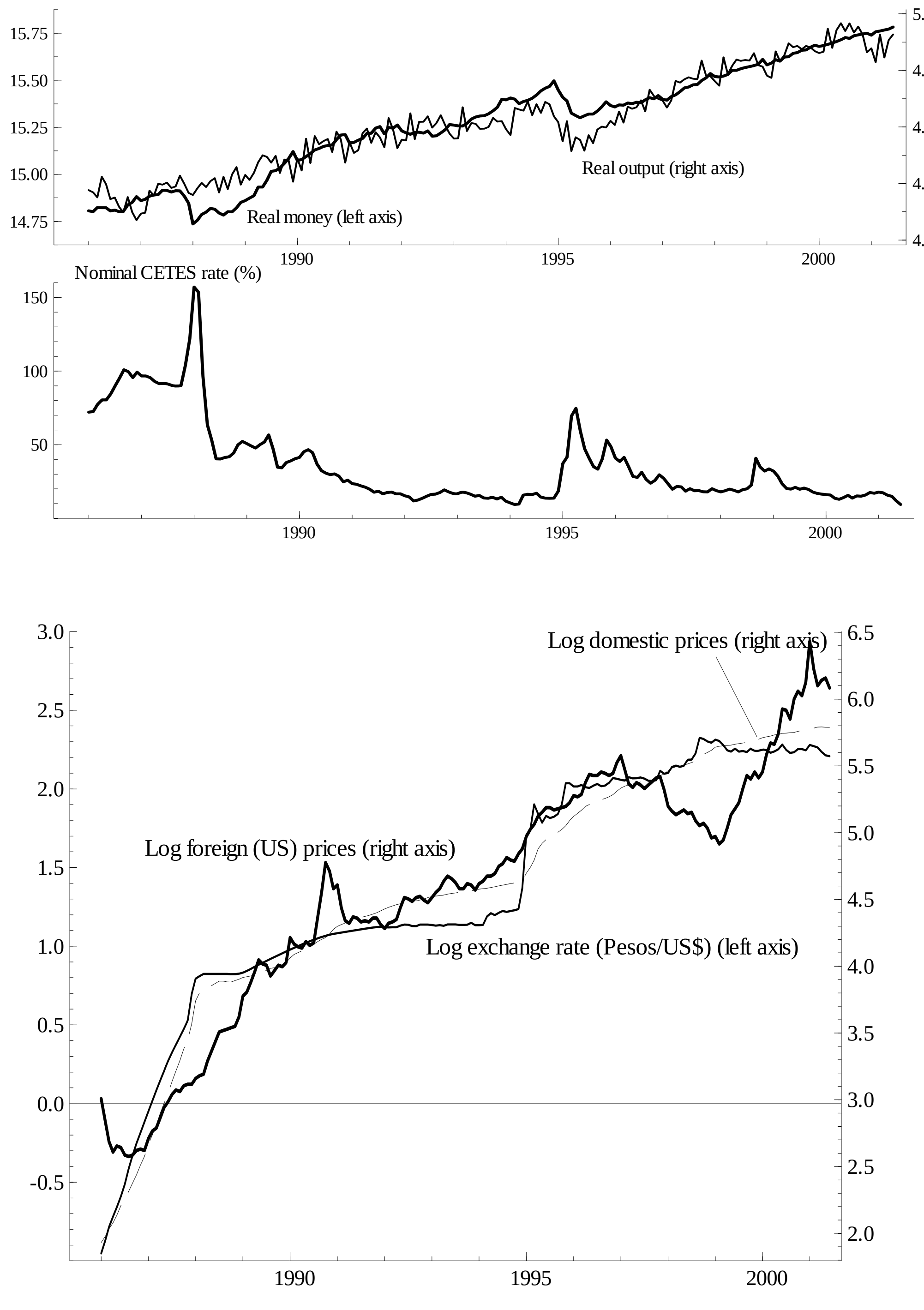


Figure 2

Mexico, 1986-2001.



In what follows $\hat{\pi}$ is the annual inflation target announced by the Central Bank of Chile and the Bank of Mexico (see Table 1). However, since the Bank of Mexico adopted a formal inflation target from January 1999, according to Mishkin and Schmidt-Hebbel (2001), the paper uses expected (forecast) inflation π^e up to that point in time in estimating the model. And that variable is proxied by the filtered series derived from π after applying a univariate structural time series model and the Kalman filter (see Harvey, 1989). This variable aims at reproducing economic agents' expected inflation by using a technically compelling technique. Appendix 2 provides further details on this procedure. Figures 3 and 4 display actual and target inflation for Chile, and actual, expected, and target inflation for Mexico, respectively.

