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Contracting Models of the Phillips Curve Empirical Estimates for Middle-Income Countries

By

Pierre-Richard Agénor and Nihal Bayraktar

Centre for Growth and Business Cycle Research, Economic Studies,
University of Manchester, Manchester, M13 9PL, UK

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Empirical Estimates for Middle-Income Countries

Pierre-Richard Agénor* and Nihal Bayraktar**

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Abstract

This paper provides empirical estimates of contracting models of the Phillips curve for eight middle-income developing countries (Chile, Colombia, Korea, Malaysia, Mexico, Morocco, Tunisia, and Turkey). Following an analytical review, a variety of models with one and more leads and lags are estimated using two-step GMM techniques. Nested and non-nested tests are used to select a specification for each country, and in-sample predictive capacity and stability are analyzed. Higher-dimension models tend to perform better than parsimonious models with one lead and one lag. Except for Colombia and Korea, backward-looking behavior has a relatively larger impact on inflation dynamics. World oil prices and relative input prices have a limited effect, whereas borrowing costs are significant for Korea and Mexico.

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*Hallsworth Professor of International Macroeconomics and Development Economics, University of Manchester, and Centre for Growth and Business Cycle Research; **Assistant Professor, Pennsylvania State University. This paper is a thoroughly revised version of an earlier study circulated as Policy Research Working Paper No. 3139 of the World Bank. We are grateful to George Bratsiotis and Kyriakos Neanidis for helpful comments on a previous draft. The views expressed in this paper are our own.

Contents

1 Introduction

The degree of wage and price stickiness plays an important role in the transmission of macroeconomic shocks. In traditional backward-looking Phillips curves, inertia in the wage- and price-contracting process is generally captured by introducing measures of past inflation. By contrast, in models of overlapping contracts with forward-looking agents, inflation is represented as a function of its expected future realizations, based on all available information about the state of the economy. Indeed, several of these models, such as the Taylor-Calvo staggered contracts approach (see Taylor (1979, 1980) and Calvo (1983)), and the quadratic price adjustment cost approach of Rotemberg (1982), have a common formulation that boils down to an expectations-augmented (or New Keynesian) Phillips curve, with current prices depending on future prices.

However, pure forward-looking models have had difficulties explaining the high degree of persistence in inflation and in inflation's response to monetary policy shocks. Figure 1 shows that inflation persistence, as measured by estimated coefficients from AR(1) and AR(2) processes, is indeed quite high in both industrial and middle-income developing countries. For their sample of 18 industrial countries, Wang and Wen (2007) found a mean value of 0.77 using an AR(1) model, and 0.84 for an AR(2) model. For the 14 developing countries shown in Figure 1, the numbers are 0.84 (with a standard deviation of 0.11) and 0.93 (with a standard deviation of 0.04), respectively.

Another difficulty with early New Keynesian models (a consequence perhaps of the fact that they were first applied to the United States) is the inadequate account of open-economy considerations. In general, there are at least two channels through which openness may influence firms' price-setting decisions via their impact on marginal production costs. First, demand for domestic products may change relative to those produced abroad. Thus, imperfectly competitive firms may take into account the price they set relative to the prices set by other firms, not only at home but also abroad. Second, prices of imported intermediate goods may change relative to other inputs engaged in the production process. Firms can therefore substitute imported intermediate goods for domestic labor in production, and changes in the price of these goods relative to wages can affect marginal costs.

Both of these issues, as well as a number of others, have been addressed in recent studies of the New Keynesian Phillips curve, which include Fuhrer (1997), Roberts (1995, 2005), Galí and Gertler (1999), Guerrieri (2002), Sbordone (2002, 2007), Guay and Pelgrin (2004), Gagnon and Khan (2005), Nason and Smith (2005), Jondeau and Le Bihan (2005), Rudd and Whelan (2005, 2007), and Zhang, Osborn, and Kim (2007). Studies that have specifically focused on open-economy considerations include Balakrishnan and López-Salido (2002), Banerjee and Batini (2004), Bardsen Jansen, and Nymoen (2004), Batini, Jackson, and Nickell (2005), Genberg and Pauwels (2005), Benigno and López-Salido (2006), and Leith and Malley (2007).

Summarizing this literature is a daunting task, but based on the review of Olafsson (2006) and the discussion in Rudd and Whelan (2007), it is fair to conclude that, in general, the evidence is mixed. Disagreements relate to a variety of issues—microfoundations for price-setting behavior, the role of openness, the measurement of underlying variables (especially the output gap), and estimation techniques. For instance, Guay and Pelgrin (2004) show that the results in Galí and Gertler (1999) for the United States are not invariant to the choice of instruments. Results based on a continuously updated GMM technique lead to a rejection of the pure forward-looking specification of the Phillips curve. Nason and Smith (2005), using a variety of estimation techniques for a hybrid equation, find limited evidence of forward-looking behavior for the case of Canada, the United Kingdom, and the United States—in contrast to some previous studies. As argued by Soderstrom, Soderlind, and Vedrin (2005), backward-looking behavior appears to be important for explaining not only inflation inertia but also the volatility and persistence of output and interest rates. More nuanced results are provided by Zhang, Osborn, and Kim (2007), who highlight changes in the structural stability of the Keynesian Phillips curve for the United States.

In addition, almost all of the existing studies focus on the United States and European countries. Given the conflicting nature of some of the results, and potentially important structural differences between developed and developing countries (regarding, in particular, the degree of openness and the role of credit markets), it is difficult to draw general inference from them regarding the behavior of inflation in the latter group of countries.

This paper attempts to fill this gap by providing empirical evidence on contracting models of inflation for eight middle-income developing countries: Chile, Colombia, Korea, Malaysia, Mexico, Morocco, Tunisia, and Turkey. These countries represent a fairly diverse experience in terms of inflation, with countries like Morocco and Tunisia experiencing relatively low and stable inflation during the sample period, whereas Chile and Turkey (as well as Korea following the Asia crisis in the late 1990s) experienced higher and more persistent episodes of inflation. Section II examines analytically various types of backward- and forward-looking models of inflation. It also highlights the role of openness, factor substitution, and the impact of borrowing costs on marginal production costs. Section III proposes several alternative empirical specifications for testing these models in a small open middle-income country. Section IV presents our econometric methodology, which is based on two-step GMM estimation. Section V presents estimation results, as well as the results of non-nested tests and predictive ability. Section VI concludes by discussing some research perspectives.

2 Models of Inflation Dynamics

This section provides a brief analytical overview of various specifications of the Phillips curve. We begin by considering pure, forward-looking (closed-economy) models, in which inflation displays no persistence. We next examine models in which inertia is introduced. Finally, we consider the specification of marginal costs when working capital needs lead to short-term borrowing and discuss open-economy considerations.

2.1 Pure Forward-Looking Models

Much recent research on the dynamics of price adjustment in closed-economy Keynesian models dwells on the staggered contracts models of Taylor (1979, 1980), Calvo (1983), and the costly adjustment model of Rotemberg (1982). Staggered contracts models à la Taylor are based on the assumption that wages are set in nominal terms at discrete periods of time in an asynchronous fashion (because they are set by different agents at different points in time) and, as a result, contracts overlap. Agents are assumed to contract a wage that reflects their anticipations of future price and output levels for the expected duration of the contract. These models typically assume that prices are a constant markup over wages and focus on the persistence induced in the aggregate price (average wage) level due to the asynchronous and overlapping nature of wage contracts.[1]

To illustrate Calvo's model of price adjustment, consider an economy characterized by a continuum of identical monopolistic competitive firms indexed by $i \in (0, 1)$, each of which producing a nonstorable differentiated good. The production function of each firm is linear in labor, N_t^i, so that

$$Y_t^i = A_t N_t^i, \tag{1}$$

where A_t denotes average productivity. If labor is paid a common nominal wage W_t, the nominal marginal cost of production—the ratio of the wage rate to the marginal product of labor—is identical across firms and given by (in logs)

$$mc_t^n = w_t - a_t, \tag{2}$$

[1]Note that this approach does not postulate that formal contracts are actually written, but rather that nominal prices (wages) are preset for some period of time.

where $x_t = \ln X_t$, with $X_t = A_t, W_t$.

In a closed economy, with no government, household consumption equals production. Each firm is assumed to face an isoelastic demand curve for its product of the form

$$Y_t^i = (\frac{P_t^i}{P_t})^{-\upsilon} Y_t, \tag{3}$$

where the parameter $\upsilon > 1$ is the Dixit-Stiglitz elasticity of substitution between differentiated goods (or the common price elasticity of demand facing each firm), $P_t = \left\{\int_0^1 P_t^{i(1-\upsilon)} di\right\}^{1/(1-\upsilon)}$ the aggregate price level, and $Y_t = \left\{\int_0^1 Y_t^{i(\upsilon-1)/\upsilon} di\right\}^{\upsilon/(\upsilon-1)}$ aggregate output.[2]

Real profits of the ith firm are given by, using (1) and (3),

$$\Phi_t^i = \frac{(P_t^i Y_t^i - W_t N_t^i)}{P_t} = (\frac{P_t^i}{P_t} - \frac{W_t}{P_t A_t})(\frac{P_t^i}{P_t})^{-\upsilon} Y_t. \tag{4}$$

In the absence of pricing frictions, the ith firm will choose its price to maximize its real profits. From (4), the first-order condition is

$$\frac{\partial \Phi_t^i}{P_t^i} = \frac{\upsilon - 1}{\upsilon}(\frac{P_t^i}{P_t})^{1-\upsilon} Y_t - \frac{W_t}{P_t A_t}(\frac{P_t^i}{P_t})^{-\upsilon} Y_t = 0,$$

which can be simplified to give the (log) equilibrium price:

$$\tilde{p}_t = \mu + mc_t^n, \tag{5}$$

where $\mu = \ln[\upsilon/(\upsilon - 1)]$ is the (log) mark-up.[3] Thus, in the absence of frictions, all firms set their price as a fixed mark-up over marginal cost. As υ increases, the mark-up falls, and in the limit of $\upsilon \to \infty$, $\mu = 0$.[4]

Calvo's random price adjustment involves assuming that in each period a random fraction $1 - \theta$ of firms reset their price (regardless of how much time

[2] Because we are only analyzing firms' price-setting decisions, we can abstract from the optimization problem underlying individuals' decisions to allocate total consumption across time periods.

[3] The index i is dropped in (5) because the equilibrium price is the same for all firms.

[4] The term μ is thus a measure of the steady-state distortion arising from monopolistic competition.

has elapsed since the last update), whereas all other firms keep their prices unchanged. The hazard rate (the probability that any given price posted at time t will be adjusted at $t+1$) is thus constant and the average duration of nominal price stickiness is $1/(1-\theta)$. The evolution of the (log) price level, p_t, is given by

$$p_t = \theta p_{t-1} + (1-\theta)\tilde{p}_t, \tag{6}$$

where $\tilde{p}_t$ is now the optimal price chosen by those who can reset their prices.

At time t, price-adjusting firms can set their price to maximize a discounted sum of current and future profits. Using (4), the maximization problem can be specified as

$$\max_{P_t^i} \left\{ \sum_{j=0}^{\infty} (\theta\beta)^j (\frac{\Lambda_{t+j}}{\Lambda_t}) \left[(\frac{P_{t+j}^i}{P_{t+j}} - \frac{W_{t+j}}{P_{t+j}A_{t+j}})(\frac{P_{t+j}^i}{P_{t+j}})^{-\upsilon} Y_{t+j} \right] \right\}, \tag{7}$$

where β is the firm's discount factor and Λ_{t+j}/Λ_t the intertemporal marginal rate of substitution of the representative household between t and $t+j$.[5] Assuming that firms take the processes driving A_t and W_t as given, it can readily be established that a firm's optimal reset price is determined by

$$\tilde{p}_t = \mu + (1-\theta\beta) \sum_{j=0}^{\infty} (\theta\beta)^j E_t mc_{t+j}^n, \tag{8}$$

where mc_{t+j}^n is nominal marginal cost at time $t+j$. Intuitively, firms take into account the fact that their prices will likely be fixed over some period of time by setting their price equal to a weighted average of expected future nominal marginal costs. The parameter θ (which is inversely related to the expected duration of the currently determined price) affects how fast the weights decline over time. If $\theta = 0$, all prices are flexible, and all firms would set their price as described in (5).

[5] Because firms distribute all profits to households at the end of each period, the discount factor depends also on the intertemporal marginal rate of substitution of the representative household.

As shown for instance by Yun (1996), Sbordone (2002), or Lawless and Whelan (2007), equations (6) and (8) can be combined to yield, in a neighborhood of a zero-inflation steady state, a New Keynesian Phillips curve (NKPC) of the form

$$\pi_t = (1-\theta)(\tilde{p}_t - p_{t-1}) = \beta E_t \pi_{t+1} + \frac{(1-\theta)(1-\theta\beta)}{\theta}(mc_t + \mu), \qquad (9)$$

which relates inflation, $\pi_t = p_t - p_{t-1}$, to next period's expected inflation rate and the current deviation of real marginal cost $mc_t = mc_t^n - p_t$ from its frictionless optimal level, μ.[6] The higher the degree of price rigidity, as measured by θ, the less sensitive is inflation with respect to movements in marginal costs. This result can easily be generalized to alternative assumptions about the production technology; for instance, with a Cobb-Douglas production function of the form $Y_t^i = A_t N_t^{i(1-\alpha)}$, where $\alpha \in (0,1)$, the coefficient of the second term on the left-hand side of (9) would simply be multiplied by $(1-\alpha)[1+\alpha(\upsilon-1)]$, as shown by Galí, Gertler, and López-Salido (2001).

Because, from (1), $a_t = y_t - n_t$, the real (average) marginal cost $mc_t^n - p_t$ is also equal to $w_t - a_t - p_t = w_t + n_t - (y_t + p_t)$, that is, the (log) labor income share, or unit labor costs.[7]

Note that in this setting, where production is linear in labor and firms hire workers in a common economy-wide labor market, all firms face the same marginal cost, which is independent of its level of output. Consequently, the firm's marginal cost is equal to the average marginal cost, which equals the labor share. By contrast, as shown by Gagnon and Khan (2005), if the

[6] In the steady state, as implied by (5) the real marginal cost corresponds to the inverse of the mark-up (in log terms $-\mu$), because then the nominal marginal cost and output are constant and so all firms charge the same price.

