Credibility and Inflation Targeting in Chile

Luis F. Céspedes
Ministry of Finance

Claudio Soto
Central Bank of Chile

After a long history of high and volatile inflation, the Central Bank of Chile began implementing its monetary policy in the early 1990s by announcing yearly targets for inflation. This new framework was the first step toward a full-fledged inflation-targeting setup, although the Central Bank continued to pursue an explicit objective for the exchange rate. One year before the first announced inflation target in 1990, the Central Bank was granted autonomy through a special law that explicitly states that the main objective of monetary policy is to ensure price stability.

Some authors argue that by enhancing the credibility of monetary policy, this new institutional framework contributed fundamentally to lowering inflation to its current level of around 3 percent (Corbo, 1998; Morandé, 2002; Schmidt-Hebbel and Werner, 2002).[1] Credibility affects the dynamics of price adjustments and, in general, the underlying process that determines inflation. Credibility may also determine the tradeoffs faced by the monetary authority when implementing its policy. In particular, a more credible monetary authority should face an improved tradeoff between inflation and output stabilization, making it less costly to carry out a stabilization process.

In this paper, we provide new evidence of changes in the dynamics of the Chilean inflationary process in recent years. Based

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1. Other possible factors in the success of the disinflationary process in Chile include favorable productivity shocks (see, for example, De Gregorio, 2003).

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on a new Keynesian Phillips curve, we show that price rigidity has increased in the last few years, while the degree of indexation in the economy—based on past inflation—has decreased. We also show that the exchange rate pass-through into traded goods inflation has decreased. Our findings are consistent with the idea that the credibility of monetary policy has increased over time. We argue that as monetary policy has become more credible, costly price adjustments have been carried out less frequently, and the prevalence of indexation based on past inflation has decreased.

These changes in the inflationary process, triggered by the enhanced credibility, may have had significant repercussions in the way monetary policy is implemented in Chile. As we show in Céspedes and Soto (2005), when credibility is low, a central bank that is concerned with the sacrifice ratio during a disinflationary process may be less aggressive in implementing its monetary policy in order to avoid large output losses. As it gains credibility, the central bank may fight inflation deviation from target strongly.

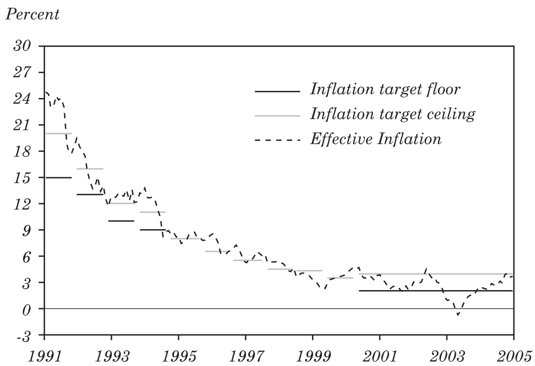
This paper presents evidence of a structural change in the policy rule that characterizes the conduction of the monetary policy. Our evidence is consistent with the idea that monetary policy in Chile has been operating in an environment of improved credibility over the last several years. We show that the monetary policy rule has become more forward-looking in terms of inflation and more aggressive in fighting deviations of inflation from the target.

The paper is organized as follows. The next section briefly describes the different phases of the monetary policy regime in Chile since the Central Bank was granted independence, with a focus on the two phases of the inflation-targeting regime that started in 1991. In section 2, we discuss what credibility means and present some preliminary evidence on the change in the Central Bank's degree of credibility. In section 3, we show how this change in the degree of credibility could have affected the inflationary process in Chile by modifying the degree of price stickiness and the extent of indexation to past inflation. The fourth section presents estimated monetary policy rules for different subsamples, and section 5 concludes.

1. Evolution of the Inflation-Targeting Framework in Chile

In the period 1991–2005, the Chilean economy grew at an average rate of 5.7 percent per year. The inflation rate fell from levels close to 30 percent in early 1991 to stationary levels around 3 percent by the end of the 1990s, and it has fluctuated around that figure since then (see figure 1). During this period, the Central Bank of Chile began to announce an explicit target for inflation, a practice that has been credited with providing a sound guide for conducting monetary policy.[2] Nevertheless, the whole monetary policy framework underwent significant changes along the way. In the first phase (1991–99), the macroeconomic framework included not only inflation targets, but also targets for the current account deficit and a managed exchange rate. In the second phase (2000 to date), the Central Bank implemented a full-fledged inflation-targeting regime in which inflation is the main policy objective, with no other explicit targets.

Figure 1. Effective Inflation and Inflation Targets in Chile, 1991–2005



Source: Authors' calculations based on Central Bank of Chile data.

The Central Bank announced the first target range for inflation in September 1990, after being granted independence in 1989. The initial target, a 15–20 percent target range for December 1991, was defined for consumer price index (CPI) inflation. The initial inflation target was set over a rather short time horizon and represented a strong commitment to reducing inflation. The short-term inflation target reflected the need to generate credibility for the new regime. Previous poor inflation performance had led to a widespread indexation

2. See Corbo (1998) on the effects of inflation targeting on Chilean inflation dynamics in the 1990s.

of the economy and high inflation expectations (see Schmidt-Hebbel and Tapia, 2002). The stabilization process was thus very gradual to avoid high output losses in the context of this widespread indexation and low credibility (Massad, 2003).

In addition to the annual inflation targets, the Central Bank implemented a target band for the exchange rate during the first phase (1991–99). The band was perceived as the appropriate instrument for achieving a normal functioning of the external payments system. The Central Bank also set targets for the current account deficit. To retain the possibility of managing the exchange rate with monetary policy independence, the Bank maintained regulations on the capital account, including a nonremunerated reserve requirement for capital inflows. Exchange rate market interventions were conducted during this period to sustain the exchange rate band. Significant modifications to the exchange rate band were also applied, including adjustments in its width and one-time realignments.

Since 2000, Chile has operated a flexible inflation-targeting regime with the objective of keeping the CPI inflation rate between 2 and 4 percent over a two-year horizon. The move to a full-fledged inflation-targeting framework was seen as the natural step after reaching the steady-state level of inflation and establishing sufficient credibility. It was triggered, in part, by the macroeconomic outcomes of the Asian crisis. GDP growth fell significantly in 1998–99, while the annual inflation rate dropped from 4.6 percent in 1998 to 2.3 percent in 1999. These events led the monetary policy authority to substantially enhance its macroeconomic framework. The main new elements were the adoption of a free-floating exchange rate regime, the deepening of the foreign exchange derivatives market, and the total opening of the capital account (see Morandé, 2002; Céspedes, Ochoa, and Soto, 2005). In addition, transparency increased significantly with the publication of a regular inflation report and the public release of policy meeting minutes.

2. Has Credibility Improved in Chile?

Answering the question of whether credibility has improved in Chile is difficult, as there are no direct measures of credibility. The literature has not yet reached a unanimous definition of the concept of credibility itself, and most of the proposed definitions have an important subjective component. Consequently, practical measures are not readily available.

