Studying the Short–Run Dynamics of Inflation:

Estimating a Hybrid New-Keynesian Phillips Curve for Argentina (1993- 2007)[1]

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A distinctive feature of the modeling of inflation dynamics in the short-run is the introduction of some nominal rigidity in the context of inter-temporal optimizing behavior by non-competitive forward-looking firms (Taylor, 1980; Calvo, 1983). Galí and Gertler (1999) extend the model, allowing for a portion of the firms to follow a backward-looking rule to set prices and obtain a "Hybrid New-Keynesian Phillips Curve". Recent developments consider a positive trend inflation in the standard Neo-Keynesian Phillips Curve (Ascari and Ropele, 2007; Kiley, 2007).

We use the Generalized Methods of Moments to estimate a "Hybrid New-Keynesian Phillips curve" for Argentina between 1993 and 2007. We extend the model to a small open economy, considering separately the influence of nominal devaluation and foreign inflation on domestic prices.

For the whole sample, the results indicate that forward and backward looking components are relevant to explain the dynamics of domestic prices, although the backward-looking term weights more in determining inflation dynamics. Nominal devaluation and foreign inflation are also significant to explain domestic inflation behavior, being the response of inflation to the second more intense. The output gap, although weak, has a significant effect on inflation.

Given the change in the monetary regime of 2002, we test for parameter stability. In line with the recent theoretical literature on trend inflation, we find that the influence of the output gap weakens in the managed float period. Along with the observed increase in trend inflation, the Phillips Curve becomes more forward looking.

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1. Introduction

Assessing the short run dynamics of inflation is a relevant issue for monetary policy. A distinctive feature of the modeling of inflation in the short-run is the introduction of some nominal rigidity in the context of inter-temporal optimizing behavior by non-competitive forward-looking firms. In these models, built on earlier work by Fischer (1977), Taylor (1980), and Calvo (1983), price stickiness could arise for different reasons. In Calvo's (1983) setting some sluggishness in price formation could be obtained by assuming that forward-looking firms face constraints on price adjustment. The resulting model is a New Keynesian, forward-looking version of the traditional Phillips Curve. The empirical relevance of inflation persistence, which imposes costs for disinflation policies has led to incorporate inflation inertia in these models, in spite of the theoretical difficulties to justify it. Galí and Gertler (1999) extend the Calvo's model, allowing for a portion of the firms to follow a backward-looking rule to set prices and obtain a "Hybrid New Keynesian Phillips Curve".

Based on these theoretical grounds, an empirical literature has developed and many issues related to theoretical and empirical aspects of these models are currently under debate. Models based on Calvo's (1983) setting have been subject to the critique of being quite unrealistic in assuming that firms should not expect to adjust prices in a finite horizon and it has been suggested that some truncation should be introduced to add them a quote of realism. The use of the output gap as a measure of marginal costs has also been questioned for both theoretical and empirical reasons. Galí and Gertler (1999) suggest using the aggregate labor income share as a a measure of marginal costs instead of the output gap.

Recent developments in the modeling of inflation dynamics (see Blake and Fernández-Corugedo, 2006; Ascari and Ropele, 2007 and Kiley, 2007) extend the standard Neo-Keynesian Phillips Curve, that assumes a zero trend inflation in the steady state, to allow for a positive trend inflation. In this context, trend inflation affects the dynamics of the standard Neo-Keynesian model. As inflation becomes less influenced by current marginal costs, the coefficient of the output gap lowers. At the same time price setting becomes more forward looking as does inflation. A higher trend inflation also leads to a stronger autoregressive component making the inflation process more persistent.

We study here how well the so-called "Hybrid New Keynesian Phillips Curve" approximates the dynamics of inflation in Argentina over the period 1993-2007. The standard model is modified to capture the role played by nominal devaluation and foreign inflation in domestic prices dynamics in a small open economy. This is a difficult task in the case of Argentina, because the economy went into structural changes after the devaluation of the peso that followed the financial and currency crises of 2001-2002. In this context, it is highly probable that the dynamics of price setting has changed after the abandonment of the currency board regime and the adoption of a dirty float. Taking into account the new literature on trend inflation and its implications for the short run dynamics of inflation, we evaluate the impact of this change on the parameters of the Phillips curve.

To estimate this New Keynesian version of the Phillips Curve we use the Generalized Method of Moments (GMM), which seems to be the appropriate method under rational expectations, since it is based on the assumption that the error in forecasting inflation by firms is orthogonal to the available information.

The paper is organized as follows: in section 2 we briefly present some theoretical developments in modeling inflation dynamics. Section 3 describes the estimation

methodology. In section 4, we present the empirical results. In section 5 we evaluate the parameter stability of our Phillips Curve. Finally, section 6 concludes.