The output gap series, $y - y^*$, are deviations of log output from its potential. Potential real output, y^*, is the smoothed series estimated from y by applying a basic structural model (BSM) and the Kalman filter (see Harvey, 1989).

Figure 3

Chile, actual and target inflation (percent), 1991-2001.

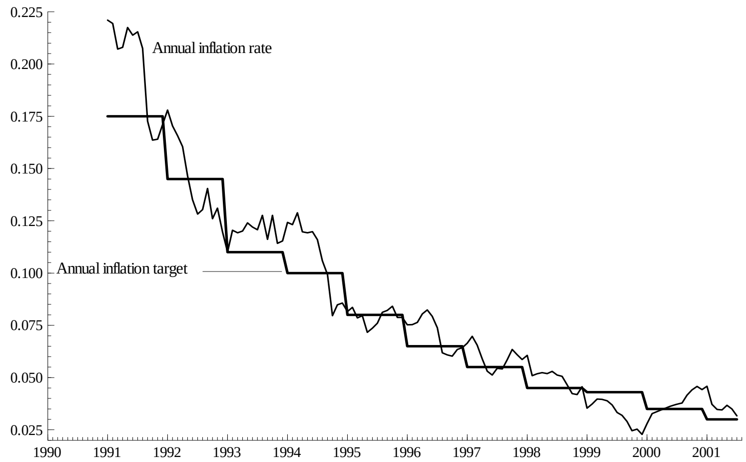
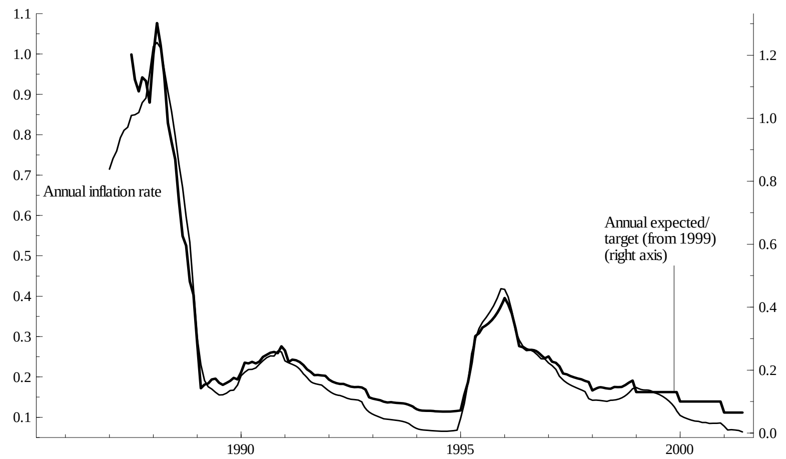


Figure 4

Mexico, Actual and expected/target inflation (percent), 1987-2001.



In estimating the exchange rate gaps the paper employs bivariate (stage three) PPP cointegration tests (see Froot and Rogoff, 1995, and Edwards and Savastano, 1999), and this specification arises after experimenting with alternative specifications. In equation form

$$e_t = \lambda(p_t - p_t^{\,f}) + \zeta_t\,, \tag{7}$$

where e is the nominal exchange rate measured in units of home currency per unit of foreign currency, while p and p^f are the domestic and foreign price levels, respectively. Equation (7) implies that the exchange rate between the currencies of two countries should equal the ratio of their price levels. Thus λ is expected to be around one.

Panel B in Table 2 exhibits the long-run solutions to the ADL PPP cointegrating relations (Hendry, Pagan and Sargan, 1984). The coefficients are statistically significant and economically interpretable, with λ displaying values of 0.92 and 0.88 for Chile and Mexico, respectively. Furthermore, the ADF tests applied to the residuals from these equations reject nonstationarity. So PPP cointegrating relations hold for Chile and Mexico (see Froot and Rogoff, 1995, and Edwards and Savastano, 1999, for wide-ranging empirical evidence on the topic). The paper employs these results in computing the real exchange rate gaps for both economies.

The study employs a simple, textbook, money demand relationship relating real monetary balances to a scale variable and to a measure of the opportunity cost of holding money, like equation (3), in estimating the real money gaps. Panel B in Table 2 displays the long-run solutions to the corresponding ADL equations. All the estimated coefficients are statistically significant and have economically sensible coefficients. It is worth noting that the estimated income elasticities, 1.72 for Chile and 1.38 for Mexico, are similar in magnitude to those obtained by Edwards (1998, Table A2, page 700).