[7] As shown by Gagnon and Khan (2005), if the production technology is Cobb-Douglas (or CES) with overhead labor costs, and if capital cannot be reallocated across firms, the relationship between real marginal costs and inflation will also depend on a "strategic complementarity" parameter, which depends essentially on the elasticity of firms' real marginal cost with respect to their own output. This parameter affects estimates of average duration of nominal price stickiness. However, we do not perform structural estimation in this paper.

production technology is Cobb-Douglas (or CES) with overhead labor costs, and if capital cannot be instantly reallocated across firms, real marginal costs will depend not only on unit labor costs but also on a "strategic complementarity" parameter, which depends essentially on the elasticity of firms' real marginal cost with respect to their own output—which is now different from zero. This parameter affects also estimates of the average duration of nominal price stickiness.

Abstracting from these complications, and under further restrictions on technology, and with a constant stock of capital, it can be shown that deviations in the labor share from its steady-state (and thus average real marginal costs), $\hat{s}_t$, are proportional to the gap between output and its potential level, that is, the output gap $\hat{y}_t$ (see Rotemberg and Woodford (1997)). With this assumption, the dynamics of inflation are given by

$$\pi_t = \beta E_t \pi_{t+1} + \gamma \hat{y}_t, \tag{10}$$

where $\gamma = (\chi + \sigma^{-1})(1-\theta)(1-\theta\beta)/\theta$, with σ denoting the elasticity of intertemporal substitution and χ the Frisch elasticity of labor supply with respect to wages (leaving constant the marginal utility of consumption).[8] Thus, the longer prices are held fixed (on average), the less responsive is inflation to cyclical fluctuations in output.

The main difference between Taylor-type contracts and Calvo's approach is that the individual firm's price-setting decision is derived from an explicit optimization problem. Nevertheless, as shown by Roberts (1995), and as discussed by Clarida, Galí, and Gertler (1999), both of these models can be formulated in a form similar to (10). The quadratic price adjustment model developed by Rotemberg (1982), in which firms are assumed to minimize the total costs of changing prices, generates also a similar relation—with γ now related to the magnitude of price adjustment costs (see Roberts (1995)).

[8] See, for instance, Kozicki and Tinsley (2002) for a derivation of (10).

2.2 Models with Inflation Inertia

Various criticisms have been addressed to the standard NKPC derived from the Taylor-Calvo-Rotemberg approach. Andersen (1998), for instance, has argued that focusing on price staggering, instead of wage staggering, does matter for output and price dynamics (namely, the degree of persistence of shocks), in contrast to some of the literature in the Taylor tradition where it has been implicitly assumed that there is no qualitative difference between the two cases. Most importantly for our purpose here, it has been pointed out that the approaches underlying the standard NKPC implies no persistence to the inflation rate (see Rudd and Whelan (2007)). All these models have the implication that a credible disinflation program can be implemented at no output cost—a somewhat implausible result, given the international experience for both industrial and developing countries.[9]

Several approaches have been proposed to generate inflation persistence in the NKPC. In Blanchard and Galí (2007), for instance, the existence of real wage rigidities implies that any change in the output gap, even purely transitory in nature, will have persistent effects on inflation—although these results appear to be valid only in log-linearized setting (see Ascari and Merkl (2007)). In Ireland (2004), nonseparability across consumption and real money holdings in household utility implies that real cash balances enter in the Phillips curve. In addition, if money demand depends on past inflation (as a measure of the rate of return on real assets), or if interest rates are set through a policy rule with a backward-looking component, inflation persistence may also emerge. In Karanassou and Snower (2007), the output gap is endogenized by relating it—somewhat arbitrarily—to real money balances; as a result, shocks to money supply generate persistent movements in inflation. Yet another approach to explaining price stickiness has been to abandon al-

[9] Not also there are differences among these models; as discussed for instance by Musy (2006), an *unexpected* disinflation in Taylor's model does entail a real cost, whereas it does not in Calvo's model.

together "time-dependent" models, in which firms set their prices for fixed periods of time, and to focus instead on "state-contingent" adjustment rules, in which firms change price when underlying determinants (such as demand and costs) reach some pre-specified upper or lower bounds.[10] In what follows, however, we will focus on two other approaches that have received widespread consideration in the empirical literature. The first involves introducing "rule-of-thumb" price setters, whereas the second involves an alternative definition of wage contracts.

2.2.1 The Galí-Gertler Hybrid Model

Galí and Gertler (1999), and subsequently Steinsson (2003), derived a hybrid Phillips curve from a price adjustment model à la Calvo and the additional assumption that although a mass of firms of measure $1-\theta$ is allowed to reset their prices, only a fraction $1-\kappa$ of them, where $\kappa \in (0,1)$, does so optimally (that is, in a forward-looking manner); the remaining fraction set their price according to a rule of thumb—perhaps as a result of information processing costs. Specifically, a fraction κ of price-resetting firms chooses the (log) price p_t^B according to the backward-looking rule

$$p_t^B = p_{t-1}^* + \pi_{t-1},$$

where p_{t-1}^* is the average set price in $t-1$, across both backward- and forward-looking firms, which depends therefore on both $\tilde{p}_{t-1}$ and p_{t-1}^B.

As derived by Galí and Gertler (1999, pp. 209-11), the hybrid equation takes the form

$$\pi_t = (\frac{\kappa}{\phi})\pi_{t-1} + (\frac{\beta\theta}{\phi})E_t\pi_{t+1} + \frac{(1-\kappa)\gamma\theta}{\phi}\hat{y}_t, \tag{11}$$

where $\phi = \theta + \kappa[1-\theta(1-\beta)]$. This equation corresponds to (10) if $\kappa = 0$.[11] A similar result would obtain if instead firms that are unable to reset their

[10] See Dotsey, King, and Wolman (1999), Romer (2000), and Bakhshi, Khan, and Rudolf (2007) for a discussion.

[11] Calvo, Celasun, and Kumhoff (2003) provide a generalization of Calvo's staggered pricing model that generates inflation inertia.

prices optimally do so by indexing them to last period's inflation, as follows:

$$p_t^B = p_{t-1} + \theta(p_{t-1} - p_{t-2}), \tag{12}$$

where, to ensure a vertical Phillips curve in the long run, the degree of indexation is set equal to θ (see, for instance, Liu (2006)).

A common criticism of the Galí-Gertler hybrid NKPC, and its variants involving backward-looking indexation as in (12), is that it dwells on an *ad hoc* behavioral assumption by some agents. However, as shown by Kozicki and Tinsley (2002), a virtually identical specification, which includes a term in lagged inflation but also an additional lead term, $E_t\pi_{t+2}$, can arise from optimizing behavior if the dynamics of pricing frictions are more complex than in the pure forward-looking version. Specifically, they show that in such conditions the NKPC takes the form

$$\pi_t = \alpha_1\hat{y}_t + \alpha_2\beta E_t\pi_{t+1} - \alpha_3\beta^2 E_t\pi_{t+2} + \alpha_3\pi_{t-1}, \tag{13}$$

where β is a discount factor.[12] Sheedy (2007) derives a similar result. He shows that, under the (plausible) assumption that newer prices are stickier than older prices, the hazard function of the Calvo model would be upward-sloping, rather than constant. Essentially, a hazard function that increases with time implies that if there has been inflation in the recent past, so that older prices are lower than more recent prices, the average price will automatically increase when prices are reoptimized—even if the new prices are no higher than the average current price. There is therefore inflation inertia and the Phillips curve takes a form similar to (13).

2.2.2 Fuhrer-Moore Contracts

Fuhrer and Moore (1995) developed a price formation equation that can also generate inflation inertia. Their model can be summarized as follows. Suppose that agents negotiate nominal wage contracts that remain in effect for

[12] If β is close to unity, Kozicki and Tinsley (2002) show that α_2 is also close to unity, and this equation can be estimated in the form $\pi_t = a_1\hat{y}_t + a_2E_t\pi_{t+1} + a_3(\pi_{t-1} - E_t\pi_{t+2}) + \varepsilon_t$,with the restriction $a_2 + a_3 = 1$.

four quarters. Unlike Taylor (1980), however, there is no fixed markup from wages to prices. This difference is essential, because it allows a meaningful distinction between prices and wages. The aggregate log price index in quarter t, p_t, is a weighted average of the log of contract prices, x_{t-i}, that were negotiated in the current and the previous three quarters and are still in effect. The weights, δ_i, are the proportions of the outstanding contracts that were negotiated in quarter $t-i$,

$$p_t = \sum_{i=0}^{3} \delta_i x_{t-i}, \tag{14}$$

where $\delta_i \geq 0$ and $\sum \delta_i = 1$. Fuhrer and Moore assume that the distribution of contract prices can be characterized by a downward-sloping linear function of contract length,

$$\delta_i = 0.25 + (1.5 - i)s,$$

with $0 < s \leq 1/6$ and $i = 0, ...3$. This distribution depends on a single slope parameter, s, and it is invertible. When $s = 0$, it is the rectangular distribution of Taylor (1980), and when $s = 1/6$, it is the triangular distribution.

Let υ_t be the index of real contract prices that were negotiated on the contracts currently in effect,

$$\upsilon_t = \sum_{i=0}^{3} \delta_i (x_{t-i} - p_{t-i}), \tag{15}$$

where δ_i is the fraction of wage contracts negotiated in period $t-i$ that are still in effect at period t.

Agents set nominal contract prices so that the current real contract price equals the average real contract price index expected to prevail over the life of the contract, adjusted for expected excess demand, measured by the output gap, $\hat{y}_t$:

$$x_t - p_t = \sum_{i=0}^{3} \delta_i E_t(\upsilon_{t+i} + \gamma \hat{y}_{t+i}) + \varepsilon_t, \tag{16}$$

where ε_t is an error term. Substituting (15) in (16) yields the "relative" (or real) version of Taylor's (1980) contracting equation:

$$x_t - p_t = \sum_{i=0}^{3} \beta_i (x_{t-i} - p_{t-i}) + \sum_{i=0}^{3} \beta_i E_t (x_{t+i} - p_{t+i})$$

$$+\gamma^* \sum_{i=0}^{3} \delta_i \, E_t(y_{t+i}) + \xi_t,$$

where

$$\beta_i = \frac{\Sigma_j \delta_j \delta_{i+j}}{1 - \Sigma_j \delta_j^2}, \qquad \gamma^* = \frac{\gamma}{1 - \Sigma_j \delta_j^2}$$

Letting $\pi_t = p_t - p_{t-1}$, the Phillips curve derived from the model is a two-sided curve defined as

$$\pi_t = \delta(L)\delta(L^{-1})[\pi_t - \gamma g^{-1}(L)\hat{y}_t],$$

where $\delta(L) = \delta_0 + \delta_1 L + \delta_2 L^2 ...$ is the lag polynomial that describes the distribution of price contracts in the model.

In the Fuhrer-Moore model, agents in their contracting decisions care about the *relative* real contract price in effect during the life of their contracts. They therefore compare the current real contract price with an average of the real contract prices that were negotiated in the recent past and those that are expected to be negotiated in the near future; the weights in the average measure the extent to which the past and future contracts overlap the current one. When output is expected to be high, the current real contract price is high relative to the real contract prices on overlapping contracts.[13] In contrast, the Taylor (1980) specification assumes that agents care about relative *nominal* contract wages (and prices) in effect during the life of their contracts.

[13] Note that γ measures the impact of the output gap on the log real contract price, not on inflation or on the price index. The inflation rate is related to the real contract price via a complex lag/lead polynomial.

The attractive feature of the Fuhrer-Moore contracting specification is that it helps to explain the high degree of persistence in inflation that is typically found in the data (as discussed in the introduction), whereas the conventional Taylor specification does not.[14] While the Taylor specification can be shown to imply that prices depend symmetrically on past and expected future prices, thus imparting significant inertia to the price *level*, it implies that the inflation *rate* is highly flexible—that is, it can jump in response to news. In contrast, the Fuhrer-Moore relative contracting specification implies that inflation depends symmetrically on past and expected future inflation, thus imparting significant inertia to *both* inflation and the price level. In addition, the relative contracting model, because it implies a link between the *inflation rate* and excess demand, can account for a positive correlation between inflation and the output gap; the Taylor model, by contrast, links the *price level* and excess demand, and is thus not able to do so.

With two periods, the Fuhrer-Moore contracting equation is

$$x_t - p_t = \frac{1}{2}[x_{t-1} - p_{t-1} + E_t(x_{t+1} - p_{t+1})] + \gamma \hat{y}_t.$$

If prices are a simple average of the nominal contract wage negotiated at t and $t-1$,

$$p_t = \frac{1}{2}(x_t + x_{t-1}),$$

and defining inflation as $\pi_t = p_t - p_{t-1}$, we have

$$\pi_t = \frac{1}{2}(\pi_{t-1} + E_t \pi_{t+1}) + \gamma \hat{y}_t^M, \tag{17}$$

where $\hat{y}_t^M$ is a moving average of current and past output. Inflation thus depends on its past value (which imparts inertia to both inflation and the

[14] As shown by Holden and Driscoll (2003), however, the presence of lagged inflation in the model depends on a somewhat arbitrary assumption regarding which other cohorts a given cohort will compare its real wage against.

price level) as well as its future value. Thus, Fuhrer-Moore contracts generate a hybrid Phillips curve similar to (11), albeit with restrictions on the parameters of the lagged and lead values of inflation.

By contrast, a two-period contracting equation of the Taylor type would imply that the contract wage is given by an average of the lagged and expected future wage contracts, adjusted for excess demand:

$$x_t = \frac{1}{2}(x_{t-1} + E_t x_{t+1}) + \gamma \hat{y}_t.$$

This specification implies that inflation is given by $\pi_t = E_t x_{t+1} + \gamma \hat{y}_t$,which implies that inflation persistence cannot result from the contracting specification *per se* but must come rather from persistence in the output gap.