2. Modeling inflation dynamics

In the hybrid version of the Phillips Curve proposed by Galí and Gertler (1999) inflation follows the process

$$\pi_t = \phi\pi_{t-1} + (1-\phi)E_t(\pi_{t+1}) + \delta mc_t + \varepsilon_t \qquad (1)$$

Where π_t is the inflation rate at time t, E_t is the expectation of inflation on $t+1$ at time t, mc_t is the marginal cost and ε_t is a random shock. The assumption that $0 < \phi < 1$, implies a vertical Phillips curve in the long run. The lagged term in inflation introduces some backwardness in price setting, an observable feature of inflation dynamics, which is quite difficult to justify from a theoretical point of view. In Calvo's framework, firms operate in a monopolistically competitive environment and face some constraints in prices setting in the form of a time dependent rule of adjustment. More specifically each firm faces a constant probability ($1-\theta$) of adjusting prices in period t and a corresponding constant probability θ of maintaining its prices unchanged.

$$p_t = (1-\theta)\sum_{j=0}^{\infty}\theta^j p^*_{t-j} = \theta\, p_{t-1} + (1-\theta)p^*_t \quad (2)$$

This implies that the price level in t is a convex combination of prices optimally set in previous periods p_{t-j} and prices optimally set in t p^*_t according to

$$p^*_t = (1-\beta\theta)\sum_{j=0}^{\infty}(\beta\theta)^j E_t\{mc_{t+j}\}(3)$$

Which assumes that firms are identical and choose the same p^*_t according to their expected marginal costs for future periods mc_{t+j}, discounted at the subjective factor β.

Combining (2) and (3) an inflation equation can be written as

$$\pi_t = \lambda mc_t + \beta E_t \pi_{t+1} \quad (4)$$

Where $\pi_t = p_t - p_{t-1}$ and $\lambda = (1-\theta)(1-\beta\theta)/\theta$.

Galí and Gertler (1999) introduce backwardness in the Calvo's price setting model (1983) and use the labor income share as a measure of marginal cost instead of the output gap, as suggested by the theoretical literature. They assume that while a fraction ($1-\omega$) of the firms that adjusts prices in t follows the optimizing behavior described by (3) a proportion ω uses a rule of thumb based on past prices to adjust prices in t. Thus, prices adjusted in t, now referred as $\bar{p}^*_t$ are set according to

$$\bar{p}^*_t = (1-\omega)p^f_t + \omega p^b_t \quad (5)$$

The fraction (*1-ω*) of the firms behaves according to (3)

$$p_t^f = (1 - \beta\theta)\sum_{j=0}^{\infty}(\beta\theta)^j E_t\{mc_{t+j}\} \quad (3')$$

where p_t^f indicates prices set according to (3) and the proportion ω behaves according to

$$p_t^b = \bar{p}_{t-1}^* + \pi_{t-1} \quad (6)$$

where p_t^b represents prices adjusted following a backward looking rule.

Combining equations *(2), (5), (3')* and *(6)* a Hybrid Phillips Curve is obtained

$$\pi_t = \lambda mc_t + \gamma_f E_t\{\pi_{t+1}\} + \gamma_b\pi_{t-1} \quad (7)$$

Where

$$\lambda \equiv (1-\omega)(1-\theta)(1-\beta\theta)\phi^{-1},$$
$$\gamma_f \equiv \beta\theta\phi^{-1},$$
$$\gamma_b \equiv \omega\phi^{-1}, \qquad\qquad (7')$$

with $\phi \equiv \theta + \omega[1 - \theta(1-\beta)]$ (7'')

We adapt Galí and Gertler specification to the case of a small open economy. As pointed out by Svensson (1998) changes in the nominal exchange rate and foreign prices have, in this context, a direct effect on domestic inflation. In addition, since the nominal exchange rate is the price of an asset, it is inherently a forward-looking variable. Thus, as a determinant of domestic inflation it contributes to make expectations play an essential role in domestic prices formation.

We then estimate an open economy version of the "Hybrid New Keynesian Phillips Curve" that modifies equation *(1)* in two directions: (i) introducing measures of nominal devaluation and foreign inflation and (ii) using a measure of the output gap as a proxy for marginal costs rather than the labor income share.

Thus, our equation is as follows:

$$\pi_t = \phi_1\pi_{t-1} + \phi_2 E_t(\pi_{t+1}) + \gamma\pi_t^* + \lambda\Delta e_t + \delta x_t + \varepsilon_t \qquad (8)$$

were π_t is domestic inflation, measured by the change in the log of the Consumer Price Index, $E_t(\pi_{t+1})$ is inflation expectation for *t+1* at time *t*, π_t^* is foreign inflation, measured by

the change in the log of the US Producer Price Index, Δe_t is nominal devaluation calculated as the change in the log of the nominal exchange rate, and x_t is the output gap.[2]

Taking into account the recent developments in the theoretical literature (Blake and Fernández-Corugedo, 2006; Ascari and Ropele, 2007 and Kiley, 2007) and given the regime change of 2002 we also evaluate the parameter stability of the estimated Phillips Curve. In line with the predictions of this literature, which suggest that a higher trend inflation leads to a weaker impact of the output gap on inflation, and to a higher forward looking component of inflation, we evaluate the constancy of the estimated coefficients for the backward and forward looking components of inflation and the output gap.