2.1 What Credibility Means

The academic literature identifies a central bank's credibility as incentive compatibility, precommitment, or strong aversion to inflation. Barro and Gordon (1983) hold that a central bank is credible if it attains higher payoffs (utility) by following its promised actions rather than reneging. Economic agents expect the central bank to do exactly what it promised because that increases its own payoff. This logic is behind Walsh's proposal for an optimal contract design for central bankers that would make the fight against inflation more efficient (Walsh, 1995). By making incentives compatible with the fight against inflation through an explicit contract, this mechanism ensures the credibility of the monetary authority's promises. In the absence of such a contract, credibility could be reached by means of a (credible) precommitment technology, although this could imply having institutional arrangements that are not always available. Finally, credibility has been associated with a strong aversion to inflation. This last definition seems tautological, however: one cannot determine whether the monetary authority really has a strong aversion to inflation unless one can determine how credible the aversion to inflation is. Moreover, having a strong aversion to inflation does not ensure that a central bank has the means to reach its objectives.

Blinder (1999) surveys a group of central bankers and academics to explore why credibility is important and how central banks can enhance their credibility. Although he does not ask participants to define credibility, he asks how close the concept of credibility is to "dedication to price stability." Nearly 90 percent of the respondents answered that the two factors are quite closely related.

The main argument for why the Central Bank of Chile may have gained credibility over the 1990s hinges precisely on this notion of credibility. The new constitutional charter of 1989 not only guaranteed the Central Bank autonomy from the government, but also explicitly established stabilizing the value of the national currency as one of the Bank's two main objectives. In other words, one of the Bank's explicit objectives is price stability.

Autonomy, however, may not be enough to guarantee credibility. Posen (1998) and Fischer (1994), for example, find a positive correlation between the sacrifice ratio and an index of central bank independence. This result seems to suggest that more autonomous central banks are not necessarily more credible. It may be more relevant "to match deeds and words" (Blinder, 1999).

Albagli and Schmidt-Hebbel (2004) undertake an international cross-section comparison of inflation performance for several inflation-targeting central banks. They find that for different measures of inflation deviations from target, Chile ranks among the most accurate inflation targeters in the sample of nineteen countries. At least from an international perspective, the Central Bank of Chile has fulfilled its promises.

The question, then, is whether this behavior has been consistent over time. Figure 2 presents two measures of inflation deviations from target for Chile since 1991. In one of the measures, we consider the center of the target range as the authority's objective, while in the other, we assume that inflation has not deviated from the target if it stays within the target range. In both cases, we report the deviation and the absolute value of the deviation. The figure clearly illustrates that the Central Bank was less accurate in hitting the target in the early 1990s.[3] It improved by the mid-1990s, when the economy entered a phase of sustained growth and low inflation. The Central Bank's performance in terms of hitting the midpoint of the target range worsens after the Russian crisis, when inflation fell more than 2 percent below the target. Over the last several years, however, the Central Bank has generally managed to keep inflation within the target range.

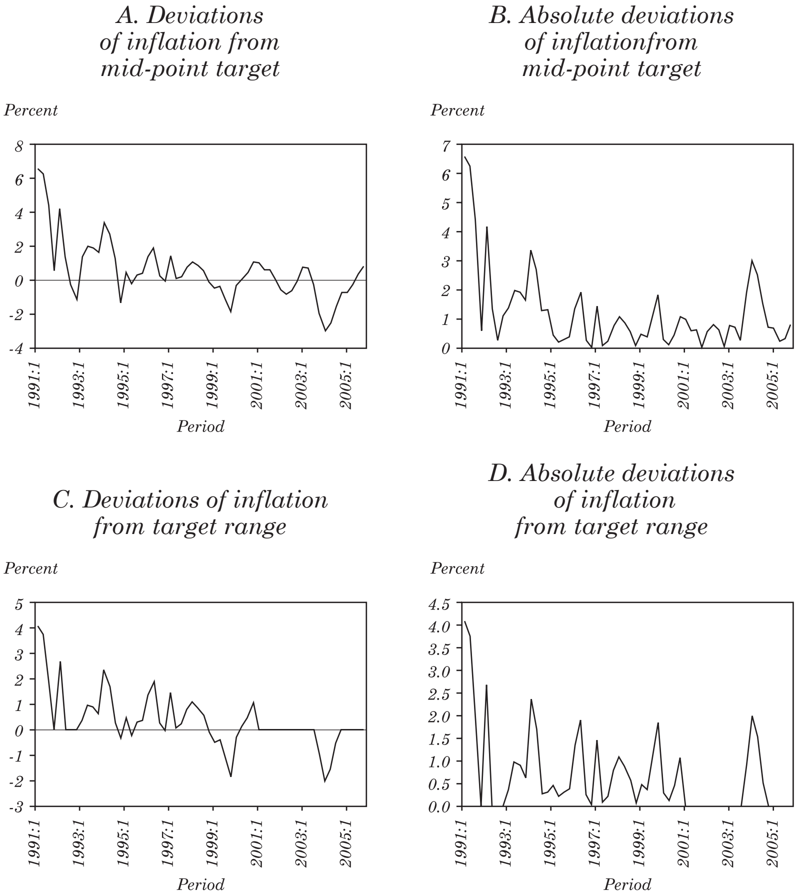
2.2 A Direct Measure of Credibility

One direct measure of credibility is the difference between private inflation expectations and the announced target. This credibility measure is consistent with the measure discussed by Faust and Svensson (2001) and Cukierman and Meltzer (1986). The latter paper defines credibility as the absolute value of the difference between the policymaker's plans and the private sector's beliefs about those plans. The smaller this difference, the higher the credibility of planned monetary policy.

To construct this measure of credibility, we consider two measures of expected inflation. The first is based on nominal-real market interest rate differentials. The second is a measure of expectation taken from survey data. Figure 3 depicts the evolution of the difference between expected inflation constructed using market interest rates and the

3. Some authors normalize the deviation of inflation from target by the inflation target level.

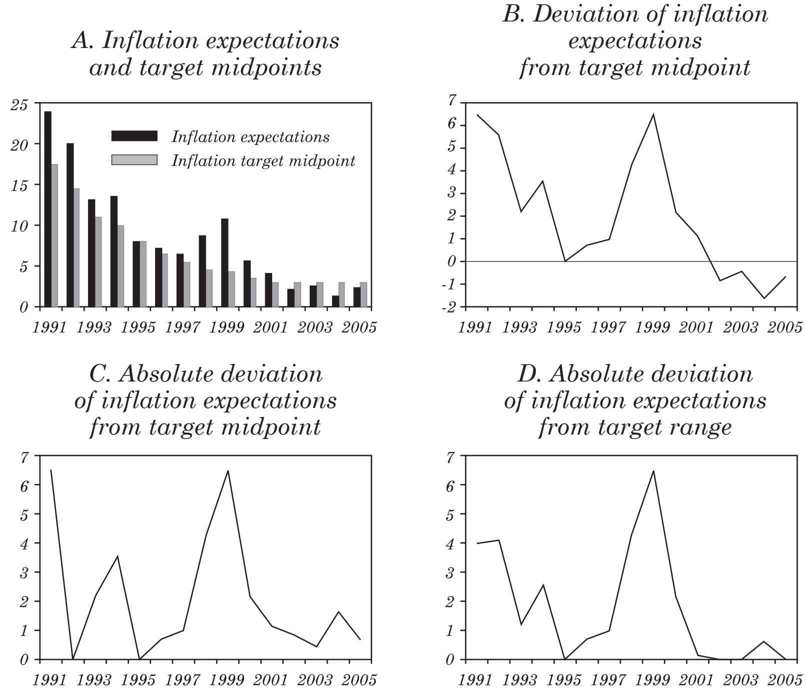
Figure 2. Inflation Deviation from Target in Chile, 1991–2005