2.3 Bank Borrowing and Marginal Costs

A key feature of the financial system in developing countries is that banks often play a critical role in financial firms' working capital needs (see Agénor and Montiel (2007, 2008)). This has important implications for the dynamics of inflation. Indeed, Ravenna and Walsh (2006) and Chowdhury, Hoffmann, and Schabert (2006) integrate a cost channel in an otherwise standard New-Keynesian model of the business cycle and show that the presence of a cost channel is tantamount to a direct effect of interest rates on inflation within a forward-looking Phillips curve. Interest rates affect firms' marginal costs of production, which in turn drive inflation dynamics. Both papers estimate these augmented NKPC equations and find evidence in support of the cost channel. Using a VAR approach, Tillmann (2006) also estimated the impact of a cost channel on inflation dynamics for the US, the UK, and the aggregate Euro area within a forward-looking Phillips curve framework. The results show that the cost channel plays a non-negligible role in explaining inflation and is able to account for certain inflation episodes that cannot be explained by the conventional New Keynesian Phillips curve.

A simple illustration of the Ravenna-Walsh model is as follows. As before, suppose that there exists a continuum of monopolistically competitive firms, of measure 1, indexed by i, producing differentiated final goods. The firm's production function is given by (1).

Workers must now be paid in advance of production. As a result, firm i must borrow an amount $W_t^i N_t^i$ from banks at the beginning of the period to cover their wage costs, and repay their loans at the end of the period, at the gross nominal interest rate $1 + R_t$. The nominal cost of labor is thus the effective wage, $(1 + R_t)W_t$. The (log) real marginal cost is again the same for all firms and equal to, instead of (2),

$$mc_t^n \equiv R_t + w_t - a_t. \tag{18}$$

If prices are flexible, the (log) equilibrium price is as before given as a markup over marginal cost, defined now as in (18). Changes in interest rates have therefore a direct effect on prices.[15] With Calvo-style sticky prices, firm i sets its price at time t to maximize, instead of (7),

$$\max_{P_t^j} \left\{ \sum_{j=0}^{\infty} (\theta\beta)^j (\frac{\Lambda_{t+j}}{\Lambda_t}) \left[(\frac{P_{t+j}^i}{P_{t+j}} - \frac{(1 + R_{t+j})W_{t+j}}{P_{t+j}A_{t+j}})(\frac{P_{t+j}^i}{P_{t+j}})^{-\upsilon} Y_{t+j} \right] \right\},$$

which again leads an equation similar to (9). However,with the cost channel, deviations in the real marginal cost from its steady-state value μ depend not only on deviations in the labor share from steady state, $\hat{s}_t$, but also on deviations in the gross nominal lending rate from its steady-state value, $\hat{R}_t$:

$$mc_t + \mu \simeq \hat{s}_t + \hat{R}_t, \tag{19}$$

Combining (9) and (19), and assuming as before a proportional relationship between $\hat{s}_t$ and the output gap $\hat{y}_t$, yields

$$\pi_t = \beta E_t \pi_{t+1} + \gamma_1 \hat{y}_t + \gamma_2 \hat{R}_t, \tag{20}$$

[15] As shown by Ravenna and Walsh (2006), however, changes in interest rates have no effect on the flexible-price equilibrium value of output if labor supply is inelastic.

or in a form similar to (11) if we again suppose that the fraction of firms that do not optimally adjust update their previous price on the basis of past inflation. Hence, the existence of a cost channel implies extending the measure of marginal costs to include bank lending rates directly into the price equation.

In the Ravenna-Walsh framework, the nominal interest rate R_t is treated essentially as predetermined, given the highly stylized treatment of the financial intermediation process. However, in a framework where credit market imperfections and monopolistic competition among financial intermediaries are explicityly accounted for, the interest rate at which firms borrow (the lending rate, R_t^L) will typically differ from the Central bank policy rate or the marginal cost of liquidity, R_t^C, by a risk premium, $\upsilon > 0$, which depends negatively on the firm's physical assets through a collateral effect (see Agénor and Montiel (2008, Chapter 6)). Assuming that the real stock of these assets is given in the short run, and that financial intermediaries set the lending rate one period in advance, implies that the nominal lending rate will vary inversely with lagged inflation, through its effect on collateral values:

$$R_t^L = R_t^C + \upsilon(\pi_{t-1}), \tag{21}$$

where $\upsilon' < 0$. Combining (9), (19) and (21) yields therefore a Phillips curve where lagged inflation also appears, as in hybrid models. However, persistence is now "extrinsic" in nature and results solely from credit market imperfections. Note also that the relevant interest rate is now the central bank policy rate. In our empirical results, however, we will use the lending rate directly (or the money market rate, as a proxy), which precludes relying on credit market imperfections as a source of inflation persistence.

2.4 Openness and Factor Substitution

Another important extension of the standard approach is to account for openness. Openness can affect the evolution of inflation in three ways: through competitive pressures from abroad, through the impact of imported final goods on consumer prices and wages, and through the real price of imported inputs.

The first channel relates to the fact that external competitive pressures, through imported substitutable final goods, may affect real marginal costs, which as a consequence depend on the *real* exchange rate. In Batini, Jackson, and Nickell (2005), this effect operates through changes in the desired markup.[16] In Guender (2006), it results from the tendency of domestic firms to adjust their prices in line with changes in the domestic-currency price of the final goods charged by their foreign competitors.

Guender (2006, Appendix) obtains a qualitatively similar result under Calvo pricing. Price-setting firms are now aware of the possibility that they may not be able to adjust prices for a while; they therefore consider the future evolution of all the factors that govern the determination of their product prices. These factors are domestic nominal marginal costs and the domestic-currency price of the foreign consumption good. The latter represents the price that prevails in world markets where domestic firms compete. Assuming a complete exchange-rate pass-through, Guender shows that the open-economy Phillips Curve takes the form

$$\pi_t = E_t \pi_{t+1} + \frac{(1-\theta)^2}{\theta}[\delta mc_t + (1-\delta) z_t], \tag{22}$$

where mc_t is real marginal cost, θ the probability (as before) that a firm cannot adjust the price of its product in a given period, δ and $1-\delta$ the relative importance of nominal marginal costs and the domestic currency price of the imported final consumption good in the process of price adjustment, respectively, and z_t the real exchange rate (or the relative price of the imported consumption good).[17] Thus, the more open the economy (the lower δ is),

[16] However, they do not find supporting evidence for this particular mechanism in their empirical results.

[17] Walsh (2003) presents an open-economy version of the model in Fuhrer and Moore (1995).

the more relative prices influence domestic inflation. In models such as those of Galí and Monacelli (2005) or Benigno and Lopez-Salido (2006), however, where consumer prices are specified directly as a weighted average of the prices of home goods and imported final goods, it is the *rate of change* of the real exchange rate, rather than its *level*, that appears in the open-economy Phillips curve.

The second channel results from the fact that nominal wages may be indexed on the consumer price index, which will normally reflect the price of both domestic and imported final goods. If labor supply depends on *consumption* real wages—because the consumer price index is the relevant deflator of spending from the perspective of workers—then through the goods market equilibrium condition inflation will also depend indirectly on the (actual or expected) relative price of foreign goods (see Walsh (1999) and Svensson (2000)).

The third channel relates to the degree of substitutability between production inputs. If the value added production function cannot be written in linear or Cobb-Douglas form, real marginal cost will be a function of the relative price of inputs, including imported ones. Put differently, if the import requirement of gross output is rising at the margin and it is not possible to substitute between labor and imported inputs, then the marginal cost of producing value added will be increasing in the real price of imported inputs (see Balakrishnan and López-Salido (2002)).

To illustrate this effect requires introducing imported intermediate goods into the production function. Suppose for instance that, instead of (1), the production function takes the CES form

$$Y_t^i = A[\alpha(N_t^i)^{(\sigma-1)/\sigma} + (1-\alpha)(J_t^i)^{(\sigma-1)/\sigma}]^{\sigma/(\sigma-1)}, \tag{23}$$

where $\alpha \in (0,1)$, J_t^i imported intermediate goods (oil, for short) by firm i, and σ the elasticity of substitution between inputs. Cost minimization implies solving for N_t^i and J_t^i to minimize $WN_t^i + P_t^J J_t^i$, where P_t^J is the domestic-currency price of imported oil, subject to (23). The optimality conditions yield

$$\frac{N_t^i}{J_t^i} = B(\frac{P_t^J}{W_t})^{\sigma}, \tag{24}$$

where $B = [\alpha/(1-\alpha)]^{\sigma}$. This equation shows that changes in the price of imported oil relative to the nominal wage results in substitution between labor and imported inputs.[18]

In terms of the consumer price index (rather than the price of output), real marginal costs of firm i are now defined by

$$MC_t^i = \frac{W_t}{P_t}(\frac{\partial N_t^i}{\partial Y_t^i}) + \frac{P_t^J}{P_t}(\frac{\partial J_t^i}{\partial Y_t^i}). \tag{25}$$

Using (23) and (24), this equation can be rewritten as (see Leith and Malley (2007, p. 410)):

$$MC_t = \left\{ \alpha^{\sigma}(\frac{W_t}{P_t})^{1-\sigma} + (1-\alpha)^{\sigma}(\frac{P_t^J}{P_t})^{1-\sigma} \right\}^{-1/(\sigma-1)}.$$

Taking logarithms, and assuming that the labor share is proportional to the output gap, yields an equation of the form

$$mc_t = \gamma_1 \hat{y}_t + \gamma_2(p_t^J - w_t) + \gamma_3(p_t^J - p_t), \tag{26}$$

where $p_t^J = \ln P_t^J$ and $\gamma_2 = 0$ if $\sigma = 0$. The third term, the relative price of imports (in terms of the price of value added), comes from the definition of GDP in the presence of intermediate goods, and is often referred to as a "terms-of-trade" effect.[19] Thus, movements in the real price of imported oil constitute an independent source of variation in marginal costs. As in Genberg and Pauwels (2005) equation (26) could be rewritten to relate marginal

[18] In the Cobb-Douglas case, the optimality conditions yield $W_t N_t^i/\alpha Y_t^i = P_t^J J_t^i/(1-\alpha)Y_t^i$, which can be rearranged to give (24) with $\sigma = 1$.

[19] If we were to assume decreasing marginal returns in the two variable factors in the short run, the level of output at the individual firm level would also affect marginal costs. As a result, as shown by Leith and Malley (2007) and Rumler (2007), average firm output would appear in the Phillips curve as well. The relationship betwen average firm output and GDP would then need to be examined.

costs to a weighted average of intermediate imported input costs and labor input costs. This equation can also be generalized to account for substitution between domestically produced and imported intermediate inputs, as in Rumler (2007).

Combining the first and third channels with Calvo pricing (that is, (22) and (26)), and using (percentage) deviations in the real exchange rate, $\hat{z}_t$, rather than its level, a general (unrestricted) open-economy NKPC can be written as

$$\pi_t = \chi_1 \hat{y}_t + \chi_2 E_t \pi_{t+1} + \chi_3 \hat{z}_t + \chi_4 (\hat{p}_t^J - \hat{w}_t) + \chi_5 (\hat{p}_t^J - \pi_t),$$

where $\hat{p}_t^J$ and $\hat{w}_t$ are deviations of p_t^J and w_t from their steady-state values. The term in the real exchange rate captures the effect of changes in the relative price of imported *final* goods, whereas the last term captures changes in the relative price of imported *intermediate* goods. This distinction, which does not appear in recent studies (for instance, Leith and Malley (2007) or Rumler (2007)) is important empirically. The coefficient χ_5 captures the "terms-of-trade" effect. Assuming that $\chi_5 < 1$, this equation can be rearranged and estimated in the alternative form

$$\pi_t = \frac{\chi_1}{1-\chi_5} \hat{y}_t + \frac{\chi_2}{1-\chi_5} E_t \pi_{t+1} + \frac{\chi_3}{1-\chi_5} \hat{z}_t + \frac{\chi_4}{1-\chi_5} (\hat{p}_t^J - \hat{w}_t) + \chi_5 \hat{p}_t^J, \quad (27)$$

which shows directly the impact of imported oil prices.[20] Note that we do not account directly for the second channel through which openness may affect the behavior of inflation (as identified earlier), given that equation (27) explicitly accounts for changes in nominal wages.

[20] With $\chi_4 = 0$, this equation is similar to the one specified and estimated by Moons et al. (2007), with the addition of a lagged term.

3 Empirical Specifications

Based on the foregoing discussion, we specify and estimate several alternative open-economy inflation specifications, in the presence of a cost channel.[21] Let $\mathbf{V}_t$ denote the vector of variables consisting of the cyclical component of the real exchange rate, deviations in the relative oil price-nominal wage ratio, deviations in imported oil prices, and the cyclical component of the lending rate:

$$\mathbf{V}_t = [\hat{z}_t \quad \hat{p}_t^J - \hat{w}_t \quad \hat{p}_t^J \quad \hat{R}_t^L]'.$$

In the first specification, which we refer to as $PC(-1,1)$, we estimate an extended Galí-Gertler hybrid model, of the form

$$\pi_t = \delta_0 + \delta_1 \pi_{t-1} + \delta_2 \pi_{t+1} + \sum_{i=0}^{m} \alpha_i \hat{y}_{t-i} + \boldsymbol{\gamma} \mathbf{V}_t + \varepsilon_t, \tag{28}$$

where $\boldsymbol{\gamma} = [\gamma_1 \quad \gamma_2 \quad \gamma_3 \quad \gamma_4]$ is a vector of parameters and $m \geq 0$. In the second specification, which we refer to as $P(-1,2)$, we further extend the model to account for two leads, as in (13):

$$\pi_t = \delta_0 + \delta_1 \pi_{t-1} + \sum_{i=1}^{2} \delta_{i+1} \pi_{t+i} + \sum_{i=0}^{m} \alpha_i \hat{y}_{t-i} + \boldsymbol{\gamma} \mathbf{V}_t + \varepsilon_t, \tag{29}$$

with $\delta_3 \leq 0$, as in (13).

The closed-economy Galí-Gertler hybrid specification (11) with no bank borrowing implies therefore $\delta_1 = \gamma_i = 0$, $\forall i$ in $P(-1,1)$, as well as $\delta_3 = 0$ in $P(-1,2)$. Equation (28) is also consistent with the Fuhrer-Moore model with two-period contracts if $\delta_1 = \delta_2 = 0.5$ and $\gamma_i = 0$, $\forall i$ (see (17)).