Finally, applying the ADF tests to the money demand equation's residuals leads to rejecting nonstationarity. Consequently, money demand cointegrating relations exist for Chile and Mexico during the periods under investigation. The investigation uses these results in calculating the real money gaps for both economies.

5 Granger Causality Analysis

As highlighted by Stock and Watson (2001), due to the rich lag dynamics involved in applying vector autoregressions (Sims, 1980) to time series data statistics derived from these estimations, Granger causality tests tend to be more informative than the usual regression coefficients. In empirically assessing the model in Section 3, the paper computes multivariate and bivariate block Granger noncausality tests.

The econometric methodology behind the block Granger noncausality test can be illustrated by using an augmented vector autoregression of order n, VAR (n), such as

$$z_t = \alpha_0 + \lambda t + \sum_{i=1}^{n} \Omega_i z_{t-i} + \Psi x_t + \xi_t , \tag{8}$$

where z_t is a $m \times 1$ vector of jointly determined (endogenous) variables, t is a linear time trend, x_t is a $p \times 1$ vector of exogenous variables, and ξ_t is a well-behaved disturbance term. If in equation (8) z_t is divided into two subsets z_{1t} and z_{2t}, which are $m_1 \times 1$ and $m_2 \times 1$ vectors, respectively, and $m = m_1 + m_2$, the following block decomposition can be written:

$$z_{1t} = \alpha_{10} + \lambda_1 t + \sum_{i=1}^{n} \Omega_{i,11} z_{1,t-i} + \sum_{i=1}^{n} \Omega_{i,12} z_{2,t-i} + \Psi_1 x_t + \xi_{1t} , \tag{9}$$

$$z_{2t} = \alpha_{20} + \lambda_2 t + \sum_{i=1}^{n} \Omega_{i,21} z_{1,t-i} + \sum_{i=1}^{n} \Omega_{i,22} z_{2,t-i} + \Psi_2 x_t + \xi_{2t} .$$

The hypothesis that the subset z_{2t} is not Granger causal on z_{1t} is given by $\Omega_{12} = 0$, with $\Omega_{12} = (\Omega_{1,12}, \Omega_{2,12}, \Omega_{3,12}, ..., \Omega_{n,12})$. The log-likelihood ratio statistic that arises from testing $\Omega_{12} = 0$ has $m_1 m_2 n$ degrees of freedom, and is χ^2 asymptotically distributed.

5.1 Chile

The study proceeds by analysing equation (6), that is, the joint impact of $y - y^*$, $\tilde{m} - \tilde{m}^*$, and $e - e^*$ on $\pi - \hat{\pi}$. Table 3 contains the outcome from the multivariate block Granger noncausality VAR tests. They support the joint significance of $y - y^*$, $\tilde{m} - \tilde{m}^*$, and $e - e^*$ on $\pi - \hat{\pi}$. However, the corresponding bivariate tests show that the hypothesis of non-causality can only be rejected for the real money gap. Still, if the real output and inflation VAR is run using 12 lags, even though the Schwarz Bayesian Criterion (SBC) selects a lag length of 2 as reported in Table 3, output is also causal on inflation [47.11(0.00)]. Therefore, for Chile the P-star model provides some rationale for using a real money gap indicator in conducting monetary policy under an inflation targeting regime.

5.2 Mexico

Following the results in Cecchetti, Flores-Lagunes and Krause (2000), showing that monetary policy in Mexico became more effective from 1991, the study adopts 1991-2001 as the baseline estimation period. Additionally, Mishkin and Schmidt-Hebbel's (2001) dating identifies formal inflation targeting adoption in Mexico starting from 1999. Accordingly, the paper analyses three periods: 1991-2001, 1991-1998, and 1999-2001. This should help in determining if the way in which money and other relevant variables affect inflationary developments changed due to inflation targeting adoption.

As for Chile, Mexico's outcomes reject the hypothesis of noncausality from $y-y^*$, $\tilde{m}-\tilde{m}^*$, and $e-e^*$ to $\pi-\hat{\pi}$ for the three time spans under analysis. However, the bivariate results indicate that amongst these indicators only $e-e^*$ is individually significant (at the 7% level) during 1991-2001. Furthermore, $e-e^*$ also seems to contain significant information (at the 1% level) on $\pi-\hat{\pi}$ in the sub-periods 1991-1998 and 1999-2001. Table 3 also shows that $y-y^*$ and $\tilde{m}-\tilde{m}^*$ help in predicting $\pi-\hat{\pi}$, but only during the post-inflation targeting regime ranging from 1999 to 2001.