Source: Authors' calculations based on Central Bank of Chile data.

announced inflation target. The figure shows that credibility increased in the early 1990s. However, expected inflation was way above the announced target during the Asian crisis, signaling a possible decrease in credibility. Since the end of the 1990s, expected inflation has been lower than the midpoint of the announced target, but inside the target range (except for 2004). In other words, the credibility of the announced target seems to have increased over the last several years.

Figure 3. Inflation Target Credibility in Chile, 1991–2005: Expected Inflation Measured as Nominal-Real Interest Rate Differentials

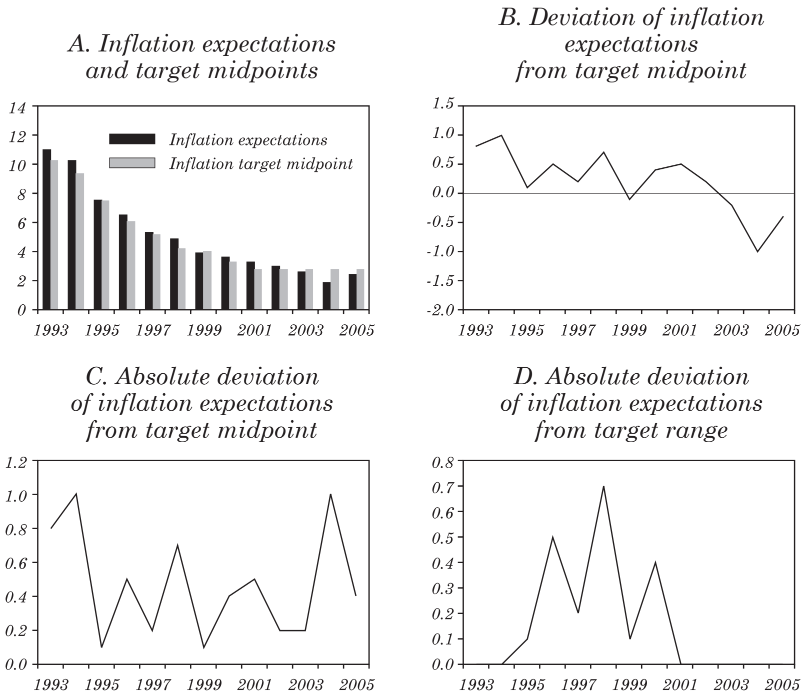


Source: Authors' own calculations.

This credibility measure, based on market expectations extracted from interest rate differentials, is subject to two important criticisms. First, the level and volatility of inflation were much higher in the early 1990s than at the end of the decade. That implies that the expected inflationary premium implicit in the nominal interest rate was also higher in the early 1990s. Consequently, the difference between expected inflation and the announced target may not have decreased over time. Second, although nominal instruments have existed for a long time, the market turnover was pretty low until 2001, when monetary policy was nominalized.[4] In other words, the market for nominal instruments was not very liquid until recently. Therefore,

4. In addition to defining an overnight nominal interest rate as its monetary policy instrument, the Central Bank introduced a set of nominal instruments in August 2001 to help set benchmarks for that market.

Figure 4. Inflation Target Credibility in Chile, 1991–2005: Expected Inflation from Consensus Forecast



Source: Authors' calculations based on Central Bank of Chile and Consensus Forecast data.

prices do not necessarily reflect market expectations on inflation.

Figure 4 charts the difference between private inflation expectations and the announced target using survey data. Survey data on analyst expectations regarding future inflation, drawn from Consensus Forecast, are available for Chile only since 1993. When using this measure of expectations, we observe that expected inflation has generally been much closer to the midpoint of the inflation target. Until 2002, expected inflation was above the midpoint of the target (except in 1999). Since that year, expected inflation has been below the midpoint of the target, but always inside the target range.[5]

5. To facilitate comparison with previous periods, we consider one-year-ahead expected inflation. However, following the introduction of a steady-state inflation target in 2001, the monetary authority explicitly stated that its goal is to keep inflation within the target range in a horizon of twelve to twenty-four months.

3. Frequency of Price Adjustment, Credibility, and Indexation

The preliminary evidence presented in the previous section hints that there may, in fact, have been gains in monetary policy credibility during the disinflationary process carried out by the Central Bank of Chile in the 1990s. In this section, we perform a more detailed analysis of how this change in the macroeconomic environment may have changed the inflationary dynamics. We start by estimating a new Keynesian Phillips curve for Chile and performing some stability tests on some key semistructural parameters. We then evaluate whether the degree of import price rigidity may have changed in a manner consistent with the hypothesis that the credibility of the Chilean monetary policy regime has risen.[6]

3.1 The Phillips Curve and the Persistence of Inflation

We estimate a new Keynesian Phillips curve (NKPC) in which the parameters have a semistructural interpretation. We emphasize that the parameters are not completely structural. In fact, changes in the macroeconomic environment seem to have had an impact on the value of some of the deep parameters that characterize the NKPC (see also Rudd and Whelan, 2003, 2005). The fraction of firms optimally adjusting prices each period decreased at the end of the 1990s, and the weight given to the announced inflation target by the firms that do not optimally adjust prices in a particular period has increased over the last few years. We argue that this evidence is consistent with the view that the Central Bank's higher credibility in its commitment to price stability (low inflation) changed the way in which firms set prices.

The theoretical NKPC comes from a standard model of monopolistic competitive firms that adjust prices infrequently.[7] We extend the basic model in Galí and Gertler (1999) to allow for passive price adjustment for all firms that do not optimize each period. Firms thus change prices every period, either optimally or by following this passive adjustment. The logic behind this dual pricing mechanism is that the menu costs for

6. García and Restrepo (2001) and Schimidt-Hebbel and Werner (2002) document the existence of a change in the pass-through coefficient for the Chilean economy using different setups.

7. See Galí and Gertler (1999) for a full derivation of the NKPC.

price adjustments are not related to the cost of changing prices itself, but rather to all the costs that an optimizing procedure involves (such as collecting information on market conditions, costs, and forecasting). Given those menu costs, firms will go through an optimization process only infrequently (see Christiano, Eichenbaum, and Evans, 2005).