[21]Some early tests of New Keynesian models attempted to take into account both backward- and forward-looking elements in price setting by estimating a restricted equation of the form $\pi_t = \mu \pi_{t-1} + (1-\mu) E_t \pi_{t+1} + \alpha_1 \hat{y}_{t-1} + \varepsilon_t$, where $\mu \in (0,1)$, ε_t is an error term, and all other variables are as defined in the text. The relative importance of backward- and forward-looking components in inflation are thus measured by μ and $1-\mu$, respectively. Chadha, Masson and Meredith (1992), for instance, strongly reject values of 0 and 1 for μ in their estimates for major industrial countries (excluding the United Kingdom), whereas Fuhrer (1997) cannot reject the assumption $\mu = 1$ for the United States.

The third price equation, which we refer to as $P(-2,3)$, is a four-period, Taylor-type equation, as discussed in Fuhrer (1997):

$$\pi_t = \delta_0 + \sum_{i=1}^{2} \delta_i \pi_{t-i} + \sum_{i=1}^{3} \delta_{i+2} \pi_{t+i} + \sum_{i=0}^{m} \alpha_i \hat{y}_{t-i} + \boldsymbol{\gamma} \mathbf{V}_t + \varepsilon_t. \qquad (30)$$

Finally, we also estimate a restricted version of a symmetric price equation, which we refer to as $P(-4,4)$, given by

$$\pi_t = \delta_0 + \delta^L (\frac{1}{4} \sum_{i=1}^{4} \pi_{t-i}) + \delta^F (\frac{1}{4} \sum_{i=1}^{4} \pi_{t+i}) + \sum_{i=0}^{m} \alpha_i \hat{y}_{t-i} + \boldsymbol{\gamma} \mathbf{V}_t + \varepsilon_t, \qquad (31)$$

with the restriction $\delta^L + \delta^F = 1$. This specification, which is in the spirit of Fuhrer (1997, p. 340) and Jondeau and Le Bihan (2005, p. 528), allows one to cope with multicollinearity between lags (or leads) of inflation, while avoiding relying too heavily on restrictions implied by a particular contracting specification.

Note that we used throughout the output gap as a measure of marginal costs. As noted earlier, under some specific assumptions about the production technology, this is totally legitimate—in particular, the output gap can be shown to be proportional to the labor share if the production function is Cobb-Douglas. In addition, as pointed out in several recent studies, including, Roberts (2005), Lawless and Whelan (2007), and Rudd and Whelan (2007), there are a number of limitations associated with the use of the labor share (or unit labor costs) as a proxy for marginal costs. Rudd and Whelan (2007), in particular, have argued that the theoretical case for this approach is quite weak, and that in practice the labor share is not a good measure of real marginal costs and inflationary pressures.[22] Moreover, it can be argued that, for the group of countries that we are focusing on—as is true for developing countries in general—labor share data are probably subject to sizable

[22] One reason for that is that unit labor costs capture only a small fraction of production costs in most small open economies. In our approach, although we maintain the assumption that labor costs are linked to the output gap, we also control for changes in the cost of imported intermediate goods.

measurement errors, given that a large fraction of the labor force is employed in the informal economy. Using these data may therefore lead to unreliable inference.

4 Econometric Methodology

The first step in our econometric methodology involves estimating deviations of output, the real exchange rate, nominal wages, oil prices, and the bank lending rate, from their steady-state values. In line with much of the literature, we approximate these deviations by the cyclical component of each variable. To estimate the trend component, we use for all variables a modified version of the "ideal" band pass filter of Baxter and King (1999), as proposed by Christiano and Fitzgerald (2003). The Baxter-King filter is a linear transformation of the data, which leaves intact the components of the data within a specified band of frequencies and eliminates all other components. But this methodology has a limitation: its application requires a large amount of data. Christiano and Fitzgerald (2003) approximate the Baxter-King filter using an optimal linear approximation. Their method can be briefly summarized as follows. Let y_t be the data created by applying the ideal band pass filter to the raw data, x_t; y_t is approximated by $\widehat{y}_t$, which is a filter of x_t. The filter weights are chosen to minimize the mean square error:

$$E\left[(y_t - \widehat{y}_t)^2 | x\right].$$

$\widehat{y}_t$ can be computed as

$$\widehat{y}_t = B_0 x_t + B_1 x_{t+1} + ... + B_{T-1-t} x_{T-1} + \widetilde{B}_{T-t} x_T + B_1 x_{t-1}$$

$$+ ... + B_{t-2} x_2 + B_{t-1} x_1, \quad \text{for } t = 1, 2, 4, ..., T,$$

where

$$B_j = \frac{\sin(jb) - \sin(ja)}{\pi j}, \ j \geq 1$$

$$B_0 = \frac{b-a}{\pi}, \quad a = \frac{2\pi}{p_u}, \quad b = \frac{2\pi}{p_l},$$

and $\widetilde{B}_{T-t}$ and $\widetilde{B}_{t-1}$ are linear functions of B_j's:

$$\widetilde{B}_{T-t} = -\frac{1}{2}B_0 - \sum_{j=1}^{T-t-1} B_j$$

and $\widetilde{B}_{t-1}$ solves

$$0 = B_0 + B_1 + ... + B_{T-1-t} + \widetilde{B}_{T-t} + ... + B_{t-2} + \widetilde{B}_{t-1},$$

with $p_u = 24$ and $p_l = 2$ in our case.[23]

The second step in our procedure involves choosing an estimation method. Equations (28) to (31) contain future price expectations; we substitute these expected values by actual future inflation and use a Generalized Method of Moments (GMM) technique for estimation. In order to apply GMM, the moment condition that the parameter set, μ say, needs to satisfy is

$$E[m(\mu, \pi_t)] = 0,$$

where π is the dependent variable (inflation), and $E(\cdot)$ stands for the estimated value.

The GMM estimator is obtained by minimizing the following equation, which is defined as the distance between $m(\cdot)$ and 0:

$$\min_{\mu} \sum_t m(\mu, \pi_t, X_t, Z_t)' \hat{\Omega}^{-1} m(\mu, \pi_t, X_t, Z_t), \tag{32}$$

where X is the set of independent variables and Z the instrumental variables. $\hat{\Omega}$ is the weighting matrix. Here we use the lagged values of X as instrumental

[23] An alternative approach would be to use the HP filter. However, as noted by Rudd and Whelan (2007), it is not valid to use an HP filtered series as an instrument for GMM estimation (as we do here) because the filter employs future information in computing the estimated trend. The modified Baxter-King filter, by contrast, is a one-sided filter.

variables.[24] The moment condition is written as an orthogonality condition between the residuals of the regression equation, u, and Z:

$$m(\mu, \pi_t, X_t, Z_t) = Z' u(\mu, \pi_t, X_t).$$

The parameter vector μ is estimated with the two-stage GMM method. In the first step, the covariance matrix $\hat{\Omega}^{(1)}$ is calculated as $T^{\ -1} \sum_{t=1}^{T} Z_t Z_t'$. Then a first estimate of $\mu^{(1)}$ is obtained using $\hat{\Omega}^{(1)}$ in equation (32). In the second step, the covariance matrix is recalculated with $u_t = y_t - X_t' \mu^{(1)}$. Then the two-stage GMM estimator $\mu^{(2)}$ is obtained again from equation (32).

We also prewhiten the sample moments $m(\cdot)$ in prior to GMM estimation. To accomplish this, we first fit a VAR(1) to the sample moments such as $m(\cdot)_t = \Psi.m(\cdot)_{t-1} + v_t$. Then the covariance matrix is $\hat{\Omega}$ is estimated as $\hat{\Omega} = (I - \Psi)^{-1} \hat{\Omega}^* (I - \Psi)^{-1}$ where $\hat{\Omega}^*$ is defined below.

The weighting matrix $\hat{\Omega}^*$ is estimated as the heteroskedasticity and autocorrelation consistent covariance matrix:

$$\hat{\Omega}^* = \hat{\Gamma}(0) + \left\{ \sum_{j=1}^{T-1} k(i,q) \left[\hat{\Gamma}(i) - \hat{\Gamma}'(i) \right] \right\},$$

where

$$\hat{\Gamma}(i) = \frac{1}{T-k} \left\{ \sum_{t=i+1}^{T} = Z_{t-i}' u_t u_{t-i}' Z_t \right\},$$

and k is the number of coefficients, q the bandwidth, and $k(\cdot)$ the Kernel function, which is included to ensure that $\hat{\Omega}^*$ is a positive semi-definite. We use the Bartlett Kernel, defined as

$$k(i,q) = \begin{cases} 1 - (i/q) & 0 \leq i \leq q \\ 0 & \text{otherwise} \end{cases}.$$

[24] In line with the literature and the overidentification test results discussed later, the instrumental variables are four lags (from $t-1$ to $t-4$) of inflation, output gap, and the cyclical components of the real exchange rate, the interest rate, the oil price, and the difference in oil prices and the wage rate.

The bandwidth is the Newey-West fixed bandwidth, based on the number of observations:

$$q = \text{integer}[4(T/100)^{2/9}],$$

which varies between 3 and 4, depending on the sample size for each country.

Note that in specifications (28) to (30), the restriction that the sum of all lead and lag coefficients is equal to one can be tested, rather than imposed, as in (31). For instance, with (30), it requires testing whether $\sum_{i=1}^{5} \delta_i = 1$. Finally, note that specification $P(-2, 3)$ nests $P(-1, 1)$, with the restrictions $\delta_2 = \delta_4 = \delta_5 = 0$, as well as $P(-1, 2)$, with the restrictions $\delta_2 = \delta_4 = 0$, as long as the set of variables in $\mathbf{V}_t$ is the same across regressions. However, a comparison between $P(-2, 3)$ and the other two specifications if $\mathbf{V}_t$ changes across regressions, or between $P(-4, 4)$ and any other specification, requires a non-nested test.

5 Estimation Results

The output gap, $\hat{y}_t$, is calculated as the log difference of the actual output level and its trend component. The number of lags for $\hat{y}_t$ is set at 4. The cyclical component of the real effective exchange rate, $\hat{z}_t$, is defined as the log difference of the real exchange rate index and its trend components. A positive value of $\hat{z}_t$ indicates that the real exchange rate is depreciating. Similarly, the difference between the actual lending rate and its trend, $\hat{R}_t$, the log difference between world oil price and its trend, $\hat{p}_t^J$, and the gap between the log difference of actual oil price and its trend and the log difference of actual wage and its trend, $\hat{p}_t^J - \hat{w}_t$, are also included in the regression equations. As noted earlier, the last variable is included to account for possible substitution between externally-produced inputs and domestic labor.

5.1 Alternative Specifications

We estimate the reduced-form equations (28) to (31) using quarterly data for eight middle-income developing countries: Chile, Colombia, Korea, Malaysia, Mexico, Morocco, Tunisia, and Turkey.[25] Preliminary testing showed that some of the variables in $\mathbf{V}_t$ had coefficients with a wrong sign; they were therefore excluded in the final regression results, which are presented in Table 1.[26]

The results show that the lagged and lead inflation rate variables are highly statistically significant and have overall a positive effect on the current inflation in all countries. In addition, for the first three specifications, the coefficient of the lagged inflation rate (or the sum of lagged coefficients) is larger than the coefficient of the lead inflation rate (or the sum of lead coefficients) in almost all of them, indicating that backward-looking behavior is a more important component in explaining inflation dynamics. Exceptions are Colombia and Korea. For Chile, our results for $P(-1,1)$ differ therefore from those of Céspedes, Ochoa, and Soto (2005), who estimated a coefficient for the backward-looking component of about 0.4, but they are close to those of Coble (2007), who found a value close to 0.6. The highest coefficients for lagged inflation pertain to Mexico, Morocco, and Turkey. By contrast, with the restricted specification $P(-4,4)$, the results change for some countries. For instance, for Mexico and Turkey, the coefficient of the sum of lagged inflation rates is lower than the one for the sum of lead inflation rates, indicating more evidence of forward-looking behavior. We address this issue later on, by performing nested and non-nested tests.

Table 1 also reports the sum of the estimated coefficients of the output

[25] See the Appendix for data sources and results of unit root tests. We chose not to report "structural" estimates because of the difficulty of choosing which parameters to estimate, and which parameters to impose, with such a diverse group of countries.

[26] Because the data are quarterly, seasonal effects are captured by including seasonal dummies, in all regression equations. To save space, the coefficients of these variables are not reported.

gap. The coefficient has the expected positive sign for each country, except Malaysia. In addition, the output gap is highly significant in determining the inflation rate is Colombia and Korea—the same two countries for which inflation appears to be more forward-looking. The cyclical component of the real exchange rate has a correct sign for all countries except Colombia, Mexico, and Morocco. The statistical significance of this variable is relatively higher in Korea and Turkey. Similarly the coefficient of the current cyclical component of the lending rate has the expected positive sign for almost all countries with the exception of Turkey, and Morocco, and Tunisia for some regression specifications. The significance level of this variable is high in Korea and Mexico, indicating that the cost channel is particularly effective in these countries. For Chile, the estimated coefficient is borderline significant and relatively small in $P(-1,2)$ and $P(-2,3)$, consistent with the single equation results in Coble (2007, Table 2).

Compared to other variables in the $\mathbf{V}_t$ set, neither $\hat{p}_t^J$ nor $\hat{p}_t^J - \hat{w}_t$ is very successful in explaining inflation. They have a wrong sign for most countries, and even when they have the expected positive sign, the statistical significance is low. While the coefficient of $\hat{p}_t^J$ is statistically significant only for Chile in two specifications, the coefficient of $\hat{p}_t^J - \hat{w}_t$ is statistically significant with a correct sign only for Chile and Turkey. These results are consistent with the recent evidence showing a low pass-through of world oil prices to inflation (see De Gregorio et al. (2007)) and a high degree of indexation of nominal wages. However, these conclusions should be tempered by the fact that the best way of measuring the effect of oil prices on inflation would be to use a *retail* oil price index, given that in many countries prices "at the pump" are either controlled or heavily subsidized and therefore do not reflect the volatility observed on international markets. But unfortunately, this series was available only for Morocco. Thus, data limitations may partially explain the failure of the oil price variable, and the relative input price, in explaining inflation in our sample of countries. It is worth noting, however, that even in the case of Morocco, changes in oil prices are not significant; it is therefore not clear that using retail prices would improve the results much in the other cases. Moreover, subsidies and price controls tend to be imposed at the retail level, not at the wholesale (or supply side) level, which is the effect that our estimation results tend to capture. It is reasonable to expect wholesale prices to be more closely correlated with world prices.

