I. INTRODUCTION

The Central Bank of Chile (CBCH) currently uses a variety of models to forecast inflation. One is the Quarterly Projection Model (Modelo Trimestral de Proyección) in two versions. The first version comprises various modules, such as the aggregate demand module and the output gap module, with some forward-looking elements. The second version is more comprehensive as it includes a well-defined supply side sector, takes into account some stock-flows relationships, and has a more disaggregated aggregate demand. The CBCH is also developing a quarterly general equilibrium model, and a real business cycle model which stresses the effects of different shocks and relates the parameters of the model to actual time series behavior. The CBCH uses VARs, filters, and leading indicators as well, and is currently exploring the possibility of developing a calibrated model for inflation projections and policy analysis.

The main objective of this paper is to add to that set of inflation-forecasting models the framework offered by state-space models, and to explore whether allowing for regime shifts seems justified by Chilean data during the period 1991–99. The state-space framework is useful not only because it provides its own forecast inflation, but also because it offers a powerful method to estimate important unobserved economic variables that are often encountered in inflation-forecasting models. Therefore, the purpose of this paper is not to enter into a horse race with other inflation-forecasting models, and thus, it only briefly compares the out-of-sample forecasts obtained using state-space models with those generated by a simple Box-Jenkins univariate time series approach. Moreover, Granger (2000) suggests that whenever there are close model specifications—as it is, for example, the case of some of the models used in the CBCH and in this research—it is optimal to find their outputs related to the purpose of the models, such as their forecasts, and pool their values.

State-space models are particularly useful for estimating relationships that might have been subject to important changes within the estimation period. In the last two decades, the Chilean economy has undergone significant structural changes that have affected the allocation of resources and its potential output growth. Those reforms have affected the output mix between tradables and nontradables, the allocation of consumption, the sources of financing of production and consumption activities, the legal framework for the allocation of leisure and work effort, and for the use of domestic and foreign capital. Similarly, the formulation and implementation of monetary policy has been changed as the country moved steadily toward an orthodox inflation targeting regime. A central feature of Chilean monetary policy in the 1990s has been the acquired autonomy of the CBCH and the pre-announcement of a 12-month point inflation target for the following calendar year starting in September of 1990. The inflation target has been attained with high precision. In September 1999, the...
CBCH announced that starting in 2001, it will target CPI inflation within an inflation target band which will range between 2 and 4 percent per annum permanently. It is likely, therefore, that the arguments, and possibly the functional form of the loss function of the CBCH, have changed over time. Similarly, these developments have probably had a large effect on the functioning of markets and on the determination of inflation expectations.

The experience of other countries that have undergone reforms as important as those underwent by Chile indicates (as theory predicts) that the parameters that describe the system's dynamics and variance change. While under normal circumstances optimizing economic agents are expected to regularly revise their estimates of the coefficients of the system when new information becomes available, in cases of large structural reforms they may also have to change the set of equations describing that economic system. As a result, macroeconomic policymakers in general, and central banks in particular, have found that in-sample re-fitting of traditional, fixed-parameter models to the data generating process becomes a regular exercise in rapidly changing economies. This notwithstanding, the out-of-sample forecasting ability of models tends to be poor. This has practical implications. For example, the structural instability of the models' parameters, as well as the uncertainty about the "true" model of the economy, may produce biased inflation forecasts which can lead to a breach of a central bank's inflation target.

The next section of the paper discusses the state-space framework proposed. Section III describes the data used, and tests for unit roots and breaks in the sample. Section IV presents the estimation results and the out-of-sample forecasts of the models. The last section concludes the paper.

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2 Wong (2000) finds that the responses of output and price levels to monetary shocks were quite variable in the U.S. in the sample period 1959:1–1994:12. Wong suggests the use of time-varying parameter models to analyze the variability in the effects of monetary policy on economic activity and prices; simple time-invariant linear VAR models may be misleading.
II. MODELS FOR THE FORECASTING OF INFLATION IN CHILE

Figure 1 displays annual inflation measured as the log difference in the average of each quarter CPI with respect to the average of the same quarter of the previous year. It is obvious that there has been a significant change in the level and in the variability of inflation over the sample period. This points to the difficulty of fitting a model of inflation in Chile during the 1990s, and thus, of forecasting inflation.

This paper starts from the premise that structural changes have altered and continue to alter the behavior of economic agents. This implies, among other things, that there will likely be instability in any econometric model that one wishes to fit to the data. The approach proposed in this paper will be, therefore, to deal with structural and regime changes by using state-space models. This opens a number of possibilities with different degrees of complexity.