This framework is appropriate for describing Chilean inflationary dynamics for two reasons. First, the notion of a passive adjustment is very plausible in a high-inflation environment—as in Chile in the early 1990s—where it is difficult to extract the information content of prices. Second, the passive adjustment setting allows us to formally introduce an indexation mechanism into the Phillips curve. In the case of Chile, indexation has been a prevalent phenomenon for many years (see Landerretche, Lefort, and Valdés, 2002).

The passive adjustment mechanism for firms that do not optimally adjust their prices consists in upgrading prices in proportion to a geometric average of past inflation and the announced target for inflation. The passive adjustment rule can be expressed as follows:

$$\Gamma_{t,i}=\prod_{j=1}^{i}\left(1+\pi_{t+j-1}\right)^{\kappa}\left(1+\pi_{t+j}^{tar}\right)^{1-\kappa}, \tag{1}$$

where $\Gamma_{t,i}$ is the percentage change in price of a firm that is not able to optimally adjust between t and $t+i$, π_{t+j-1} is the inflation rate in $t+j-1$, and π_{t+j}^{tar} is the inflation target set by the authority for the period $t+j$. This updating rule implies that whenever firms do not receive a signal, they adjust their prices by a geometric average of the inflation target set by the authority and past inflation. Parameter κ is a measure of the degree of persistency of inflation and can be associated with the credibility of the target set by the authority. The larger this parameter, the larger the weight given to past inflation and the lower the weight given to the inflation target—and thus the lower the credibility of the announcement. The Phillips curve with the Calvo model and this updating rule is given by

$$\hat{\pi}_t=\lambda\xi mc_t+\gamma_f E_t\hat{\pi}_{t+1}+\gamma_b\hat{\pi}_{t-1}+\zeta_t, \tag{2}$$

where $\hat{\pi}_t=\pi_t-\pi_t^{tar}$ corresponds to the deviation of inflation from the target, mc_t represents marginal costs, $\lambda(\kappa, \theta, \beta) = (1-\theta)(1-\theta\beta)/[\theta(1+\kappa\beta)]$, $\gamma_f(\kappa, \theta, \beta) = \beta/(1+\kappa\beta)$, and $\gamma_b(\kappa, \theta, \beta) = \kappa/(1+\kappa\beta)$. Parameter β is the subjective discount factor and ξ defines whether capital is mobile across firms ($\xi=1$) or firm specific ($\xi\neq 1$). The term ζ_t is a function of changes in the inflation target.[8] Notice that parameter κ defines the backward-looking component of the Phillips curve. In other words, this parameter is also a measure of the degree of indexation in the economy.

Parameter θ is the probability that a firm will keep the current price until the next period. It also corresponds to the share of firms that do not optimally adjust prices in a particular period. We assume that this parameter captures, to some extent, the credibility of monetary policy. In the standard Calvo model, the probability of adjusting prices—or the share of firms optimally setting prices each period—has no economic interpretation other than as a measure of the degree of price stickiness. In our case, we assume that firms will adjust prices more often (that is, a larger share of firms will optimally adjust prices each period) when future inflation is more uncertain and the central bank's commitment regarding inflation is perceived to be weak.

Parameters θ, β, and κ are estimated using a generalized method of moments (GMM) estimation with quarterly data from 1991:1 to 2005:4, based on the following orthogonality condition derived from equation (2):

$$E_t\left\{\begin{bmatrix}\theta(1+\kappa\beta)\hat{\pi}_t-(1-\theta)(1-\theta\beta)\xi mc_t\\ -\theta\beta\hat{\pi}_{t+1}-\theta\kappa\hat{\pi}_{t-1}+\theta(1+\kappa\beta)\zeta_t\end{bmatrix}\mathbf{z}_t\right\}=0\,, \tag{3}$$

where $\mathbf{z}_t$ is a vector of instruments that includes three lags of the inflation deviation from target, the real marginal cost deviation from trend, and the output gap (from t – 3 to t – 5).[9]

Benchmark results for the estimated hybrid model are presented in table 1. We give the estimated values of parameters θ, β, and λ under four different specification for the marginal cost—derived from a Cobb-Douglas technology, a technology with labor hoarding, a CES technology and a CES technology for an open economy—, and assuming alternatively that capital is freely mobile and firm specific. We also report the estimated value of parameter κ, which measures the degree to which firms index their prices to past inflation.

8. In the steady state, with a constant inflation target, $\zeta_t = 0$.

9. To check the relevance of the instrument set used in our regressions, we test the null hypothesis that the coefficients on all the instruments are jointly zero in the first stage of the estimation. The F statistic, the associated p value, and the adjusted R^2 from the first stage of these regressions allowed us to reject the null hypothesis that the instruments are jointly irrelevant. The adjusted R^2 is generally over 0.5.

Table 1. Benchmark Phillips Curve Estimation[a]

	Model											
	Cobb-Douglas		*Overhead labor*		*CES, σ=0.5*		*CES, σ=1.5*		*CES for an open economy, σ=0.5*		*CES for an open economy, σ=1.5*	
Parameter	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$
θ	0.910	0.679	0.927	0.721	0.916	0.659	0.930	0.809	0.921	0.669	0.932	0.802
	(0.05)	(0.04)	(0.04)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
β	0.973	0.973	0.966	0.966	0.973	0.973	0.970	0.970	0.969	0.969	0.951	0.951
	(0.16)	(0.16)	(0.13)	(0.13)	(0.15)	(0.15)	(0.15)	(0.15)	(0.14)	(0.14)	(0.12)	(0.12)
κ	0.776	0.776	0.730	0.730	0.732	0.732	0.760	0.760	0.727	0.727	0.737	0.737
	(0.11)	(0.11)	(0.09)	(0.09)	(0.08)	(0.08)	(0.11)	(0.11)	(0.10)	(0.10)	(0.09)	(0.09)
λ	0.006	0.090	0.005	0.068	0.005	0.108	0.004	0.029	0.005	0.102	0.004	0.034
	(0.00)	(0.03)	(0.00)	(0.03)	(0.00)	(0.06)	(0.00)	(0.02)	(0.00)	(0.06)	(0.00)	(0.02)
γ_f	0.554	0.554	0.567	0.567	0.568	0.568	0.558	0.558	0.568	0.568	0.559	0.559
	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
γ_b	0.442	0.442	0.428	0.428	0.427	0.427	0.437	0.437	0.426	0.426	0.433	0.433
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
τ_1	0.430	0.430	0.414	0.414	0.416	0.416	0.424	0.424	0.413	0.413	0.412	0.412
	(0.07)	(0.07)	(0.06)	(0.06)	(0.06)	(0.06)	(0.07)	(0.07)	(0.07)	(0.07)	(0.06)	(0.06)

Table 1. (continued)

