Lack of Credibility, Inflation Persistence and Disinflation in Colombia *

Andrés González G Franz Hamann †

Banco de la República

Bogotá, Colombia

First draft: October 2006

Please do not cite or quote without author's permission

Abstract

Empirical evidence shows that it is important for central banks to be transparent and to build credibility. *Ex ante* there is a great deal of uncertainty about the degree of commitment of the central bank to meet its targets. Theory and practice have shown that as credibility grows, disinflation costs fall. This paper evaluates the ability of Erceg and Levin's (2003) lack of credibility model to account for the observed disinflation process in Colombia during the 1990-2005. We estimate the model using Bayesian techniques and compare the posterior odds of this model against one with ad-hoc price indexation. The odds are in favor of the imperfect credibility model. Our results suggest that the main source of inflation persistence in Colombia has been the inability of the central bank to anchor inflation expectations. Finally we find that the sacrifice ratio of achieving the long term inflation target in Colombia is 2.09.

Keywords: credibility, inflation, persistence, disinflation costs, Colombia

JEL Classification Codes: C52, E31, E52

*Paper prepared for the Second Monetary Policy Workshop in Latin America and the Caribbean on Monetary Policy, Uncertainty and the Business Cycle, in Lima Peru. The views expressed in this document are those of the authors and not necessarilly those of the Banco de la República.

†Dirección autor correspondiente: *agonzago@banrep.gov.co* Banco de la República, Cr 7a No 14-78

1 Introduction

Colombia used to be a country of high and volatile inflation. At the beginning of the nineties the newly independent Banco de la República, began announcing end of year inflation targets along with other explicit targets for other macroeconomic variables, like the nominal exchange rate[1]. These targets should guide monetary policy in order to meet the Constitutional mandate of achieving price stability. Although no long run inflation target was set, central bank officials publicly claimed that their goal was to reduce inflation to a "single digit". In 1991, in line with other central banks in the world, the Banco de la República established its first quantitative inflation target of 22% and so a gradual disinflation process began.

At that time, it was unclear whether the central bank was truly committed to reduce inflation. It was also unclear what the speed of disinflation would be, since there was no announcement in this regard. There was also little evidence of the costs associated with a disinflation. The capital inflows to Emerging Economies during the first half of the decade translated into higher liquidity and the economic boom made it difficult to reduce inflation. The central bank missed the target for six years in a row. With this record at hand, one can easily guess that, during this period, credibility was not one of the main assets in the balance sheet of the central bank.

It was until 1998, in the midst of a mayor financial and economic crisis, that the Banco de la República was able to meet its target (18% in that year). This, presumably, affected the public's credibility on the ability and/or the willigness of the bank to commit to an inflation target. The effects of the economic crises of the end of the nineties were protracted. Inflation declined steadily from about 9% in 1999 to about 5% in 2005. This decline in inflation occurred even in the presence of historically low real interest rates. Nowadays, Colombian inflation, measured as the variation in the CPI, is under control and close to target (but not there yet!). How costly has been this process? How cheaper would have been to disinflate under full credibility? Given the current degree of credibility and the current level of inflation, what is the sacrifice ratio of achieving the long run target? In this paper we try to address these questions.

Standard macro models usually take the inflation target as a known parameter to private agents. The inflation target is observable and perfectly known. A movement from high inflation to low

[1] A crawling band was implemented during the first years of the 90s.

inflation implies that agents firmly believe in the central bank's target. In practice, as the Colombian experience shows, this is usually not the case.

Erceg and Levin (2003) in an influential paper, have studied the episode of the Volcker disinflation in the US using a model in which agents learn about the ultimate intentions of the central bank. They calibrate a standard staggered contracts neokeynesian model to the US economy and find that the output cost of the Volcker disinflation was about 1.6 percentage points for each percentage point reduction in the inflation rate. This number is similar to other results found in the literature for the US. They show that their results are consistent with the idea that most of the inflation persistence found in the US inflation data is attributable to lack of credibility instead of to the contract structure or the existence of adaptative expectations.

Typically staggered contracts models have been critiziced for not being able to reproduce the observed inflation persistence present in the data. Many modelers add ad hoc lags to induce persistence. One of the main implications of Erceg and Levin's (2003) work is that inflation persistence is not only an inherent characteristic of the economy, but also that it can vary with the stability and transparency of the monetary policy regime.

In our work, we use a simplified version of Erceg and Levin's (2003) imperfect credibility model to estimate the speed at which agents learn about the ultimate intentions of the Colombian central bank. In addition we estimate how conventional estimates of the monetary policy rule changes when the central bank lacks full credibility on its commitment to reduce inflation. Finally we test whether the colombian data supports the adaptative expectations model or the lack of credibility model. As Schorfheide (2000) has shown, Bayesian analysis can help to discriminate between alternative models. We compare the posterior distributions of two models: one that captures Erceg and Levin's (2003) imperfect information idea and another that introduces persistence by ad-hoc indexation.

The paper proceeds as follows: in the next section we describe the main facts about inflation persistence in Colombia for the period 1990-2005. Then, in section 3, we describe the model. Then we briefly describe the methodology used to estimate it. In the fifth section we report the results. In the sixth we compare the two models while in the seventh we compute the cost of disinflation in Colombia under imperfect credibility. The last section concludes.

2 Measuring inflation persistence

Research on inflation dynamics has been very active, focusing mainly on the existence of persistence defined as the speed at which inflation converges to its long run value after a given shock. The fact that inflation persistence is a highly attractive area of research can be explain by: First if persistence is a stylized fact it should be incorporated within the theoretical macroeconomic models. On the contrary, if the degree of persistence is not constant but varies over time it can be used as a measure of success of a given monetary policy. For instance, Sargent (1999) argues that the decline of US inflation persistence during the 90's has been associated with an increase in the credibility of the monetary policy. In the sense that inflation expectations have been anchored at a low level and so they are unlikely to adjust to temporary increases in the inflation rate.

Measuring persistence is also important for the design of monetary policy. In particular, if inflation is consider to be non persistent the monetary authority can react mildly to inflationary shocks. However, if persistence is erroneously underestimated delays in response to inflationary shocks will create relatively large deviations from the targeted inflation.

Even though inflation persistence constitutes a key feature of inflation dynamics, its mesurement is controversial. In particular, standard definitions of persistence such as the sum of the autoregressive coefficients, the spectrum at zero frequency and half life are all concepts that assume convergence to a given mean. Marques (2004) makes this point clear and argues that measures of inflation persistence should be based on a time varying mean as it may reflect exogenous factors such as inflation drivers and/or the inflation target. In fact, Levin and Piger (2004), Corvoisier and Mojon (2005), Ciccarelli and Mojon (2005) show that persistence in some European countries has been stable when computed over small samples or when the mean of inflation is allowed to change. In these papers inflation persistence is measured as sum of the autoregressive coefficients in a linear model that allows for breaks in mean. The number of breaks and the break dates are estimated using either Bai and Perron (1988) or Altissimo and Corradi (2003).

We take a different approach and consider a smooth trend rather than the abrupt one. The smooth nonlinear trend of inflation can be associated with the adjustment of inflation expectations to new inflation targets. If this is the case, the adjustment process is more likely to happen gradually rather than abruptly. However, it can be argued that inflation trend reflects the inflation targets and consequently the abrupt trend may give a better approximation.

