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THE NEW KEYNESIAN PHILLIPS CURVE IN AN EMERGING MARKET ECONOMY: THE CASE OF CHILE

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Resumen

En este trabajo se presenta una estimación de la Nueva Curva de Phillips Keynesiana para la economía chilena utilizando el Método Generalizado de Momentos. Nuestra estimación tiende a favorecer una versión que no sólo incluye las expectativas de inflación sino que también incorpora un componente de inflación rezagada. Nuestra evidencia indica que el coeficiente asociado a la inflación rezagada es cercano a 0.4. El coeficiente que captura el grado de rigidez de precios se estima alrededor de 0.65. Este resultado implica que, en promedio, los precios permanecen fijos alrededor de 3 trimestres. Finalmente, presentamos evidencia que la existencia de un cambio estructural en la curva de Phillips en torno al período de convergencia a una inflación estable (alrededor de 2000). La evidencia sugiere que tanto la frecuencia de ajuste óptimo de precios como el grado de indexación a la inflación pasada habrían caído.

Abstract

This paper presents GMM empirical estimations of the New Keynesian Phillips curve (NKPC) for Chile. Our results tend to support the hybrid version of the NKPC, with an estimated backward-looking coefficient of about 0.4. The estimated Calvo coefficient, that captures the degree of price rigidity, assuming firm specific capital is about 0.65. This implies that prices are optimally adjusted on average every 3 quarters, approximately. Our results also indicate the existence of a structural break in the NKPC, which occurred when the inflation target converged to its long-run level (around 2000). We find evidence that the frequency of optimal price adjustment and the degree of indexation to past inflation have decreased over time.

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1 Introduction

The New Keynesian Phillips Curve (NKPC) has attracted broad attention by academics and policy makers in recent years. As opposed to the old tradition, the NKPC highlights that expected future inflation plays a significant role in explaining current inflation. The new approach also emphasizes that real marginal costs, rather than conventional measures of the output gap, are the main driving force behind the inflationary process. On the empirical side, the evidence has tended to confirm that both the forward-looking and the backward-looking components of an hybrid version of the NKPC are important to explain inflation dynamics (see Galí and Gertler, 1999, GG henceforth, and Sbordone, 2002).[1]

While most of the empirical work has been devoted to estimate versions of the NKPC for developed economies, little empirical literature exists for emerging market economies.[2] In this paper, we estimate a NKPC for the Chilean economy using quarterly data for the period 1990:1-2004:4. The Chilean economy has particular features that make it an interesting case of analysis. It has successfully gone through a disinflationary process over the 90s, converging from relatively high inflation rates to a stationary single digit inflation rate.[3] However, historically high inflation rates led to widespread use of contracts with explicit indexation clauses based on previous inflation (see Lefort and Schmidt-Hebbel, 2002). Moreover, these high inflation rates could have affected the credibility of the monetary authority at the beginning of the inflation-targeting period. For both reasons, we would expect an important role for the backward-looking component of the Phillips curve.[4] Another important element in the case of Chile is that, being an small open economy, external terms of trade shocks may have played an important role determining the dynamics of the relevant marginal cost. Therefore, the explicit use of real marginal costs measures –instead of the output gap–, may be particularly important to obtain more accurate estimations of the Phillips curve.

Estimating the Phillips curve is not only interesting from an academic point of view. An assessment of the existence of a relationship between inflation and some measure of economic activity, and an adequate characterization of this relation, has important implications for policy analysis. Moreover understanding the inflationary process is particularly important for a country like Chile that conducts its monetary policy within an inflation-targeting framework.

We consider a version of the basic Calvo (1983) price setting model, where firms adjust prices optimally according to the expected evolution of their marginal costs. As in Christiano, Eichenbaun and Evans (2005) we modify the basic model by allowing passive

[1] The current empirical debate hinges on the extent by which rational expectation forecast of inflation can actually account for the inflation dynamics (see Rudd and Whelan, 2003).

[2] One of the few exceptions is Agenor and Bayraktar (2003) who estimate Phillips curve equations for middle-income countries, including Chile.

[3] This disinflationary process has been characterized by a declining target for the CPI inflation rate set by the monetary authority and the active use of the interest rate as the main monetary policy instrument.

[4] Erceg and Levin (2003), Rudd and Whelan (2003), and Collard and Dellas (2004) show that when credibility of the monetary authority is low, the response of the inflation rate to certain types of shocks is slow and persistent.

price adjustments in order to account for the trend in inflation observed over the 90s. Therefore, every period firms adjust prices either optimally or by following an indexing rule. Price rigidity in this setup is associated to the infrequent *optimal* price adjustment. We also consider an hybrid version of the NKPC where not only forward-looking inflation determines the current evolution of inflation, but also a backward-looking component plays a role. We consider four measures of marginal costs, each consistent with an alternative specification for the technology utilized by firms.

Our results using the Generalized Method of Moments (GMM) tend to support the hybrid version of the NKPC. The evidence shows that the backward-looking coefficient in the NKPC is approximately 0.4. This figure is larger than the corresponding one for the Euro area, as estimated by Galí, Gertler and López-Salido (2001, 2002) (GGL henceforth), Gagnon and Khan (2005) and Jondeau and Le Bihan (2005). It is also larger than the estimated value for the U.S. by GG. The estimated Calvo coefficient that captures the degree of price rigidity when capital is specific is around 0.65, indicating that prices remain unchanged on average for about 3 quarters. This figure are, in general, slightly larger than in the case of the U.S. and similar to the ones for the Euro Area.

We evaluate the goodness of fit of our NKPC in tracking the evolution of the actual inflation rate by computing a measure of *fundamental* inflation as in GG and Sbordone (2002). We find that the model with the lowest mean square error between actual and *fundamental* inflation is the one that utilizes real marginal cost derived from a Cobb-Douglas technology. We also analyze the predictive performance of each model using the approach proposed by Diebold and Mariano (1995) based on a one step ahead forecast. The results of the test suggest that no model has a significant superiority against the others, but on average the specification based on a Cobb-Douglas technology seems to marginally outperform the alternative three models. The results of the test also suggests that an AR(1) model for inflation does not statistically dominate any of our estimated specifications in terms of forecasting accuracy.

As a robustness check we investigate whether changes in the macroeconomic environment may have affected some of the parameters of the NKPC. Given the important change in the level of inflation over the 90s, it is likely that the frequency of price adjustment may have changed. As Taylor (2000) and Devereux and Yetman (2001) have argued, a lower and more stable inflation rate could give rise to less frequent optimal price adjustments. We consider a predictive test for structural change with unknown breakpoint developed by Ghysels and Hall (1990), Ghysels et. al. (1997) and Guay (2003). One of the advantage of this approach is that it allows us to test the presence of a break even when the second subsample contains a few observations and parameter estimates are not feasible, as it is our case. Our results show that for the four specifications of marginal costs we cannot reject the existence of a breakpoint which would have occurred around 2000. Moreover, we find that the backward-looking component in the first subsample is larger than the estimated one using the whole sample, and that the duration of price stickiness is smaller before the breakpoint. This shows that inflation has become less persistent by the end of the sample period and also that the response of inflation to marginal cost fluctuation is now smaller. In other words, the Phillips curve has become "flatter".

The paper is organized as follows. In the next section we review the theory behind the new Phillips curve in the New Keynesian tradition and describe the different measures of marginal costs. In the third section we present the estimation of the NKPC for Chile, together with an assessment of its goodness of fit and an analysis of parameter stability. Finally, the fourth section concludes.

2 Theoretical Framework

2.1 Price setting

We follow the standard Calvo (1983) price setting setup. A fraction $1-\theta$ of the firms in the economy, randomly picked, adjust optimally prices each period. The probability that a particular firm receives a "signal" to update its price at time t is also $1-\theta$, which is independent from the history of the firm.

We assume that a firm that does not receive a signal follows a simple updating rule to reset its price. In particular, if the firm does not receive a signal between t and $t+i$, then the price it charges in $t+i$ is given by $\Gamma_t^i P_t$, where Γ_t^i is a function defined below. Let $R_{t,t+i}$ be the relevant discount factor for the firm between period t a and $t+i$. The maximization problem faced by a generic firm z in t is the following

$$\max_{P_t(z)} E_t \sum_{i=0}^{i} \theta^i R_{t,t+i} \left[\frac{\Gamma_t^i P_t(z) - MC_{t+i}(z)}{P_{t+i}} Y_{t+i}(z) \right] \tag{1}$$

subject to the demand for its good which is given by $Y_{t+i}(z) = \left[\Gamma_t^i P_t(z)/P_{t+i}\right]^{-\epsilon} Y_{t+i}$.[5]

Lets define $Q_t = \frac{P_t^{new}}{P_t}$, where P_t^{new} corresponds to the optimal price that a firm receiving a signal in t would charge, and where P_t is the average price level. From the the first order condition we can express Q_t as:

$$Q_t = \mu \frac{\sum_{i=0} \theta^i R_{t,t+i} \left[\frac{MC_{t+i}(z)}{P_{t+i}} \left(\frac{\Gamma_t^i}{P_{t+i}}\right)^{-\epsilon} Y_{t+i} \right]}{\sum_{i=0} \theta^i R_{t,t+i} \left[\left(\frac{P_t}{P_{t+i}}\right) \Gamma_t^i \left(\frac{\Gamma_t^i}{P_{t+i}}\right)^{-\epsilon} Y_{t+i} \right]} \tag{2}$$

where $\mu = \frac{\epsilon}{\epsilon - 1}$ corresponds to the steady-state gross mark-up.[6] When prices are rigid, the optimal price depends on the expected aggregate output, aggregate prices and real marginal costs. To obtain a linear expression we utilize a first order Taylor expansion of this equation around the steady state:

$$\widehat{q}_t = (1-\theta\beta) \sum_{i=0} (\theta R)^i \left[\widehat{mc}_{t+i} + \sum_{j=1}^{i} \widehat{\pi}_{t+j} - \widehat{\Gamma}_t^{i+1} \right] \tag{3}$$

[5] This demand is obtained from a utility function where households choose optimally the composition of a consumption bundle among a continuum of varieties with an elasticity of substitution ϵ.