	Model											
	Cobb-Douglas		*Overhead labor*		*CES, σ=0.5*		*CES, σ=1.5*		*CES for an open economy, σ=0.5*		*CES for an open economy, σ=1.5*	
Parameter	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$	$\xi=1$	$\xi\neq1$
τ_2	-0.442	-0.442	-0.428	-0.428	-0.427	-0.427	-0.437	-0.437	-0.426	-0.426	-0.433	-0.433
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
D	11.169	3.121	13.674	3.591	12.022	2.937	14.372	5.226	12.693	3.020	14.632	5.056
	(6.75)	(0.38)	(6.86)	(0.44)	(6.39)	(0.35)	(8.98)	(0.10)	(5.93)	(0.33)	(6.70)	(0.90)
J statistic	4.505	4.505	4.060	4.060	4.227	4.228	4.063	4.063	4.487	4.487	4.464	4.464
P value	(0.98)	(0.98)	(0.99)	(0.99)	(0.98)	(0.98)	(0.99)	(0.99)	(0.98)	(0.98)	(0.98)	(0.98)

Source: Authors' calculations.

a. Standard errors based on a Newey-West covariance matrix robust to serial correlation up to twelve lags are in parentheses. Parameter ξ defines whether capital is mobile across firms (ξ=1) or firm specific ($\xi\neq$1). Row D reports the estimated duration of price stickiness. The *J* statistic is the Hansen test of overidentifying restrictions (we report the *p* values in parentheses). The set of instruments includes five lags of inflation deviation from its target and detrended output, three lags of real marginal costs, and five lags of detrended terms of trade.

The estimated share of the backward-looking component, γ_b, is statistically significant in all specifications, and it is about 0.45. This figure is slightly smaller than Agénor and Bayraktar (2003) find for Chile in their study of the inflation dynamics in middle-income countries. These authors estimate a nonstructural Phillips curve that includes both backward- and forward-looking components; they find that the backward-looking component is about 0.52. Unlike our case, Agénor and Bayraktar use several lags of the output gap (up to three for Chile) as the driving force for inflation.

The estimated values for parameter θ under this hybrid specification of the NKPC lie in the range of 0.8 to 0.9. This implies that firms adjust prices every seven quarters, on average. In the case of discount factor, β, our empirical results indicate that the estimated value for this parameter is somewhat low. Our specification for marginal cost with firm-specific capital delivers a lower value for θ, implying shorter average duration of price stickiness, as in Galí, Gertler, and López-Salido (2001). Finally, the overidentifying restrictions are satisfied for all cases. Table 1 also presents the results for the specification for marginal cost that assumes firm-specific capital (the results assuming capital mobility are similar).[10]

Parameter κ is consistently estimated in the range of 0.75 to 0.82, implying that within the sample period, firms gave more weight to past inflation than to inflation targets when passively adjusting their prices. These figures suggest a much stronger role for the backward-looking component of inflation in the Chilean case than is estimated by Galí and Gertler (1999) and Galí, Gertler, and López-Salido (2001) for the euro area. They are quite similar, however, to the estimates for the United States in those same two papers.

We turn now to the issue of parameter stability. [11] Following Céspedes, Ochoa, and Soto (2005), we consider a predictive test for structural change with an unknown breakpoint developed by Ghysels and Hall (1990), Ghysels, Guay, and Hall (1997), and Guay (2003). This test consists in estimating the parameter vector for a first subsample and then evaluating the moment conditions for the second subsample

10. We perform a robustness exercise to address the importance of the backward-looking component of the Phillips curve, in which we incorporate additional lags of inflation to the hybrid model. Additional lags of inflation turn out to be insignificant, as does the sum of the three additional lags of inflation.

11. This issue has not been formally analyzed in the literature. Jondeau and Le Bihan (2005) test the stability of their Phillips curve estimates, but they only examine the reduced-form (linear) parameters and use Wald-type tests, which have some important drawbacks (as they point out).

at these parameter values. The subsample is then increased by one observation at the time. Each time, the PR statistic (or predictive test) is constructed.[12]

Table 2 presents the estimated predictive tests (supremum: supPR; average: avgPR; and exponential: expPR), along with the date for which the largest PR test is obtained. The PR-type tests can be divided into a test of structural change for the vector of parameters and a test of the stability of the overidentifying restrictions (Sowell, 1996). We report the PR1-type statistic, which tests both parameter and overidentifying restriction stability, and the PR2-type tests for parameter variations only. The results show that we cannot reject the existence of a breakpoint for the four marginal cost specifications. Furthermore, the PR1 and PR2 tests consistently estimate the date of the breakpoint around the first quarter of 2001, which is close to the date that the inflation target reached its stationary annual value of 3 percent. We also report the estimated values of the parameter κ and the duration of price stickiness before the breakpoint date, $\kappa 1$ and D1 respectively in the table. In most cases, the first subsample value of κ is larger than the estimated parameter using the whole sample, suggesting that after the breakpoint in 2001 firms have given a larger weight to inflation targets when updating their prices. Finally, we find that the estimated duration of price stickiness is smaller before the breakpoint, as expected.

These results are consistent with the findings of Hutchison and Walsh (1998), who provide evidence of a possible shift in the output-inflation tradeoff in New Zealand after the implementation of the 1989 Reserve Bank Act. In their case, both direct tests of parameter instability and inference from inflation forecast errors indicate that the 1989 Act may have altered some fundamental economic relationships by changing the degree of central bank independence and the Reserve Bank's commitment to price stability.

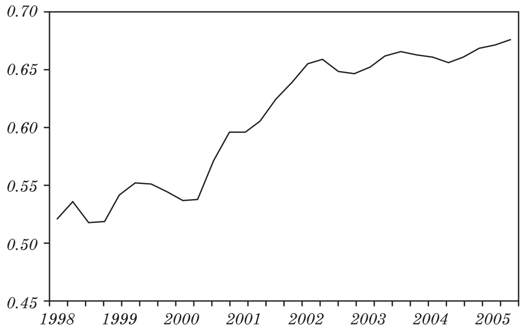
12. For our purposes, this approach has several advantages over alternative approaches like the Wald-type tests proposed by Andrews (1993) and Andrews and Ploberger (1994). First, we only use first subsample estimates of the parameters, which allows us to test for the presence of a break even when the second subsample contains few observations and parameter estimates are not feasible. A common drawback of Wald-type tests is that they cannot be applied to detect structural instability at the end of the sample. Second, we do not set a priori orthogonality conditions equal to zero in the second subsample, so we avoid rejecting stability when the parameters are, in fact, stable but there are certain types of misspecification (for example, omitted variables). A brief description of the test can be found in Céspedes, Ochoa, and Soto (2005).