This paper will estimate two models of inflation for Chile. The first model is a time-varying Phillips curve model of inflation and the second model is a reduced form model of inflation in a small open economy that does inflation targeting. In turn, the first model will be estimated excluding the pre-announced official inflation target—henceforth, version one—and including the pre-announced inflation target—henceforth, version two. The time-varying Phillips curve model and the small open economy model will also be estimated allowing for a two-state Markov-switching process.

A. A Time-Varying Phillips Curve Model of Inflation

The first model of inflation is based on an expectations-augmented Phillips curve derived from Lucas' (1973) supply function in the usual manner. In contrast to the standard expectations-augmented Phillips curve, the model allows the parameters to vary over time (in agreement with Lucas' (1973) well-known conclusion). In Chile, the variation of parameters over time could be interpreted as reflecting the learning process of economic agents as reforms unfolded, and the monetary policy framework approached its steady state.

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3 Appendix I briefly describes state-space models.

4 As shown in the literature (e.g., McCallum (1989) and Turnovsky (1997)), Lucas' (1973) supply function (his equation (7)) can be transformed into a standard expectations-augmented Phillips curve.

5 Whether the time-varying Phillips curve is consistent with the suggestion that the Phillips curve is nonlinear and asymmetric (e.g., Clark et al (1995) and Razzak (1995)) depends on the rationale given for that nonlinearity and asymmetry. For instance, the view that the nonlinearity and asymmetry of the Phillips curve is mostly due to time-varying, central bank's weights on inflation and output variance, would be consistent with the rationale for the time-varying Phillips curve given in this paper. In that case, the institutional changes that made the (continued...
The time-varying Phillips curve is:

\[ \pi_t = E_{t-1}\pi_t + \beta_t(L)x_t + \varepsilon_t, \quad (1) \]

where \( \pi_t \) is inflation at time \( t \); \( E_{t-1} \) is the mathematical expectation operator based on the information set available at time \( t-1 \); \( \beta_t \) is parameter that is allowed to vary over time; \( x_t \) is a measure of the output gap, and \( L \) is the lag operator; \( \varepsilon_t \) is a stochastic process zero mean and variance \( \sigma^2 \). It is assumed that the roots of \( \beta_t(L)x_t \) lie outside the unit circle.

Notice that the regressors of equation (1) are unobservable variables. To deal with that feature of the model, the strategy followed is the following. First, based on the assumption (econometrically tested below) that the inflation series has been subject to changes in its intercept as well as in its slope, the unobserved expected inflation is assumed to follow a random walk. Normally, structural shifts are best modeled as discrete shifts. However, in a context in which it is assumed that agents adjust their forecasts only when new information is received, modeling discrete changes using a random walk is a good approximation.\(^6\) Thus,

\[ E_{t-1}\pi_t = E_{t-2}\pi_{t-1} + \tau_t, \quad (2) \]

where \( \tau_t \) is a stochastic process zero mean and variance \( \sigma^2 \). Second, following Clark (1987), the output gap is estimated assuming a local linear trend and an autoregressive process of order two for output behavior. The output gap model is described in Appendix II. Equations (1)–(2) can be used to calculate expected inflation as an unobserved variable, and to forecast inflation in periods \( t+s \) for \( s \leq 1 \).

Given the significant reforms underwent by Chile (including changes to the monetary policy framework of the CBCH), the time-varying parameter \( \beta_t \) represents the learning process of economic agents. Therefore, the time-varying Phillips curve model captures the uncertainty introduced into the inflationary process by those changes. The time-varying Phillips curve model is also estimated allowing for another source of uncertainty, i.e., the uncertainty due to future random shocks. As explained in Appendix I, the error term \( \varepsilon_t \) in equation (1) is a discrete variable \( S_t \) which evolution depends on \( S_{t-1} \) only, i.e., \( S_t \) follows an order 1 Markov process.\(^7\) The model becomes a time-varying, Markov-Switching, model of inflation.

CBCH independent, and made price stability its primary goal, should minimize the importance of that cause of the Phillips curve's non-linearity and asymmetry over time. However, there are other rationalizations for the non-linearity and asymmetry of the Phillips curve that may be more difficult to reconcile with the rationale for the time-varying Phillips curve given in this paper.

\(^6\) This insight is owed to Kim.