[6] Under fully flexible prices –i.e $\theta = 0$– the optimal resetting price for a firm would be $\left.\frac{P_t^{new}}{P_t}\right|_{flex} = \mu \frac{MC_t}{P_t}$. In other words, under flexible prices, firms define an optimal price that is a constant markup over its marginal cost of production.

Given that a fraction θ of the firms adjusts prices passively –i.e. following the simple updating rule Γ_t^i– and the remaining $1-\theta$ fraction sets prices to P_t^{new}, the aggregate price index is $P_t = \left[(1-\theta)\left(P_t^{new}\right)^{1-\epsilon} + \theta\left(\Gamma_t^1 P_{t-1}\right)^{1-\epsilon}\right]^{\frac{1}{1-\epsilon}}$.

2.1.1 The New Keynesian Phillips curve

As a baseline case we assume that the updating rule for those firms that can not optimally adjust prices consist in resetting their prices according to the inflation target the authority defines for the period:[7]

$$\Gamma_t^i = \prod_{j=1}^{i}\left(1+\pi_{t+j}^*\right)$$

Assuming that capital is freely mobile across firms –such that its marginal productivity is the same across firms– we obtain the modified version of the New Keynesian Phillips curve:

$$\widehat{\pi}_t = \lambda\widehat{mc}_t + \beta E_t\left\{\widehat{\pi}_{t+1}\right\} \tag{4}$$

where $\widehat{\pi}_t = \pi_t - \pi_t^*$ corresponds to the difference between inflation and the inflation target set by the authority for the period, $\widehat{mc}_t$ represents the log-deviation of the real marginal cost from its steady-state value, and where $\lambda = \frac{(1-\theta)(1-\theta\beta)}{\theta}$.

Notice that assuming a passive updating rule implies that firms adjust prices every period. Nominal rigidity, in this setup, thus refers to the extent to which firms optimally adjust prices. If θ is large then the frequency of *optimal* price adjustment is low.

2.1.2 Hybrid model

We considered an alternative specification for the Phillips curve, where inflation exhibits persistency. This alternative specification is close in spirit to the hybrid Phillips curve in GG, and it is based on the formulation by Christiano, Eichenbaun and Evans (2005). For concreteness, we assume that the passive updating rule for those firms that can not optimally adjust prices is given by

$$\Gamma_t^i = \prod_{j=1}^{i}\left(1+\pi_{t+j-1}\right)^{\kappa}\left(1+\pi_{t+j}^*\right)^{1-\kappa}$$

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 to the extent to which indexation clauses are present in the economy or the

[7] Notice that we are assuming that the inflation target in any period may differ from steady-state inflation. Thus, under our formulation not only we eliminate the possibility of having a non-vertical Phillips curve in the long-run but also we are able to address the trend observed in inflation for the Chilean case over the 90s.

credibility of the target set by the authority (see Erceg, and Levin, 2003).[8] The Phillips curve under this updating rule is given by

$$\widehat{\pi}_t = \lambda\xi\widehat{mc}_t + \gamma_f E_t\widehat{\pi}_{t+1} + \gamma_b\widehat{\pi}_{t-1} + \zeta_t \tag{5}$$

where $\lambda = \frac{(1-\theta)(1-\theta\beta)}{\theta(1+\kappa\beta)}$, $\gamma_f = \frac{\beta}{1+\kappa\beta}$, $\gamma_b = \frac{\kappa}{1+\kappa\beta}$. The term ζ_t is a function of changes in the inflation target and it is given by $\zeta_t = \tau_1 E_t\Delta\pi^*_{t+1} + \tau_2\Delta\pi^*_t$, where $\tau_1 = \beta\gamma_b$, $\tau_2 = -\gamma_b$.[9].

2.2 Real marginal cost

2.2.1 Benchmark formulation

The benchmark formulation for the marginal cost corresponds to the case where the production function is Cobb-Douglas with two inputs, capital and labor. The nominal marginal cost that is obtained under the assumption of competitive factor markets is given by:

$$MC_t = \frac{1}{1-\alpha}\frac{W_t L_t}{Y_t} \tag{6}$$

where W_t corresponds to the nominal wage and L_t to labor. The linearized version of the real marginal cost is given by

$$\widehat{mc}_{t+i} = s_{t+i} - s \tag{7}$$

where s_t corresponds to the logarithm of the labor share, and where s is the logarithm of its steady-state (long-run) value.

2.2.2 Alternative specifications

We consider three alternative specifications for the underlying technology as in Gagnon and Khan (2005). These alternative specifications define different formulations for marginal costs.

Overhead labor First, we consider the inclusion of overhead labor in the production function. Let $\overline{L}$ be the (fixed) quantity of labor devoted to cover fixed cost. The production function in this case can be expressed as,

$$Y_t = K_t^{\alpha}\left(A_t\left(L_t - \overline{L}\right)\right)^{1-\alpha} \tag{8}$$

The nominal marginal cost consistent with this technology is given by the following expression:

$$MC_t = \frac{1}{1-\alpha}\frac{W_t L_t}{Y_t}\frac{\left(L_t - \overline{L}\right)}{L_t} \tag{9}$$

[8] If credibility of the monetary authority is low, permanent reductions in the inflation target could be initially perceived as transitory. In this case, private agents will learn about the true intentions of the policy makers gradually given rise to persistent inflation.

[9] Notice that in steady-state $\zeta_t = 0$.

Again, if capital can be freely allocated across firms, the marginal cost is independent from the scale of production –it depends only on the prices of factors– and, therefore, is the same for all firms. The log-deviation of real marginal costs from steady-state in this case is given by:

$$\widehat{mc}_{t+i} = s_{t+i} - s + \frac{\overline{L}}{L - \overline{L}} \widehat{l}_{t+i} \tag{10}$$

where $L > \overline{L}$ represents the steady-state level of employment for any particular firm. Therefore, under this specification the log-deviations of marginal costs do not only depend on the deviations of the unitary labor costs from steady-state but also on the deviations of employment. Thus, if the unitary labor cost (labor share) remains constant, an increase in employment leads to an increase in the marginal cost.

Constant Elasticity of Substitution We also consider a more general production function where we allow for a non-unitary elasticity of substitution between inputs. In particular we consider the following *CES* technology:

$$Y_t = \left[K_t^{\frac{\sigma-1}{\sigma}} + (A_t L_t)^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}} \tag{11}$$

where σ is the elasticity of substitution across inputs. Under this specification for the technology the nominal marginal cost is given by:

$$MC_t = \frac{W_t L_t}{Y_t \gamma_t} \tag{12}$$

where $\gamma_t = \frac{(A_t L_t)^{\frac{\sigma-1}{\sigma}}}{K_t^{\frac{\sigma-1}{\sigma}} + (A_t L_t)^{\frac{\sigma-1}{\sigma}}}$. As in the previous case, if capital is perfectly mobile across firms the marginal cost for all firms will be the same. Linearizing the marginal cost about the steady state we obtain the following expression:

$$\widehat{mc}_{t+i} = (s_{t+i} - s) + \omega \widehat{yk}_{t+i} \tag{13}$$

where $\widehat{yk}_{t+i}$ represents log-deviations of the output-capital ratio, $\frac{Y_t}{K_t}$, with respect to its steady state value. Parameter $\omega = \frac{1-\sigma}{\sigma} \left(\frac{1}{\mu s} - 1 \right)$ may be positive or negative depending on the degree of complementarity between inputs. If the degree of complementarity between capital and labor is high (a low σ) then decreases in the capital-output ratio are associated with increases in the marginal cost along the business cycle. Parameter μ corresponds to the steady-sate gross mark-up.

If the firm utilizes an imported input instead of capital in the production function (with a *CES* technology) then the real marginal cost can be expressed as follows

$$\widehat{mc}_{t+i} = (s_{t+i} - s) + \varphi \omega \left(\widehat{p}_{m,t+i} - \widehat{w}_{t+i} \right) \tag{14}$$

where $\widehat{p}_{m,t} - \widehat{w}_t$ corresponds to log-deviations of the relative price of foreign inputs with respect to the nominal wage, and where $\varphi = \sigma \mu s$.

Firm specific capital When capital is specific to firms, changes in factors' relative prices and changes in output level do not affect the amount of capital utilized in production by a particular firm. In this case, there would be a difference between specific marginal costs and the observable average marginal cost. Sbordone (2002) shows that under this circumstances it is possible to establish a relationship between firm specific marginal cost and average marginal cost.[10]

If we consider a production function with a *CES* technology the real marginal cost in period $t+i$ for firms that adjusted their prices in period t is given by:

$$\widehat{mc}_{t,t+i} = (s_{t+i} - s) + \omega\widehat{yk}_{t+i} - \left[\frac{\epsilon}{\sigma}\left(\frac{1}{\mu s} - 1\right)\right]\widehat{q}_t + \frac{\epsilon}{\sigma}\left(\frac{1}{\mu^* s} - 1\right)\sum_{j=1}^{i}\widehat{\pi}_{t+j} \tag{15}$$

Combining equations (3), (13) and (15) we obtain the following expression for the Phillips curve:

$$\widehat{\pi}_t = \Omega\,(s_t - s) + \Omega\omega\widehat{yk}_{t+i} + \beta E_t\,\{\widehat{\pi}_{t+1}\} \tag{16}$$

where $\Omega = \lambda\xi$, and where $\xi = \frac{s\sigma(\mu-1)}{s(\sigma(\mu-1)-\mu)+1}$ is a scaling parameter.