5.3 Robustness Check

In checking the robustness of the results, the study applies bivariate block Granger noncausality VAR tests using alternative money, output, and exchange rate indicators. The output gap proxy is achieved by passing y through the Hodrick-Prescott filter, HPF (Hodrick and Prescott, 1997). The exercise uses a value of 129,600 for the HPF's smoothness parameter λ, as suggested by Ravn and Uhlig (2002) for analysing monthly data. Also, the study considers two money growth measures, $\Delta\tilde{m}$ and $\Delta\tilde{m}-\Delta y$, and a standard real exchange rate indicator ($REXR$). The analysis defines $\Delta\tilde{m}$ and Δy as annual changes in the log of the real money and real output variables, respectively, whereas $REXR$ is $e-p+p^f$.

Table 4 reports the outcome from estimating the bivariate block Granger noncausality tests. For Chile only the output gap and money growth indicators are significant. Mexico's results endorse the robustness of the previous findings. The alternative exchange rate indicator *REXR* contains statistically significant information on inflation in the three periods under scrutiny. Thus it can be argued that the exchange rate is useful in predicting inflation in Mexico during the last decade. This comes as no surprise, particularly in the time span under scrutiny, encompassing the well-documented Mexican exchange rate crisis of 1995. Additionally, for Mexico the analysis rejects $\Delta\tilde{m}$ as being noncausal on $\pi - \hat{\pi}$ during both 1991.01-1998.12 and 1999.01-2001.06. In contrast, $HPF(y - y^*)$ is so only for the post IT period.

Why, in contrast to Mexico's case, in Chile is the exchange rate not statistically significant? This could reflect the effectiveness of the exchange rate band regime implemented during most of the sample under study – as explained in Section 2 (see also Morandé and Tapia, 2002). In this respect, Reinhart and Rogoff's (2004) novel approach to reclassifying exchange rate regimes identifies Chile's exchange rate regimes during 1989-2001 as a pre-announced crawling band around the US Dollar (June 1, 1989 - January 22, 1992), *de facto* announced crawling band around the US Dollar (January 22, 1992 - June 25,1998), pre-announced crawling band to US Dollar (June 25, 1998 - September 2, 1999), and managed floating (September 2, 1999 - December 2001).

5.4 Sensitivity Analysis

Table 5 provides the results from running alternative specifications. The first row in this Table displays a closed economy model omitting the exchange rate from the Phillips curve/P-star model, i.e. using equation (1) rather than equation (6). For Chile the outcome from this exercise shows that the hypothesis that the output and real money gaps (equation 1 in Table 5) are noncausal on inflation can be rejected. However, Granger's test statistic is larger for the model including the exchange rate, thus supporting equation (6) over equation (1). More interestingly, for Mexico the analysis shows that equation (1) is not significant in all periods except the one using data for 1999.01-2001.06. This somehow reinforces the findings elsewhere in the paper, revealing that the exchange rate is a crucial variable in predicting inflationary developments in Mexico.

Additionally, even though the intention of the paper has been to explore the role of money in inflation targeting economies using a nested P-star/Phillips curve framework, the second row in Table 5 reports the outcome from estimating a more conventional (benchmark) model, namely one including output, the interest rate, and the exchange rate, but not money.[6] For Chile and Mexico the hypothesis that these variables are noncausal on inflation is rejected.

What is more, in both cases the values of the Granger statistic are larger for the model including the interest rate instead of the real money gap. The exception is Mexico during the period 1991-1998. But this is reasonable since, as explained in Section 2, from the beginning of 1995 the Mexican authorities switched to a policy in which the monetary base played a major role. In pondering the last set of findings it is worth recalling that equation (5a) underlines the fact that the real money gap measure in the baseline model actually accounts for the stance of monetary policy via the interest rate gap.

5.5 Future Research

The research effort towards understanding the mechanics behind inflation targeting is rife. Future studies could explore various related pressing issues, particularly in the context of emerging market economies. For instance, better understanding monetary policy behaviour under inflation targeting could follow from applying Martin and Milas's (2004) nonlinear approach. Also, better understanding the link amongst macroeconomic uncertainty and macroeconomic performance in countries adopting inflation targeting, e.g. Fountas, Karanasos, and Kim (2006), is likely to be fruitful. Finally, the institutional changes taking place as a result of introducing a new monetary policy framework like inflation targeting could be suitably approached by testing for structural breaks and moving in such a direction would probably benefit from approaches like that of Bai and Perron (2003a, b).