To characterize the evolving changes in the mean of inflation, we use the following model

$$y_t = \mu_t + x_t \tag{1}$$
$$x_t = \rho x_{t-1} + \nu_{3t}$$
$$\mu_t = \beta_{t-1} + \mu_{t-1} + \nu_{1t}$$
$$\beta_t = \beta_{t-1} + \nu_{2t}$$

with $\nu_{it} \sim N(0, \sigma^2_{\nu_i})$, $i = 1, 2, 3$ and $\mathsf{E}[\nu_{it}\nu_{js}] = 0$ for $i \neq j$ and $t \neq s$. Equation (1) decomposes the inflation process into a evolving mean component and the fluctuation around it. The latter component is defined by the stationary AR(1) process $x_t = \rho x_{t-1} + \nu_{3t}$ where $\rho \in (-1, 1)$ form a persistence measure. The trend component is given by μ_t and its specification resembles the standard local linear model. The model (1) can be easily estimated using Kalman filter and the standard error decompositions. See Harvey (1990), West and Harrisson (1999) and Durbin and Koopman (2001) for details. The advantage of (1) compare with other approaches is that trend and persistence are modelled simultaneously rather that sequentially[2].

To measure inflation we use alternative indexes. One is the percentage change of the seasonally adjusted quarterly Consumer Prices Index (CPI). In addition to CPI inflation, we present results for the following inflation rates: $\pi_{\text{IPC-SA}}$ that excludes food, $\pi_{\text{IPC-T}}$ and $\pi_{\text{IPC-NT}}$ that includes only traded goods and non traded goods, π_{IPCR} with only regulated goods and $\pi_{\text{IPC-B}}$ that excludes food and regulated prices from the CPI. The sample consists on quartlerly data for the period 1988:1 to 2006:2.

We report the results in Table 1 and Figures 1 to 5. These results confirm most of our priors: all measures of inflation display a significant amount of persistence. The only exception being the non traded inflation; for which the estimated trend component of inflation follows closely the observed inflation. So, deviations from trend quickly revert to the mean [3]. Notwithstanding this, most

[2] A similar model has being used by Clark (1987, 1989) to decompose the US real GDP between trend and cycle

[3] One explanation could be that price formation in the nontraded sector could be better anchored to the inflation target. The nontraded sector comprises mainly service oriented businesses and construction firms. Real activity in this sector was depressed during the financial crisis and relative prices adjusted quickly. During the crisis the central bank tightened monetary policy to defend the exchange rate band and so, tighter monetary policy was associated with a significant real exchange rate depreciation (i.e. a collapse in the relative price of the nontraded goods). We speculate that this event may have caused a price setting behavior that puts an important weight on inflation expectations.

Table 1: Estimated persistence for the different inflation rates.

	π_{IPC}	π_{IPCSA}	π_{IPCT}	π_{IPCNT}	π_{IPCR}	π_{IPCB}
$\hat{\rho}$	-0.40	0.67	0.86	0.09	-0.86	0.91

Note: We report the absolute value of the estimated coefficient. Negative values were found for π_{IPC}, π_{IPCR}.

Table 2: Credibility and inflation targeting in Colombia

Year	Observed (1)	Expectation (2)	Target (3)	Mistake (1)-(3)	Surprise (1)-(2)	Anchoring (2)-(3)	Credibility
1997	17.68	18.45	18.00	-0.32	-0.77	0.45	
1998	16.70	17.95	16.00	0.70	-1.25	1.95	
1999	9.23	15.79	15.00	-5.77	-6.56	0.79	
2000	8.75	9.89	10.00	-1.25	-1.14	-0.11	33.00
2001	7.64	8.85	8.00	-0.36	-1.21	0.85	46.90
2002	6.99	6.95	6.00	0.99	0.04	0.95	35.00
2003	6.49	6.58	5.50	0.99	-0.09	1.08	42.00
2004	5.50	6.13	5.50	0.00	-0.64	0.63	69.10
2005	4.85	5.41	5.00	-0.15	-0.56	0.41	77.80
2006	4.58	4.60	4.50	0.08	-0.02	0.10	90.10

Note: Expectation: refers to the expected value of the end of year inflation measured at the beginning of the year. Credibility: refers to the number of people that believed (at the beginning of the year) that the target would be met for that year Inflation expectations for the year 1997 corresponds to June Survey. Observed inflation for 2006 corresponds to annual yoy inflation in September 2006.

measures of inflation display high persistence. In the next section we use a model in which lack of credibility plays a key role in explaining inflation persistence.

3 The model

We consider a standard closed economy neokeynesian model commonly used in many central banks. The main elements of the model are: a Phillips an IS curve and a monetary policy rule, described in equations (2)-(4).

$$\pi_t = \beta E_t \pi_{t+1} + \lambda x_t + u_t \tag{2}$$

$$y_t = E_t y_{t+1} + \sigma \left(r_t + E_t \pi_{t+1} + E_t g_{t+1} - g_t \right) \tag{3}$$

$$i_t = \gamma_i i_{t-1} + (1 - \gamma_i) \left[\gamma_\pi \left(\pi_t - \overline{\pi}_t \right) + \gamma_x y_t \right] + z_t \tag{4}$$

This is an open question.

where π_t is the inflation rate, $\overline{\pi}_t$ is the inflation target, x_t is the marginal cost, i_t is the nominal interest rate, r_t is the real interest rate. The variables g_t and z_t are a preference and a policy shock respectively, which we later describe. An important element of the model is the variable u_t. As we will see later, this variable captures the deviation of the Phillips curve under imperfect information with respect to the perfect information. The expectation operator E_t denotes the rational expectation of private agents if they use all available information at time t. The parameters of this set of equations are: $\lambda = \frac{(1-\alpha)(1-\beta\theta)(1-\theta)}{\theta(1+\alpha(\epsilon-1))}$ where α is the share of labor factor in production, θ the probability of keeping prices fixed during the period, ϵ the elasticity of substitution between slightly differentiated types of goods and $\beta \in (0, 1]$ the discount factor. λ measures the slope of the Phillips curve. $\sigma > 0$ is the elasticity of intertemporal substitution and γ_π, γ_x and γ_i measure the degree of responsiveness of the monetary authority to deviations from target, the output gap and past interest rate.

The rest of equations of the model describe the technology, the marginal cost, the marginal rate of substitution and the real wage:

$$y_t = a_t + (1 - \alpha)n_t \tag{5}$$

$$x_t = w_t + n_t - y_t \tag{6}$$

$$m_t = \frac{1}{\sigma}y_t + \gamma n_t - g_t \tag{7}$$

$$w_t = m_t \tag{8}$$

where y_t is output, a_t is a productivity shock, n_t is the number of hours worked, w_t is the real wage per hour and m_t is the marginal rate of substitution. The parameters are: α the labor share in the production function and γ the inverse of the elasticity of labor supply to the real wage.

We specify the shocks to follow the following stochastic processes:

$$a_t = \rho_a a_{t-1} + \epsilon_t^a$$

$$g_t = \rho_g g_{t-1} + \epsilon_t^g$$

$$z_t = \epsilon_t^i$$

where each of the innovations ϵ^j follows a normal distribution with zero mean a standard deviation σ_j for $j = a, g, i$. We assume that the innovations are uncorrelated with each other.

When information about the inflation target is perfect, $\overline{\pi}_t$ is known for all t. The interesting case is when information is not perfect. A special case, described in Erceg and Levin (2003), is when $\overline{\pi}_t$ varies over time due to a combination of a white noise shock, ϵ_t^w, and a shock ϵ_t^p with permanent effects on the inflation target. The target is observable to agents, but its individual components are not. In this case, they show that by defining the variable u_t as:

$$u_t = \beta \left(E_t \pi_{t+1} - \widehat{\pi}_{t+1} \right) \tag{9}$$

where $\widehat{\pi}_{t+1}$ is the rational forecast given all information available to private agents at time t in the case of perfect information. By replacing (9) in (2) we can see that in the case of $u_t = 0$, we obtain the standard neokeynesian Phillips curve in the case of perfect information. If there are discrepancies between private agents expectations under imperfect information and perfect information then $u_t \neq 0$. So, u_t will contribute to inflation persistence in the case in which private agents do not have perfect information about the evolution over time of the inflation target.