\(^7\) See equations A1.3', A1.6, A1.7a and A1.7b.
Note that equation (1) is not identified because it is not possible to generate simultaneously an estimate of the output gap and an estimate of the time-varying parameters $\beta_t$. The alternative of estimating simultaneously the unobserved components of real GDP (i.e., a stochastic trend and a cyclical component) and a standard expectations-augmented Phillips curve with constant $\beta_t$ was not feasible. Estimates either displayed significant serial correlation or the information matrix was singular, an indication that the model may not be identified.\(^8\) Therefore, a two-step approach is followed by which first a series for the output gap is estimated, and then, the output gap so generated is used to estimate the time-varying parameter model (1)–(2). The price to pay for this approach is an efficiency loss due to the use of the generated regressor $x_t$ when estimating the expectations-augmented Phillips curve with time-varying parameters.

**B. A Small Open Economy Model of Inflation**

An alternative to specifying a random walk process for expected inflation as in equation (2), is to substitute the set of state variables (predetermined) suggested by a structural model for expected inflation. The set of state variables is determined from a rational expectations model of a small open economy that does inflation targeting. The model is briefly described in Appendix III.\(^9\) The inflation process can thus be represented by the following reduced form equation:

$$\pi_t = F(g_t, g^*_t, r_t, \pi^*_t, d_t, f_t, c_t) + \tau_t, \quad (3)$$

where $g_t$ is domestic productivity, $g^*_t$ is the rest of the world's productivity; $r_t$ is the cost of foreign financing faced by the Chilean economy (including the risk premium); $\pi^*_t$ is an index of the country's terms of trade; $d_t$ is a measure of fiscal impulse; $f_t$ is the nominal exchange rate, $c_t$ is the pre-announced official inflation target, and $\tau_t$ is a white noise process.

Note, however, that the estimation will be done without imposing the set of cross-equations restrictions that result from the solution of the model of Appendix III. There are simply not

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\(^8\) Different versions of the model were tried, e.g., using different lags in the output gap, or using an ARMA representation for the output gap, or using two constants different from zero for the processes describing the cyclical parts of real GDP and inflation. During estimation of the models, difficulties were encountered either in inverting the matrix of second derivatives of the log likelihood function, or the estimation did not converge. According to Rothenberg (1971), this may indicate that a model is not locally identified, i.e., that more than one set of values for the parameters can give rise to the same value of the log likelihood function; the data cannot discriminate among the possible values.

\(^9\) This draws on Nadal-De Simone (1999), which contains the analytical solution and simulations of a similar model using New Zealand data.
enough data to estimate all the parameters. Most importantly, the model does not provide guidance on the time-varying combinations of parameters implied by the cross-equation restrictions.

III. DATA ANALYSIS

The data used in this paper were provided by the CBCH. The data set comprises the following variables: Chilean real GDP ($y_t$), an index of real economic activity in partner countries ($y_t^*$), Chilean annual inflation as measured by the CPI ($\pi_t$), annual inflation in partner countries ($\pi_t^*$), the cost of financing for the Chilean economy, i.e., the three-month LIBOR rate plus a country risk premium ($r_t^*$), an index of terms of trade ($p_t^*$), the nominal exchange rate defined as the number of pesos exchanged for one U.S. dollar ($f_t$), a measure of the fiscal impulse ($d_t$), and the pre-announced point inflation target for the following calendar year since 1991 ($c_t$). The data are quarterly. The sample starts in 1986:1 and finishes in 1999:3. The data used in the estimations are always in natural logarithms with the exception of interest rates. Series $\pi_t$, $\pi_t^*$, $r_t^*$, and $f_t$ are quarter averages. The seasonal component of the series has been removed using X-11.


Given that inflation and real output seem to have a stochastic trend, the unit-root test used is the modified Dickey-Fuller t-test (DFGLS) proposed by Elliott, Rothenberg, and Stock (1996), a point-optimal invariant test which has a substantially improved power when an unknown mean or trend is present in the data. Table 1a shows that, in the period 1986:1–1999:3, the null of a unit root with a constant and a linear trend cannot be rejected for any variable. Changes in all the variables are stationary. In the period 1990:1–1999:3, the null of unit root with a constant and a linear time trend cannot be rejected for any of the variables except annual inflation and the terms of trade. Changes in all the other variables are stationary.