3. The estimation methodology

Under rational expectations economic agents are supposed to use current and past information efficiently. In terms of equation (8) this implies that the error in forecasting future inflation (π_{t+1}) is uncorrelated to the set of information z_t available at date t, that is

$$E\{(\pi_t - \phi_1 \pi_{t-1} - \phi_2 \pi_{t+1} - \gamma \pi_t^* - \lambda \Delta e_t - \delta x_t) z_t\} = 0 \qquad (9)$$

Where z_t is a vector of variables (instruments) dated at t and earlier. A natural way to deal with the estimation of equation (1) is to use the Generalized Method of Moments (GMM), developed by Hansen (1982) which is a generalization of the method of moments. In what follows we present a brief description of GMM and some methodological issues related to time series estimation using this method. We stress two main advantages of the GMM estimation: (i) it does not require imposing a certain probability distribution to the variables and (ii) it is consistent with the assessments of inter-temporal optimizing behavior by economic agents.

Suppose we have a set of observations of a random variable y, whose probability function depends on a vector of k unknown parameters denoted by θ. We can then define

$$E(g(y_t, \theta)) = 0 \quad for \ \theta = \theta_0 \ (10)$$

as a vector of the moment conditions of y.

The sample counterpart of the population moment condition is

$$g_T(\theta) = \frac{\sum_{t=1}^{T} g(y_t, \theta)}{T} \qquad (11)$$

If the number of moment conditions is equal to the number of parameters to be estimated, $a=k$, we have a system of k equations and k unknowns, which can be perfectly identified.

The Method of Moments estimator $\hat{\theta}$ can be defined as that which equals the sample moment with the population moment.

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[2] The nominal exchange rate corresponds to the multilateral exchange rate with the three main trade partners of Argentina: Brazil, the United States and the European Union. The output gap is calculated according to the Production Function methodology.

$$g_t(\overset{\wedge}{\theta})=\frac{\sum\limits_{t=1}^{T} g(y_t,\overset{\wedge}{\theta})}{T}=0 \qquad (12)$$

If the number of moment conditions exceeds the number of unknown parameters, *a>k,* the system is over-identified, since there does not exist a unique $\overset{\wedge}{\theta}$ satisfying (12). The Generalized Method of Moments proposes to use $\overset{\wedge}{\theta}$

$$\overset{\wedge}{\theta}_{GMM} \equiv arg\, min\, g_t(\theta)' C_T g_t(\theta) \quad (13)$$

where C_T is a symmetric positive definite matrix, known as the "weighting matrix" that weights the moment conditions as to solve *(13)*.

Hansen (1982) proposes a method to chose C_T optimally, that is, to obtain the $\overset{\wedge}{\theta}$ with the minimum asymptotic variance

$$C_T \overset{p}{\longrightarrow} \partial E[g_T(\theta_0)g_T(\theta_0)']$$

where ∂ is constant.

Hansen shows that given S

$$S = \lim_{T\to\infty} T.E[g_T(\theta_0)g_T(\theta_0)']$$

the optimum value of the matrix C_T is given by S^{-1}, the inverse of the asymptotic variance covariance matrix. Then, the minimum variance estimator of θ is obtained by choosing $\overset{\wedge}{\theta}$ as to minimize

$$Q(\theta)=[g_T(\theta)]' S^{-1}[g_T(\theta)] \qquad (14)$$

Assuming that $g_T(\theta_0)$ is not serially correlated, $\overset{\wedge}{\theta}$ is a consistent estimator of θ_0.

$$\hat{S} \equiv (1/T)\sum_{t=1}^{T} g_t(\overset{\wedge}{\theta})g_t(\overset{\wedge}{\theta})' \overset{p}{\to} S \quad (15)$$

The estimation of $\hat{S}$ requires having a previous estimation of $\overset{\wedge}{\theta}$. Thus, substituting C_T in *(13)* by the identity matrix *I*, an initial estimation of $\overset{\wedge}{\theta}$ is obtained and then used in *(15)* to obtain an initial $\hat{S}_0$. The expression *(14)* is minimized using $S^{-1} = \hat{S}_0{}^{-1}$, to obtain a new estimation of $\overset{\wedge}{\theta}$. The process can be repeated until $\overset{\wedge}{\theta}{}^{j} \cong \overset{\wedge}{\theta}{}^{j+1}$.