Erceg and Levin (2003) derive an analytic expression for u_t for the case of the standard neokeynesian model when the inflation target is the sum of two shocks: a random walk component and a white noise component. So,

$$u_t = (1 - \kappa) u_{t-1} + (1 - \kappa) \eta \epsilon_t^p$$

where κ is the Kalman gain parameter that determines the speed at which agents learn to distinguish between the two components of the inflation target, $\eta < 0$ is a parameter and ϵ_t^p is a normally distributed zero mean shock with standard deviation σ_ϵ, that affects the permanent component of the inflation target. Notice that in the case of full information $\kappa = 1$, agents learn at the highest possible rate and so $u_t = 0$ and therefore $E_t \pi_{t+1} = \widehat{\pi}_{t+1}$.

Of particular interest is the value of the learning parameter κ. We are interested in assessing the information contained in the data about the speed at which agents have learned during the disinflation period using a standard neokeynesian model augmented by learning about the inflation target. Unlike Erceg and Levin (2003), we estimate a simplified version of their model using Bayesian methods. The advantages of estimating models using a Bayesian approach is discussed formally in

Lubik and Schorfheide (2003). A review of the Bayesian tools for macroeconomists is presented in An and Schorfheide (2006).

4 Bayesian estimation

One of the advantages of estimating economic models using a Bayesian approach is that we can incorporate additional information on parameters through the use of priors. To perform the Bayesian estimation of each model we follow Schorfheide (2000) and proceed in five steps which we summarize briefly. First, for a given set of parameters we solve the model using Klein's (2000) method to find the state transition equation. The solution defines the way in which the system evolves around the deterministic steady state. The state-space representation is completed by adding a measurment equation to the model dynamics. The next step consists on computing the likelihood through Kalman filtering and to combine it with the prior distribution of the parameters to get the posterior density. Draws from the posterior density are obtained using the random walk Metropolis-Hastings algorithm as described in Schorfheide (2000). The algorithm is started at the posterior mode found by numerical optimization. In this section we report the data used in the estimation, our priors and estimation results.

4.1 Data

We seek to explain the behavior of inflation, output, nominal interest rate and real wages. We use quarterly HP-detrended data from 1987:1 to 2005:4. As a proxy of the nominal interest rate we use the interest rate on 90-day certificates of deposits. Our inflation measure is the quarterly (annualized) growth rate of the CPI. Output is measured as the real GDP and real wages are the median of the real wage per hour of wage earners who work 40 or more hours per week.[4] Let $d_t = (\pi_t, y_t, i_t, w_t)$ denote the observed data and $\Phi = (\beta, \alpha, \epsilon, \theta, \sigma, \gamma, \theta_i, \theta_\pi, \theta_y, \rho, \sigma_a, \sigma_g, \sigma_z, \sigma_\epsilon)$ the vector of parameters to be estimated.

[4] This time series is deseasonalized and calculated by the Statistics section of the Banco de la República based on the DANE's National Households Survey (ENH) and the Continuous Households Survey (ECH).

4.2 Priors

We impose strong priors on β, α and ϵ. $\beta = 0.98$ and implies a real annual return close to 4%. To replicate the labor factor compensation share in real GDP we set $\alpha = 0.36$. $\epsilon = 6$ which is a standard value in the literature[5].

As explained in Rabanal and Rubio-Ramirez (2005), there is an identification problem in the model between the probability of adjusting prices and the elasticity of substitution. That is, θ and ϵ cannot be identified sepparately. In order to circumvent the identification problem we choose to estimate θ for a given markup, since the estimated parameter tells us about the implicit frequency at which firms adjust prices in Colombia. To our knowledge, there is no empirical evidence on this topic.

The prior distributions for the rest of the parameters in vector Φ are reported in Table 4. The inverse of the elasticity of subsitution follows a normal distribution with mean 4 and standard deviation 1. We use these prior because evidence for the US shows that its value is higher than 1 but not much larger than 2. However, evidence for emerging markets shows that it should be between 2 and 5. As for the probability of adjusting prices we choose to set a gamma distribution for parameter θ with mean 4 and standard deviation 2. This is so because informal evidence points at price contracts being adjusted once a year (every four quarters). We do not put a very tight prior on this parameter. For the Taylor rule coefficients we use the priors that are commonly used in the literature: $\theta_\pi = 1.5$ and $\theta_y = 0.125$. We use a normal distribution for both with standard deviations of 0.25 and 0.125 respectively. There is little evidence about the elasticity of labor supply to the real wage in Colombia, and so we set the same prior for this parameter as the one used by Rabanal and Rubio-Ramírez for the EU: a normal distribution with mean 1 and standard deviation of 0.5. For all autoregressive parameters we use a uniform prior between [0,1). We do not have strong priors about these values, since to our knowledge this is the first joint estimation of a micro-founded neokeynesian model. For all standard deviations of shocks we use an inverse gamma distribution with mean 0.1.

As for the credibility parameters, results for the US suggest that κ is around 0.13. We use this value as mean prior since there are no previous studies for Colombia. For the parameter η we use

[5] We use strong priors for β and α because the model does not have capital and so the likelihood does not have information to estimating them.

a uniform distribution [-1,0).

5 Results

We report the posterior statistics in the last three columns of Table 4. We use a Random Walk Metropolis Hastings algorithm to draw a two-block chain of size 100.000 from the posterior distribution of Φ. The acceptance rates are 59 percent for the first block and 61 percent for the second.

The data supports the idea of lack of credibility as κ, the sppeed at which agents learn in the economy is very close to zero. So, we reject the hypothesis of perfect credibility. Our findings show that the posterior mean of the speed of learning is $\kappa = 0.07$, while the average duration of price contracts is strikingly low: less than a quarter. This result implies that firms adjust prices every every month and is at odds with other results in the literature. Erceg and Levin use a calibrated imperfect credibility model for the US and Rabanal and Rubio-Ramírez estimate a perfect credibility model for the EU. Erceg and Levin calibrate $\kappa = 0.13$, while Rabanal and Rubio-Ramírez find that that firms adjust prices every 5.8 quarters. We argue that, since we are estimating instead of calibrating both parameters using an imperfect credibility model, the parameter κ captures part of the price dynamics that is embedded in the parameter θ when there is perfect credibility. So, a significant part of the persistence in the Colombian data is captured by a slow speed of learning.

For the policy rule parameters, we find that a the posterior mean response to inflation of $\lambda_\pi = 2.01$, to the output gap of $\lambda_y = 0.06$ and a smoothing desire of $\lambda_s = 0.27$. The response to inflation and the output gap differ from our priors, that were set accordingly to the Taylor principle. Our results show an active Central Bank when responding to deviations from long run inflation and a passive one when responding to output. This result is in line with recent theoretical developments of Schmitt-Grohe and Uribe (2004). They find that social welfare increases when the central bank only responds to inflation. Bernal (2002), using a frequentist approach to estimate Taylor rules in a partial equilibrium model for Colombia, finds that $\lambda_\pi = 1.34$, $\lambda_y = 0.19$ and $\lambda_s = 0.10$. We interpret our results as supporting the idea that, given the lack of credibility on monetary policy during a large part of the sample, the response of the Central Bank to inflation has to be higher than in environments with higher credibility.

The posterior mean of the risk aversion coefficient lies between 4.3 and 6.9 with a 95 percent confidence. The point estimate is 5.6, which is a standard result in the International Macroeconomics literature. This estimate reflect the higher variability of the macroeconomic time series in Emerging Markets.