3 Estimation of the NKPC

3.1 Specification

We estimate two specifications for the Phillips curve. The baseline case corresponds to the standard NKPC without a backward-looking component (4). The following alternative orthogonality conditions were used:

$$E_t\left\{\left(\widehat{\pi}_t - \frac{(1-\theta)(1-\theta\beta)}{\theta}\xi\widehat{mc}_t - \beta\widehat{\pi}_{t+1}\right)\mathbf{z}_t\right\} = 0 \tag{a}$$

$$E_t\left\{\left(\widehat{\pi}_t - \frac{(1-\theta)^2}{\theta}\xi\widehat{mc}_t - \widehat{\pi}_{t+1}\right)\mathbf{z}_t\right\} = 0 \tag{b}$$

where $\xi = 1$ under the assumption that capital is freely mobile across firms, and $\xi \neq 1$ under the assumption that capital is firm specific.

All estimations where made using the Generalized Method of Moments (GMM). Vector $\mathbf{z}_t$ contains a set of instruments. The instruments list includes four lags of: the deviation of inflation from target, the deviation of real marginal cost from trend, and the output gap (lags from $t-2$ to $t-5$); two lags of the monetary policy interest rate (from $t-5$ to $t-6$); three lags of nominal wages growth relative to trend (from $t-4$ to $t-6$); and four lags of terms of trade deviations from trend (from $t-1$ to $t-4$). Notice that specification (b) normalizes β to 1.

[10] This assumption has helped reconciling the degree of inertia implied by the empirical estimation of the Phillips curve base on the Calvo model with the micro evidence regarding the frequency of price adjustments in the US.

For this hybrid specification we utilize the following orthogonality condition to estimate parameters θ, β and κ:

$$E_t\left\{\left(\widehat{\pi}_t-\frac{(1-\theta)(1-\theta\beta)}{\theta(1+\kappa\beta)}\xi\widehat{mc}_t-\frac{\beta}{1+\kappa\beta}\widehat{\pi}_{t+1}-\frac{\kappa}{1+\kappa\beta}\widehat{\pi}_{t-1}-\zeta_t\right)\mathbf{z}_t\right\}=0 \qquad \text{(c)}$$

where $\mathbf{z}_t$ is a vector of instruments similar to the one considered previously. However, in this case we include only three lags of the inflation deviation from target, the real marginal cost deviation from trend, and the output gap (from $t-3$ to $t-5$).

3.2 Data

We use the change in the Consumer Price Index (CPI) as our measure of inflation. There are several reasons why using the CPI instead of the GDP deflator —which is the most common measure of inflation used in empirical studies of the Phillips curve– is better in our case. First, the GDP deflator for the Chilean economy is measured with considerable noise. Second, the GDP deflator in a commodity-intensive economy like Chile is subject to strong variations due to changes in the terms of trade. Then, a significant fraction of the changes in the GDP deflator reflects changes in relative prices rather than persistent changes in the general level price. Finally, the inflation target set by the authority is defined in terms of CPI inflation. To avoid including regulated prices and prices that fluctuate significantly –whose dynamics are not well represented by the Phillips curve– we use a measure of *core* CPI inflation that removes those items from the CPI basket.

Figure 1 shows the evolution of CPI inflation and *core* CPI inflation over the period 1986-2004. The first characteristic of the inflationary process is that it exhibits a strong decreasing tendency over the period 1991-1999. This trend is the consequence of the disinflationary process that the monetary authority started in 1990 with the introduction of the first inflation target by the Central Bank.[11]

Another relevant feature of the Chilean inflationary process is the high volatility of quarterly CPI inflation. To cope with potential problems arising from this high volatility we decompose the inflation rate in two different components: irregular and non-irregular (trend plus cycle). After removing the irregular component, quarterly core inflation resembles closely yearly moving-average inflation (figure 2). We "de-trend" inflation by computing the difference between the non-irregular component and the interpolated inflation target set by the authority. This de-trending procedure is consistent with our theoretical framework, in which firms update their prices using the inflation target.

We consider four measures of real marginal costs consistent with the different specifications for the production technology presented above. The first is a measure of labor share –total labor income as a fraction of GDP– that excludes mining, fishing, energy,

[11] There have been two clear phases in the implementation of the inflation targeting regime in Chile. In the first phase, when gaining credibility was a key issue, the Central Bank set short-term horizon CPI inflation targets, and actively managed the exchange rate. In the second phase, that started in 1999, the Central Bank moved to a fully flexible exchange rate system with a stationary lon-run target for inflation (see Cespedes and Soto, 2005).

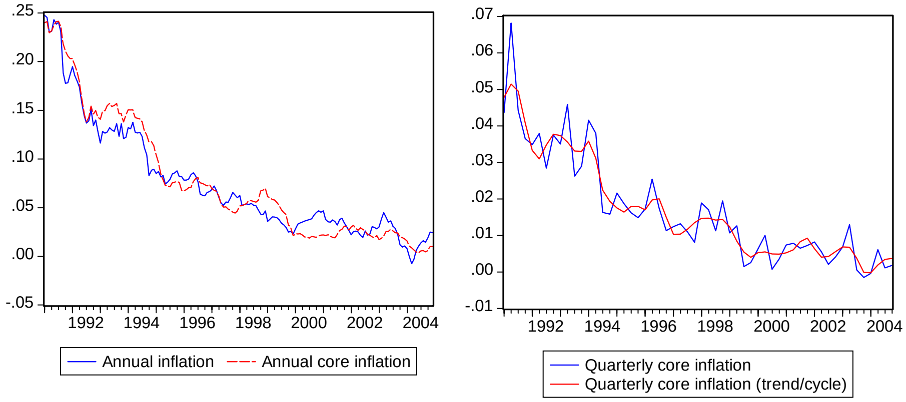

Figure 1: Evolution of CPI and core CPI inflation (left panel) and core inflation against its trend/cyclical component (right panel)

agriculture, and the public sector from both employment and output.[12] In the case of the technology with overhead labor we adjust marginal cost by adding log-deviation of employment from a deterministic linear trend.

When the technology is characterized by a CES production function–with labor and capital–, marginal costs is adjusted to consider the capital-output ratio. Finally, we use a measure of real marginal costs that allows the inclusion of imported inputs in the production function. In particular, we construct a measure of the ratio of the price of imported inputs over wages. Figure 2 we presents the evolution of our four measures of marginal cost together with the deviation of inflation from target. We observe that all of these measures of marginal cost and inflation present a slight decreasing trend during the nineties.

Figure 3 displays different components of the marginal cost. Notice that the labor share exhibits a trend over the 90s. This trend could be explained by changes in relative sizes of different sectors with different markups (as in McAdam and Willman, 2003), or by changes in the average markup in the all economy. It could also be reflecting developments in the labor market, such as permanent changes in the wage markups. We let for further work a more exhaustive investigation on the evolution of the marginal cost, in particular the labor share.

Figure 5 displays the correlation pattern of real marginal cost –labor share– and inflation. These two variables are positively and significantly correlated contemporaneously. Also, leads of real marginal costs are positively correlated with current inflation. This implies that current inflation is positively associated with current and future real marginal costs, as the theoretical framework would suggest. On the other hand, the output gap is a poor proxy for the inflationary pressures underlying the Phillips curve. In fact,

[12] GDP is measured at factor prices., i.e., we exclude indirect taxes. The reason to remove these sectors is that they are associated to commodities, regulated or non market-determined prices.

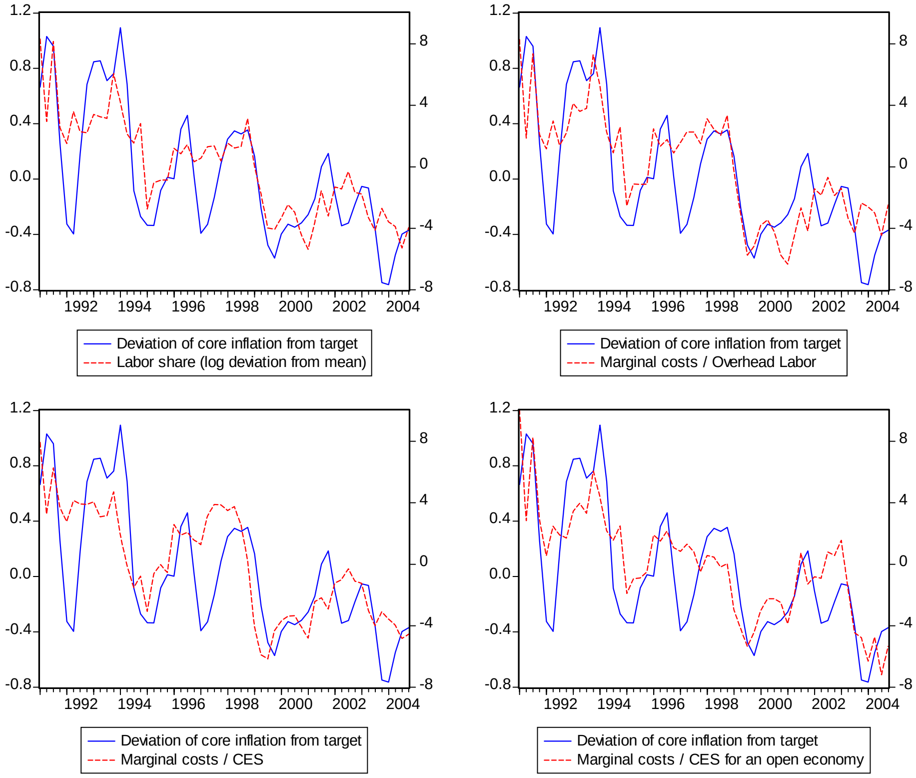

Figure 2: Evolution of our four different measures of marginal costs and inflation deviations from its target