Given the nonstationarity of inflation in the period 1986:1–1999:3, the inflation series as well as changes in it, are tested using Perron (1997) test which allows for a shift in the intercept of the trend function and/or a shift in the slope; the date of the possible change is not fixed a priori but it is endogenously determined. Table 1b shows the results for 2 models. The

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10 Quarterly figures for the pre-announced official inflation target were calculated assuming a constant rate of decline from each yearly inflation target to the following one.

11 The lags used in the unit-root tests are chosen using the Schwarz Information Criterion and checking that the residuals are white noise using the Box and Pierce Q statistics.
"innovational outlier model" (model 1) allows only a change in the intercept under both the null and the alternative hypothesis, and the "additive outlier model" (model 2) allows a change in both the intercept and the slope. Two methods are used to determine the break point (\(T_b\)): the first method selects as breaking point the one that minimizes the t-statistic for testing the null of unit root (\(t_{u}\)) while the second one minimizes the t-statistic on the parameter associated with the change in the intercept (model 1) (\(t_{i}\)), or the change in the slope (model 2) (\(t_{s}\)). The lag parameter is chosen following a general-to-specific recursive procedure so that the coefficient on the last lag in an autoregression of order \(k\) is significant, and that the last coefficient in an autoregression of order greater than \(k\) is insignificant, up to a maximum order \(k_{\text{max}}\).

Model 1 does not reject the null hypothesis of a unit root either in the inflation series or in its changes using any of the two methods for choosing the break point \(T_b\). The tests show a change in the intercept of inflation in 1988, and a change in the intercept of changes in inflation either in 1992, or in 1990, depending on the method chosen to estimate the break point.

Model 2 strongly rejects the null hypothesis of a unit root in the inflation series and in its change independently of the method used to choose the break point \(T_b\). The break point in the intercept and/or slope of inflation is in 1988 while the break point in the intercept and/or slope of changes in inflation is in 1990.

Therefore, the Perron (1997) test indicates that the inflation rate is stationary in the entire sample period when allowance is made for changes in the intercept and the slope.

### IV. The State-Space Representation of the Models and Results

#### A. Estimation of the Output Gap Series Using an Unobserved Components Model of Output

As stated above, estimation of the model (1)–(2) requires an estimate of the output gap. This is done using the entire sample period 1986:1–1999:3. The state-space representation of the estimated unobserved components model of output is:

\[
y_t = \begin{bmatrix} T_t \\ X_t \\ X_{t-1} \\ g_t \end{bmatrix},
\]

\[
\begin{bmatrix} T_t \\ X_t \\ X_{t-1} \\ g_t \end{bmatrix} = \begin{bmatrix} 1 \cdots 0 \cdots 1 \\ 0 \cdots \eta_1 \cdots \eta_2 \cdots 0 \\ 0 \cdots 1 \cdots 0 \cdots 0 \\ 0 \cdots 0 \cdots 0 \cdots 1 \end{bmatrix} \begin{bmatrix} T_{t-1} \\ X_{t-1} \\ X_{t-2} \\ g_{t-1} \end{bmatrix} + \begin{bmatrix} h_t \\ I_t \\ X_{t-2} \\ 0 \end{bmatrix},
\]

\[
(5)
\]
Once the model is in the state-space form, it can be estimated using the Kalman filter.\textsuperscript{12} Table 2 shows the estimated variances $\sigma_1^2$ and $\sigma_2^2$, as well as the fixed parameters of the autoregressive process of order 2 assumed for the cyclical component of real GDP, i.e., $\theta_1$ and $\theta_2$. The estimation was constrained such that the roots of the characteristic equation of the process $\phi(L)x_t$ lie outside the unit circle, and that the variances $\sigma_1^2$ and $\sigma_2^2$ are positive numbers. All parameters are significant at the usual significance levels. The Q-statistic tests for serial correlation as well as the Kolmogorov-Smirnov periodogram test of the standardized forecast errors and the squared of the standardized forecast errors, cannot reject the white noise null hypothesis.