We also obtain an estimation of the labor supply elasticity. The posterior mean of η is 1.5. Rabanal and Rubio-Ramirez (2005) find that the value of this parameter is model dependent: with flexible wages and perfect credibility tends to one. When sticky wages are introduced, it gets closer to two. Our estimation lies in between. This may reflect the fact that imperfect credibility induces a value of η higher than one.[6]

A quick inspection of the smoothed shocks obtained in our estimation shows that the variability of all shocks is lower since the last quarter of 1999, period in which the inflation targeting regime was implemented.[7]. In this sense we could argue that the implementation of IT in Colombia is associated with a higher degree of macroeconomic stability.

5.1 Model Comparison and Inflation Persistence

The neokeynesian model is a standard model in many central banks. To induce inflation persistence many modelers introduce ad-hoc indexation, but keep the perfect information assumption. As a result, the hybrid Phillips curve encompasses the effects of inflation expectations, lagged inflation and the real marginal cost. We estimate such a model, using the same priors, to compare the performance of the imperfect credibility model against the conventional NKM. Table 5 reports the results of the estimation of the NKM with indexation.

To compare the two models we perform a posterior odds test. The posterior odds ratio is the ratio of the posterior model probabilities. Consider the two models $\mathcal{M}_p$ and $\mathcal{M}_i$ with two associated parameters θ_p and θ_i where p refers to the model with perfect information and i refers to the model with imperfect information. Both models were estimated using the sample Y_T. The fit of each model $m = p, i$, is given by its marginal density of the data $p\left(Y_T \left| \mathcal{M}_m\right.\right)$. We compute the marginal

[6]The rest of the parameters are the autocorrelation and standard deviations of the shocks (productivity, preferences, monetary policy and the target shock). There is a significant amount of persistence in the productivity and preference shocks. The posterior mean of the autocorrelation coefficients is 0.46 and 0.73, with a standard deviation of 1.8 percent and 8.9 percent. The posterior mean of the volatility of the interest rate rule shock and the inflation target is 3.6 percent and 5.7 percent. We attribute this high value of the target shock to the period of high and volatile inflation in Colombia that characterized the first part of our sample.

[7]See figure 8. The inflation targeting regime starts from position 50 in the graph

density of the data conditioned on the model:

$$p\left(Y_T \left|\mathcal{M}_m\right.\right) = \int_{\Theta_m} p\left(\theta_m \left|\mathcal{M}_m\right.\right) p\left(Y_T \left|\theta_m, \mathcal{M}_m\right.\right) d\theta_m$$

by integrating out the parameters θ_m from the posterior kernel. Using Bayes theorem, we can compute the posterior distribution over models as:

$$p\left(\mathcal{M}_m \left|Y_T\right.\right) = \frac{p(M_m) p\left(Y_T \left|\mathcal{M}_m\right.\right)}{\sum_{m=p,i} p(M_m) p\left(Y_T \left|\mathcal{M}_m\right.\right)}$$

where $p\left(\mathcal{M}_m\right)$ is the prior that we have on each of the models. So, the posterior odds ratio is:

$$\frac{p\left(\mathcal{M}_p \left|Y_T\right.\right)}{p\left(\mathcal{M}_i \left|Y_T\right.\right)} = \frac{p(\mathcal{M}_p) p\left(Y_T \left|\mathcal{M}_p\right.\right)}{p(\mathcal{M}_i) p\left(Y_T \left|\mathcal{M}_i\right.\right)}.$$

If we had the same prior on each model, the posterior odds ratio is the ratio of the marginal likelihoods:

$$F_{p,i} = \frac{p\left(Y_T \left|\mathcal{M}_p\right.\right)}{p\left(Y_T \left|\mathcal{M}_i\right.\right)}$$

also known as the Bayes factor. As the Bayes factor gets larger, the higher the support for model $\mathcal{M}_p$.

We find that $p\left(Y_T \left|\mathcal{M}_p\right.\right) = 522.06$ while $p\left(Y_T \left|\mathcal{M}_i\right.\right) = 646.7667$ implying a Bayes factor of 0.81, so that the odds are in favor of the imperfect credibility model.[8]

An interesting result of this exercise is the impact of the imperfect information assumption on the estimation of the degree of stickiness in the economy. According to the literature on inflation persistence, there are three sources of inflation persistence: extrinsic, intrinsic and expectations-driven. The first can be associated with the coefficient that accompanies the real marginal cost, λ, in the New Keynesian Phillips curve (NKPC), equation (2). The second with the lagged inflation term. The third, in a full information model, with the inflation expectations term. In an imperfect credibility model the expectations-driven persistence is affected by parameter κ.

To analyze the degree of extrinsic persistence the key parameter is the probability of keeping prices fixed during a quarter. In both models, the degree of price stickiness is very low. In the

[8] We use the modified harmonic mean. The Laplace approximation is 646.7734 for the imperfect information model.

imperfect information model price adjustments occur every month, while in the full information every two months. So, the assumption of price indexation coupled with price indexation is rejected by the colombian data.

To see this we draw from the distribution of the elasticity of inflation to the real marginal cost, λ, and find that its mode on the case of price indexation is 0.08, while in the case of imperfect credibility is 0.37. A higher elasticity means a lower degree of price stickiness and so a lower degree of extrinsic persistence.

The degree of intrinsic persistence is similar in both models. Altough we did not estimate parameter β in the imperfect credibility model, its value obtained in the estimation of the full information model is very similar.

Finally, the expectations-driven persistence of the full information model is closely linked to the intrinsic one through the discount factor. In the imperfect information model not only the discount factor matters but also the speed of learning, κ. The posterior odds test favors the idea that this is our main source of inflation persistence in Colombia. Agents learn at a very slow pace. About half the rate at with agents learn in the US, a country that has no explicit inflation target. We could associate this result to the fact that the disinflation process in Colombia has been long and gradual. This has created difficulties to the central bank in anchoring inflation expectations to a long run inflation target.

The estimation of the policy rule is also consistent with the lack of credibility result. The responsiveness of the central bank to deviations of inflation from target is higher under imperfect information than under full information and price indexation -see Table 5. This means that the central bank has to exert more effort when its policy lacks credibility than when agents have full information but the price setting behavior is to adjust prices according to past inflation. In the next section we estimate how costly has been to disinflate without well anchored expectations.

5.2 Disinflation Costs

How large are the disinflation costs under imperfect credibility in Colombia? One way to answer this question is to use the estimated model and compute the sacrifice in terms of the output gap of reducing the inflation target by 100 bp. We also compute the effort of monetary policy in terms of the increment of the policy rate.

Table 3: Disinflation Costs and Monetary Policy Effort (basis points).

β	0.95	0.96	0.97	0.98	0.99
Sacrifice Ratio	182	191	200	209	219
Policy Effort	129	135	141	148	155

Under full credibility a central bank can disinflate at little or no output cost. By relaxing this assumption, the disinflation cost depends on the degree of credibility of the monetary authority. In the model, agents learn gradually about the permanence of the target shock. So, the speed at which agents learn provides a natural measure of the degree of credibility of the central bank. If agents learn quickly the disinflation process will resemble the perfect credibility case. The slower the speed of learning the greater the output costs and the effort of monetary policy.

Given the estimated speed of learning, we compute the macroeconomic effects of a 100 bp disinflation focusing on the sacrifice ratio and the monetary policy effort. We measure the sacrifice ratio as the present value of the output gap during the disinflation period. To measure the monetary policy effort we compute the present value of the nominal interest rate gap during the disinflation period. Both present values depend on the discount factor used. In the estimation of the model we used a 0.98 discount. We vary this parameter between 0.95 and 0.99. Table 3 reports the sacrifice ratio and the monetary policy effort under alternative discount factors.