The other result that can be seen in Table 1 is that at least two of the variables given in $\mathbf{V}_t$ enter the specifications with the expected positive sign. Morocco is the only exception. For this country, we see only the lending rate in estimation results for $P(-2,3)$ and $P(-4,4)$.

We performed two tests for our set of instrumental variables: the first for overidentification and the second for potential weakness of these variables—that is, the possibility that the selected instruments are only weakly correlated with the included endogenous variables, implying that standard GMM point estimates, hypothesis tests, and confidence intervals are unreliable (see Stock, Wright, and Yogo (2002) and Mavroeidis (2005)). Because the number of instrumental variables is larger than the number of independent variables in our regressions, we run the J test to check for a possible overidentification problem. In this test, the null hypothesis indicates that the overidentifying restrictions hold. The test results are given in Table 1. For each country and specification, we fail to reject the null hypothesis.

Table 1 also reports weak instrument test results for each country. It is the first-stage F statistics advocated by Stock and Yogo (2005). The test is based on the first-stage regression of instrumental variables on endogenous variables which are the lead inflation in our specifications. Then the F statistic is calculated for the collective rejection of the estimated coefficients of instrumental variables in the first-stage regression. Thus, the null hypothesis is that instruments are weak. In the table, we present this F statistic which converges to the "first-stage F statistics" plus one when the number of observations is large enough. There are specific critical values for this test, as reported in Table 1 of Stock and Yogo (2005). When we compare our test results with the critical values, we infer that the null hypothesis of weak instruments is rejected for each country when the desired maximal bias of the instrumental variable estimator relative to ordinary least square is taken as 0.20, given that 28 instrumental variables are included in our specifications.

In all of these regression equations, except specification (31), we do not restrict the sum of the estimated coefficients for past and future inflation to be equal to unity. For specification (31), we test for the validity of the null hypothesis that the sum of the estimated coefficients of past and future inflation is unity, in order to ensure that no long-run trade-off exists between inflation and other variables. Specifically, we use a Wald test. The results, which are also reported in Table 1, indicate that the null hypothesis is accepted for all countries.

5.2 Nested and Non-Nested Tests

In order to decide on which price formation equation is most suitable for each country, we run nested or non-nested tests, depending on specifications. Note that specification $P(-2,3)$ nests $P(-1,1)$, with the restrictions $\delta_2 = \delta_4 = \delta_5 = 0$, as well as $P(-1,2)$, with the restrictions $\delta_2 = \delta_4 = 0$, as long as the set of variables in $\mathbf{V}_t$ is the same across regressions. When we check the regression results in Table 1, we can see that $P(-2,3)$ nests $P(-1,1)$ as well as $P(-1,2)$, and $P(-1,2)$ nests $P(-1,1)$ for all countries except Chile and Mexico. For these two countries, $P(-2,3)$ nests $P(-1,2)$, but not $P(-1,1)$; $P(-1,2)$ does not nest $P(-1,1)$ either.

The results for the nested tests is given in the second panel of Table 2. The first null hypothesis H_0: $\delta_2 = \delta_4 = \delta_5 = 0$ tests whether we can reject specification $P(-2,3)$ against $P(-1,1)$. If we reject H_0, it indicates that specification $P(-2,3)$ is better specified. The second null hypothesis H_0: $\delta_2 = \delta_4 = 0$ tests whether we can reject specification $P(-2,3)$ against $P(-1,2)$, so that rejection of H_0 implies that $P(-2,3)$ is better specified. In the last nested test, the null hypothesis is taken as H_0: $\delta_4 = 0$. Here we test specification $P(-1,2)$ against $P(-1,1)$, where rejection of H_0 indicates that $P(-1,2)$ is better specification.

A comparison between $P(-2,3)$ and the other two specifications, if $\mathbf{V}_t$ is *not* the same across regressions, or between $P(-4,4)$ and any other specification, requires a non-nested test. We use for this purpose Davidson and McKinnon's non-nested J-test, despite its well-known weaknesses in small samples (see Gourieroux and Monfort (1994)). The results are given in the first panel of Table 2. The specifications given in the null hypothesis are the ones that we are testing against other specifications given in the alternative hypotheses. For instance, when H_0: $P(-1,1)$ and H_a: $P(-1,2)$, we test whether $P(-1,1)$ is well specified against $P(-1,2)$. If $P(-1,1)$ is well specified, the estimated coefficient of the fitted value of specification $P(-1,2)$ should be equal to zero when that value is plugged into $P(-1,1)$ as an independent variable. The other hypotheses are defined in a similar way. In the table, we report the t-statistic and p-value of the estimated coefficient of the fitted value of the alternative specification in the null hypothesis model.

The results for Chile indicate that the correct specification is $P(-4,4)$. While the nested test shows that we fail to reject $P(-2,3)$ against $P(-1,2)$, the non-nested test results suggest that the fitted values of all other specifications enter $P(-4,4)$ insignificantly. For Colombia, Korea, Malaysia, and Turkey $P(-2,3)$ is the best specification. For all these countries, $P(-4,4)$ fails because the fitted values of all other specifications enter with a statistically significant coefficient, and according to the nested test results we fail to reject $P(-2,3)$ against $P(-1,1)$ and $P(-1,2)$. For Mexico, the results show that both $P(-2,3)$ and $P(-4,4)$ may be a correct specification. But we accept $P(-2,3)$ because it is superior to $P(-1,2)$ according to the results of the nested tests, and the fitted value of $P(-4,4)$ is less significant when plugged in $P(-2,3)$. Similarly, for Morocco, the test results show that $P(-2,3)$ is not very different from $P(-4,4)$. But we accept $P(-4,4)$, given that the significance level of the fitted value of $P(-2,3)$ is lower in $P(-4,4)$. Tunisia is another country with two possible specifications: $P(-4,4)$ and $P(-1,2)$. Similar to the other two countries with two possible specifications, we accept $P(-4,4)$ based on the significance level of the fitted values in the non-nested tests.

Overall, the results indicate that $P(-1,1)$ and $P(-1,2)$ are rejected for each country. They suggest that we need longer lagged and lead values of inflation to better specify inflation equations for the set of developing countries that we investigate in the paper.

5.3 In-Sample Predictive Capacity

In this section we compare actual and predicted inflation based on the best specification given in Table 2. This comparison helps us to better assess the goodness of fit of chosen specifications. Figure 2 shows the results. Overall, the estimated models seem to perform quite well. For Korea, Malaysia, Mexico, and Turkey, predicted values follow actual values fairly closely and the turning points (or inflation hikes) corresponding to various crises—the oil shock of the early 1980s and the Asia crisis in late 1997-98 for Korea and Malaysia; the debt crisis of the early 1980s, the stabilization program of the late 1980s, and the currency crisis in late 1994 for Mexico; and the early 1994 and 2001 currency crises in Turkey—are predicted well. Results are less impressive for Tunisia and Morocco; in both countries, however, inflation is strongly affected by the behavior of food prices, which depend on weather conditions.

5.4 Structural Stability

To further evaluate the performance of the best specifications reported in Table 2, we perform within-sample parameter stability tests. This is important because for instance the coefficient of the output gap, may itself depend on the expected inflation rate (Ball, Mankiw, and Romer (1988)).

More generally, the countries considered in our sample have undergone significant structural adjustment (including greater trade openness) during the estimation period, and this may have affected the inflation process.

To test for a structural breakpoint in the sample, we run the Quandt-Andrews test (see Andrews (1993) and Andrews and Ploberger (1994)). Essentially, the Quandt-Andrews test is based on a single Chow breakpoint test and is performed at every observation between two dates. The test statistics from those Chow tests are then summarized into a single test statistic for a test against the null hypothesis of no breakpoints between two dates. This test checks whether there is a structural change in all of the original equation parameters. For each individual Chow breakpoint test, we retain the Wald F-statistic, which is based on a standard Wald test of the restriction that the coefficients in the equation are the same in all subsamples.

Table 3 reports the maximum value of the individual Chow F-statistics, its p-value and corresponding period.[27] Because the distribution of this test statistics is non-standard, we use approximate asymptotic p-values provided by the Hansen p-values (see Hansen (1997)). Note that the distribution of the statistic becomes degenerate as the first period tested approaches the beginning of the equation sample, or the end period approaches the end of the equation sample; to compensate for this behavior, we trimmed 25 percent of the ends of the equation sample, where we excluded the first and last 12.5 percent of the observations. The results in Table 3 show that we fail to reject the null hypothesis for each country where the null hypothesis is no structural breakpoints available. Interestingly enough, the results can be traced to particular events in some cases; in Mexico, for instance, the break occurs right after the announcement of the Pact for Economic Solidarity in December 1987, a "heterodox" stabilization program that involved fixing the exchange rate and a temporary freeze on wages, public sector prices, and the

[27]We cannot run this test for Tunisia and Malaysia, because we use a dummy variable for the wage series.

prices of commodities in a basket of basic goods and services (see Agénor and Montiel (2008, Chapter 11)). For Turkey, the break occurred concurrently with the currency crisis of January 1994.

To check the stability of estimated coefficients or changes in these coefficients over time, the "preferred" equations from Table 2 could also be estimated using rolling regressions. Because this procedure is somewhat cumbersome in the present case, we illustrate its application by focusing on one country, Turkey. In addition, we concentrate on one parameter—the forward-looking component of inflation, that is, the sum of the lead coefficients of inflation in $P(-2,3)$ specification. The starting point is the estimated coefficient of the sum of lead inflation between 1982:Q1 up to 1999:Q4 and the following coefficients are obtained by adding one quarter at a time and reestimating, up to 2006:Q2. The choice of Turkey and the focus on the most recent period are not innocuous; as noted earlier, Turkey went through a major financial crisis in 2001 (see Yilmaz and Boratav (2003)). It then adopted (in May of that year) an inflation stabilization program with announced inflation targets, and eventually a formal inflation targeting regime in early 2006. If the program generated quick success, the forward-looking component of inflation should have increased.

Figure 3 presents the results, with a 2-standard-error band. The value of the forward-looking coefficient increases up to 0.4 from 0.3 until 2001:Q2. After that, it starts dropping and stays stable at 0.3. Overall the coefficient looks stable. There is therefore no evidence of an "early credibility gain" associated with inflation stabilization.

6 Concluding Remarks

This paper provided empirical estimates of contracting models of the Phillips curve for eight middle-income developing economies—Chile, Colombia, Korea, Malaysia, Mexico, Morocco, Tunisia, and Turkey. The first part reviewed a variety of models, with a particular focus on the cost channel and open-economy considerations. The second part presented the econometric methodology (based on two-step GMM techniques) and the third reduced-form estimation results for several alternative specifications. We also presented results of nested and non-nested tests for choosing among these specifications, as well as in sample-predictive capacity and parameter stability tests. We found that, in general, parsimonious models with one lead and one lag of inflation are rejected for higher-dimension models. Except for Colombia and Korea, backward-looking behavior is a more important component in explaining inflation. Moreover, in contrast to some studies, we do not find that the weight of the forward-looking component in a model with several lags and leads falls significantly when compared to the specification with a single lag and lead. World oil prices and relative input prices appear to have a limited effect on inflation, whereas bank borrowing costs are particularly significant for some countries (Korea and Mexico).

The analysis in this paper can be extended in various directions. In particular, evidence for industrial countries suggests that the inflation process may be asymmetric, in that excess demand tends to have a larger effect on inflation than an equivalent degree of excess supply. Contributions include Chadha, Masson and Meredith (1992), Laxton, Meredith, and Rose (1995), Dupasquier and Ricketts (1998), and more recently Clark et al. (2001) and Dolado et al. (2005). Changes in prices of imported goods, as well as changes in interest rates, may also have an asymmetric effect on inflation. Although some preliminary work on nonlinearities in the Phillips curve has been done for developing countries (see Agénor (2002)), the scope for further research for these countries is significant.

Appendix
Data Sources and Unit Root Tests

The dataset is an update of the quarterly database compiled by Agénor, McDermott, and Prasad (2000). It covers the following years for the countries in the sample: 1979:Q1-2006:Q1 for Chile, 190:Q1-2005:Q3 for Colombia, 1979:Q3-2006:Q3 for Korea, 1978:Q3-2006:Q3 for Malaysia, 1978:Q3-2006:Q3 for Mexico, 1982:Q1-2006:Q2 for Morocco, 1979:Q1-2006:Q2 for Tunisia, and 1981:Q1-2006:Q2 for Turkey. The variables and the sources of them are as follows.

- π_t is the annual log difference in the consumer price index. Source: International Monetary Fund (IMF).
- $\hat{y}_t$ is the log difference of output to the trend component of output, where output is the real industrial production index for all countries except for Morroco for which real GDP series are used. The trend component is calculated using a generalized version of the Baxter-King filter, as explained in the text. Source: IMF for all countries except Morroco, for which the source is Bank al Maghrib.
- $\hat{z}_t$ is the log difference of the real exchange rate to the trend component of the real exchange rate. A rise is a depreciation. The trend component is calculated using a generalized version of the Baxter-King filter. Source: IMF.
- $\hat{R}_t$ is the difference of interest rate to the trend component of the interest rate, where the interest rate is the lending rate for Chile and Korea, the deposit rate for Turkey, Mexico, and Malaysia, the discount rate for Colombia, and the money market rate for Morocco and Tunisia. The trend component is calculated using a generalized version of the Baxter-King filter. Source: IMF for all countries except Morocco, for which the source is Bank al-Maghrib.
- $\hat{p}_t^J$ is the log difference of the oil price index to the trend component of the oil price index, where the oil price index is 3-spot index for all countries except Morocco for which the domestic oil price index is used. The trend component is calculated using a generalized version of the Baxter-King filtering method. Source: IMF for all countries except Morocco, for which the source is Bank al-Maghrib.