Table 2. PR Test for Structural Breaks in the Hybrid Phillips Curve[a]

Model	*Sup PR1*	*Avg PR1*	*Exp PR1*	*Sup PR2*	*Avg PR2*	*Exp PR2*
Cobb-Douglas						
PR test statistic	152.63*	32.56*	73.12*	85.86*	17.29*	39.63*
Estimated breakpoint	2001:3					
κ1	0.921					
D1	2.41					
Overhead labor						
PR test statistic	117.84*	15.48**	55.62*	81.52*	9.97	37.46*
Estimated breakpoint	2001:4					
κ1	0.868					
D1	3.18					
CES						
PR test	79.40*	20.45**	37.10*	46.98*	10.01	20.21*
Estimated breakpoint	2001:4					
κ1	0.765					
D1	2.46					
CES for an open economy						
PR test	141.42*	34.50*	67.49*	89.77*	18.17**	41.67*
Estimated breakpoint	2001:4					
κ1	0.768					
D1	2.59					

Source: Authors' calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
a. The table reports predictive tests for the null hypothesis of structural stability, along with the estimated breakpoint date, the value of parameter κ, and the duration of price stickiness before the breakpoint date (κ1 and D1 respectively). The tests are estimated using Monte Carlo simulations with 10,000 replications.

Figure 5. Price Rigidity: Phillips Curve (θ)



Source: Authors' calculations.

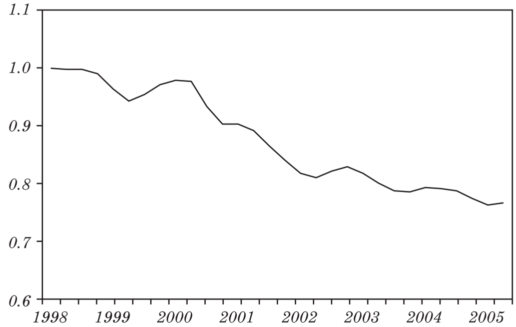
To further assess the potential direction of the structural change in the inflationary process in the Chilean economy, we run a set of regressions for the NKPC starting with the period 1991:1–1997:4 and then adding one more observation at a time. For these calculations, we use real marginal costs computed using a production function with overhead labor and firm-specific capital. The results indicate that price rigidity, captured by the parameter θ, has increased in recent years (see figure 5). Specifically, the estimated duration of price rigidity went from two quarters in the period previous to 2000 to around three quarters for the whole period. We also find that the parameter that measures the degree by which firms adjust their current prices based on previous inflation (κ) has decreased over time (see figure 6).

The evidence presented here is consistent with the view that some parameters are policy dependent (Rudd and Whelan, 2005). The increase in the price rigidity parameter and the decrease in the degree of indexation to past inflation support the hypothesis that monetary policy in Chile has become more credible. These results are also in line with the empirical and theoretical view that a more credible monetary policy is negatively correlated with inflation persistence (see Taylor, 2000; Sargent, 1999).

3.2 Changes in the Exchange Rate Pass-Through

Broad empirical literature shows that the pass-through from the exchange rate to prices is not full in the short run. Import prices in various countries deviate from the law of one price (Campa and

Figure 6. Weight of Past Inflation in the Passive Price Adjustment Rule (κ)



Source: Authors' calculations.

Goldberg, 2002), and the response of CPI inflation to changes in the exchange rate is often less than one to one in the short run (Borensztein and De Gregorio, 1999; Goldfajn and Werlang, 2000). In this subsection, we provide new evidence on the evolution of the pass-through in Chile in the last fifteen years using the new Keynesian Phillips curve as our analytical framework (see Monacelli, 2003).

The pass-through from the exchange rate to prices may be imperfect in the short run because of structural features of the market, as in Dornbusch (1977), or because of nominal rigidities. Given pricing-to-market behavior (Betts and Devereux, 1996), if a fraction of import prices are sticky in the domestic currency, then changes in the exchange rate will not be completely passed on to prices. In other words, the pass-through from the exchange rate to import prices will be incomplete. Therefore, changes in price rigidity will also imply changes in the exchange rate pass-through.

Since we are interested in the pass-through as a macroeconomic phenomena, we estimate the effect of changes in the nominal exchange on imported goods inflation based on a sticky-price model similar to the one used to derive the Phillips curve in the previous section. Following Monacelli (2003), we assume that domestic retail firms buy different varieties of imported goods in the international market and later sell them domestically. We assume that each retail firm has monopoly power over a particular imported variety, and it adjusts its prices infrequently with probability of $1 - \theta_M$ each period. Coefficient θ_M is a measure of nominal price rigidity in the import sector; it determines the degree of imperfect pass-though. The larger this coefficient, the less frequently prices are changed and the lower is the pass-through from nominal exchange rate movements into imported price inflation (see the appendix).

A particular retail firm selling variety z_M that is able to adjust its price in period t chooses a new price, $P_{M,t}^{new}(z_M)$, to maximize its expected profits:

$$E_t \sum_{i=0}^{\infty} \theta_M^i \Lambda_{t,i} \left[\frac{P_{M,t}^{new}(z_M) - S_{t+i} P_{M,t+i}^{*}(z_M)}{P_{t+i}} C_{M,t+i}(z_M) \right],$$

subject to the domestic demand for that variety, given by $C_{M,t+i}(z_M)$. Variable S_t is the nominal exchange rate, and $P_{M,t}^{*}(z_M)$ is the international price of variety z_M expressed in foreign currency. From the first-order condition for this problem, we can establish the following relation for imported goods inflation, $\pi_{M,t}$:

$$\pi_{M,t} = \frac{(1-\theta_M)(1-\theta_M\beta)}{\theta_M}(s_t + p_{M,t}^{*} - p_{M,t}) + \beta E_t \pi_{M,t+1}. \tag{4}$$

where $s_t + p_{M,t}^{*}$ is the logarithm of the price of imported goods abroad expressed in domestic currency, and $p_{M,t}$ is the domestic price of imported goods in the domestic market.

We estimate equation (4) by GMM, using an imports price index (the *índice de valor unitario de importaciones*, or *IVUM*) as a proxy for the price of imported goods in foreign currency. The domestic price of imported goods is proxied by the price index corresponding to tradable goods included in the CPI basket. Our measure of inflation thus corresponds to the quarterly change in the tradable goods price index. The specification in equation (4) assumes that the distribution service provided by retail firms does not require the use of any inputs.[13] Our results indicate that the relative price of imports abroad is a strong driver of the tradables inflation rate, and the forward-looking component is also very relevant. Moreover, the coefficient associated with price rigidity, θ_M, is positive and significant (see table 3). This value yields an estimated duration of price stickiness of around five quarters.

13. Results under alternative specifications, in which labor is required to distribute the imported goods, are similar to the ones reported here.

Table 3. Tradable Inflation Phillips Curve

Parameter	*(1)*	*(2)*	*(3)*
β	0.972	0.907	0.919
	(0.01)*	(0.00)*	(0.01)*
θ_M	0.816	0.648	0.846
	(0.05)**	(0.00)*	(0.04)**
λ	0.047	0.223	0.040
	(0.03)**	(0.01)*	(0.02)**
Sample	1991:1–2005:4	1991:1–2000:4	1995:1–2005:4
J statistic	4.870	3.229	3.533
P value	(0.99)	(1.00)	(1.00)
SupPR test	1287.22	276.01	2396.41
Estimated breakpoint	2000:1	1999:4	2003:1

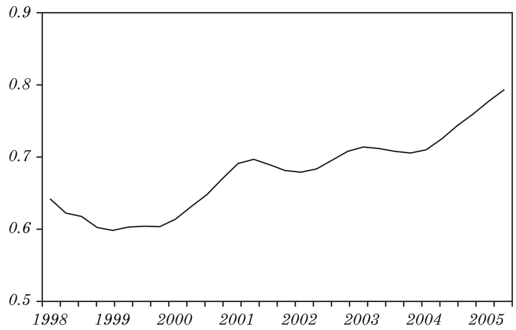
Source: Authors' calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
Standard errors in parentheses.