In the benchmark estimation, a 100 basis points disinflation during five years requires the central bank to keep interest rates by nearly 150 bp above average generating a sacrifice of 209 bp in output during a disinflation period of 10 years. In the model, to effectively reduce inflation in 100 bp a shock to the target of equal magnitude is not enough. The central bank needs to exert more effort in order to reduce inflation in 100 bp effectively. Our sacrifice ratio is higher than the 170 bp obtained by Erceg and Levin for the US economy and other similar estimates for Canada and the United Kingdom. Part of the difference can be explained due to the lower speed of learning, the higher response of the central bank to deviations of inflation from target and the sensitivity of inflation to the real marginal cost estimated in the Colombian data.

The sensitivity of this result depends on the discount factor used compute the present value. Using a discount factor of $\beta = 0.95$, the disinflation cost falls to 182 bp, closer to the international estimates. Interest rates have to be 129 bp higher on average during the disinflation period.

Table 4: Prior and Posterior Distributions of Imperfect Credibility Model

Parameter	Prior Distribution	Prior mean	Prior Std Dev	Posterior Mean	Posterior Conf. Interval	
$\frac{1}{\sigma}$	Normal	4	1	5.34	4.07	6.63
γ_s	Uniform [0,1)	0.5	0.2887	0.24	0.06	0.40
γ_π	Normal	1.5	0.25	2.12	1.80	2.43
γ_y	Normal	0.125	0.125	0.06	-0.13	0.25
θ	Gamma	4	2	0.27	0.17	0.37
γ	Normal	1	0.5	1.74	1.11	2.34
κ	Beta	0.13	0.0650	0.07	0.02	0.12
ρ_a	Uniform [0,1)	0.5	0.2887	0.44	0.31	0.56
ρ_g	Uniform [0,1)	0.5	0.2887	0.74	0.64	0.84
η	Uniform [0,1)	-0.5	0.2887	-0.4406	-0.6875	-0.1082
σ_a	Inv. Gamma	0.04	∞	0.016	0.012	0.020
σ_g	Inv. Gamma	0.11	∞	0.088	0.067	0.108
σ_z	Inv. Gamma	0.04	∞	0.038	0.029	0.047
σ_ϵ	Inv. Gamma	0.09	∞	0.053	0.026	0.077

Table 5: Prior and Posterior Distributions of Full Credibility with Indexation Model

Parameter	Prior Distribution	Prior mean	Prior Std Dev	Posterior Mean	Posterior Conf. Interval	
$\frac{1}{\sigma}$	Normal	4	1	5.92	4.66	7.17
γ_s	Uniform [0,1)	0.5	0.2887	0.51	0.42	0.61
γ_π	Normal	1.5	0.25	1.66	1.32	2.00
γ_y	Normal	0.125	0.125	0.20	0.00	0.38
θ	Gamma	4	2	0.85	0.57	1.10
γ	Normal	1	0.5	1.75	1.05	2.40
β	Beta	0.98	0.01	0.98	0.97	0.99
ρ_a	Uniform [0,1)	0.5	0.2887	0.019	0.013	0.043
ρ_g	Uniform [0,1)	0.5	0.2887	0.74	0.64	0.84
σ_a	Inv. Gamma	0.04	∞	0.05	0.03	0.07
σ_g	Inv. Gamma	0.11	∞	0.10	0.08	0.12
σ_z	Inv. Gamma	0.04	∞	0.026	0.022	0.030

6 Final Remarks

To be written.