If the vector $g_T(\theta_0)$ is serially correlated, the matrix $\hat{S}$ will have the following structure

$$\hat{\Omega}_{HAC} = \hat{\Gamma}(0) + \left(\sum_{j=1}^{T-1} k(j,q)(\hat{\Gamma}(j) + \hat{\Gamma}'(-j)) \right) \quad (16)$$

where

$$\hat{\Gamma}(0) = \frac{1}{T}\left(\sum_{t=1}^{T} g_t(\hat{\theta})g_t(\hat{\theta})' \right)$$

is White's heteroskedasticity consistent covariance matrix and

$$\hat{\Gamma}(j) = \frac{1}{T}\left(\sum_{t=j+1}^{T} g_t(\hat{\theta})g_{t-j}(\hat{\theta})' \right)$$

describes the autocovariances and $k(j,q)$ is a kernel

The matrix $\hat{\Omega}_{HAC}$ is known as the Heteroskedasticity and Autocorrelation Consistent (HAC) Covariance Matrix. The estimation of $\hat{\Omega}_{HAC}$ needs to specify a kernel, used to weight the covariances so that $\hat{\Omega}_{HAC}$ is positive semi-definite and a bandwidth which is a lag truncation parameter for the autocovariances.

Two type of kernel are commonly used in the estimation of $\hat{\Omega}_{HAC}$, Barlett and quadratic spectral.[3]

With regards to the bandwidth selection, different methods have been developed. The E-View program provides three methods: Fixed Newey-West, Variable Newey-West (1994) and Andrews (1991).

The use of the GMM estimator implies that number of orthogonality conditions exceeds the number of parameters to be estimated, thus the model is overidentified, since more orthogonality conditions than needed are being used to estimate the parameters. Hansen (1982) suggests a test of whether all of the sample moments are close to zero as would be expected if the corresponding population moments were truly zero.

Hansen's test of over-identifying restrictions can be conducted using the *J-statistic* reported in E-Views and using it to construct the following statistic:

$$T.J - statistic \sim \chi^2(p-q)$$

where *p* represents the number of orthogonality conditions and *q* the number of parameter to be estimated.

4. Empirical results

4.1. The Argentine framework: a brief descriptive analysis

We estimate equation *(8)* for the period 1993.1-2007.12, using monthly information. This period includes two very different exchange and monetary regimes: a currency board, known as Convertibility, at place between 1993 and 2001 and a dirty float from then on. The Convertibility was a successful attempt to anchor inflation expectations by fixing the peso to the dollar by law. By 1993 inflation had stabilized at very low levels. Although this change

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[3] See E-Views User's Guide for specific functional forms (pgs. 498, 499).

was perceived as being quite permanent, and inflation remained very low, the fiscal reform was rather incomplete. Monetary financing of fiscal disequilibrium was replaced to some extent by external financing. Government and private sector external debt increased over time and began to be perceived as unsustainable once the economy entered a long recession in 1998. There was even a deflationary period during this prolonged recession. In 2001 an external and financial crises unchained leading to the abandonment of the Convertibility regime, a sharp devaluation of the currency and the adoption of a managed float. The devaluation of the currency provoked a dramatic change in relative prices and a jump in the inflation rate, which reached a peak in April 2002. It then returned to low levels, close to those of the Convertibility period, but began to accelerate slightly by the end of 2004, once the economy entered a period of strong growth after the prolonged recession in which it had been immersed for several years.

Recent empirical evidence indicates the presence of a structural break in 2002 (see D'Amato, Garegnani and Sotes Paladino, 2007). Both the mean and the autoregressive dynamics of inflation experienced a break since January 2002. In particular, inflation became more persistent and its trend experienced a slight increase. As mentioned above the recent literature on trend inflation suggests that the Phillips Curve is not neutral to changes in trend inflation. Then, it is highly probable that the relation between inflation and its regressors could have changed with the new regime. To give some intuition about the relation between inflation in and its determinants in both regimes we use graphic analysis.
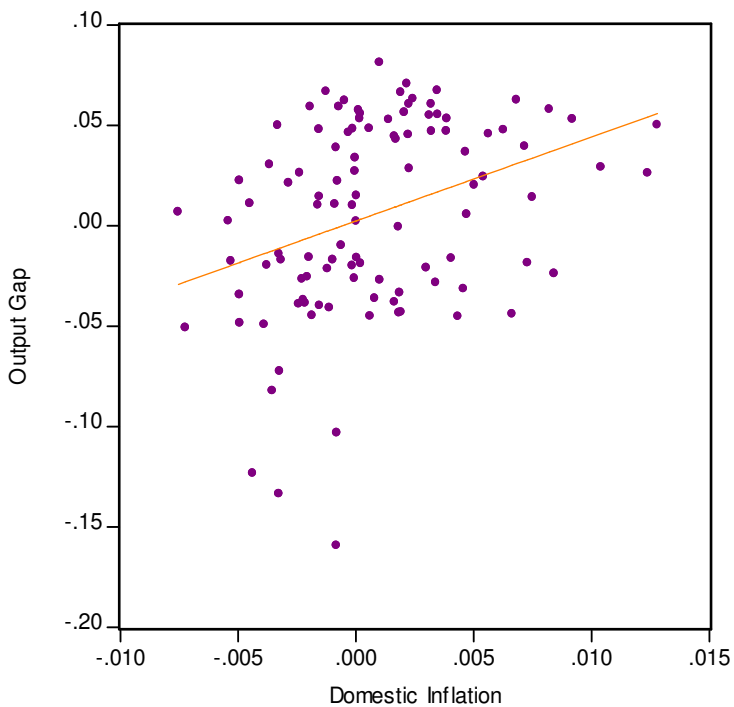
Figure 1