[6] The interest rate gap $\left(R - R^{*}\right)$ is the difference between the actual interest rate and the series resulting from passing that variable through an HPF - using a value of 129,600 for the HPF's smoothness parameter λ as suggested by Ravn and Uhlig, 2002, for analysing monthly data.

Table 3: Multivariate and bivariate block Granger noncausality VAR tests

Null hypotheses		LR test statistic (probability)							
		Chile		Mexico					
		1992.09-2001.06	Lags	1991.01-2001.06	Lags	1991.01-1998.12	Lags	1999.01-2001.06	Lags
$y-y^*$ $\tilde{m}-\tilde{m}^*$ $e-e^*$	Noncausal on $\pi-\hat{\pi}$	76.90 (0.00)**	10	51.302 (0.047)*	12	98.816 (0.000)**	12	33.483 (0.004)**	5
$y-y^*$	Noncausal on $\pi-\hat{\pi}$	0.77 (0.678)	12	5.752 (0.569)	7	5.664 (0.685)	8	13.763 (0.017)*	5
$\tilde{m}-\tilde{m}^*$	Noncausal on $\pi-\hat{\pi}$	26.64 (0.009)**	12	11.081 (0.522)	12	11.244 (0.423)	11	24.733 (0.010)**	11
$e-e^*$	Noncausal on $\pi-\hat{\pi}$	8.24 (0.766)	12	20.257 (0.062)†	12	47.948 (0.000)**	12	37.452 (0.000)**	12

Notes: The block Granger noncausality statistic is calculated through an LR test, and has a χ^2 distribution. The test statistics are displayed with corresponding probability values inside parentheses. For both countries an unrestricted VAR of an initial lag order of 12 was considered. For Mexico, the multivariate estimations for 1999.01-2001.01 start with 6 lags due to data limitations. The lag lengths were determined through the SBC. †, * and ** denote rejection of the null at the 10%, 5%, and 1% levels, respectively.

Table 4: Bivariate block Granger noncausality VAR tests

Null hypotheses		LR test statistic (probability)							
		Chile		Mexico					
		1992.09-2001.06	Lags	1991.01-2001.06	Lags	1991.01-1998.12	Lags	1999.01-2001.06	Lags
HPF $(y-y^*)$	Noncausal on $\pi-\hat{\pi}$	21.572 (0.010)**	9	10.050 (0.526)	11	15.811 (0.148)	11	21.291 (0.011)*	9
$\Delta\tilde{m}$	Noncausal on $\pi-\hat{\pi}$	19.995 (0.067)†	12	15.807 (0.200)	12	25.757 (0.012)*	12	26.012 (0.004)**	10
$\Delta\tilde{m}-\Delta y$	Noncausal on $\pi-\hat{\pi}$	15.137 (0.234)	12	5.625 (0.897)	11	13.724 (0.249)	11	0.476 (0.490)	1
$REXR$	Noncausal on $\pi-\hat{\pi}$	12.042 (0.442)	12	22.624 (0.031)*	12	60.235 (0.000)**	12	23.539 (0.009)**	10

Notes: The block Granger noncausality statistic is calculated through an LR test, and has a χ^2 distribution. The test statistics are displayed with corresponding probability values inside parentheses. For both countries an unrestricted VAR of an initial lag order of 12 was used. The lag lengths were determined through the SBC. †, * and ** denote rejection of the null at the 10%, 5%, and 1% levels, respectively.

Table 5: Estimates of alternative specifications
Block Granger noncausality VAR tests

Null hypotheses		LR test statistic (probability) Chile 1992.09-2001.06	Lags	Mexico 1991.01-2001.06	Lags	1991.01-1998.12	Lags	1999.01-2001.06	Lags
(i) Equation 1									
$y-y^*$ $\tilde{m}-\tilde{m}^*$	Noncausal on $\pi-\hat{\pi}$	57.03 (0.00)**	10	25.127 (0.196)	10	26.756 (0.142)	10	22.672 (0.031)*	6
(ii) Benchmark model									
$y-y^*$ $R-R^*$ $e-e^*$	Noncausal on $\pi-\hat{\pi}$	81.70 (0.00)**	10	55.50 (0.003)**	10	65.95 (0.00)**	8	42.45 (0.00)**	5

Notes: The block Granger noncausality statistic is calculated through an LR test, and has a χ^2 distribution. The test statistics are displayed with corresponding probability values inside parentheses. For both countries an unrestricted VAR of an initial lag order of 12 was used –excepting the VARs for Mexico during 1999.01-2001.06, for which a lag length of 6 was considered. The lag lengths were determined through the SBC. †, * and ** denote rejection of the null at the 10%, 5%, and 1% levels, respectively.