Figure 1: Trend and stationary component for π_{IPC}

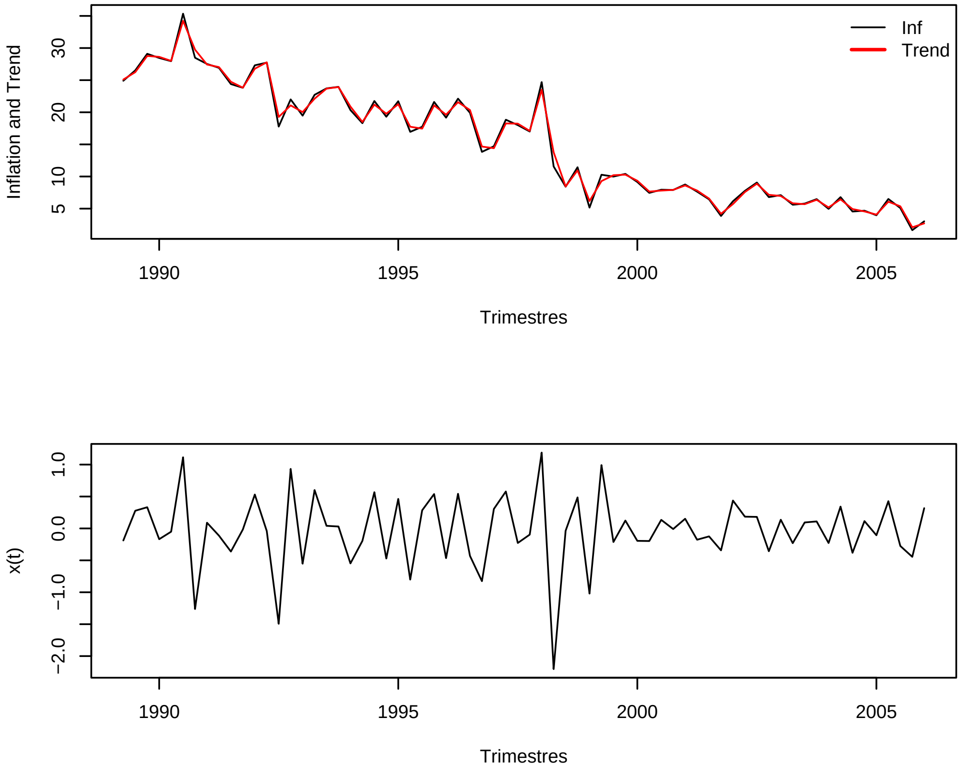


Figure 2: Trend and stationary component for π_{IPC}

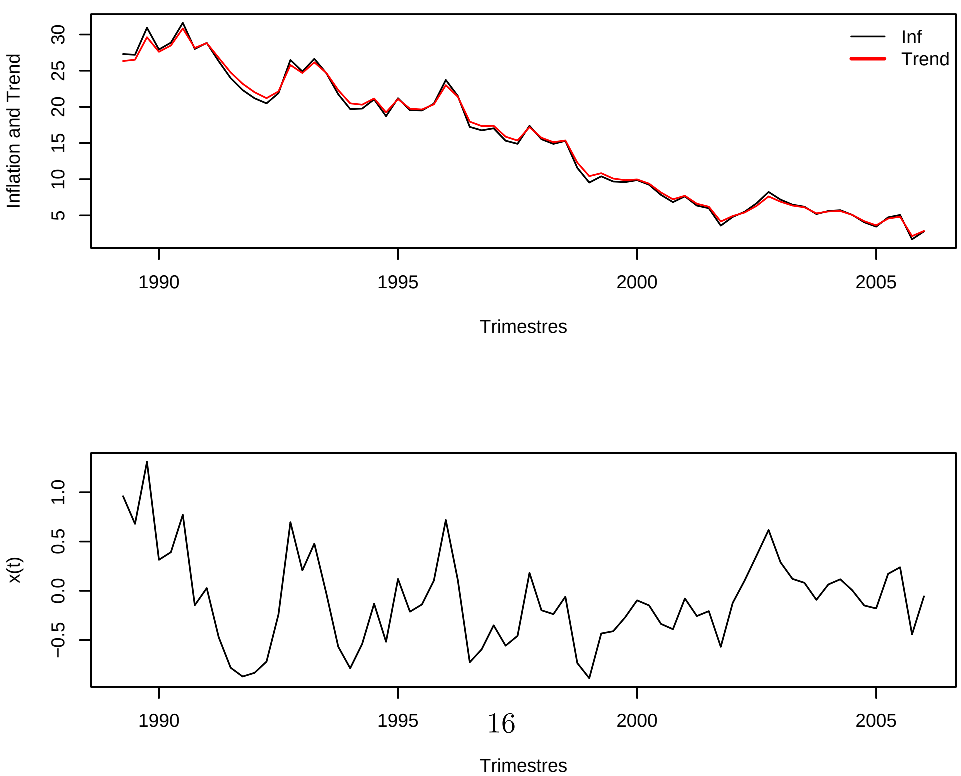


Figure 3: Trend and stationary component for π_{IPC-T}

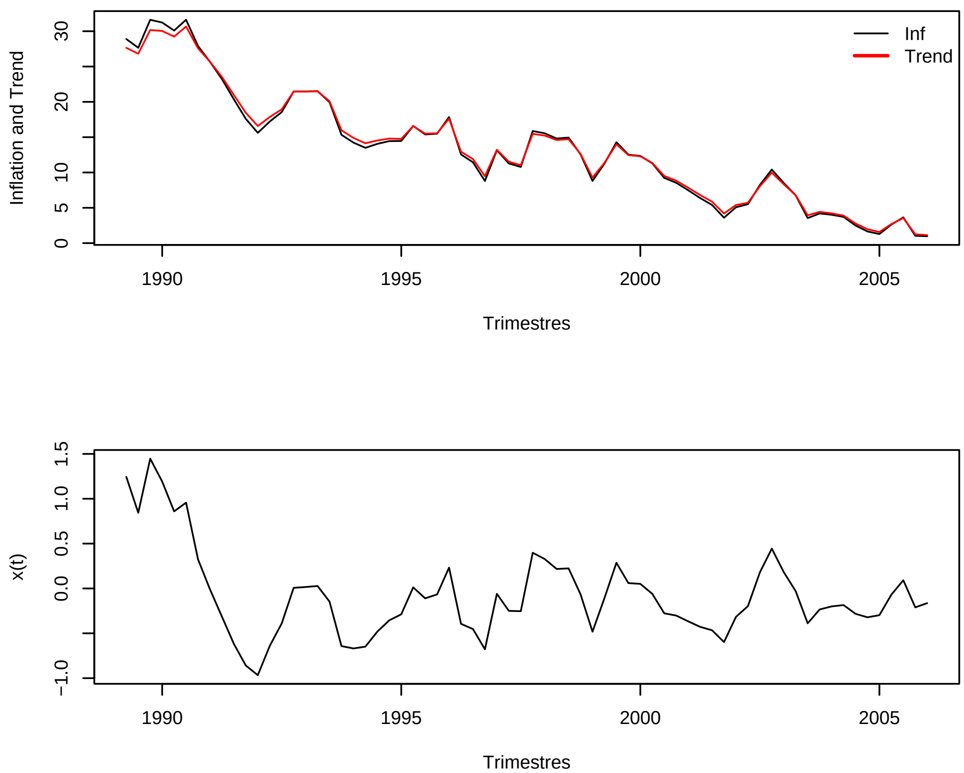


Figure 4: Trend and stationary component for π_{IPC-NT}