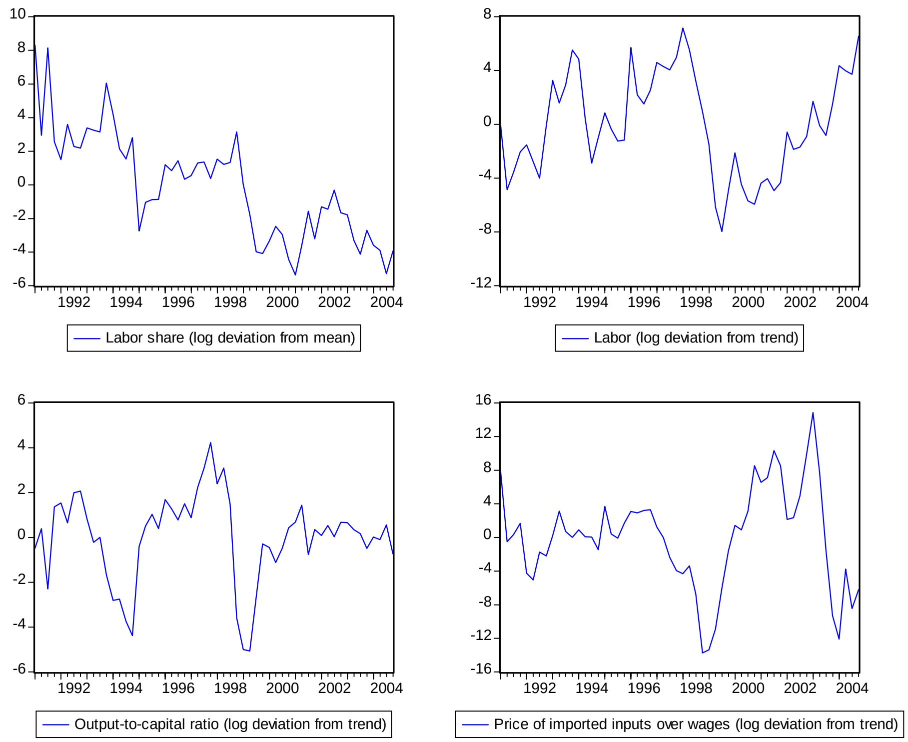

Figure 3: Marginal costs components

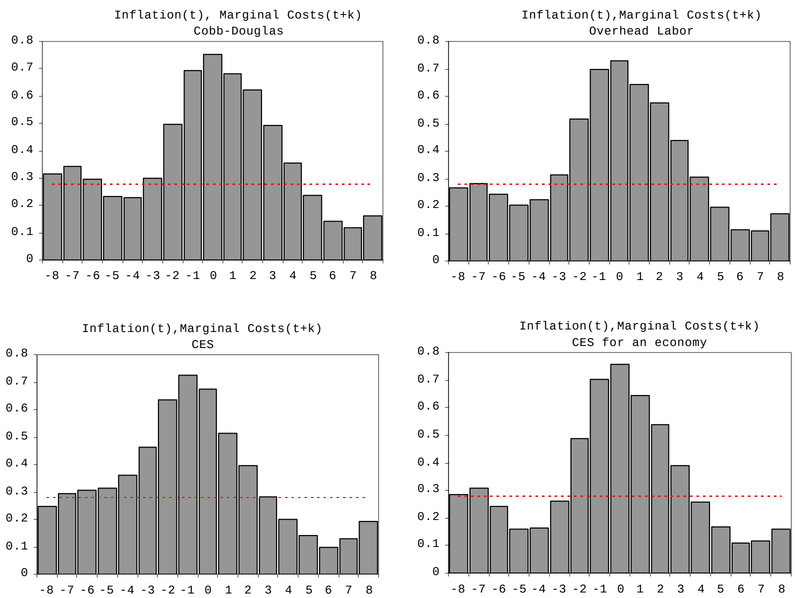

Figure 4: Dynamic cross correlations of Inflation(t) and Marginal Costs $(t+k)$ assuming four different technologies. The dotted lines are the approximate two standard error bounds.

as in GG, the correlation pattern between the output gap and the real marginal cost is negative.[13] Moreover, lags of the output gap are positively correlated with inflation which contradicts the theory.

3.3 Results

Tables 1 to 3 present the estimated values of parameters θ, β, and λ for the baseline Phillips curve (4) under the four alternative measures of marginal costs. For each case we report results under the two alternative normalizations (specifications (a) to (b)), and considering both the case when capital is freely mobile across firms, $\xi = 1$, and when capital is firms specific. In this case, the proportionality factor ξ is assume to be known with certainty and computed from the sample average of the labor share and the gross markup.[14] For the cases were we assume a CES technology we consider two alternative values for the elasticity of substitution across inputs: $\sigma = 0.5$ and $\sigma = 1.5$.

Together with the estimated values of the two structural parameters and λ, we also report the implied average number of quarters prices are not optimally adjusted, and the J-test for overidentified restrictions. According to this test the overidentified restrictions

[13] The output gap is computed using the HP filter.

[14] The proportionality factor is computed as $\overline{\xi} = \frac{\overline{s}\sigma(\overline{\mu}-1)}{\overline{s}(\sigma(\overline{\mu}-1)-\overline{\mu})+1}$, where $\overline{s}$ is the sample average labor share, and where $\overline{\mu}$ is the sample average gross markup. The value for these two magnitudes were $\overline{s}$ =0.5 and $\overline{\mu}$ =1.1.

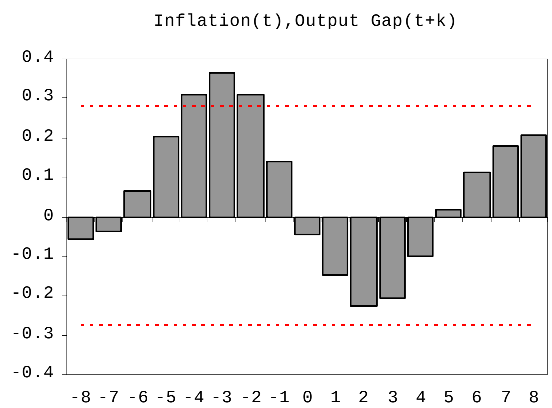

Figure 5: Dynamic cross correlations of Inflation(t) and the Output gap ($t+k$). The dotted lines are the approximate two standard error bounds.

are satisfied for all specifications of the model.

Notice first that coefficient λ is statistically significant in all specifications. This implies that marginal cost is in fact relevant to explain the inflationary process, as emphasized by the NKPC formulation. In general, the estimation of parameter θ is robust to the two normalizations (a)-(b). For the case were capital is assumed to be freely mobile across firms, the estimated value for this parameter lies in the range of $0.85 - 0.91$. This implies an estimated average duration of prices in the range of 6.7 up to 11 quarters, approximately. Therefore, our figures imply a similar price rigidity for Chile than for the Euro area as estimated by GGL, and coincident with the price rigidity in the U.S. as estimated by same authors. When we assume that capital is firm specific the estimated value of θ decreases. Under this assumption the point estimate of this parameter lies in the range $0.55 - 0.80$, which implies durations for price stickiness in the range of 2.3 up to 5 quarters. These figures are also consistent with those for the U.S. estimated by GGL when assuming firm specific capital (Table 5).

The estimated discount factor β is somewhat not very precise: it lies between 0.93 and 1.03 . Again, these results are in line with those reported by GGL for the Euro area and the U.S.

Results for the estimated hybrid model are presented in Table 4 . Again, we report the estimated values of parameters θ, β, and λ under four different specification for the marginal cost, and assuming alternatively that capital is freely mobile and firm specific. We also report the estimated value of parameter κ, which measures the extent to which firms index their prices to past inflation.

Parameter κ is consistently estimated in the range $0.65 - 0.76$ implying that within the sample period firms weighted more past inflation rather than the announced inflation targets to passively adjust their prices. The estimated backward-looking component γ_b is statistically significant in all specifications and it is about 0.4. This figure is slightly smaller than the one reported by Agenor and Bayraktar (2003) for Chile in their study of the inflation dynamics in middle income countries. These authors estimate a non-structural Phillips curve that includes both a backward and a forward-looking component, and found that the backward-looking component is about 0.52. Unlike our case, Agenor and Bayraktar use several lag of the output gap (up to 3 for Chile) as the driving force for inflation. Our estimated value for coefficient γ_b implies a stronger role for the backward-looking component of inflation in the Chilean case than the estimated by GGL for the Euro area and the U.S. However, it is quiet similar –or even smaller– than the estimates for some of the countries of the Europe area (Spain and Italy).[15]

The estimated values for parameter θ under the hybrid specification do not significantly change from the baseline case. The same holds true for the discount factor β which, again, is somewhat low. As in the baseline case, marginal cost specification with firm-specific capital tend to give lower values for θ implying shorter average duration of price stickiness. In general, all our estimates are in line with the range of values found in the literature, while we find duration of price stickiness in line with the results found for France and Italy, we obtain estimates of the forward-looking component similar to those found for the US. Finally notice that for all cases the overidentifying restrictions are satisfied.

Despite the important role of the backward-looking component in our estimations, our results could be biased against this component if additional lags of inflation enter directly in the true Phillips curve.[16] Therefore, we follow GG and GGL, and perform a robustness exercises to address the importance of the backward-looking component. In particular, we add additional lags of inflation to the hybrid model. Table 6 presents the results for the specification for marginal cost that assumes firm-specific capital (results assuming capital mobility are similar). Additional lags of inflation turn out to be non-significant. Neither is the sum of the three additional lag of inflation. Therefore, our results are robust to the inclusion of additional lags of inflation.

In order to check the relevance of the instrument set used in our regressions we test null hypothesis that the coefficients on all the instruments are jointly zero in the first stage of the estimation.[17] Table 7 reports the F-statistic, the associated p-value and the adjusted R^2 from the first stage regressions. As can be seen, the null hypothesis that the instruments are jointly irrelevant is soundly rejected in all cases, and the adjusted R^2 is over 0.5 in most of them. Therefore, we do not find evidence of weak instruments in our estimations.