- $\hat{p}_t^J - \hat{w}_t$ is the log difference of the nominal wage index to the trend component of the wage index. The trend component is calculated using a generalized version of the Baxter-King filter. Source: IMF for all countries except for Morocco (Bank al-Maghrib), Turkey (Central Bank of Turkey), Malaysia (Bank Negara), Colombia (ILO Database), and Tunisia (Central Bank of Tunisia).

Each series used in the regressions is tested for stationarity. Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests are employed. The null hypothesis is that a unit root exists and the alternative hypothesis is that the series is trend stationary. Results of these tests are reported in Table A1.

In general, both test statistics give similar results. The presence of a unit root is rejected for almost all series, with most series being stationary at a 1 percent significance level. Only Colombia and Turkey appear to have a unit root problem for inflation. However, given the relatively short size of the sample, and to maintain comparability with other countries, estimation was also performed in the levels of inflation for both countries.

References

Agénor, Pierre-Richard, "Monetary Policy under Flexible Exchange Rates: An Introduction to Inflation Targeting," in *A Decade of Inflation Targeting in the World*, ed. by Norman Loayza and Raimundo Soto, Central Bank of Chile (Santiago: 2002).

Agénor, Pierre-Richard, C. John McDermott, and Eswar Prasad, "Macroeconomic Fluctuations in Developing Countries: Some Stylized Facts," *World Bank Economic Review*, 14 (May 2000), 251-86.

Agénor, Pierre-Richard, and Peter J. Montiel, "Monetary Policy Analysis in a Small Open Credit-Based Economy," Working Paper No. 90, Centre for Growth and Business Cycle Research, University of Manchester (October 2007). Forthcoming, *Open Economies Review*.

——, *Development Macroeconomics*, 3rd. ed., forthcoming Princeton University Press (Princeton, New Jersey: 2008).

Andersen, Torben M., "Persistency in Sticky Price Models," *European Economic Review*, 42 (May 1998), 593-603.

Andrews, Donald W. K., "Tests for Parameter Instability and Structural Change with Unknown Change Point," *Econometrica*, 61 (July 1993), 821-56.

Andrews, Donald W. K., and Werner Ploberger, "Optimal Tests when a Nuisance Parameter is Present only under the Alternative," *Econometrica*, 62 (November 1994), 1383-414.

Ascari, Guido, and Christian Merkl, "Real Wage Rigidities and the Cost of Disinflations," IZA Working Paper No. 3049 (September 2007).

Bakhshi, Hasan, Hashmat Khan, and Barbara Rudolf, "The Phillips Curve under State-Dependent Pricing," *Journal of Monetary Economics*, 54 (November 2007), 2321-45.

Balakrishnan, Ravi, and J. David López-Salido, "Understanding UK Inflation: The Role of Openness," Working Paper No. 164, Bank of England (March 2002).

Ball, Laurence, N. Gregory Mankiw, and David Romer, "The New Keynesian Economics and the Output-Inflation Trade-off," *Brookings Papers on Economic Activity*, No 1 (June 1988), 1-65.

Banerjee, Ryan, and Nicoletta Batini, "Inflation Dynamics in Seven Industrialised Open Economies," unpublished, Bank of England (October 2004).

Bardsen, Gunnar, Eilev S. Jansen, and Ragnar Nymoen, "Econometric Evaluation of the New Keynesian Phillips Curve," *Oxford Bulletin of Economics and Statistics*, 66 (September 2004), 671-86.

Batini, Nicoletta, Brian Jackson, and Stephen Nickell, "Inflation Dynamics and the Labour Share in the UK," *Journal of Monetary Economics*, 52 (September 2005), 1061-71.

Baxter, Marianne, and Robert G. King, "Measuring Business Cycles: Approximate Band-Pass Filters for Economic Time Series," *Review of Economics and Statistics*, 81 (November 1999), 575-93.

Benigno, Pierpaolo, and J. David López-Salido, "Inflation Persistence and Optimal Monetary Policy in the Euro Area," *Journal of Money, Credit, and Banking*, 38 (April 2006), 587-614.

Blanchard, Olivier, and Jordi Galí, "Real Wage Rigidities and the New Keynesian Model," *Journal of Money, Credit and Banking*, 39 (February 2007), 35-65.

Calvo, Guillermo A., "Staggered Contracts in a Utility-Maximizing Framework," *Journal of Monetary Economics*, 12 (September 1983), 383-98.

Calvo, Guillermo A., Oya Celsasun, and Michael Kumhoff, "Inflation Inertia and Credible Disinflation—The Open Economy Case," Working Paper No. 9557, National Bureau of Economic Research (March 2003).

Céspedes, Luis F., Marcelo Ochoa, and Claudio Soto, "The New Keynesian Phillips Curve in an Emerging Market Economy: The Case of Chile," Working Paper No. 355, Central Bank of Chile (December 2005).

Chadha, Bankim, Paul R. Masson, and Guy Meredith, "Models of Inflation and the Costs of Disinflation," *IMF Staff Papers*, 39 (June 1992), 395-431.

Chowdhury, Ibrahim, Mathias Hoffmann, and Andreas Schabert, "Inflation Dynamics and the Cost Channel of Monetary Transmission," *European Economic Review*, 50 (May 2006), 995-1016.

Christiano, Lawrence J., and Terry J. Fitzgerald, "The Band Pass Filter," *International Economic Review*, 44 (May 2003), 435-65.

Clarida, Richard, Jordi Galí, and Mark Gertler, "The Science of Monetary Policy: A New Keynesian Perspective," *Journal of Economic Literature*, 37 (December 1999), 1661-707.

Clark, Peter, Douglas Laxton, and David Rose, "An Evaluation of Alternative Monetary Policy Rules in a Model with Capacity Constraints," *Journal of Money, Credit, and Banking*, 33 (February 2001), 42-64.

Coble, David, "Dinámica de inflación y el Canal de Costos: Una Aplicación para Chile," Working Paper No. 431, Central Bank of Chile (October 2007).

De Gregorio, José, Oscar Landerretche, and Christopher Neilson, "Another Pass-Through Bites the Dust? Oil Prices and Inflation," Working Paper No. 417, Central Bank of Chile (May 2007).

Dolado, Juan D., Ramon María-Dolores, and Manuel Naveira, "Are Monetary-Policy Reaction Functions Asymmetric? The Role of Nonlinearity in the Phillips Curve," *European Economic Review* 49, (February 2005), 485-503.

Dotsey, Michael, Robert G. King, and Alexander L. Wolman, "State-Dependent Pricing and the General Equilibrium Dynamics of Money and Output," *Quarterly Journal of Economics*, 114 (May 1999), 655-90.

Dupasquier, Chantal, and Nicholas Ricketts, "Non-Linearities in the Output-Inflation Relationship: Some Empirical Results for Canada," Working Paper No. 98-14, Bank of Canada (August 1998).

Fuhrer, Jeffrey C. "The (Un)Importance of Forward-Looking Behavior in Price Specifications," *Journal of Money, Credit, and Banking*, 29 (August 1997), 338-50.

Fuhrer, Jeffrey C., and Geoffrey R. Moore, "Inflation Persistence," *Quarterly Journal of Economics*, 110 (February 1995), 127-59.

Gagnon, Edith, and Hashmat Khan, "New Phillips Curve under Alternative Production Technologies for Canada, the United States, and the Euro area," *European Economic Review*, 49 (August 2005), 1571-602.

Galí, Jordi, and Mark Gertler, "Inflation Dynamics: A Structural Econometric Analysis," *Journal of Monetary Economics*, 44 (October 1999), 195-222.

Galí, Jordi, Mark Gertler, and J. David López-Salido, "European Inflation Dynamics," *European Economic Review*, 45 (May 2001), 1237-70. Erratum: *European Economic Review*, 47 (August 2003), 759-60.

Galí, Jordi, and Tommaso Monacelli, "Monetary Policy and Exchange Rate Volatility in a Small Open Economy," *Review of Economic Studies*, 72 (July 2005), 707-34.

Genberg, Hans, and Laurent Pauwels, "An Open Economy New Keynesian Phillips Curve: Evidence from Hong Kong," *Pacific Economic Review* 10 (June 2005), 261-77.

Gourieroux, Christian, and Alain Monfort, "Testing Non-Nested Hypotheses," in *Handbook of Econometrics*, ed. by Robert Engle and Daniel McFadden, Vol. 4, Elsevier (Amsterdam: 1994).

Guay, Alain, and Florian Pelgrin, "The U.S. New Keynesian Phillips Curve: An Empirical Assessment," Working Paper No. 2004-35, Bank of Canada (September 2004).

Guender, Alfred V., "Stabilising Properties of Discretionary Monetary Policies in a Small Open Economy," *Economic Journal*, 116 (January 2006), 309-26.

Guerrieri, Luca, "The Inflation Persistence of Staggered Contracts," International Finance Discussion Paper No. 734, Board of Governors of the Federal

Reserve System (August 2002).

Hansen, Bruce E., "Approximate Asymptotic P Values for Structural-Change Tests," *Journal of Business and Economic Statistics*, 15 (January 1997), 60-67.

Holden, Steinar, and John Driscoll, "Inflation Persistence and Relative Contracting," *American Economic Review*, 93 (September 2003), 1369-72.

Ireland, Peter N., "Money's Role in the Monetary Business Cycle," *Journal of Money, Credit, and Banking*, 36 (December 2004), 969-83.

Jondeau, Eric, and Hervé Le Bihan, "Testing for the New Keynesian Phillips Curve: Additional International Evidence," *Economic Modelling*, 22 (May 2005), 521-50.

Karanassou, Marika, and Dennis J. Snower, "Inflation Persistence and the Phillips Curve Revisited," IZA Working Paper No. 2600 (February 2007).

Kozicki, Sharon, and P. A. Tinsley, "Dynamic Specifications in Optimising Trend-Deviation Macromodels," *Journal of Economic Dynamics and Control*, 26 (August 2002), 1585–611.

Laxton, Douglas, Guy Meredith, and D. Rose, "Asymmetric Effects of Economic Activity on Inflation," *IMF Staff Papers*, 2 (June 1995), 344-74.

Lawless, Martina, and Karl Whelan, "Understanding the Dynamics of Labor Shares and Inflation," Working Paper No. 785, European Central Bank (July 2007).

Leith, Campbell, and Jim Malley, "Estimated Open Economy New Keynesian Phillips Curves for the G7," *Open Economies Review*, 18 (September 2007), 405-26.

Liu, Philip, "A Small New Keynesian Model of the New Zealand Economy," Discussion Paper No. 2006/03, Reserve Bank of New Zealand (May 2006).

Mavroeidis, Sophocles, "Identification Issues in Forward-Looking Models Estimated by GMM, with an Application to the Phillips Curve," *Journal of Money, Credit, and Banking*, 37 (June 2005), 421-48.

McKinnon, James, "Critical Values for Cointegration Tests," in Robert Engle and Clive W. Granger, *Long-Run Economic Relationships: Readings in Cointegration*, Oxford University Press (Oxford: 1991).

Moons, Cindy, Harry Garretsen, Bas van Aarle, and Jorge Fornero, "Monetary Policy in the New-Keynesian Model: An application to the Euro Area," *Journal of Policy Modeling*, forthcoming December 2007).

Musy, Olivier, "Inflation Persistence and the Real Costs of Disinflation in Staggered Prices and Partial Adjustment Models," *Economics Letters*, 91 (April 2006), 50-55.

Nason, James M., and Gregor W. Smith, "Identifying the New Keynesian Phillips Curve," Working Paper No 2005-01, Federal Reserve Bank of Atlanta (January 2005).

Olafsson, Thorvaldur T., "The New Keynesian Phillips Curve: In Search of Improvements and Adaptation to the Open Economy," unpublished, Central Bank of Iceland (September 2006).

Ravenna, Federico and Carl E. Walsh, "Optimal Monetary Policy with the Cost Channel," *Journal of Monetary Economics*, 53 (March 2006), 199-216.

Roberts, John M., "New Keynesian Economics and the Phillips Curve," *Journal of Money, Credit, and Banking*, 27 (November 1995), 975-84.

——, "How Well does the New Keynesian Sticky-Price Model Fit the Data?," *Contributions to Macroeconomics*, 5 (March 2005), 1206-.

Romer, David, *Advanced Macroeconomics*, 2nd ed., McGraw Hill (New York: 2000).

Rotemberg, Julio J., "Monopolistic Price Adjustment and Aggregate Output," *Review of Economic Studies*, 49 (October 1982), 517-31.

Rotemberg, Julio., and Mark Woodford, "An Optimization-based Econometric Framework for the Evaluation of Monetary Policy," in *NBER Macroeconomic Annual*, ed. by Ben Bernanke and Julio Rotemberg, MIT Press (Cambridge, Mass.: 1997).

Rudd, Jeremy and Karl Whelan, "New Tests of the New Keynesian Phillips Curve," *Journal of Monetary Economics*, 52 (September 2005), 1167-81.

——, "Modeling Inflation Dynamics: A Critical Review of Recent Research," *Journal of Money, Credit and Banking*, 39 (February 2007), 155-70.

Rumler, Fabio, "Estimates of the Open Economy New Keynesian Phillips Curve for Euro Area Countries," *Open Economies Review*, 18 (September 2007), 427-51.

Sbordone, Argia M., "Prices and Unit Labor Costs: A New Test of Price Stickiness," *Journal of Monetary Economics*, 49 (March 2002), 265-92.

——, "Inflation Persistence: Alternative Interpretations and Policy Implications," *Journal of Monetary Economics*, 54 (July 2007), 1311-39.

Sheedy, Kevin D., "Intrinsic Inflation Persistence," unpublished, London School of Economics (May 2007).

Soderstrom, Ulf, Paul Soderlind, and Anders Vedrin, "New-Keynesian Models and Monetary Policy: A Re-examination of the Stylized Facts," *Scandinavian Journal of Economics*, 107 (September 2005), 521-46.

Steinsson, Jon, "Optimal Monetary Policy in an Economy with Inflation Persistence," *Journal of Monetary Economics*, 50 (October 2003), 1425-56.

Stock, James H., Jonathan H. Wright, and Motohiro Yogo, "A Survey of Weak Instruments and Weak Identification in Generalized Method of Moments," *Journal of Business and Economic Statistics*, 20 (October 2002), 518–28.