6 Conclusion

This paper has analysed the role of monetary and open economy indicators in conducting monetary policy in two inflation targeting Latin American economies: Chile and Mexico. A nested P-star/Phillips curve model was developed and estimated. The main findings reveal that for Chile and Mexico real money gap and real output indicators contain significant information on deviations of inflation from target. In contrast, for Mexico diverse real exchange rate indicators are consistently relevant in predicting the inflation gap.

Appendix 1

Table A1: Data definitions and sources

Variables	Chile	Mexico
m	M_2A, real private money, monthly averages in thousands of millions of 1986 Chilean Pesos.	$M_2 = M_1$ + internal financial assets in the hands of residents in thousands of Mexican Pesos.
y	Monthly indicator of economic activity, IMAE, 1986 = 100.	Industrial production index, 1993 = base.
R	PRBC, 90 days in annual percentage points.	Interest rate of 28 days CETES, in annual percentage points.
e	Observed monthly exchange rate, Chilean Pesos per United States Dollar.	Exchange rate, Mexican Pesos per United States Dollar, monthly average.
p^D	Producer price index, June 1992 = 100.	Wholesale price index, excluding oil, 1994 = base.
p	Consumer price index, December 1998 = 100.	Consumer price index, 1994 = base.
p^f	Foreign (US) producer price index, 1982 = base, from the FRED of the Federal Reserve Bank of St Louis (http://www.stls.frb.org/fred/data/ppi/ppiaco).	

Sources
Chile: Unless otherwise indicated, Central Bank of Chile, http://www.bcentral.cl/Indicadores/eindicadores.htm.
Mexico: Unless otherwise indicated, Central Bank of Mexico, http://www.banxico.org.mx/eInfoFinanciera/FSinfoFinanciera.html.

Appendix 2: Econometric Estimation of the Inflation Target for Mexico, 1987-2001

Since Mexico adopted a formal inflation targeting strategy from January 1999, the paper estimates a proxy for an 'implicit' inflation target ($\hat{\pi}$) using a univariate structural time series model (Harvey, 1989). Table A2 provides details on the specification and on related statistics.

Table A2: Structural time series model for inflation Mexico, 1987-2001

Model	
$\hat{\pi}_t$ = Fixed level + cycle + irregular + interventions	
Statistics	
Level (final state vector) = 0.0052	T-ratio = 5.0692 RMSE = 0.00103
Standard error of equation	0.0084
RPEV	0.000071
R_d^2	0.8726
R_s^2	0.8724

Notes: RMSE = root mean square error. RPEV = residuals prediction error variance. R_d^2 and R_s^2 are goodness of fit statistics that compare the results with a random walk plus drift and a random walk with fixed seasonal dummies, respectively.

1. The baseline specification employed for the case at hand is based on a 'local linear trend' model that can be written as

$$\hat{\pi}_t = \lambda_t + \zeta_t,$$
$$\zeta_t \sim NID\left(0, \sigma_\zeta^2\right). \qquad \text{(A1)}$$

In this application, the model search process leads to an equation in which the level (λ) is fixed, i.e. does not contains a stochastic element.

Furthermore,

2. Two lags of the dependent variable are used in fitting (A1).
3. Cycle and irregular components are also added to equation (A1).
4. Interventions affecting λ in (A1) are included in the periods 1987.12, 1988.01, and 1988.02 to account for the impact of 'The Pacto' economic program.
5. All the elements outlined in points 2. to 4. are statistically significant at the 1% level, excepting the intervention for 1988.01 which is so at the 2% level.
6. The fitted model is used to calculate the *filtered* estimate of the trend in (A1) at all points in the sample, using only data available up to the previous period, i.e. $\hat{\pi}_{\langle t|t-1\rangle}$. Therefore, the 'econometric inflation target' intends to reproduce the economic agents' expected central bank inflation objective. Empirically, this is achieved by using the *Kalman filter* in estimating the state (λ).
7. Note that the λ implied by the final state vector, reported in Table A2, is 0.0052, which amounts to an annual inflation rate of 0.0624 (6.24%) at the end of the sample period.

Further details on the above exercise can be obtained from the author upon request.

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