Finally, it is important to notice that some studies on the Phillips curve for OECD countries use raw data on inflation and marginal cost, despite clear trends in those series (see for example Galí and López-Salido (2001) for estimations of the Spanish Phillips curve). Our results presented so far were obtained using de-trended data (inflation minus inflation target and detrended marginal costs). To analyze whether these results

[15] See Table 5.

[16] As noted by Rudd and Whelan (2001), if the instrument set includes variables that cause inflation directly but are not included in the hybrid specification of the Phillips curve, the estimation of the model may be biased in favor of the forward looking component.

[17] Recent literature has questioned inference using GMM methods in the presence of weak instruments (e.g., Stock, Wright and Yogo, 2002).

were driven by the de-trending methodology we re-estimate the hybrid model using undetrended data. Results do not significantly change. The new estimates for parameter λ, γ_f, γ_b are 0.04, 0.57 and 0.43, respectively.[18]

3.4 Actual versus fundamental inflation

Following GG and Sbordone (1999), we assess the "goodness-of-fit" of our estimations by comparing the extent to which the inflation rate and the deviation of the inflation rate from the inflation target implied by our model lines up against actual data. The model-based measure of inflation or fundamental inflation, as termed by GG, can be obtained by iterating the pricing equation (5),

$$\widehat{\pi}_t = \delta_1 \widehat{\pi}_{t-1} + \frac{1}{\delta_2 \gamma_f} E_t \sum_{i=0}^{\infty} \left(\frac{1}{\delta_2} \right)^i \left[\lambda \xi \widehat{mc}_{t+i} + \tau_1 \Delta \pi^*_{t+1+i} + \tau_2 \Delta \pi^*_{t+i} \right] \tag{17}$$

where $\delta_1 \leqslant 1$ and $\delta_2 \geqslant 1$ are the stable and the unstable root associated with the stationary solution to the difference equation given by (5) and π^*_t is the inflation target. In the pure forward looking specification, the lagged term in (17) disappears, thus, fundamental inflation reduces to a discounted stream of expected future real marginal costs.

As argued by GG, we do not observe future marginal costs –nor variations in the inflation target. However, under certain assumptions we can construct an estimate of the right-hand side as follows. Let, $X_t = [\widehat{mc}_t, \widehat{mc}_{t-1}, \ldots, \widehat{mc}_{t+1-q}]$ for some finite value of q. We can use an unrestricted autoregressive process of order q to forecast future inflation using the fact that,

$$E_t\{\widehat{mc}_{t+i}\} = \mathbf{A}^i X_t$$

where $\mathbf{A}$ is the companion matrix of the $AR(1)$ representation of X_t. Analogously to marginal costs, expected future values of the inflation target can be obtained by estimating a first-order autoregressive process, $E_t \left\{ \pi^*_{t+i} - \pi^* \right\} = \phi^i \left(\pi^*_t - \pi^* \right)$ where π^* corresponds to the steady-state (long run) inflation target. Accordingly, we can re-write (17) as,

$$\begin{aligned} \pi_t &= \pi^*_t + \delta_1 \widehat{\pi}_{t-1} + \frac{\lambda \xi}{\delta_2 \gamma_f} \left(\mathbf{I} - \delta_2^{-1} \mathbf{A} \right)^{-1} X_t \\ &\quad - \frac{\tau_2}{\delta_2 \gamma_f} \Delta \pi^*_t + \frac{\left(\tau_1 + \delta_2^{-1} \tau_2 \right) (1 - \phi)\, \delta_2}{(\delta_2 - \phi)\, \delta_2 \gamma_f} \left(\pi^*_t - \pi^* \right) \end{aligned} \tag{18}$$

Figures 6 to 9 depict the actual inflation and our measure of fundamental inflation together with the observed difference between inflation and the inflation target versus the deviations of inflation from the target implied by our estimated model. Overall, under all specifications, our measure of fundamental inflation tracks actual inflation and its deviations from the inflation target quite well. The lowest mean square error between actual and fundamental inflation is obtain when assuming a Cobb-Douglas technology (see Table 8).

[18] Structural estimate with raw data are not directly comparable with the ones obtained with de-trended data as the underlying model implies removing the inflation target from inflation.

In order to explore the adequacy and discriminate among the four specifications, we formally compare the predictive performance of each model using the approach proposed by Diebold and Mariano (1995). Particularly, we compare the mean squared forecast error between models i and j for a sequence of n-steps forecast,

$$d_{t+\tau} = \left(\hat{\pi}_{t+\tau} - \hat{\pi}^{i}_{t+\tau}\right)^2 - \left(\hat{\pi}_{t+\tau} - \hat{\pi}^{j}_{t+\tau}\right)^2 \tag{19}$$

where $\tau = 1, \ldots n$ is the forecast horizon, $\hat{\pi}_{t+\tau}$ is the observed deviation of inflation from the target and, $\hat{\pi}^{i}_{t+\tau}$ ($\hat{\pi}^{j}_{t+\tau}$) is the deviation predicted by model i (j). Our test is based on the observed sample mean, $d = n^{-1} \sum_{t=T_0}^{T_1} d_{t+\tau}$ and the Diebold and Mariano statistic given by,

$$DM = \frac{d}{\sqrt{V_d}} \tag{20}$$

where $V_d = \gamma_0 + 2\sum_{v=1}^{q}\left(1 - \frac{v}{q+1}\right)\gamma_v$ is the consistent estimate of the variance of d. With this at hand, we can test the null hypothesis that model j has no predictive superiority than model i (i.e., $H_0 : E\left(d_{t+\tau}\right) < 0$). As a first exercise we compare the forecast performance of each of the four specifications against the three alternative models for $n = 1$. The results presented in Table 9 suggest that no model has a significant superiority against the others, but on average the specification using a Cobb-Douglas technology seems to outperform the alternative three models. Furthermore, we compare the performance of each model against the forecast from an AR(1) specification. We evaluate out-of-sample forecasting accuracy using three alternative forecasting horizons. We obtain for the relevant horizon starting from the first quarter of 2000 until the end of the sample (i.e., one-step, two-steps and four-steps forecasts). The evidence presented in Table 10 suggests that an AR(1) model does not dominate any of our estimated specifications in terms of forecasting accuracy.[19]

3.5 Stability of the hybrid Phillips Curve

One issue that has not been formally analyzed in the literature is the temporal stability of the hybrid Phillips curve estimated parameters.[20] This concern is particularly important in an emerging country like Chile that has experienced a significant decrease of the inflation rate accompanied with an increase in the credibility of the inflation targeting regime over the last fifteen years (Céspedes and Soto, 2005). The lower inflation rate might have been reflected in a decrease in the average length between price adjustments (a lower value of D) while the gain of credibility by the monetary authority could have translated in an increased number of firms updating their prices according to the inflation target (lower values for κ).

[19] These results remain if we increase the forecast length to eight quarters.

[20] The paper of Jondeau and Le Gihan (2005) test the stability of their Phillips curve estimates but examine only the reduced (linear) form parameters and use Wald-type tests which have some important drawbacks as they point out.

In order to examine the stability of the parameters and the occurrence of a break point, we consider a predictive test for structural change with unknown breakpoint developed in the papers of Ghysels and Hall (1990), Ghysels et. al. (1997) and Guay (2003). This test consists on estimating the parameter vector for the first subsample and then evaluating the moment conditions for the second subsample at these parameter values.[21]

In our particular case, this approach has several advantages over alternative approaches like the Wald-type tests proposed in the work of Andrews (1993) and Andrews and Ploberger (1994). Firstly, we only use first subsample estimates of the parameters which allow us to test the presence of a break even when the second subsample contains a few observations and parameter estimates are not feasible. Indeed, it is a common drawback of Wald-type tests that they cannot be applied to detect structural instability at the end of the sample. Secondly, we do not set a priori orthogonality conditions equal to zero in the second subsample thus, we avoid rejecting stability when in fact the parameters were stable but there was certain type of misspecification (e.g., omitted variables).

Table 11 presents the estimated predictive tests (supremum $\sup PR$, average $avgPR$ and exponential $\exp PR$) 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 (Sovell, 1996). We only report the PR-type statistic that tests parameter variation. The results show that for the four specifications of marginal costs, we cannot reject the existence of a breakpoint around 2000, which is close to the date that the inflation target reached its stationary annual level of 3%. The results show that for the four specifications of marginal costs, we cannot reject the existence of a breakpoint. Furthermore, the PR_1 and the PR_2 tests estimate consistently the date of the breakpoint around 2000 which is close to the date that the inflation target reached its stationary annual value of 3%. We also report the estimated values of the parameter κ and the duration of price stickiness before the breakpoint date. It is noteworthy that in most cases the first subsample value of κ is larger than the estimated parameter using the whole sample. This result suggests that after the break the degree to which firms update their prices according to the inflation target has increased. Regarding price stickiness, we find that the estimated duration of price stickiness is smaller before the breakpoint. In other words, the frequency of optimal price adjustment has decreased over time.

4 Conclusion and directions for further work

In this paper we estimate a NKPC for Chile, using quarterly data for the period 1990:1-2004:4. Our results using the Generalized Method of Moments (GMM) tend to support the NKPC. The evidence shows that the backward-looking coefficient in a hybrid specification of the Phillips curve is about 0.45. This figure is larger than the corresponding one for the Euro area and the U.S. as estimated by GG and GGL. In general, different specifications for the marginal costs lead to similar estimates. The estimated Calvo co-

[21] A brief description of the test can be found in Appendix A.

efficient that captures the degree of price rigidity lies, in the baseline case, in the range of 0.85 to 0.91 indicating that prices remain unchanged on average for about 6.7 to 11 quarters. When firm specific capital is considered then the Calvo coefficient falls to a range 0.55 to 0.80, implying average price durations in the range of 2.3 up to 5 quarters. These results do not significantly change when considering the hybrid specification for the Phillips curve.

Regarding parameter stability, our results support the hypothesis of the existence of a structural break in the NKPC, which occurred when the inflation target converged to its long-run level (around 2000). Moreover, our evidence supports the idea that the inflationary process became more forward looking in recent years, which is also consistent with an increased credibility in the inflation target.