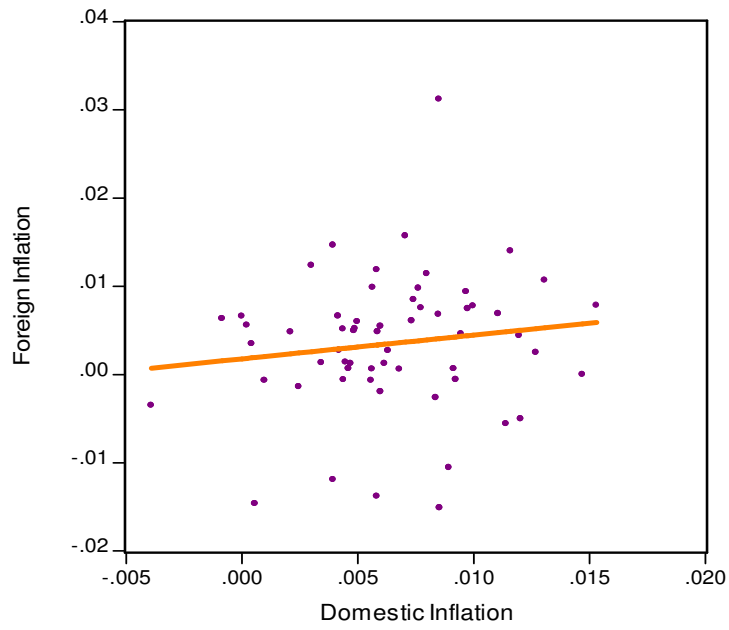Cross-plot Domestic and Foreign Inflation
1993-2001