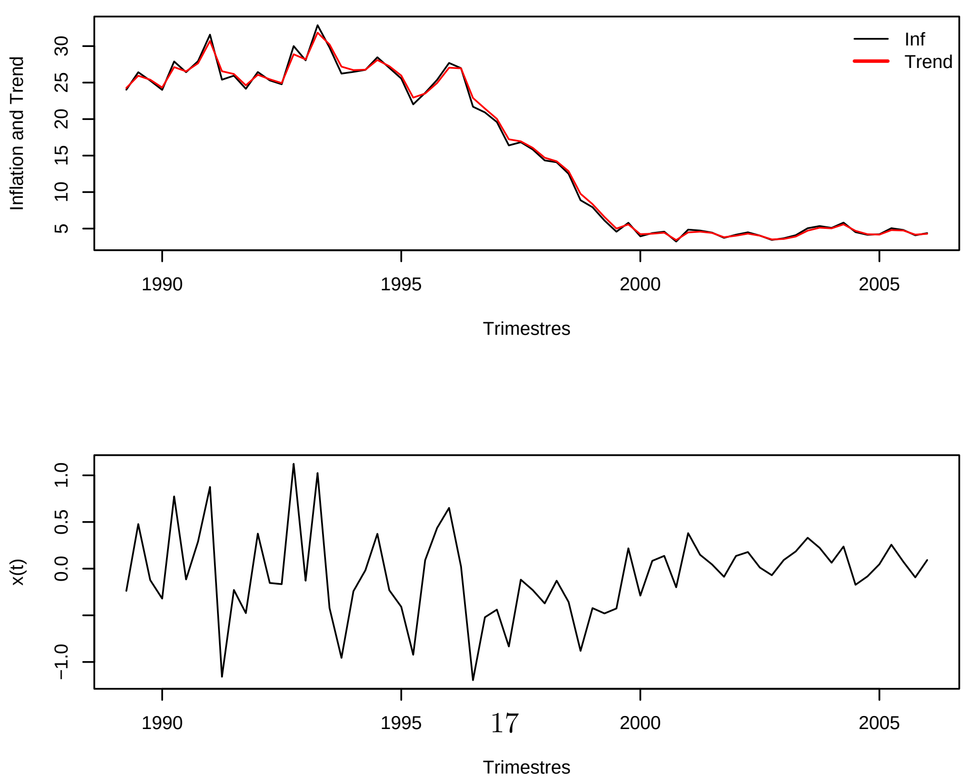


Figure 5: Trend and stationary component for π_{IPC-B}

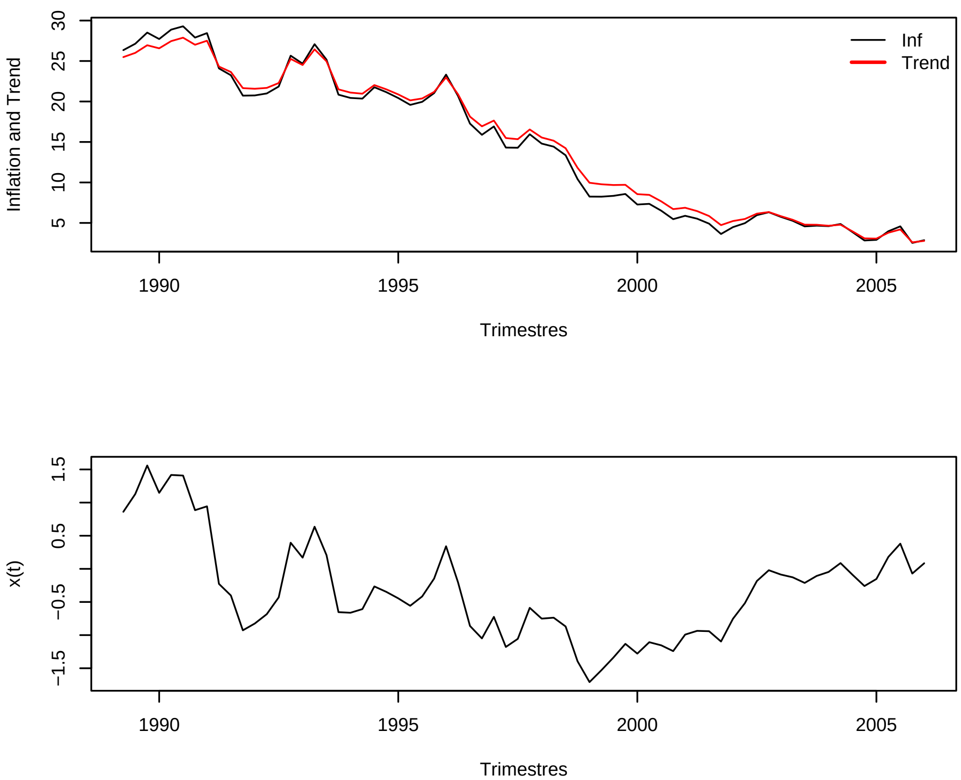


Figure 6: Prior and Posterior Distributions of Imperfect Credibility Model

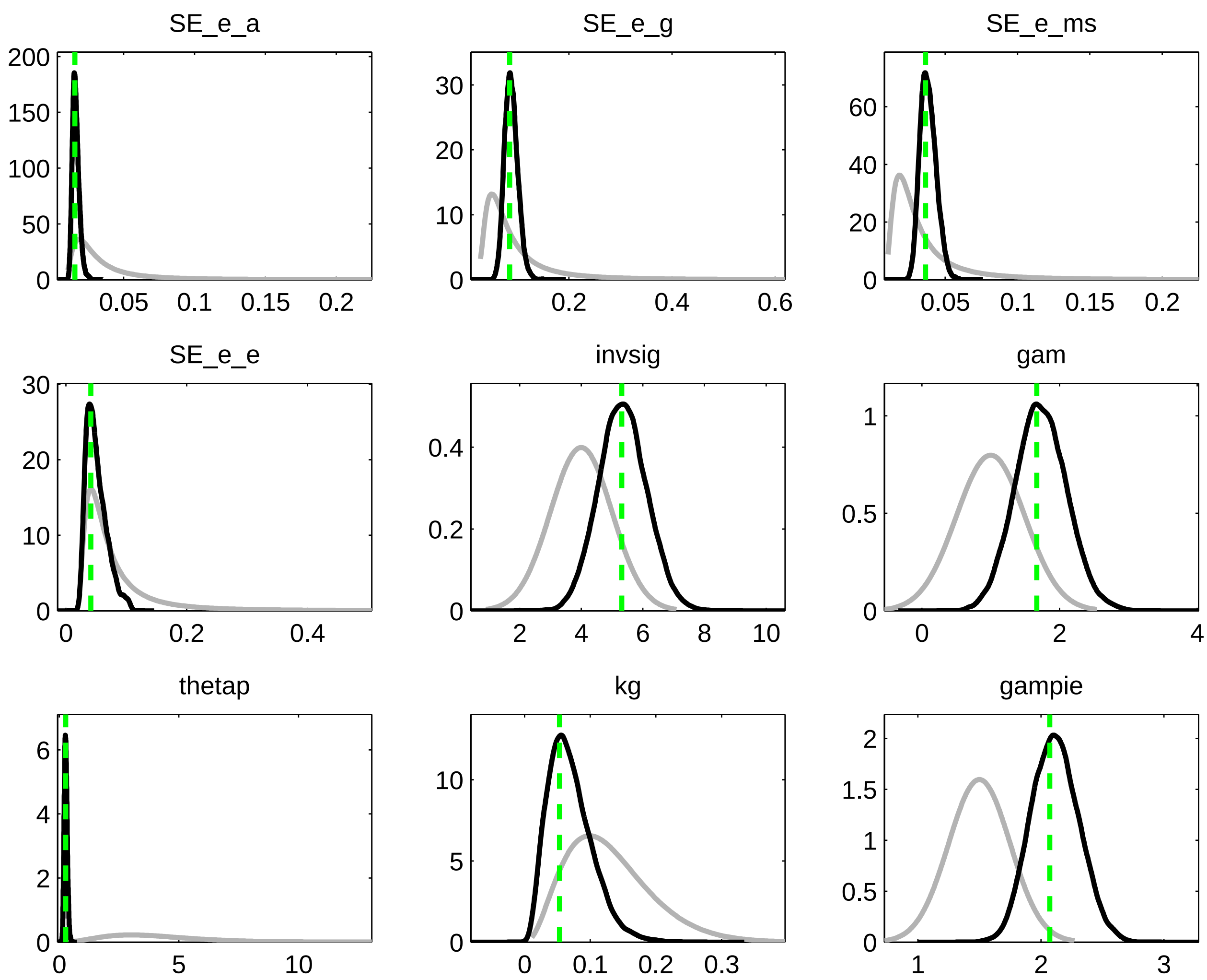


Figure 7: Prior and Posterior Distributions of Imperfect Credibility Model (cont.)