There are still some issues that need to be address. First, it its necessary to analyze more in detail how robust are our estimates to weak identification. As it has been put forward by Ma (2002), Kurmann (2004) and Nason and Smith (2005), GMM estimation of the structural parameters of the NKPC may be inaccurate because of weak instruments. Therefore, it is necessary to test whether our instruments are not only valid but also relevant. Second it is necessary to perform more research on the determinant of the evolution marginal cost. We showed that labor unit cost and marginal cost in Chile have exhibit a slight trend over the 90s. Factors behind that trend may include changes in the optimal markups by firms, composition effects associated to changes in the relative share of different sectors in the economy, and changes in wages markups. As emphasized by the NKPC, understanding the dynamics of marginal costs is key to understand the dynamics of inflation.

A Variable definitions

- **Inflation rate**: Cyclical component of the quarterly variation of the core CPI (IPCX1), $\pi_t = \frac{P_t}{P_{t-1}} - 1$.

- **Output gap**: Log-deviation of output from its long-run value, $\widehat{y}_t = 100 * (y_t - \bar{y}_t)$. The long-run value of output, $\bar{y}_t$, is approximated by using a quadratic trend.

- **Inflation target**: Quarterly linear interpolation of the annual inflation targets, $\widehat{\pi}_t^*$.

- **Labor share**: Ratio of the nominal labor compensation and the nominal output, $S_t = \frac{W_t L_t}{P_t Y_t}$. The log-deviation of the labor share is: $\widehat{s}_t = 100 * (s_t - \bar{s}_t)$ where $s_t = \log S_t$, and where the long run value of this variable, $\bar{s}_t$, is approximated by using a quadratic trend. Nominal output is computed by using real GDP and the core CPI index used to compute inflation.

- **Output-Capital ratio**: We compute the log-deviation of the output-capital ratio from its long-run trend by using quarterly data on capital stock and GDP. The long-run value for this ratio is computed by using a quadratic trend.

- **Relative price of imports**: We utilize an imports price index (IVUM) that is computed by the Central Bank, and a nominal wages index by INE (Instituto Nacional de Estadísticas).

B Predictive tests for structural change

This appendix briefly describes the predictive tests for structural change with unknown breakpoint presented in the papers of Ghysels and Hall (1990), Ghysels et. al. (1997) and Guay (2003). Recall that the GMM estimator is based on a set of moment conditions,

$$E\left[f\left(x_t, \theta\right)\right] = 0$$

where $f(\cdot)$ is a $q \times 1$ vector of continuous differentiable functions of a vector of data (x_t) and, the model parameters (θ). When this moment conditions do not hold throughout the whole sample $(t = 1, \ldots, T)$, the model is said to be structurally unstable.

If we presume that there is a break at some date $[\pi T]$ for $\pi \in (0, 1)$, Ghysels and Hall (1990) propose to estimate the model parameters θ using the observations in the first subsample $T_1(\pi) = \{1, 2, \ldots, [\pi T]\}$ and then, evaluate the moment conditions for the observations in the second subsample, $T_2(\pi) = \{[\pi T] + 1, \ldots, T\}$, at these parameter values. The idea behind the predictive tests is to examine whether parameter estimates of one subsample can be used to predict over the other subsample. Particularly, they propose to test whether these estimated moment conditions are approximately zero (i.e., $H_0 : E\left[f\left(x_t, \theta_1\right)\right] = 0$ for $T_1(\pi)$ and $T_2(\pi)$ while $H_1 : E\left[f\left(x_t, \theta_1\right)\right] = 0$ for $T_1(\pi)$ but $E\left[f\left(x_t, \theta_1\right)\right] \neq 0$ for $T_2(\pi)$).

The predictive tests used in this paper are based on the following statistics,

$$\begin{aligned} PR_1(\pi) &= \left[(T - [\pi T])^{-\frac{1}{2}} \sum_{t=[\pi T]+1}^{T} f\left(x_t, \theta_1\right)\right]' V_2^{-1}(\pi) \left[(T - [\pi T])^{-\frac{1}{2}} \sum_{t=[\pi T]+1}^{T} f\left(x_t, \theta_1\right)\right] \\ PR_2(\pi) &= \frac{\pi}{1-\pi} \left[T^{-\frac{1}{2}} \sum_{t=[\pi T]+1}^{T} f\left(x_t, \theta_1\right)\right]' S_T^{-\frac{1}{2}} P_2(\pi)\, S_T^{-\frac{1}{2}} \left[T^{-\frac{1}{2}} \sum_{t=[\pi T]+1}^{T} f\left(x_t, \theta_1\right)\right] \end{aligned}$$

where $V_2(\pi) = S_2(\pi) + \frac{1-\pi}{\pi} F_2(\pi) \left[F_1(\pi)' S_1^{-1}(\pi) F_1(\pi)\right]^{-1} F_2(\pi)'$ and $P_2(\pi) = S_1^{-\frac{1}{2}} F_2(\pi) \times \left[F_1(\pi)' S_1^{-1}(\pi) F_1(\pi)\right] F_2(\pi)' S_1^{-\frac{1}{2}}$ with $S_1(\pi)$ being the covariance estimator involving data from the first subsample and, $S_2(\pi)$ the estimator using the second subsample data, while the matrices F_1 and F_2 denote the jacobian of the moment conditions evaluated in the first and second subsample, respectively. Sovell (1996) shows that the test $PR_1(\pi)$ is divided into a test of structural change for the vector of parameters and a test of stability of the overidentifying restrictions. The statistic $PR_2(\pi)$ accounts for the test for parameter variation.

Statistics for optimal predictive tests with unknown breakpoint can be obtained by computing the average supremum, average and exponential form,

$$
\begin{aligned}
\sup PR_i &= \sup_{\pi\in\Pi} PR_i(\pi) \\
avgPR_i &= \int_{\Pi} PR_i(\pi) dJ(\pi) \\
\exp PR_i &= \log \int_{\Pi} \exp\left[0.5 \times PR_i(\pi)\right] dJ(\pi)
\end{aligned}
$$

where $i = 1, 2$ depending on the PR statistic employed, $J(\pi)$ is the probability function specified for π and is assumed to be the uniform distribution (see Andrews and Ploberger, 1995).

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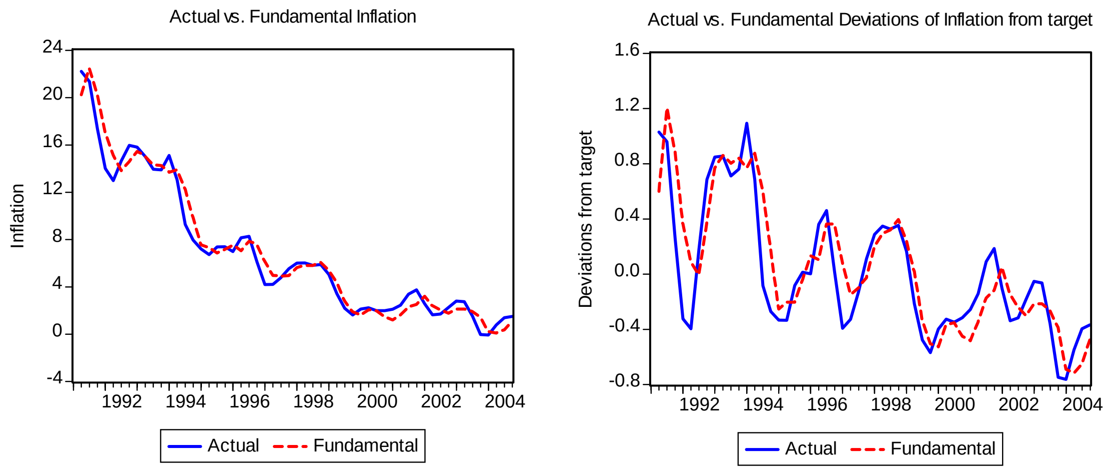

Figure 6: Actual versus fundamental inflation implied by the hybrid Phillips curve assuming a Cobb-Douglas technology

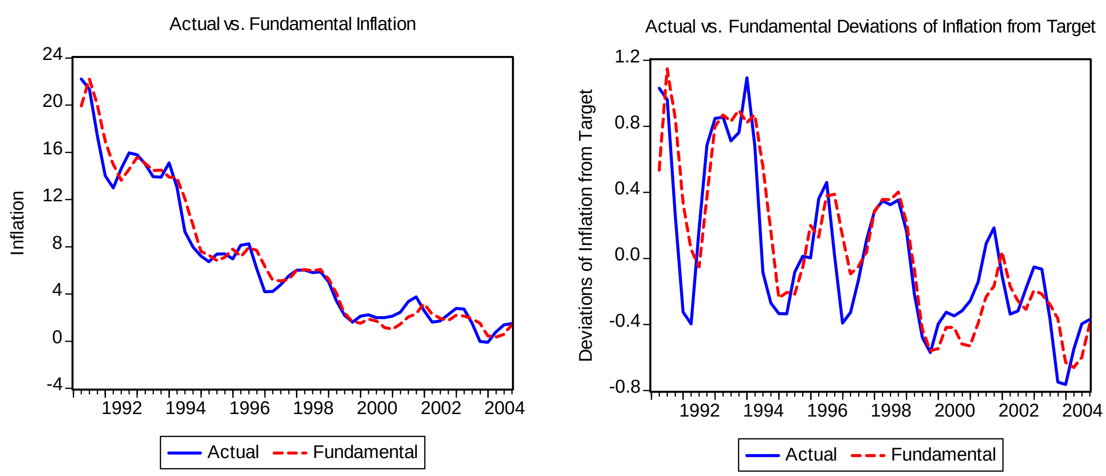

Figure 7: Actual versus fundamental inflation implied by the hybrid Phillips curve assuming a Cobb-Douglas technology with overhead labor