To test for potential structural breaks, we consider the predictive test for structural change discussed in the previous section. The supPR test points to a structural break around the first quarter of 2001. Again, to assess the scope of this change, we estimate a set of regressions starting with the period 1991:1–1997:4 and then adding one observation at a time. The results from this exercise indicate that the average price rigidity in the 1990s was close to 2.5 quarters, half the value for the whole sample (see figure 7). Also in table 3 we report results obtained when splitting the sample. Again, we find evidence that the frequency of imports price adjustments has decreased over time.[14]

These results indicate that the short-run pass-through was lower after 1999, during the Chilean economy's low-inflation period, than in the previous period. Taylor (2000) argues that the pass-through coefficient should be lower in low-inflation environments, as low inflation may be associated with less persistent changes in costs. From the perspective of our theoretical framework, we argue the higher credibility of monetary policy, interpreted as a credible commitment

14. Although the resulting number of observations is low, we divided the sample into two parts: from 1991:1 to 1997:4 and from 1998:1 to 2005:4. We find results consistent with the hypothesis of a change in price rigidity. In particular, the average duration of price stickiness in the second period is almost twice as large as in the first period.

Figure 7. Price Rigidity: Tradables Phillips Curve (θ_M)



Source: Authors' calculations.

to low and stable inflation rate around 3 percent, has reduced the frequency of price changes.[15]

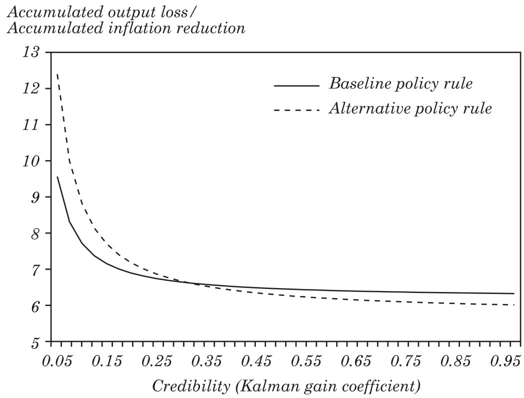
One special feature of the change in the pass-through coefficient in Chile is that it seems to have occurred not at the moment in which inflation rates reached low values (around 1998), but sometime later. We interpret this result as indicating that only when monetary policy was credible did the pass-through coefficient change. This evidence is related to the findings of Bailliu and Fujii (2004), who report that the exchange rate pass-through coefficients for a group of industrialized economies declined following the inflation stabilization in the early 1990s, but they did not decline following a similar episode in the early 1980s. One interpretation for this result is that the disinflationary process of the 1990s was perceived as more credible than the previous one.

4. Credibility and Monetary Policy

The previous section showed that the frequency of price adjustments, the degree of indexation to past inflation, and the exchange rate pass-through have decreased over time, especially after 2000. We interpreted this evidence as support for the view that monetary policy has become more credible over the last several years.

15. Céspedes and Valdés (2006) provide evidence that countries with more independent central banks, which are associated with a more credible monetary policy, have lower pass-through coefficients.

Figure 8. Sacrifice Ratio and Credibility



Source: Céspedes and Soto (2005).

A broad literature shows that when the monetary authority is not fully credible, its tradeoff between output and inflation stabilization is worse than when credibility is larger. In fact, in the simplest version of the new Keynesian Phillips curve, a fully credible stabilization process can be carried out without any output losses (Ball, 1995).

In Céspedes and Soto (2005), we present a model for comparing the sacrifice ratio implied by different monetary policy rules under alternative assumptions regarding the credibility of a disinflationary shock. We show that if the central bank is strict in fulfilling its inflation target, then inflation stabilization may generate significant output losses and large sacrifice ratios for low credibility levels. Figure 8 depicts the sacrifice ratio—that is, accumulated output losses relative to the accumulated reduction in inflation—for a disinflationary shock as a function of credibility, measured by a Kalman gain coefficient, under two alternative policy rules: a baseline Taylor rule and an alternative policy rule with a strong feedback response of the interest rate to deviation of inflation from the target.[16] The figure clearly illustrates

16. The model in Céspedes and Soto (2005) follows Erceg and Levin's (2003) approach to model imperfect credibility, assuming that private agents must solve a signal extraction problem in order to sort out whether a shock to the inflation target is permanent or transitory. They use the Kalman filter to solve for the signal extraction problem. The larger the Kalman gain coefficient, the faster they understand the nature of the shock to the target (that is, the more credible is a permanent shock to the inflation target).

how the sacrifice ratio falls as the credibility of the inflation target rises (that is, with a higher Kalman gain coefficient). The tougher rule has a higher sacrifice ratio than the baseline Taylor rule when credibility is low. As credibility increases, however, the sacrifice ratio of the alternative rule drops below the sacrifice ratio of the baseline policy rule. In other words, as credibility increases, a central bank can reduce inflation faster without having to sacrifice too much output.

In what follows, we investigate whether the way monetary policy is conducted in Chile has changed in line with the previous analysis. We estimate monetary policy interest rate rules for two periods that roughly coincide with the two phases described earlier: the first quarter of 1991 to the fourth quarter of 1997; and the first quarter of 1998 to the fourth quarter of 2005. As mentioned, the first phase was characterized by short-run horizons for the inflation targets, a managed exchange rate, and a target for the current account. We capture these features by assuming that the real interest rate (r_t) set by the Central Bank during the first period of inflation targeting was a function of current inflation, the output gap, and the current difference between the real exchange rate and its equilibrium level. We estimate the following relation:

$$\begin{aligned} r_t = {} & (1-\rho_0)r + \rho_0 r_{t-1} + (1-\rho_0)\omega_{\pi,0}(\pi_t - \pi_t^{tar}) \\ & + (1-\rho_0)\omega_{y,0} g_{t-1} + (1-\rho_0)\omega_{e,0} q_t^{mis} + u_t, \end{aligned} \tag{5}$$

where g_{t-1} corresponds to the output gap in $t - 1$, q_t^{mis} is the real exchange rate misalignment, and u_t is a linear combination of prediction errors and policy innovations.[17] Since the term u_t could be correlated with actual values of inflation and the output gap, we estimate this relation using GMM, as suggested by Clarida, Galí, and Gertler (2000).