Figures 2-4 show the log of real GDP and its stochastic trend component, its cyclical component, and its productivity growth component, respectively.\textsuperscript{15} The cyclical component profile seems to match the standard description of the Chilean business cycles of the 1990s. Three points are noteworthy. First, during the 1990s, the area covered by the negative part of the cyclical component of output was larger than the area covered by the positive part of the cyclical component of output. This is consistent with the steady decline in inflation sought, and successfully obtained, by the monetary authorities during that period. Second, the cyclical component of output seems to have peaked in 1998:1, i.e., before the monetary policy tightening of the second half of 1998. Finally, average productivity growth\textsuperscript{14} seems to have declined steadily from a quarterly average growth rate of 2.12 percent in 1992–94 to 2.05 percent in 1995–96, to 2.0 percent in 1997, and to 1.9 percent in 1998. This trend seems consistent with the view held by some observers that the potential output growth of the Chilean economy in this decade may not reach the levels of the 1990s due to the completion of one-time gains from past structural reforms. Recent policy measures to widen and deepen the domestic capital and money markets, to further liberalize the capital account, as well as to continue the unilateral trade liberalization and education reforms, may reverse that downward trend. In any case, the relevant point is that productivity growth will be shown to be an important factor in the forecasting of inflation in Chile.

\textbf{B. Estimation of the Time-Varying Phillips Curve Model of Inflation}

The state-space representation of the time-varying parameter Phillips curve model of inflation is:

\textsuperscript{12} Appendix IV contains a brief description of the Kalman filter. For a thorough description of the Kalman filter, see Hamilton (1994).

\textsuperscript{13} The first 16 observations were used to eliminate the influence of the "wild guess" made for the nonstationary $\beta_{00}$ (initial values). Large values were given to the diagonal elements of the covariance matrix of $\beta_1(P_{00})$ so as to assign most of the weight in the updating equation to the new information contained in the forecast error.

\textsuperscript{14} As mentioned in Appendix II, productivity growth includes changes in factor endowments.
The model is both locally and globally identified. It seems relevant to recall, however, that identification should not be approached in a rigid manner as it is possible for an equation (or system of equations) to be identified according to the strict rules of identification but the equation may have very little predictive power if the predetermined variables in the equation have little variance. From this point of view, the relatively good forecasting performance of the models (see below) is encouraging.

Table 3 shows two sets of estimates of the time-varying Phillips curve model: the first one, excludes the pre-announced official inflation target while the second one includes it. In the second one, the dependent variable is measured as inflation deviations from the official target. The results suggest that although the estimation of the model that includes the pre-announced official inflation target displays some additional serial correlation, this version is to be preferred because it achieves a better forecasting performance than the model in terms of actual inflation without the information provided by the official pre-announced inflation target.

In both versions of the time-varying Phillips curve model, all parameters are significant at the usual confidence levels. The log-likelihood function of version two of the model is relatively higher. However, there is serial correlation in the standardized forecast errors of version two of the model. In contrast, the Q-statistic tests for serial correlation of the standardized forecast errors of version one of the model show some serial correlation at the 90 percent significance level only for Q(16) and Q(24). The Q-statistic tests for serial correlation of the squares of the standardized forecast errors show no serial correlation in any of the two versions.

Table 3 also shows the out-of-sample forecasts for 1, 2, 3, and 4 quarters and the actual values of inflation. Version two of the model achieves a considerable reduction in Theil's

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Rothenberg (1971) showed that local identification at $\beta_0$ requires that the information matrix be nonsingular in a neighborhood around $\beta_0$. This criterion was used for testing for local identification of the model described by equations (1)–(2), and also of the model described by equation (3). In no case there was any difficulty in inverting the matrix of the second derivatives of the log likelihood functions. Global identification was tested working directly with the state-space representation of the model, as suggested by Burmeister et al (1986).
inequality coefficient as the value of the statistics for version two is between 19 to 64 percent of its value for version one of the model. A similar picture is offered by the root mean-square forecast percent error or the mean forecast percent error. Figure 5 shows actual and expected inflation, and Figure 6 shows actual and expected deviations from the pre-announced official inflation target, during the sample period. From both figures two observations seem important. First, the pre-announced official inflation target point was below the predicted level of inflation over two thirds of the time. Second, actual inflation was closer to the target point than to the predicted level of inflation just over half of the time.\textsuperscript{16} While the first observation is consistent with Morandé and Schmidt-Hebbel (2000), the second is not necessarily. However, given that the framework used in this paper explicitly considers the expectations-updating process by economic agents, all results can be interpreted in the same manner the authors do, i.e., as evidence that the inflation targeting regime contributed to enhance the credibility of the CBCH and, thus, played a role in reducing inflation in Chile. The inflation target announced, and always accomplished with a great deal of accuracy, offset inflationary inertia over time.