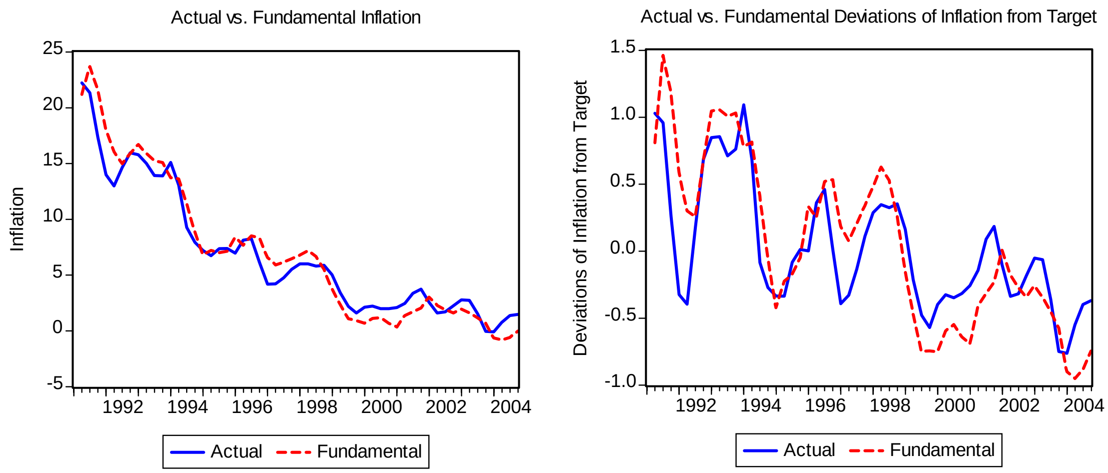

Figure 8: Actual versus fundamental inflation implied by the hybrid Phillips curve assuming a CES technology with $\sigma = 0.5$

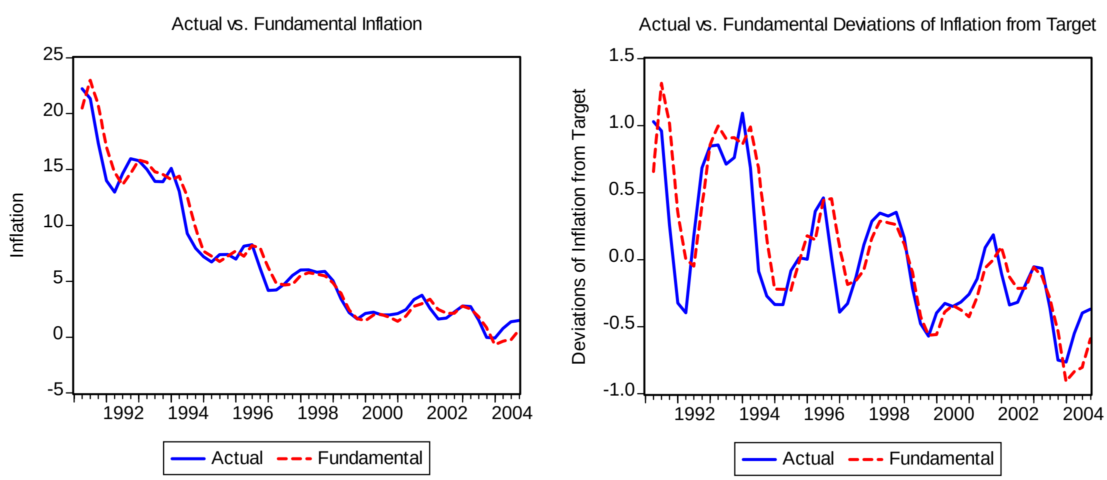

Figure 9: Actual versus fundamental inflation implied by the hybrid Phillips curve assuming a CES technology for an open economy with $\sigma = 0.5$

Table 1: Structural estimates

Model	θ	β	λ	D	$J-Test$ $(p-value)$
Cobb-Douglas					
$\xi = 1$					
(a)	0.8882 (0.015)	0.9780 (0.045)	0.0165 (0.007)	8.950 (1.226)	4.168 (0.939)
(b)	0.8872 (0.016)	1.0000	0.0143 (0.004)	8.866 (1.281)	4.268 (0.961)
$\xi \neq 1$					
(a)	0.6242 (0.010)	0.9780 (0.045)	0.2345 (0.103)	2.661 (0.401)	4.168 (0.939)
(b)	0.6392 (0.043)	1.0000	0.2036 (0.062)	2.771 (0.332)	4.268 (0.961)
Overhead Labor					
$\xi = 1$					
(a)	0.8944 (0.012)	0.9891 (0.046)	0.0135 (0.005)	9.475 (1.131)	4.163 (0.939)
(b)	0.8936 (0.013)	1.0000	0.0126 (0.003)	9.400 (1.156)	4.260 (0.962)
$\xi \neq 1$					
(a)	0.6494 (0.012)	0.9891 (0.046)	0.1930 (0.082)	2.852 (0.412)	4.163 (0.939)
(b)	0.6564 (0.035)	1.0000	0.1798 (0.046)	2.910 (0.300)	4.260 (0.962)

Note: Standard errors based on a Newey-West covariance matrix robust to serial correlation up to 12 lags within parenthesis. Column D reports the estimated duration of price stickiness, and J the Hansen test of the overidentifying restrictions (below in parenthesis we report the p-value). The set of instruments includes three lags of inflation deviation from its target, real marginal costs and, detrended output, and, three lags of detrended terms of trade.

Table 2: Structural estimates

Model	θ	β	λ	D	$J-Test$ $(p{-}value)$
CES, $\sigma = 0.5$					
$\xi = 1$					
(a)	0.9065 (0.017)	0.9276 (0.097)	0.0163 (0.011)	10.701 (1.927)	3.915 (0.951)
(b)	0.9125 (0.023)	1.0000	0.0083 (0.004)	11.435 (3.092)	4.116 (0.966)
$\xi \neq 1$					
(a)	0.5915 (0.093)	0.9276 (0.097)	0.3115 (0.215)	2.448 (0.558)	3.915 (0.951)
(b)	0.6727 (0.074)	1.0000	0.1592 (0.090)	3.055 (0.696)	4.116 (0.966)
CES, $\sigma = 1.5$					
$\xi = 1$					
(a)	0.8571 (0.013)	0.9565 (0.050)	0.0299 (0.005)	7.002 (0.639)	4.920 (0.896)
(b)	0.8504 (0.011)	1.0000	0.0263 (0.004)	6.686 (0.474)	4.925 (0.934)
$\xi \neq 1$					
(a)	0.6455 (0.019)	0.9565 (0.050)	0.2100 (0.037)	2.821 (0.154)	4.920 (0.896)
(b)	0.6532 (0.021)	1.0000	0.1841 (0.028)	2.883 (0.175)	4.925 (0.934)

Note: Standard errors based on a Newey-West covariance matrix robust to serial correlation up to 12 lags within parenthesis. Column D reports the estimated duration of price stickiness, and J the Hansen test of the overidentifying restrictions (below in parenthesis we report the p-value). The set of instruments includes three lags of inflation deviation from its target, real marginal costs and, detrended output, and, three lags of detrended terms of trade.

Table 3: Structural estimates

Model	θ	β	λ	D	$J-Test$ ($p-value$)
CES for an open economy, $\sigma = 0.5$					
$\xi = 1$					
(a)	0.8874 (0.010)	0.9466 (0.039)	0.0202 (0.004)	8.882 (0.801)	4.228 (0.980)
(b)	0.8819 (0.011)	1.000	0.0158 (0.003)	2.2364 (0.149)	4.269 (0.961)
$\xi \neq 1$					
(a)	0.5528 (0.030)	0.9466 (0.039)	0.3855 (0.080)	2.236 (0.149)	4.228 (0.980)
(b)	0.5817 (0.033)	1.000	0.3006 (0.065)	2.391 (0.189)	4.269 (0.961)
CES for an open economy, $\sigma = 1.5$					
$\xi = 1$					
(a)	0.9072 (0.024)	1.0322 (0.038)	0.0064 (0.006)	10.781 (2.855)	4.110 (0.9422)
(b)	0.9082 (0.020)	1.000	0.0092 (0.004)	10.899 (2.401)	4.239 (0.962)
$\xi \neq 1$					
(a)	0.7964 (0.069)	1.0322 (0.038)	0.0454 (0.043)	4.913 (1.688)	4.110 (0.9422)
(b)	0.7756 (0.045)	1.000	0.0648 (0.030)	4.458 (0.901)	4.316 (0.987)

Note: Standard errors based on a Newey-West covariance matrix robust to serial correlation up to 12 lags within parenthesis. Column D reports the estimated duration of price stickiness, and J the Hansen test of the overidentifying restrictions (below in parenthesis we report the p-value). The set of instruments includes three lags of inflation deviation from its target, real marginal costs and, detrended output, and, three lags of detrended terms of trade.