Stock, James, and Motohiro Yogo, "Testing for Weak Instruments in Linear IV Regression," in *Identification and Inference for Econometric Models*, ed. by Donald W. K. Andrews and James H. Stock, Cambridge Univesity Press (Cambridge: 2005).

Svensson, Lars E. O., "Open-Economy Inflation Targeting," Journal of International Economics, 50 (February 2000), 155-83.

Taylor, John B., "Staggered Contracts in a Macro Model," *American Economic Review*, 69 (May 1979), 108-13.

——, "Aggregate Dynamics and Staggered Contracts," *Journal of Political Economy*, 88 (February 1980), 1-23.

Tillmann, Peter, "Does the Cost Channel of Monetary Transmission Explain Inflation Dynamics?," unpublished, University of Bonn (May 2006).

Walsh, Carl E., "Monetary Policy trade-offs in the Open Economy," unpublished, University of California at Santa Cruz (November 1999).

——, *Monetary Theory and Policy*, 2nd ed., MIT Press (Cambridge, Mass.: 2003).

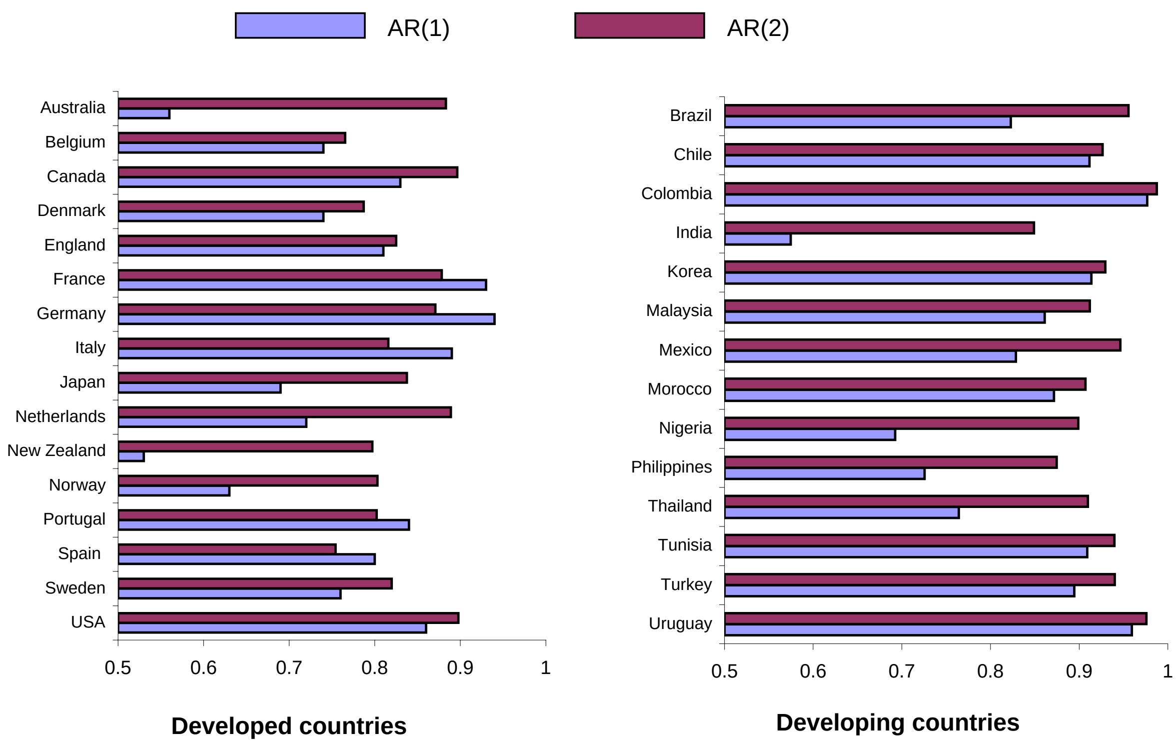
Wang, Pengfei, and Yi Wen, "Inflation Dynamics: A Cross-Country Investigation," *Journal of Monetary Economics*, 54 (October 2007), 2004-31.

Yilmaz, Akyuz, and Korkut Boratav, "The Making of the Turkish Financial Crisis," *World Development*, 31 (September 2003), 1549-66.

Yun, Tack, "Nominal Price Rigidity, Money Supply Endogeneity, and Business Cycles," *Journal of Monetary Economics*, 37 (April 1996), 345–70.

Zhang, Chengsi, Denise R. Osborn, and Dong Heon Kim, "The New Keynesian Phillips Curve: From Sticky Inflation to Sticky Prices," Working Paper No. 07-15, Institute of Economic Research, Korea University (July 2007).

Figure 1 - Measures of Inflation Persistence



Source:For developed countries Wang and Wen (2007) and for developing countries authors' calculation.
Note: The degree of persistence is measured by the first-order autoregressive coefficient in the AR(1) model of inflation, and the sum of the first- and second-order autoregressive coefficients in the AR(2) model.

Figure 2
Goodness of Fit: Predicted and Actual Inflation

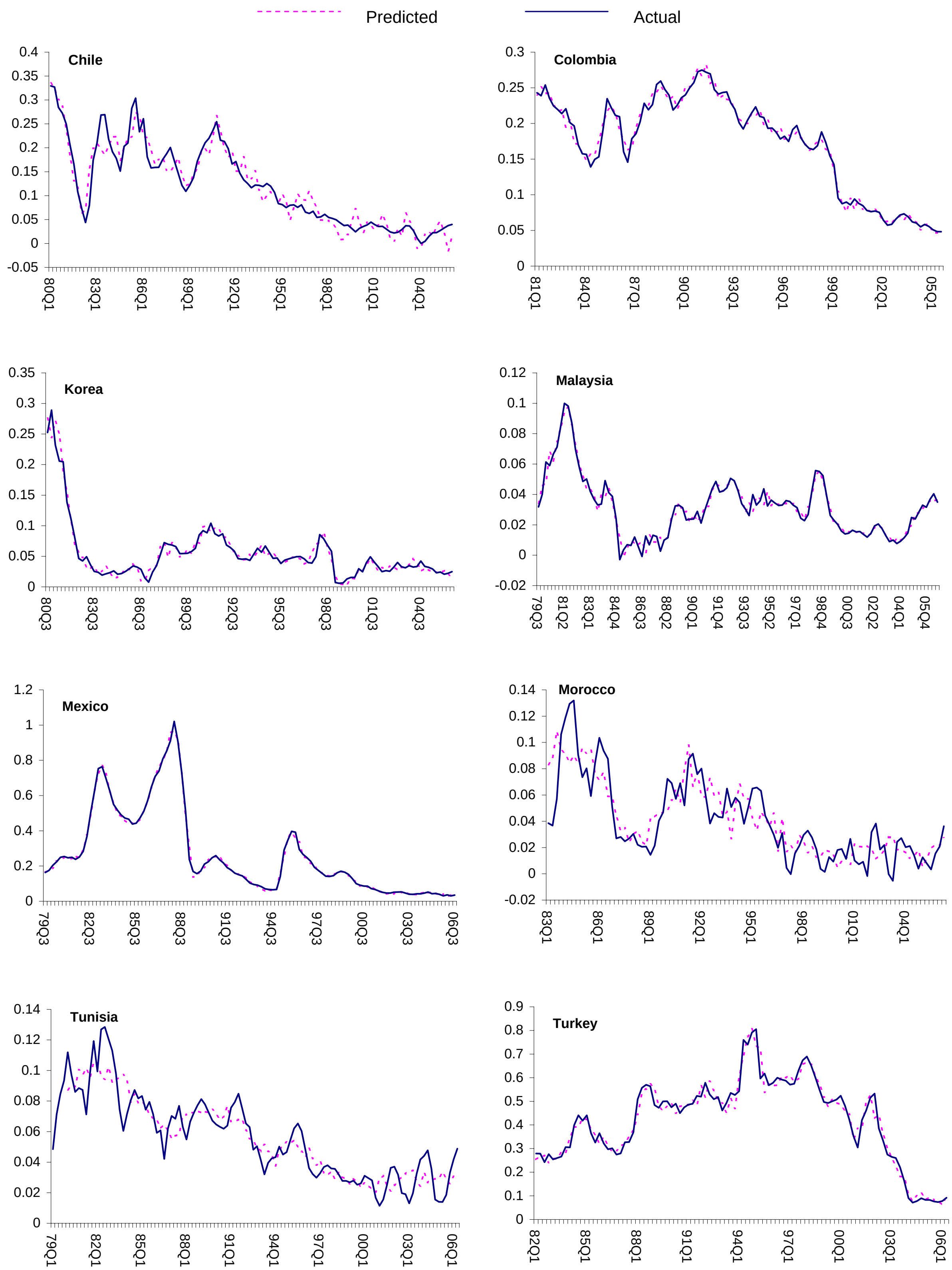


Figure 3
Turkey: Recursive Test for lead inflation

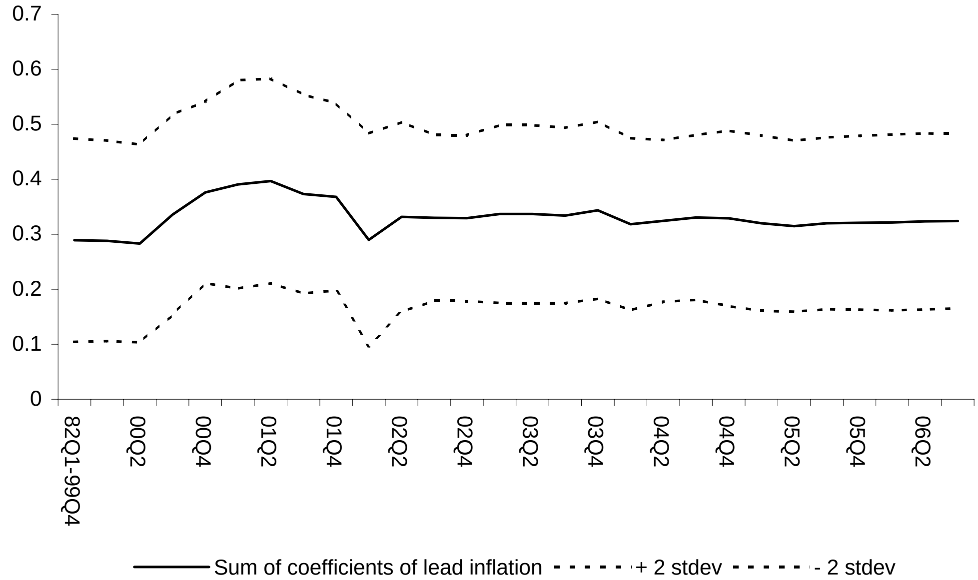


Table 1 - Estimation Results

	Chile				Colombia				Korea				Malaysia			
	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)
π_{t-1}	0.598	0.586	0.421		0.468	0.474	0.619		0.474	0.485	0.117		0.532	0.514	0.699	
	(9.98)	(11.363)	(2.191)		(4.982)	(4.353)	(4.264)		(12.056)	(10.602)	(0.704)		(12.611)	(11.846)	(9.555)	
π_{t-2}			0.131				-0.129				0.257				-0.141	
			(0.917)				(-1.448)				(2.251)				(-2.468)	
π_{t+1}	0.397	0.532	0.515		0.537	0.501	0.548		0.570	0.512	0.450		0.481	0.562	0.521	
	(6.343)	(5.746)	(4.295)		(5.585)	(2.246)	(2.836)		(10.018)	(4.592)	(3.214)		(9.876)	(4.429)	(5.457)	
π_{t+2}		-0.129	0.074			0.031	-0.253			0.055	0.573			-0.068	-0.171	
		(-1.955)	(0.315)			(0.219)	(-0.875)			(0.716)	(1.708)			(-0.651)	(-1.525)	
π_{t+3}			-0.154				0.219				-0.343				0.086	
			(-0.965)				(1.08)				(-1.432)				(1.257)	
$\Sigma\pi_{t-i}$			0.552	0.331			0.490	0.389			0.374	0.530			0.558	0.504
			(7.231)	(4.598)			(5.14)	(3.93)			(5.469)	(7.603)			(14.956)	(4.908)
$\Sigma\pi_{t+i}$		0.404	0.435	0.669		0.532	0.513	0.611		0.567	0.679	0.470		0.493	0.436	0.496
		(7.635)	(5.762)	(4.598)		(4.797)	(5.238)	(3.93)		(9.443)	(7.364)	(7.603)		(11.044)	(10.901)	(4.908)
$\Sigma\hat{y}_{t-i}$	0.087	0.063	0.070	0.390	0.206	0.209	0.215	0.514	0.072	0.067	0.098	0.129	-0.008	-0.009	0.001	-0.008
	(1.144)	(0.863)	(1.041)	(1.888)	(2.446)	(2.225)	(2.454)	(2.564)	(2.018)	(2.001)	(2.462)	(1.837)	(-0.605)	(-0.665)	(0.089)	(-0.167)
$\hat{z}_t$	0.082	0.055	0.069	0.499					0.024	0.024	0.022	0.071	0.002	0.001	0.005	0.050
	(1.306)	(0.845)	(1.068)	(3.756)					(1.322)	(1.333)	(0.863)	(1.504)	(0.206)	(0.142)	(0.416)	(1.821)
$\hat{R}_t$	0.051	0.075	0.071		0.032	0.034	0.029	0.115	0.329	0.326	0.466	0.767	0.036	0.035	0.056	0.218
	(1.077)	(1.601)	(1.633)		(1.081)	(1.014)	(0.934)	(1.422)	(3.437)	(3.298)	(2.94)	(2.864)	(0.747)	(0.802)	(1.253)	(1.178)
$\hat{p}_t^J$	0.012			0.073	0.010	0.019	0.008	0.046				0.013	0.002	0.002	0.001	0.003
	(2.021)			(3.105)	(1.198)	(1.084)	(0.961)	(2.204)				(0.08)	(0.506)	(0.555)	(0.192)	(0.286)
$\hat{p}_t^J - \hat{w}_t$		0.010	0.012						0.008	0.007	0.008	0.002				
		(1.933)	(1.448)						(0.85)	(0.777)	(0.698)	(0.012)				
Adj-R2	0.974	0.973	0.973	0.898	0.978	0.977	0.978	0.944	0.925	0.924	0.895	0.828	0.942	0.942	0.942	0.810
No of obs	105	105	105	105	99	99	99	99	103	102	101	100	109	108	107	106
J-test	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Test for weak IV	51.116	51.116	51.116	51.116	47.607	47.607	47.607	47.607	9.956	9.956	9.956	9.956	14.960	14.960	14.960	14.960
F-test for restriction				0.004				1.108				0.781				1.678