The results indicate that the specification of the monetary policy rule in equation (5) is a good description of monetary policy actions for the period 1991–97 (see table 4). The estimated coefficients have the

17. The output gap measure corresponds to the difference between effective output and a trend measure for output. This trend measure is obtained using the Hodrick-Prescott (HP) filter. The real exchange rate misalignment corresponds to the difference between the effective real exchange rate and the equilibrium real exchange rate. The equilibrium real exchange rate is computed by applying the HP filter to the effective real exchange rate. A positive value for q_t^{mis} implies that the real exchange rate is undervalued relative to its equilibrium level.

expected sign and are significant in all cases. In addition to responding to deviations of inflation from the target, the Central Bank of Chile responded to deviations of output from its trend value and to the exchange rate. As discussed previously, the Central Bank also targeted the current account during this period.[18]

Table 4. Monetary Policy Rules: Chile 1991–2005[a]

Parameter	*(1)*	*(2)*
ρ_0, ρ_1	0.63 (0.06)*	0.63 (0.04)*
$\omega_{\pi,0}$, $\omega_{\pi,1}$	0.35 (0.17)***	1.84 (0.28)*
$\omega_{y,0}$, $\omega_{y,1}$	0.35 (0.04)*	0.65 (0.24)**
$\omega_{q,0}$	0.15 (0.05)**	
Period	1991:1–1997:4	1998:1–2005:4
R^2	0.81	0.94
J statistic	1.92	3.52
P value	(0.98)	(0.99)

Source: Authors' calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
*** Statistically significant at the 10 percent level.
a. The dependent variable is the monetary policy real interest rate. The regressions are estimated using GMM. Regression 1 estimates parameters with subscript 1; regression 2 estimates parameters with subscript 2. The *J* statistic is the Hansen test of overidentifying restrictions (we report the *p* values in parentheses).Standard errors in parentheses.

When we include 1998–2005 in the sample, our estimations of the policy reaction function in equation (5) yield poor results. Most coefficients become insignificant or have the wrong sign (or both). This may reflect the change in the inflation-targeting framework starting in 1999. To explore the possibility of a structural change, we use the

18. To test how this could have been reflected in the policy rule (other than through the exchange rate), we included the current account in some of our estimations. The coefficient associated with the current account balance turned out to be negative and significant, indicating that the Central Bank responded actively to changes in the current account. Increases in the current account deficit beyond the target level triggered increases in the interest rate to cool down domestic expenditure.

predictive test proposed by Ghysels, Guay, and Hall (1997) and Guay (2003) and discussed in the context of the Phillips curve estimation above. The results of the test indicate that there is a structural break and that this break occurs around the first quarter of 2000.[19]

To account for this structural change, we estimate the following policy reaction function:

$$r_t = (1-\rho_1)r + \rho_1 r_{t-1} + (1-\rho_1)\omega_{\pi,1}\left[E_t(\pi_{t+4}) - \pi_{t+4}^{tar}\right] + (1-\rho_1)\omega_{y,1} g_{t-1} + u_t, \tag{6}$$

using quarterly data from 1998:1 to 2005:4. We do not split the sample exactly at the breakpoint detected with the PR test in order to have enough data to estimate the monetary policy rule for the second phase.

The estimations of this last policy rule indicate that in the full-fledged inflation-targeting phase, the monetary authority has become more forward looking in terms of inflation. In contrast to the previous phase, the evidence indicates that the Central Bank is willing to allow deviation of current inflation from the target as long as it does not have an impact on future inflation. Additionally, the coefficient associated with inflation, $\omega_{\pi,1}$, increased relative to the corresponding coefficient in the interest rate rule for the previous inflation-targeting phase. This stronger response of the interest rate to inflation deviations is consistent with a central bank that takes advantage of an improved trade-off between inflation and output, possibly as a consequence of higher credibility or commitment to controlling inflation.

5. Conclusions

In this paper, we have provided new evidence of changes in the dynamics of the Chilean inflationary process in recent years. Based on a new Keynesian Phillips curve, we showed that price rigidity has increased in recent years, while the degree of indexation in the economy—based on past inflation—has decreased. We also showed that the exchange rate pass-through into tradable goods inflation has decreased. Our findings are consistent with the idea that the

19. The PR test results were as follows: supPR: 174.48; avgPR: 71.27; and exp PR: 83.94. The tests were calculated using a Newey-West covariance matrix robust to serial correlation up to twelve lags.

credibility of monetary policy has increased over time. As monetary policy has become more credible, costly price adjustments have been carried out less frequently, and the prevalence of indexation based on past inflation has decreased.

These changes in the inflationary process, triggered by an enhanced credibility of monetary policy, may have had significant repercussions on the way monetary policy is implemented in Chile. In particular, as the Central Bank improves its credibility, it may fight inflation deviation from target strongly without having to sacrifice much output. In this context, the paper also presented evidence on a structural change in the policy rule that characterizes the conduction of the monetary policy. Our results are consistent with the idea that over the last several years, Chilean monetary policy has been operating in an improved credibility environment. The monetary policy rule has become more forward-looking in terms of inflation and more aggressive in fighting deviations of inflation from the target.

Appendix
Frequency of Price Adjustment and Pass-Though

From the main text, we have that imported goods inflation is given by,

$$\pi_{M,t} = \frac{(1-\theta_M)(1-\theta_M\beta)}{\theta_M}(s_t + p^*_{M,t} - p_{M,t}) + \beta E_t \pi_{M,t+1}.$$

When we normlaize $p^*_{M,t} = 1$, the law of motion for the domestic price is given by

$$p_{M,t} = \frac{\lambda_M}{1+\lambda_M+\beta}s_t + \frac{1}{1+\lambda_M+\beta}p_{M,t-1} + \frac{\beta}{1+\lambda_M+\beta}p_{M,t+1},$$

where $\lambda_M = [(1-\theta_M)(1-\theta_M\beta)]/\theta_M$. For simplicity, let us assume that the exchange rate follows a random walk process. Let δ_1 and $\delta_2 = (\beta/\delta_1)$ be the unstable and stable roots of the solution for $p_{M,t}$.[20] We can write the solution for the domestic currency price of imported goods as

$$p_{M,t} = \frac{1}{\delta_1}p_{M,t-1} + \frac{\lambda_M}{\delta_1}\frac{1}{1-\delta_2}s_t.$$

The short-run exchange rate pass-through coefficient is defined as

$$\frac{\partial p_{M,t}}{\partial s_t} = \lambda_M \frac{1}{\delta_1 - \beta}.$$

Under flexible prices, $\lambda_M \to \infty$, $\delta_1 \to \infty$, and $(\delta_1/\lambda_M) \to 1$. Then, the pass-through elasticity is unitary. If prices are completely sticky, however, then $\lambda_M = 0$ and $\delta_1 = 1$. In this case, the pass-through coefficient is zero. By a continuity argument, the pass-through coefficient is increasing in λ_M and, therefore, decreasing in θ_M.

20. Coefficient δ_1 is the solution to $\delta_1^2 - \delta_1(1 + \lambda_M + \beta) + \beta = 0$.

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