Table 4: Structural estimates hybrid NKPC

Model	θ	β	κ	λ	γ_f	γ_b	τ_1	τ_2	D	$J-Test$ $(p-value)$
Cobb-Douglas										
$\xi = 1$										
(c)	0.8940 (0.056)	0.9919 (0.182)	0.6600 (0.114)	0.0081 (0.004)	0.5994 (0.035)	0.3988 (0.020)	0.3956 (0.083)	-0.3988 (0.020)	9.435 (5.027)	4.544 (0.919)
$\xi \neq 1$										
(c)	0.6505 (0.036)	0.9927 (0.183)	0.6606 (0.114)	0.1148 (0.066)	0.5995 (0.035)	0.3989 (0.020)	0.3956 (0.083)	-0.3989 (0.020)	2.862 (0.299)	4.545 (0.919)
Overhead Labor										
$\xi = 1$										
(c)	0.9008 (0.049)	0.9622 (0.167)	0.6498 (0.095)	0.0090 (0.003)	0.5920 (0.037)	0.3998 (0.0175)	0.3847 (0.073)	-0.3998 (0.0175)	10.087 (4.959)	4.435 (0.9554)
$\xi \neq 1$										
(c)	0.6455 (0.027)	0.9622 (0.167)	0.6498 (0.095)	0.1280 (0.052)	0.5920 (0.037)	0.3998 (0.0175)	0.3847 (0.073)	-0.3998 (0.0175)	2.821 (0.219)	4.435 (0.9554)
CES, $\sigma = 0.5$										
$\xi = 1$										
(c)	0.8923 (0.058)	0.9510 (0.196)	0.7180 (0.116)	0.0108 (0.003)	0.5651 (0.040)	0.4266 (0.019)	0.4058 (0.086)	-0.4266 (0.019)	9.288 (5.015)	4.607 (0.970)
$\xi \neq 1$										
(c)	0.5693 (0.026)	0.9511 (0.196)	0.7181 (0.116)	0.2061 (0.075)	0.5651 (0.040)	0.4266 (0.019)	0.4058 (0.086)	-0.4266 (0.019)	2.321 (0.140)	4.607 (0.970)
CES, $\sigma = 1.5$										
$\xi = 1$										
(c)	0.8841 (0.052)	1.0075 (0.088)	0.6742 (0.075)	0.0085 (0.005)	0.5910 (0.024)	0.4015 (0.020)	0.4045 (0.044)	-0.4015 (0.020)	8.632 (3.861)	5.640 (0.958)
$\xi \neq 1$										
(c)	0.7273 (0.074)	1.0075 (0.088)	0.6742 (0.075)	0.0597 (0.035)	0.5910 (0.024)	0.4015 (0.020)	0.4045 (0.044)	-0.4015 (0.020)	3.667 (0.998)	5.640 (0.958)
CES for an open economy, $\sigma = 0.5$										
$\xi = 1$										
(c)	0.8962 (0.067)	0.9740 (0.194)	0.7615 (0.136)	0.0084 (0.003)	0.5592 (0.028)	0.4372 (0.017)	0.4258 (0.091)	-0.4372 (0.017)	9.638 (6.199)	4.4897 (0.993)
$\xi \neq 1$										
(c)	0.5987 (0.038)	0.9740 (0.194)	0.7611 (0.136)	0.1605 (0.061)	0.5592 (0.028)	0.4372 (0.017)	0.4258 (0.091)	-0.4372 (0.017)	2.492 (0.241)	4.4897 (0.993)
CES for an open economy, $\sigma = 1.5$										
$\xi = 1$										
(c)	0.9144 (0.056)	0.9912 (0.155)	0.7097 (0.108)	0.0051 (0.002)	0.5818 (0.027)	0.4166 (0.017)	0.4129 (0.073)	-0.4166 (0.017)	11.683 (7.662)	4.341 (0.976)
$\xi \neq 1$										
(c)	0.7841 (0.045)	0.9912 (0.155)	0.7097 (0.108)	0.0051 (0.002)	0.5818 (0.027)	0.4166 (0.017)	0.4129 (0.073)	-0.4166 (0.017)	4.6315 (0.964)	4.341 (0.976)

Note: Standard errors based on a Newey-West covariance matrix robust to serial correlation up to 12 lags within parenthesis. Column D reports the estimated duration of price stickiness, and J the Hansen test of the overidentifying restrictions (below in parenthesis we report the p-value). The set of instruments includes five lags of inflation deviation from its target, four lags of real marginal costs, detrended output and three lags of detrended terms of trade.

Table 5: Structural estimates: An international comparison

Model	*Chile*	*US**	*Euro Area**	*Spain***	France†	Italy†	Canada‡
Baseline Model							
β	0.946	0.924	0.914	0.759	-	-	-
θ	0.553	0.627	0.771	0.743	-	-	-
D	2.2	2.7	4.4	3.9	-	-	-
Hybrid Model							
θ	0.651	0.569	0.787	0.671	0.710	0.570	0.640
γ_f	0.600	0.599	0.689	0.487	0.653	0.409	0.574
γ_b	0.400	0.364	0.272	0.488	0.300	0.516	0.349
D	2.9	2.3	4.7	3.0	3.5	2.3	2.8

* Galí, Gertler and López-Salido (2001).
** Galí and López-Salido (2000).
† Benigno and López-Salido (2002).
† Gagnon and Khan (2004).
Note: to be consistent we re-calculate θ and D France and Italy assuming $\mu = 1.1$

Table 6: Hybrid Model: Further inflation lags

Model	λ	γ_f	γ_b	ϕ_1	ϕ_2	ϕ_3	$\sum \phi_i$
Cobb-Douglas							
	0.0299 (0.055)	0.5601 (0.068)	0.4530 (0.064)	-0.1357 (0.047)	0.1039 (0.033)	0.0199 (0.018)	-0.0118 (0.029)
Overhead Labor							
	0.0320 (0.036)	0.5072 (0.050)	0.4925 (0.046)	-0.0482 (0.0487)	0.0193 (0.038)	0.0256 (0.015)	-0.0032 (0.026)
CES, $\sigma = 0.5$							
	0.1569 (0.131)	0.5134 (0.082)	0.5811 (0.073)	-0.2029 (0.113)	0.1109 (0.087)	-0.0487 (0.402)	-0.1408 (0.0570)
CES for an open economy, $\sigma = 0.5$							
	0.1149 (0.202)	0.5805 (0.107)	0.4638 (0.067)	-0.1640 (0.151)	0.0458 (0.097)	-0.0228 (0.046)	-0.1410 (0.091)

Note: Standard errors based on a Newey-West covariance matrix robust to serial correlation up to 12 lags within parenthesis. The set of instruments includes three lags of inflation deviation from its target, three lags of real marginal costs and detrended output and, three lags of detrended terms of trade.

Table 7: Results from F-test from first-stage regressions

	Variable			
	mc_t		π_{t+1}	
	$F-stat$ $(p-value)$	$Adj.\ R^2$	$F-stat$ $(p-value)$	$Adj.\ R^2$
Cobb-Douglas				
Non-hybrid	14.188 (0.00)	0.81	5.637 (0.00)	0.63
Hybrid	17.839 (0.00)	0.86	3.300 (0.00)	0.53
Overhead Labor				
Non-hybrid	13.273 (0.00)	0.80	5.754 (0.00)	0.63
Hybrid	15.04 (0.00)	0.85	2.377 (0.00)	0.48
CES				
Non-hybrid	9.180 (0.00)	0.74	4.132 (0.00)	0.56
Hybrid	11.032 (0.00)	0.82	4.320 (0.00)	0.65
CES open economy				
Non-hybrid	13.06 (0.00)	0.80	5.529 (0.00)	0.62
Hybrid	8.935 (0.00)	0.79	2.028 (0.00)	0.47

Table 8: Sum of squared residuals

Model	*SSR*
Cobb-Douglas	3.610
Overhead Labor	3.626
CES, $\sigma = 0.5$	6.372
CES open economy, $\sigma = 0.5$	3.885
AR(1)	3.676

Note: The statistic is calculated using residuals from actual inflation and inflation implied by the model.

Table 9: Diebold and Mariano Test of predictive accuracy

model i \model j	*Cobb-Douglas*	*Overhead Labor*	*CES*	*CES open economy*
Cobb-Douglas	–	−0.431 (0.33)	−0.365 (0.36)	−0.393 (0.35)
Overhead Labor	0.431 (0.67)	–	−0.352 (0.36)	−0.321 (0.38)
CES, $\sigma = 0.5$	0.365 (0.64)	0.352 (0.64)	–	0.052 (0.53)
CES open economy, $\sigma = 0.5$	0.393 (0.65)	0.321 (0.62)	−0.052 (0.47)	–

Note: p-values for the Diebold-Mariano test of equal forecast accuracy are reponented within parenthesis. The test was performed using one-step forecasts beginning at 2000q1.

Table 10: Diebold and Mariano Test of predictive accuracy

$i \diagdown j$	*AR(1)*		
	one-step forecast	2-steps forecast	4-steps forecast
Cobb-Douglas	0.113 (0.54)	0.114 (0.54)	0.467 (0.68)
Overhead Labor	0.361 (0.64)	0.288 (0.61)	0.454 (0.68)
CES, $\sigma = 0.5$	0.371 (0.65)	0.470 (0.68)	0.484 (0.69)
CES open economy, $\sigma = 0.5$	0.398 (0.65)	0.358 (0.64)	0.360 (0.64)

Note: p-values for the Diebold-Mariano test of equal forecast accuracy are reporented within parenthesis. The test was performed using forecasts beginning at 2000q1.

Table 11: Predictive tests for structural change

	Sup PR	Avg PR	Exp PR
Cobb-Douglas			
PR *test*	295.29*	59.68*	144.5*
Estimated beak point date 2001:4			
$\kappa_{t1} =$	0.847 (0.133)	$D_{t1} =$	2.049 (0.289)
Overhead labor			
PR *test*	196.29*	52.89*	95.01*
Estimated beakpoint date 2000:3			
$\kappa_{t1} =$	0.925 (0.091)	$D_{t1} =$	1.808 (0.161)
CES			
PR *test*	66.75*	32.50*	30.24*
Estimated beakpoint date 2000:3			
$\kappa_{t1} =$	0.997 (0.090)	$D_{t1} =$	1.715 (0.118)
CES for an open economy			
PR *test*	42.33*	7.77	18.12**
Estimated beakpoint date 1999:3			
$\kappa_{t1} =$	0.969 (0.093)	$D_{t1} =$	1.820 (0.258)

Note: The table reports predictive tests for the null hypothesis of structural stability along with the estimated breakpoint date and the value of κ and the estimated duration of price stickiness before the break. The critical values were calculated using Monte Carlo simulations.
*Indicates that the statistic is significant at the 1% level
.**Indicates that the statistic is significant at the 5% level.

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