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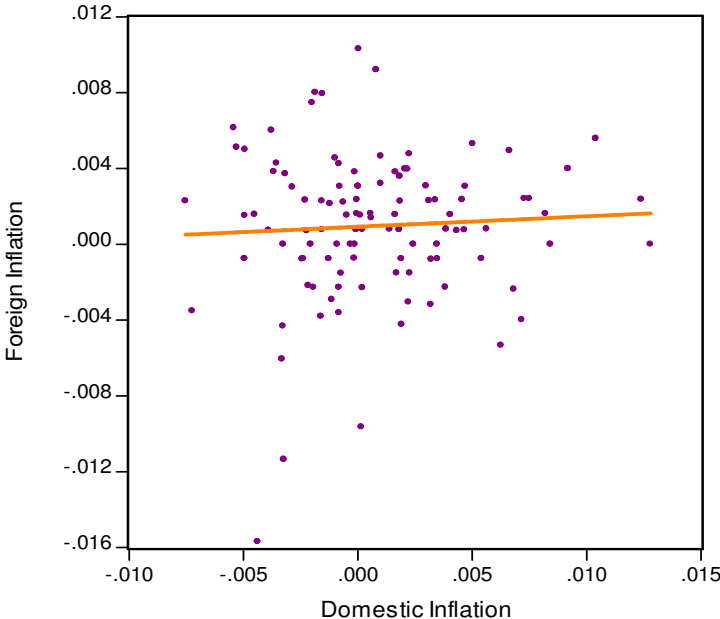
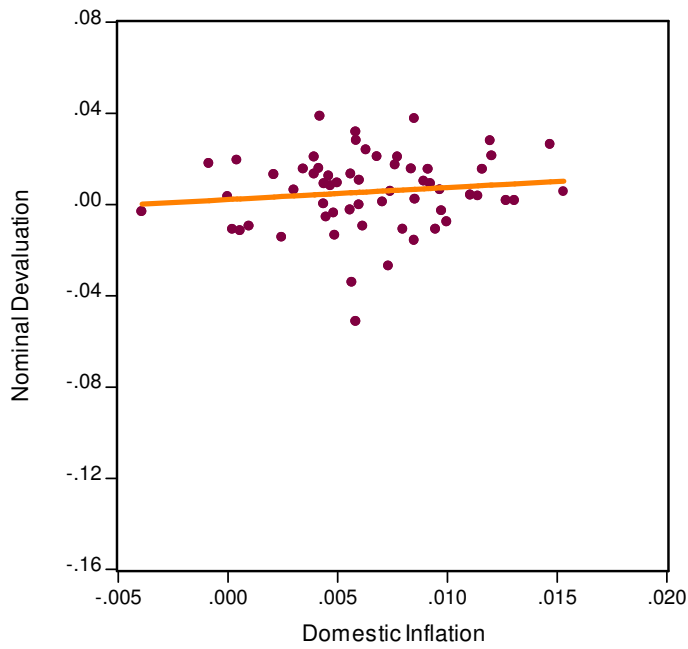Cross-plot Domestic Inflation and Nominal Devaluation
1993-2001
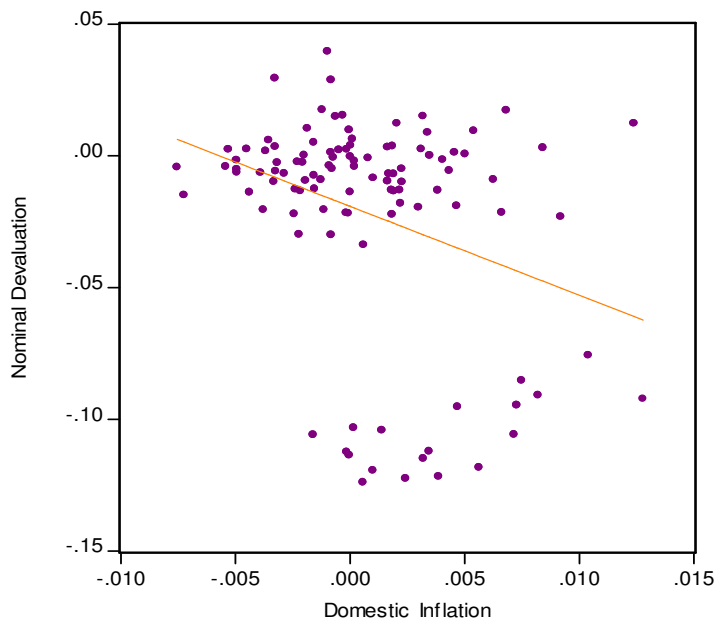
Cross-plot Domestic Inflation and Output Gap
1993-2001

Cross-plot Domestic and Foreign Inflation
2003-2007

Cross-plot Domestic Inflation and Nominal Devaluation
2003-2007

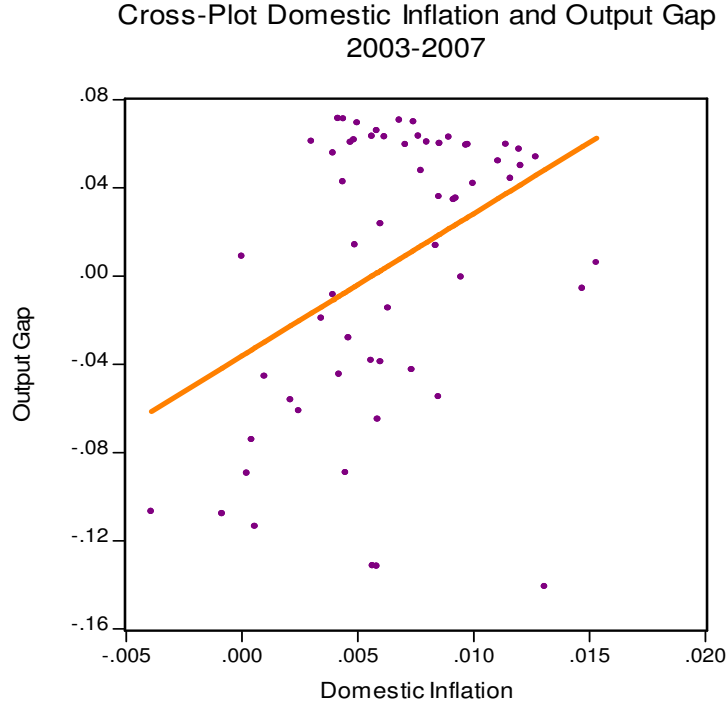Cross-Plot Domestic Inflation and Output Gap
2003-2007

Figure 1 presents the relationship between inflation and its determinants for both sub-periods: Convertibility and the post-crisis period (the managed float regime). The cross-plots show, as suggested by the theory, a positive relation between domestic inflation and its regressors, except for nominal devaluation in the Convertibility period. It can be observed from the cross-plot of domestic inflation and nominal devaluation for this period, that the outliers corresponding to the Brazilian devaluations of 1994 and 1999 force the relationship to be negative.