Figures 7 and 8 show the smoothed\textsuperscript{17} time-varying parameters of the contemporaneous and first lag of the output gap of version two of the model. There seems to be a tendency for the variability of the parameter values to fall over time. The absolute value of the contemporaneous parameter of the output gap also falls over time. Given that the CBCH was successful in its inflation targeting strategy over the sample period, this result should not come as a surprise because hitting the inflation target implies that the output gap variance should be reduced over time. For given private sector’s expectations, inflation targeting implies inflation forecast targeting (Svensson, 1997). As the implicit loss function of the independent Chilean monetary authority includes the inflation target as the primary objective of monetary policy, one of the sources for a nonzero output gap (i.e., changes in the relative weights attached by the monetary authority to output and inflation variance), and thus for inflation uncertainty, is removed. All the other forces that determine the output gap are, certainly, still operational.\textsuperscript{18}

\textsuperscript{16} This is not statistically different from 50 percent.

\textsuperscript{17} As the state vector is given a structural interpretation in model (1)-(2), it is important to form an inference about the value of the state vector based on the full sample. Therefore, the value of the contemporaneous and lagged coefficients of the output gap have been calculated based on the full set of data collected by moving through the sample backward starting with $t = T-1$. Therefore, the time-varying parameters have been smoothed.

\textsuperscript{18} Given the profession’s lack of agreement on the relationship between nominal and real variables, other interpretations are certainly possible. For instance, efficiency wage theories of wage determination would suggest that given a constant markup, the effect of the output gap on wages and hence on inflation would be smaller with about 3 percent average inflation in 1997 than with about 15 percent average inflation in 1992. However, the sum of the estimated output gap coefficients falls until mid-1995, to increase thereafter. As indicated by
The state-space representation of the solution of the open economy model is:

\[ \pi_t = x_t \beta_{t-1} + e_t, \]  \hspace{1cm} (8)

\[ \beta_{t} = I_k \beta_{t-1} + r_{t}, \]  \hspace{1cm} (9)

where the \( x_t \) are the \( k-1 \) state variables of the open economy model, i.e., domestic productivity, foreign productivity, the real cost of financing of the Chilean economy, an index of the terms of trade, a fiscal impulse measure, the nominal exchange rate with respect to the U.S. dollar, and the pre-announced official inflation target. \( I_k \) is a \( k \)-order identity matrix.

Table 4 shows the results of the estimation of the open economy model represented by equations (8)–(9). The estimated small open economy model of inflation has a somewhat higher likelihood function than version two of the time-varying Phillips curve model. Serial correlation is only present between lags 16 and 24.

All variables are highly significant at standard confidence levels with the exception of foreign productivity. The out-of-sample forecasts are less accurate than those produced by the time-varying Phillips curve model but are more stable (the Theil’s inequality coefficient deteriorates less as the out-of-sample forecasting period is lengthened).

Figures 9–14 show the smoothed time-varying parameters of the model. There is a number of interesting observations. First, the coefficient on the cost of foreign capital is negative, and its absolute value increases until 1995. In standard models of exchange rate determination, a reduction in the cost of foreign financing is expected to increase domestic expenditure thereby pushing up inflation in non-tradable goods and in the CPI, other things equal. This is the real-sector channel of the monetary transmission mechanism. However, a reduction in the cost of foreign financing is expected to also appreciate the domestic currency, and translate into lower CPI inflation, other things equal. This is the asset-market channel of the monetary transmission mechanism. In an inflation targeting regime, in contrast, the cost of foreign financing is negatively correlated with domestic CPI inflation. Briefly, under inflation targeting, because liquidity is endogenous, the asset-channel part of the transmission mechanism is weakened and thus, the negative correlation between changes in the cost of

Wong (2000) in his study for the U.S., no single theory of the monetary transmission mechanism seems capable of explaining the changing response of output and prices to monetary policy; one should rather search for a combination of economic and institutional factors.

\[ \text{Footnote:} 19 \text{ For the reasons explained in Appendix III, the model is estimated without imposing the cross-equation restrictions that result from solving the model.} \]
foreign financing and domestic CPI inflation that result from the real-sector channel may prevail. There seems to be some indication that this happened in Chile.²⁰

Second, the fiscal impulse coefficient is positive, as expected. Consistent with the relatively expansionary fiscal stance adopted after 1995, the weight of the coefficient in explaining inflation variance doubled between the end of 1996 and the mid-1998. It remained at that level thereafter.