Note: The estimation method is the two-stage GMM. Instruments are inflation (π), cyclical components of output($\hat{y}$), interest rate ($\hat{R}$), real effective exchange rate ($\hat{z}$), oil price ($\hat{p}_t^J$) and the difference between oil prices and wages ($\hat{p}_t^J - \hat{w}_t$), all from t-1 to t-4. t-statistics are given in parenthesis. P(-1,1), P(-1,2), P(-2,3) and P(-4,4) are different equation specifications as explained in the text. $\Sigma\pi_{t-i}$ stands for the sum of lagged inflation coefficients for i = 1 to 3 for P(-2,3). $\Sigma\pi_{t+i}$ stands for the sum of lead inflation coefficients for i = 1 to 2 for P(-1,2), and for i = 1 to 3 for P(-2,3). $\Sigma\hat{y}_{t-i}$ stands for sum of coefficients for output gap for i = 0 to 4. (it continues …)

Table 1 (continued) - Estimation Results

	Mexico				Morocco				Tunisia				Turkey			
	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)	PC(-1,1)	PC(-1,2)	PC(-2,3)	PC(-4,4)
π_{t-1}	0.575	0.694	0.806		0.661	0.657	0.922		0.578	0.584	0.832		0.638	0.635	0.769	
	(10.441)	(6.367)	(8.239)		(11.403)	(10.917)	(6.57)		(7.736)	(7.324)	(4.502)		(9.816)	(9.918)	(4.563)	
π_{t-2}			-0.205				-0.281				-0.170				-0.102	
			(-4.323)				(-2.031)				(-1.258)				(-0.837)	
π_{t+1}	0.433	-0.188	0.536		0.285	0.217	0.242		0.429	0.438	0.281		0.358	0.394	0.382	
	(8.116)	(-2.934)	(5.145)		(3.498)	(2.023)	(1.912)		(5.224)	(3.005)	(2.122)		(5.605)	(3.493)	(2.734)	
π_{t+2}		0.085	-0.219			0.087	-0.280			-0.020	0.046			-0.035	-0.278	
		(0.592)	(-2.009)			(1.137)	(-1.203)			(-0.21)	(0.264)			(-0.423)	(-1.528)	
π_{t+3}			0.080				0.400				-0.007				0.221	
			(1.216)				(2.045)				(-0.047)				(1.873)	
$\Sigma\pi_{t-i}$			0.602	0.294			0.642	0.635			0.662	0.512			0.667	0.451
			(10.288)	(1.744)			(10.042)	(5.2)			(7.81)	(6.126)			(8.361)	(5.337)
$\Sigma\pi_{t+i}$		0.506	0.396	0.706		0.304	0.362	0.365		0.418	0.320	0.488		0.359	0.324	0.549
		(9.44)	(6.645)	(1.744)		(3.559)	(3.422)	(5.2)		(4.898)	(3.269)	(6.126)		(5.753)	(4.077)	(5.337)
$\Sigma \hat{y}_{t-i}$	0.127	0.222	0.136	0.354	0.150	0.179	0.276	0.439	0.018	0.011	0.053	0.098	0.160	0.170	0.088	-0.078
	(0.955)	(1.632)	(0.912)	(0.379)	(0.986)	(1.155)	(1.898)	(1.544)	(0.439)	(0.254)	(0.999)	(0.864)	(0.97)	(0.953)	(0.46)	(-0.137)
$\hat{z}_t$									0.002	0.001	0.028		0.193	0.194	0.254	0.408
									(0.038)	(0.023)	(0.54)		(2.23)	(2.377)	(2.63)	(2.744)
$\hat{R}_t$	0.133	0.001	0.106	0.656			0.081	0.194			0.041					
	(2.513)	(0.211)	(2.369)	(2.024)			(0.449)	(0.496)			(0.508)					
$\hat{p}_t^J$	0.003			0.027					0.001	0.001	0.001	0.004				
	(0.27)			(0.402)					(0.372)	(0.189)	(0.298)	(0.578)				
$\hat{p}_t^J - \hat{w}_t$									0.007	0.006	0.008	0.013	0.026	0.028	0.038	0.071
									(0.755)	(0.635)	(0.741)	(0.333)	(1.475)	(1.67)	(2.005)	(1.241)
Adj-R2	0.994	0.995	0.996	0.934	0.863	0.857	0.847	0.650	0.937	0.937	0.918	0.742	0.960	0.958	0.952	0.885
No of obs	105	104	103	102	93	92	91	90	92	91	90	89	100	99	98	97
J-test	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Test for weak IV	89.848	89.848	89.848	89.848	9.300	9.300	9.300	9.300	12.126	12.126	12.126	12.126	26.955	26.955	26.955	26.955
F-test for restriction				0.302				0.782				0.028				0.671

Note (continued): In the regressions, no restriction used except in P(-4,4). In this specification, it is assumed that lagged inflation as well as lead inflation have common coefficients and the sum of these coefficients is one. F-test for restriction is given in the last row of the table where H_0:$\delta^L + \delta^F = 1$ in P(-4,4) specification. All test result indicates that we fail to reject H_0. J-test is for overidentification problem where H_0: there is no overidentification problem. We fail to reject in each case. F-test for weak instrumental variables (test for weak IV) is also reported, where H_0: instruments are weak. We reject weakness for each country given critical values reported by Stock and Yogo (2004) in Table 1 of their paper when the desired maximal bias of the IV estimator relative to OLS is taken as 0.20. Note that the test result is the same for each country since the set of instrumental variables used is the same for each specification.

Table 2 - Nested and Non-nested Specification Tests

Non-nested Test

	H_0: P(-1,1)			H_0: P(-1,2)			H_0: P(-2,3)			H_0: P(-4,4)			Best specification
	H_a: P(-1,2)	H_a: P(-2,3)	H_a: P(-4,4)	H_a: P(-1,1)	H_a: P(-2,3)	H_a: P(-4,4)	H_a: P(-1,1)	H_a: P(-1,2)	H_a: P(-4,4)	H_a: P(-1,1)	H_a: P(-1,2)	H_a: P(-2,3)	
Chile	1.855	2.300	-2.106	-0.170	…	-1.468	-1.197	…	-2.444	1.106	1.206	1.186	**P(-4,4)**
	0.067	0.024	0.038	0.866	…	0.146	0.235	…	0.017	0.271	0.231	0.239	
Colombia	…	…	-0.857	…	…	-1.206	…	…	-0.811	1.813	1.798	1.787	**P(-2,3)**
	…	…	0.394	…	…	0.231	…	…	0.420	0.073	0.076	0.077	
Korea	…	…	1.045	…	…	0.978	…	…	-1.424	2.667	2.598	2.407	**P(-2,3)**
	…	…	0.299	…	…	0.331	…	…	0.158	0.009	0.011	0.018	
Malaysia	…	…	-2.790	…	…	-3.142	…	…	-1.424	2.971	3.051	3.158	**P(-2,3)**
	…	…	0.006	…	…	0.002	…	…	0.158	0.004	0.003	0.002	
Mexico	2.734	7.182	-4.062	-0.447	…	-2.744	-0.344	…	1.131	1.400	1.492	1.542	**P(-2,3)**
	0.008	0.000	0.000	0.656	…	0.007	0.732	…	0.261	0.165	0.139	0.127	
Morocco	…	…	-1.042	…	…	-2.459	…	…	-3.386	3.054	2.872	2.674	**P(-4,4)**
	…	…	0.301	…	…	0.016	…	…	0.001	0.003	0.005	0.009	
Tunisia	…	…	-3.259	…	…	-3.902	…	…	-2.837	2.003	2.039	2.233	**P(-4,4)**
	…	…	0.002	…	…	0.000	…	…	0.006	0.049	0.045	0.028	
Turkey	…	…	-0.635	…	…	-0.618	…	…	-0.678	2.143	2.173	2.130	**P(-2,3)**
	…	…	0.527	…	…	0.538	…	…	0.500	0.035	0.033	0.036	

Nested Test

	H_0: $\delta_2=\delta_4=\delta_5=0$ in P(-2,3)		H_0: $\delta_2=\delta_5=0$ in P(-2,3)		H_0: $\delta_4=0$ in P(-1,2)	
Chile	…		9.241	(reject Ho)	…	
	…		0.000		…	
Colombia	4.858	(reject Ho)	7.273	(reject Ho)	3.674	(reject Ho)
	0.004		0.001		0.059	
Korea	11.717	(reject Ho)	7.572	(reject Ho)	5.565	(reject Ho)
	0.000		0.001		0.021	
Malaysia	15.334	(reject Ho)	19.463	(reject Ho)	0.003	(fail to reject Ho)
	0.000		0.000		0.955	
Mexico	…		13.815	(reject Ho)	…	
	…		0.000		…	
Morocco	3.479	(reject Ho)	2.526	(reject Ho)	3.538	(reject Ho)
	0.020		0.087		0.064	
Tunisia	1.638	to reject Ho)	2.393	(reject Ho)	0.339	(fail to reject Ho)
	0.188		0.099		0.562	
Turkey	4.288	(reject Ho)	5.122	(reject Ho)	0.198	(fail to reject Ho)
	0.007		0.008		0.658	

Note: The non-test test is J-test. The null hypothesis is defined by different specification. The rejection rule is that reject H_0 if the fitted value from H_a specification enter significantly in H_0 specification. t-statistics of the fitted values are reported in the table. p-values are reported under t-statistics. Nested test is a wald test (F-test). p-values are reported under F-statistics. H_0: $\delta_2=\delta_4=\delta_5=0$ tests P(-2,3) against P(-1,1), H_0: $\delta_2=\delta_5=0$ tests P(-2,3) specification against P(-1,2), and H_0: $\delta_4=0$ tests specification P(-1,2) against P(-1,1).

Table 3
Stability Test: Quandt-Andrews Unknown Breakpoint Test

Ho: No breakpoints

	Maximum Wald F-statistics	**P-value**	**Breakpoint**
Chile	5.905	1.000	1987 Q1
Colombia	1.680	1.000	1997 Q1
Korea	4.361	1.000	1988 Q4
Mexico	1.793	1.000	1988 Q1
Morocco	1.730	1.000	1993 Q1
Turkey	1.281	1.000	1994 Q1

Note: The Hansen's p-values are used in the table (Hansen, 1997).

Table A1
Order of Integration: Unit Root Test Statistics

Country	Variabl	k	ADF test test statistic	PP test	Country	Variat	k	ADF test test statistic	PP test
Chile	$\hat{\pi}$	2	-2.87*	-1.58	Mexico	$\hat{\pi}$	2	-3.78**	-1.91
	$\hat{y}$	0	-9.60***	-9.58***		$\hat{y}$	0	-5.70***	-5.89***
	$\hat{z}$	0	-4.92***	-4.55***		$\hat{z}$	0	-3.69***	-4.25***
	$\hat{R}$	0	-8.20***	-8.21***		$\hat{R}$	0	-4.67***	-4.93***
	$\hat{p}_t^J$	0	-4.44***	-4.66***		$\hat{p}_t^J$	0	-4.44***	-4.66***
	$\hat{p}_t^J - \hat{w}_t$	0	-4.20***	-4.52***		$\hat{p}_t^J - \hat{w}_t$	0	-4.58***	-4.87***
Colombia	$\hat{\pi}$	3	-2.28	-0.49	Morocco	$\hat{\pi}$	2	-4.59***	-2.13
	$\hat{y}$	0	-9.42***	-10.33***		$\hat{y}$	0	-8.01***	-15.03***
	$\hat{z}$	0	-4.60***	-4.42***		$\hat{z}$	0	-5.45***	-5.12***
	$\hat{R}$	0	-4.54***	-3.91***		$\hat{R}$	0	-4.98***	-5.12***
	$\hat{p}_t^J$	0	-4.44***	-4.66***		$\hat{p}_t^J$	0	-3.97***	-4.57***
	$\hat{p}_t^J - \hat{w}_t$	0	-4.03***	-4.19***		$\hat{p}_t^J - \hat{w}_t$	0	-4.81***	-5.15***
Korea	$\hat{\pi}$	3	-4.34***	-2.17	Tunisia	$\hat{\pi}$	2	-4.83***	-1.65
	$\hat{y}$	0	-3.78***	-3.79***		$\hat{y}$	0	-10.93***	-10.99***
	$\hat{z}$	0	-4.33***	-4.56***		$\hat{z}$	0	-4.63***	-4.63***
	$\hat{R}$	0	-3.48**	-2.73*		$\hat{R}$	0	-3.49**	-3.49**
	$\hat{p}_t^J$	0	-4.44***	-4.66***		$\hat{p}_t^J$	0	-4.44***	-4.66***
	$\hat{p}_t^J - \hat{w}_t$	0	-5.34***	-4.55***		$\hat{p}_t^J - \hat{w}_t$	0	-3.83***	-4.07***
Malaysia	$\hat{\pi}$	2	-3.23*	-2.37	Turkey	$\hat{\pi}$	3	-2.35	-1.76
	$\hat{y}$	0	-4.41***	-4.68***		$\hat{y}$	0	-9.69***	-10.05***
	$\hat{z}$	0	-3.86***	-4.34***		$\hat{z}$	0	-6.56***	-7.46***
	$\hat{R}$	0	-3.31**	-3.79***		$\hat{R}$	0	-6.57***	-6.56***
	$\hat{p}_t^J$	0	-4.44***	-4.66***		$\hat{p}_t^J$	0	-4.44***	-4.66***
	$\hat{p}_t^J - \hat{w}_t$	0	-4.16***	-4.45***		$\hat{p}_t^J - \hat{w}_t$	0	-4.89***	-4.77***

Notes: Variables and estimated period are as defined in the text. k denotes the number of lags in the ADF test. Asterisks *, ** and *** denote rejection of the null hypothesis of a unit root at the 10%, 5%, and 1% significance levels. Critical values are from McKinnon (1991).