This graphic analysis is a first approximation to the data. As the new literature on trend inflation suggests, a higher mean inflation in the post-devaluation period could weaken the relation between inflation and the output gap and make the Phillips Curve more forward looking. This descriptive analysis doesn't allow us to test for the presence of these changes in the relation between the variables across the two periods. For this, we need to evaluate these relationships in a multivariate regression framework. In the following section we address this issue through the estimation of a Phillips Curve.

4.2. Estimation results

In this context we first estimate a reduced form of the "Hybrid New Keynesian Phillips Curve", given by equation *(8)*, which provides interesting information about the dynamics of inflation. As we consider the whole period, dummy variables are introduced to control for the 2002 crisis. Rather than imposing the verticality of the Phillips Curve in the long run, we test for it, specifying *(8)* as follows

$$\pi_t = \phi_1 \pi_{t-1} + \phi_2 E_t(\pi_{t+1}) + \gamma \pi_t^* + \lambda \Delta e_t + \delta x_t + \varepsilon_t \qquad (8')$$

We then estimate equation (8') using GMM for the period 1993:1 to 2007:12. We use at most twelve lags of each variable as instruments. To test for the robustness of our results, we conducted several estimations of (8') using the different specifications for matrix $\hat{\Omega}_{HAC}$

described in section 3. As can be seen from Table 1 the estimations are quite robust to changes in the specification of $\hat{\Omega}_{HAC}$. Tests for over-identifying restrictions, applied to each of the estimations confirm that the instruments are valid in all cases.

Table 1. Estimation results

GMM estimates	Newey-West Fixed (4)	Andrews (4.55)	Variable Newey-West (4)
ϕ_1	0.65116	0.63783	0.67572
Std. Error	0.04998	0.05423	0.04309
ϕ_2	0.13492	0.14189	0.12777
Std. Error	0.04010	0.04196	0.03701
$\delta*$	0.01086	0.01096	0.01023
Std. Error	0.00279	0.00280	0.00255
$\gamma*$	0.18577	0.19009	0.18913
Std. Error	0.05619	0.05893	0.04772
$\lambda*$	0.02802	0.02715	0.02743
Std. Error	0.00991	0.01034	0.00935
J-statistic	0.11916	0.12337	0.11212

* These coefficients correspond to the first lag of each variable

A first important finding is that there is a significant forward-looking component in price formation. The backward looking component is also relevant, but the relative values of ϕ_1 and ϕ_2 indicate more weight of the backward looking component.[4] We checked for the validity of imposing verticality in the long run, and we couldn't reject the null hypothesis (see *Table 2*).

Table 2: Testing for linear restrictions

Linear Restriction:			
$\phi_1 + \phi_2 + \gamma + \lambda = 1$			
Test Statistic	Value	df	Probability
Fixed Newey-West			
F-statistic	0.000007	(1, 158)	0.9979
Andrews			
F-statistic	0.0032	(1, 158)	0.9548
Variable Newey-West			
F-statistic	0.2294	(1, 158)	0.6326

Given that we are extending the model to the case of a small open economy it is interesting to observe that changes in foreign inflation and nominal devaluation have a significant effect on domestic inflation. Inflation responds to lagged values of both, nominal devaluation and foreign inflation. While the response of domestic prices to changes in foreign inflation is quite important, of around 0.19, its response to nominal devaluation, although significant, is much weaker (0.03). These results appear to be rather counterintuitive, since nominal devaluation is expected to have a rather significant effect on domestic inflation in a small open economy, where tradable goods are supposed to represent a quite significant portion of domestic output and consumption. A possible explanation for these findings is that, although we are considering a multilateral exchange rate, given the high weight of the dollar in this basket, the

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[4] We tested for equal weights of the backward and forward looking components and the hypothesis was strongly rejected by the Wald test at the conventional significance levels (p-value:0.0000).

nominal exchange rate exhibits a low variability during the Convertibility period, it jumps after the devaluation of January 2002 and then remains quite stable after a few months of high volatility. Thus, the weak response of domestic inflation to nominal devaluation is consistent with the fact that the sample corresponds to a period in which the nominal exchange rate was kept fixed or rather administrated. It is important to note that the very different responses of domestic inflation to nominal devaluation and foreign inflation do not allow to impose the same coefficient to both variables, as it is usually done in the empirical literature. We also find a weak response of domestic inflation to changes in the output gap. This is a frequent empirical finding in the literature on short-run inflation dynamics.

The results suggest that a hybrid representation of the "New Keynesian Phillips Curve" could adequately describe inflation dynamics in Argentina over the period 1993-2007. The estimates indicate that both components, forward and backward-looking appear to be significant in price formation decisions. We also find strong evidence of verticality in the long run.

In the next section we will investigate the presence of a structural break in 2002, along with the abandonment of the Convertibility regime and the adoption of a managed float.