Third, the pass-through of changes in the nominal exchange rate to domestic CPI inflation varied over time significantly. As the inflation targeting regime acquired more credibility, the value of the pass-through coefficient fell. After mid-1995, and until the last year of the sample, the pass-through coefficient was stable at about 10 percent. It seems that when annual real GDP growth fell from 8 percent in the first quarter of 1998 to 5.9 percent in the second quarter of 1998, the pass-through coefficient also started to fall. It reached a value of less than 5 percent at the end of the sample period. Collins and Nadal-De Simone (1996) show that the “pass-through coefficient” depends at a minimum, on the structure of the economy, on the nature of the shocks affecting the economy, on the composition of the CPI regimen, and on the central bank's operating procedure. One implication they draw is that pass-through coefficients are likely to be econometrically unstable across policy regimes as well as across time within the same policy regime. This econometric study seems to validate that conclusion.

Finally, judging from Figure 14, the smoothed inflation target coefficient remained quite stable after 1994. Given the framework used in this study, this seems to be compelling evidence that the credibility of the Chilean inflation target framework was well established since the mid-1990s.

D. Does it Matter to Allow for Regime-Switching in the Models of Inflation?

The squared standardized forecast errors of version two of the time-varying Phillips curve model do not seem to suggest the presence of heteroskedasticity. However, it was considered useful to explore this possibility further by estimating the time-varying Phillips curve model with a Markov-switching process for the disturbance terms. Therefore, version two of the time-varying Phillips curve model was estimated allowing for uncertainty derived not only from the economic agents' updating of the model's parameters but also for uncertainty derived from heteroskedasticity of the disturbance terms. This is the model described by equations (6)-(7) plus equations (A1.3'), (A1.6), (A1.7a) and (A1.7b) from Appendix I.

In this model, the conditional variance of the forecast error can be decomposed into conditional variance due to the unknown regression coefficients, and conditional variance due to the heteroskedasticity of the disturbance term. The first conditional variance depends on the state of the world at time t-1, S_t-1, while the second one depends on the state of the

²⁰ Nadal-De Simone (1999) confirms this point using New Zealand data.
world at time $t$, $S_t$. Figure 15 shows that during most of the sample period, the first source of conditional variance was far more important than the second source. In the period 1992:1–1998:2, the average conditional variance of inflation was 0.0025, and about 75 percent of it was accounted for by the learning process of economic agents. It is only between the third quarter of 1998 and the third quarter of 1999 that the heteroskedasticity of the disturbance term is significant; in that period, it explained about 82 percent of the 0.0048 conditional variance of inflation.

The small open economy model of inflation was also estimated allowing for Markov switching, i.e., estimating equations (8)-(9) together with equations (A1.3), (A1.6), (A1.7a) and (A1.7b). Even more forcefully than in case of the time-varying Phillips curve model, the decomposition of variance in figure 16 shows that the conditional variance of the small open economy model forecast error was mostly due to the unknown regression coefficients.

One possible explanation for these results is suggested by the presence of the pre-announced official inflation target among the regressors of the models. To the extent that the inflation target was credible, it reflected changes in the Chilean inflation process; the inflation target proxied the change in the Chilean monetary policy regime from high and variable inflation to low and stable inflation. The implication of these results is that during the sample period of this study, there is little to gain from including in the model uncertainty due to heteroskedasticity of future random shocks. However, once the inflation targeting regime is in its steady state, this is an issue that will require revisiting.

### E. Estimation of Selected Time Series Models

A number of models based on Box-Jenkins techniques was also estimated. A reduced set of selected estimation results is reported in Table 5. The reported results refer to an AR(1), an AR(2), and an AR(1) transfer function model. The models are:

\[
\begin{align*}
\pi_t &= \phi_1 \pi_{t-1} + \epsilon_t, \quad (10) \\
\pi_t &= \phi_1 \pi_{t-1} + \phi_2 \pi_{t-2} + \epsilon_t, \quad (11) \\
\pi_t &= \phi_1 \pi_{t-1} + \theta_1 z_t + \epsilon_t, \quad (12)
\end{align*}
\]

where $\pi_t$ is measured as the deviation of annual inflation with respect to the pre-announced official inflation target in equations (10) and (11), and it is measured as annual inflation in equation (12); $z_t$ is the pre-announced official inflation target. As before, model (12) assumes that $z_t$ is exogenous.

The AR(1) and AR(2) models show serial correlation in the first 8 lags while the AR(1) transfer function model also shows serial correlation in the first 16 lags. The AR(1) and AR(2) models have low $R^2$ while the AR(1) transfer function model has a reasonable $R^2$ (this is obviously biased upward by the presence of serial correlation in the residuals).