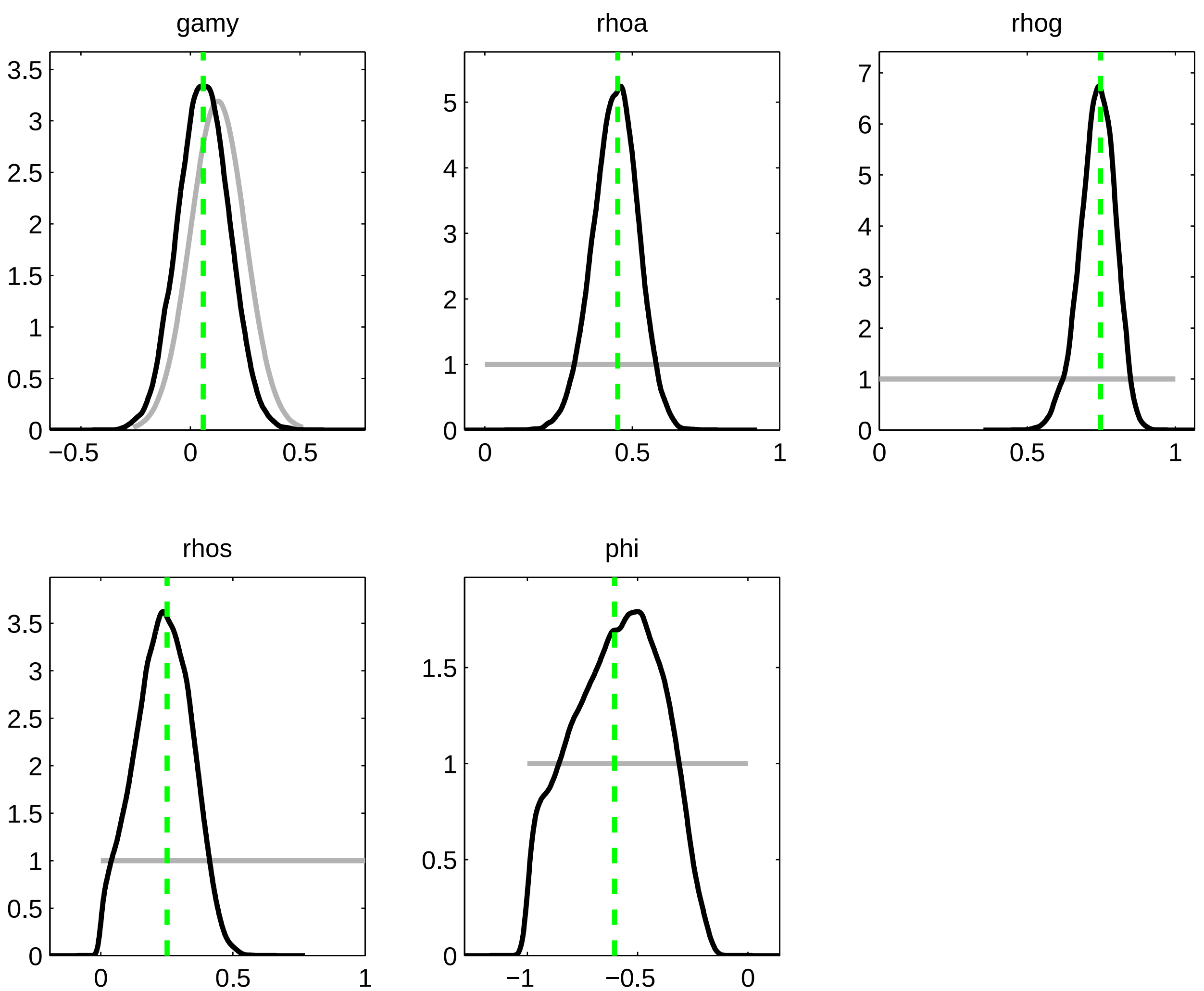
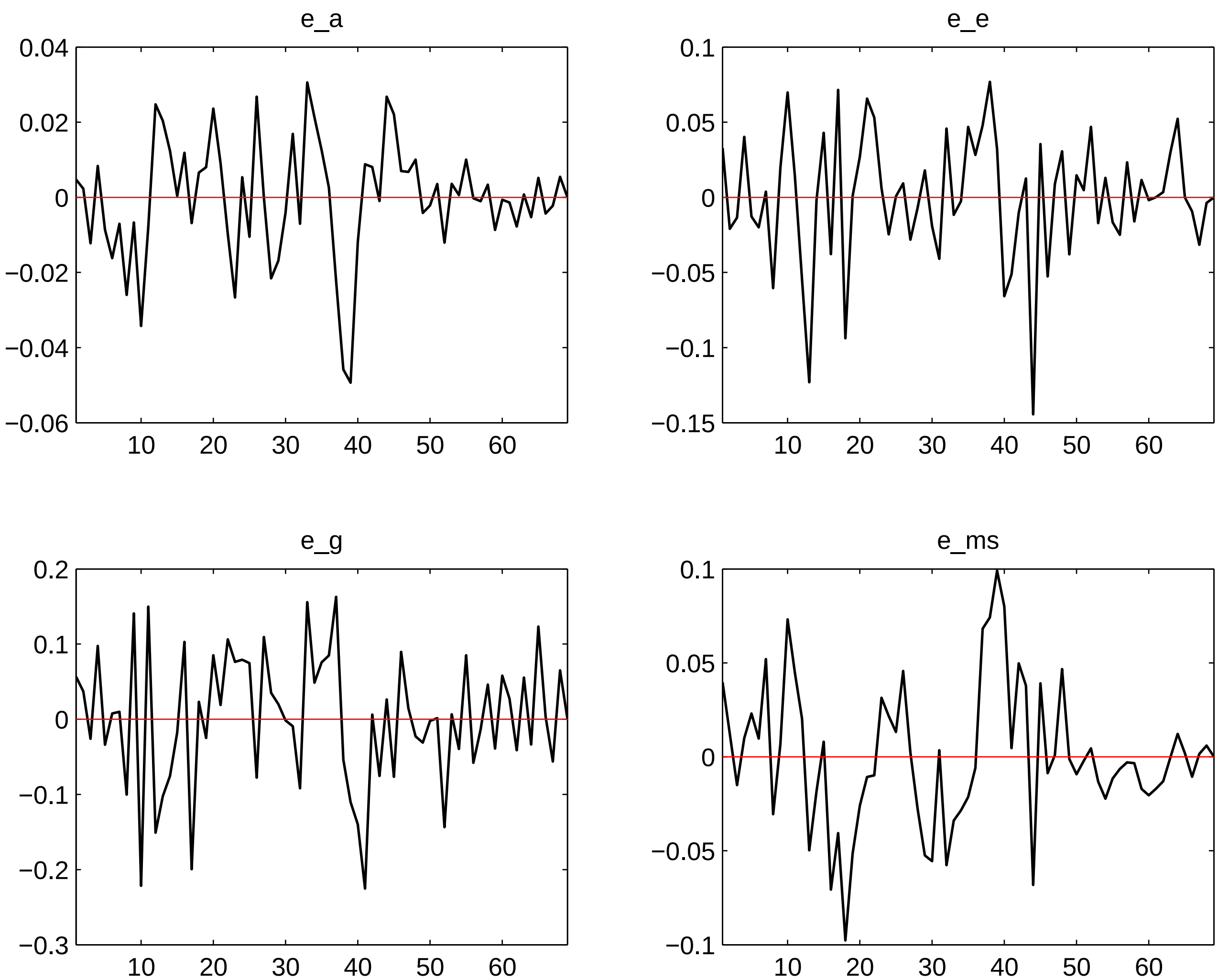


Figure 8: Smoothed Shocks (cont.)



References

ALTISSIMO, F., AND V. CORRADI (2003): "Strong rules for detecting the number of breaks in a time series," *Journal of Econometrics*, 117, 207–244.

AN, S., AND F. SCHORFHEIDE (2006): "Bayesian analysis of DSGE models," *Econometric Reviews*, forthcomming.

BAI, J., AND P. PERRON (1988): "Estimating and testing linear model with multiple structural changes," *Econometrica*, 66, 47–78.

BERNAL, R. (2002): "Monetary policy rules in Colombia," Documento cede, CEDE Universidad de los Andes.

CICCARELLI, M., AND B. MOJON (2005): "Global inflation," *Working paper series Central Banck of Chile*, 357.

CLARK, P. (1987): "The cyclical component of the U.S. economic activity," *Quarterly Journal of Economics*, 102, 797–814.

——— (1989): "Trend reversion in real output and unemployment," *Journal of Econometrics*, 40, 15–32.

CORVOISIER, S., AND B. MOJON (2005): "Breaks in the mean of inflation How the happen and what to do with them?," *Working paper series European Central Bank*, 451.

DURBIN, J., AND S. J. KOOPMAN (2001): *Time series analysis by state space methods*, Oxford Statistical science series. Oxford University press, Oxford, UK.

ERCEG, C., AND A. LEVIN (2003): "Imperfect credibility and inflation persistence," *Journal of Monetary Economics*, 50, 915–944.

HARVEY, A. C. (1990): *Forecasting, structural time series models and the Kalman filter*. Cambridge University press, Cambridge, UK.

KLEIN, P. (2000): "Using the generalized Schur form to solve a multivariate linear rational expectations model," *Journal of Economic Dynamics and Control*, 24, 1405–1423.

LEVIN, A. T., AND J. M. PIGER (2004): "Is inflation persistence intrinsic in industrial economies," *Working paper series European Central Bank*, 334.

LUBIK, T., AND F. SCHORFHEIDE (2003): "Computing sunspot equilibria in linear rational expectations models," *Journal of Economic Dynamics and Control*, 28, 273–285.

MARQUES, C. R. (2004): "Inflation persistence: facts or artefacts?," *Working paper European Central Bank*, 371.

RABANAL, P., AND J. RUBIO-RAMIREZ (2005): "Comparing wew keynesian models in ihe Euro Area: a bayesian approach," Discussion Paper WP2003-30a, Federal Reserve Bank of Atlanta.

SARGENT, T. (1999): *The conquest of American inflation*. Princeton University press, Princeton.

SCHMITT-GROHE, S., AND M. URIBE (2004): "Optimal simple and implementable monetary and fiscal rules," Working Paper 10253, NBER.

SCHORFHEIDE, F. (2000): "Loss function-based evaluation of DSGE models," *Journal of Applied Econometrics*, 15, 645–670.

WEST, M., AND J. HARRISSON (1999): *Bayesian forecasting and dynamic models*. Springer-Verlag, New York.