5. Change in trend inflation and parameter stability

Recent developments in the modeling of inflation dynamics (see Blake and Fernández-Corugedo, 2006; Ascari and Ropele, 2007 and Kiley, 2007) extend the standard Neo-Keynesian Phillips curve, that assumes a zero trend inflation in the steady state, to allow for a positive trend inflation. In this context, trend inflation affects the dynamics of the standard Neo-Keynesian model. As inflation becomes less influenced by current marginal costs, the coefficient of the output gap lowers. At the same time price setting becomes more forward looking as does inflation. A higher trend inflation also leads to a stronger autoregressive component making the inflation process more persistent. In the case of Argentina, D'Amato, Garegnani and Sotes Paladino (2007) were able to identify changes in both, mean and persistence of inflation, in the period 1993-2007 which appear to be associated to the introduction of the new monetary regime.

Given the findings mentioned above and the fact that the whole period of analysis 1993-2007 includes a currency board, at place between 1993 and 2001, the crisis of 2002 and a dirty float from then on, we consider of high relevance to evaluate the constancy of the parameter estimations for 1993-2001 and 2003-2007, without taking into account the 2002 crisis.

In order to evaluate the parameter stability a Wald test (proposed by Andrews and Fair, 1988) is used (see also Hamilton, 1994). The following chi-square statistic under the null evaluates the hypothesis, $H_0 : \theta_1 = \theta_2$ where θ_1 (θ_2) is a (q x 1) parameter vector that characterises the first T_0 (the last T-T_0) observations.

$$\lambda_T = T(\hat{\theta}_{1,T_0} - \hat{\theta}_{2,T-T_0})'\{\pi^{-1}\hat{V}_{1,T_0} + (1-\pi)^{-1}\hat{V}_{2,T-T_0}\}^{-1}(\hat{\theta}_{1,T_0} - \hat{\theta}_{2,T-T_0}) \sim \chi^2(q)$$

where π is the fraction of observations contained in the first sub-sample T_0/T, $\hat{\theta}_{1,T0}$ ($\hat{\theta}_{2,T-T0}$) is the parameter vector estimates with the first T_0 (the last T-T_0) observations and $\hat{V}_{1,T_0}$ ($\hat{V}_{2,T-T_0}$) is the estimated coefficient covariance matrix with the first T_0 (the last T-T_0) observations.

The test requires the definition of a break-point, denoted as T_0, and it was set so that the second period started in January 2003, excluding the atypical observations corresponding to the crises of 2002. Results for the two samples are presented in Table 3.

Table 3: Wald test for parameter stability

GMM estimates	1993:1 -2001:12	2003:1-2007:12
ϕ_1	0.25324	0.50256
Std. Error	0.06103	0.01885
ϕ_2	0.09872	0.22665
Std. Error	0.04545	0.02808
$\delta*$	0.01737	0.00708
Std. Error	0.00380	0.00089
$\gamma*$	0.10689	0.07113
Std. Error	0.05485	0.01685
$\lambda*$	0.06283	0.05965
Std. Error	0.01285	0.00540
λ_T=2834.1 (5)		

The parameter constancy over this two sub-samples is clearly rejected. The changes in the estimated parameters confirm the predictions in the literature on trend inflation: the Phillips Curve becomes more forward looking and the effect of the output gap on inflation lowers. Comparing these results with the ones for the whole sample, we can conclude that including the Convertibility and the crisis period observations can bias the estimation considerably.[5]

6. Conclusions

The empirical modeling of the short-run dynamics of inflation assumes inter-temporal optimizing behavior by non-competitive firms. The relevance of persistence in inflation dynamics has led to introduce backwardness in these models by assuming that a portion of the firms could follow a backward-looking rule. The resulting model is known as the "Hybrid New Keynesian Phillips Curve". Using GMM, we estimate a "Hybrid New Keynesian Phillips Curve" for Argentina over the period 1993-2007 and test for parameter stability, given the regime change of 2002. We extend the basic model to the case of a small open economy, allowing nominal devaluation and foreign inflation to play a role in domestic prices setting. For the whole sample, we find that both components, forward and backward are relevant to explain the dynamics of domestic prices, although the backward-looking component weights more in determining inflation dynamics. Nominal devaluation and foreign inflation are also significant to explain domestic inflation behavior, being the response of inflation to the second more intense. The output gap, although weak, has a significant effect on inflation. We cannot reject verticality of the Phillips Curve in the long run.

When testing for parameter stability, we find that, in line with the recent theoretical literature on trend inflation, there are significant differences in the estimated parameters between the Convertibility period and the dirty float. While the influence of the output gap on domestic inflation weakens in the post-Convertibility period, along with the observed increase in trend inflation, the Phillips Curve becomes more forward looking compared with the Convertibility period in which trend was virtually zero.

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[5] A Wald test for the null hypothesis of equal estimated coefficients between the whole period and the managed float rejects it at the 1% significance level.

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