Based on the Theil's inequality coefficient for the first-step forecast, the AR(1) and AR(2) models do better than version two of the time-varying Phillips curve model and the open
economy model (Table 6 has a summary of forecast results). However, the relative performance of models AR(1) and AR(2) deteriorates rapidly for subsequent forecast periods so that the time-varying Phillips curve model outperforms them. The time-series models still do better than the open economy model of inflation.

The AR(1) transfer function model does better than the time-varying Phillips curve model and the open economy model up to the third-period forecast. It does still better than the open economy model for the fourth-period forecast.

However, although the AR(1) transfer function model does better for short-run forecasts than the AR(1) and AR(2) models, it is likely that this ranking will change when the monetary policy regime enters its steady state during 2001 because the inflation target variance will become constant, i.e., its information content will fall. It will then be relatively important that the deterioration of the out-of-sample forecasts of the different models of inflation is not significant as subsequent periods are added to the forecast. On the basis of this criterion, i.e., for medium term forecasting, it seems that the time-varying Phillips curve model and the open economy model do relatively better than the AR(1) transfer function model. The open economy model forecasts actually show quite a remarkable stability in terms of bias and variance. The second version of the time-varying parameter Phillips curve model, given that it is relatively parsimonious, also does a good forecasting job. However, as with the AR(1) transfer model, it remains to be seen whether the time-varying Phillips curve model will still be the best specification of the inflation process in Chile after 2000. It is possible that a richer structure becomes then necessary.

V. CONCLUSIONS

The objective of this study is to estimate and forecast inflation in Chile using a state-space framework. Two models of inflation are estimated and used for out-of-sample forecasting of Chilean inflation. The first model is a time-varying Phillips curve model estimated in two versions; version one excludes the pre-announced official inflation target point and version two includes it. The second model is a reduced form model of a small open economy that does inflation targeting. The results of those estimations are compared with those of simple Box-Jenkins specifications of the inflation process in Chile. The two models of inflation are also estimated allowing for regime changes by using a two-state Markov-switching model of order 1. The sample period comprises 1990:1–1999:3, a period which is one year short of the steady state of the monetary policy regime, i.e., a regime in which the authorities target annual CPI inflation within an inflation target band which will range between 2 percent and 4 percent on a permanent basis.

For the United States, Stock and Watson (1999) have shown that a stand-alone conventional Phillips curve generally produces more accurate forecasts than other macroeconomic variables, including interest rates, money, and commodity prices. However, it is inferior to a generalized Phillips curve model based on measures of real activity other than unemployment such as an index based on a large number of real economic indicators.
Models that include the pre-announced official inflation target point are to be preferred to those that exclude this variable. Although including the pre-announced official inflation target introduces some serial correlation in the residuals, it also reduces forecasts errors significantly.

The out-of-sample performance of the time-varying Phillips curve model that includes the official pre-announced inflation target point is more favorable than the out-of-sample performance of small open economy model—which includes the pre-announced official inflation target point. In contrast, the statistics for the out-of-sample forecasts of the small open economy model are relatively more stable.

For the first step of the out-of-sample forecast, the Box-Jenkins models of inflation tend to do better than the time-varying Phillips curve model that includes the pre-announced official inflation target. However, their relative forecasting superiority deteriorates rapidly for forecasts further out in time.

An AR(1) transfer function model does better than both the time-varying parameter Phillips curve model and the open economy model, up to the third out-of-sample forecast. It still does better than the open economy model for the fourth out-of-sample forecast.

The addition of a Markov-switching model to the time-varying models of inflation does not improve the fit of the models. During most of the sample period, the conditional variance of the forecast error due to the unknown regression coefficients was far more important than the conditional variance due to the heteroskedasticity of the disturbance term.

A note of caution is necessary, however. It is quite likely that the ranking of models performance will change when the monetary policy regime enters its steady state in 2001. The variance of inflation will probably depend on terms of trade shocks, productivity shocks, the fiscal position, the exchange rate, and the like; the variance of the inflation target will in contrast be constant, and will thus have, ceteris paribus, no power in explaining actual inflation. It will then become important to reassess the forecasting performance of the models paying particular attention that the quality of the forecasts in the medium term does not deteriorate rapidly. In that case, it is likely that models less parsimonious than the time series models of Box-Jenkins will become necessary.