

Lack of Credibility, Inflation  
Persistence and Disinflation in  
Colombia

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# Lack of Credibility, Inflation Persistence and Disinflation in Colombia \*

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## Abstract

This paper measures inflation persistence in Colombia for the period 1990-2010 and estimates the implied speed at which agents learn about the central bank's inflation target. We estimate Erceg and Levin's (2003) imperfect credibility model using Bayesian techniques and compare the posterior odds of this model against a conventional Neoknesian model with ad-hoc price indexation. The odds are strongly in favor of the imperfect credibility model, suggesting that lack of credibility on the inflation target is an important source of inflation persistence. We use the model to compute the sacrifice ratio associated to 100 basis points inflation target shocks and find that it is (0.83%) in line with previous estimates for Colombia. We also find that the speed at which agents learn in the model has increased, albeit marginally, since the central bank implemented its inflation targeting strategy. Although during this period macroeconomic volatility has fallen, inflation persistence has remained roughly constant suggesting that so far, the impact of those credibility gains has been modest.

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

**JEL Classification Codes:** C52

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# 1 Introduction

Inflation is a persistent process. For long, this fact has called the attention of researchers and policy makers because it has important economic and policy implications. For instance, the central bank's policy response can vary depending on the degree of persistence of inflation shocks. If inflation shocks have short-lived effects, the monetary authority could react mildly to them, and inflation would stabilize soon around a given target, without much impact on credibility and macroeconomic volatility. However, if persistence is erroneously underestimated, delays in response to inflationary shocks could create relatively large deviations from the central bank's objectives, undermine the credibility of the central bank and create additional instability.

[Fuhrer \(2009\)](#), in a comprehensive survey on the persistence of inflation, shows that even though it constitutes a key feature of inflation dynamics, its definition and measurement are controversial. The study provides a “taxonomy” of the body of research on inflation persistence and distinguishes “reduced-form persistence” from “structural persistence”. The first refers to an empirical property of inflation without economic interpretation. The second refers to persistence that arises from identified economic structures that produce it. Fuhrer surveys a variety of methodologies to measure inflation persistence, from the estimation of simple autocorrelation functions to sophisticated filtering techniques, and concludes that regardless of how inflation persistence is defined and measured, it has declined somewhat in recent years. However, it is still a subject of debate how much it has declined.

Colombia has not been immune to this debate. The majority of local studies are econometric studies that characterize inflation persistence as some measure of the degree of “mean-reversion” of the inflation rate. To our knowledge, [Birchenall \(1999\)](#) is the first effort in Colombia to characterize inflation dynamics (See [Table 1](#)). Using data for the period 1965-1996 he finds that the estimate of the autoregressive component of the consumer price inflation is 0.6. Thus, the study characterized inflation as a persistent stationary process. More than ten years later, [Capistran and Ramos-Francia \(2009\)](#) using a similar approach (the sum of the estimated coefficients of an autoregressive process) find similar results for Colombia (0.67), but using a larger sample of the 10 largest Latin American economies, albeit for a shorter period (2000-2006).

More recently, [Echavarría, López, and Misas \(2010\)](#) and [Echavarría, Rodríguez, and Rojas \(2010\)](#), following the recent international literature, measure inflation persistence with the relative contributions

of permanent and transitory components of inflation, argue that the adoption of inflation targeting in 1999 caused an important reduction in mean and variance, but has not significantly modified the persistence of inflation.

Table 1: Recent estimates of inflation persistence in Colombia

Authors	Methodology	Estimated	
		Persis- tence	Sample
Birchenall (1999)	A recursive estimation of the autoregressive component of CPI inflation is carried using data from 1965 to 1996 as an additional exercise to analyze inflation dynamics in Colombia.	0.6	1965-1996
Capistran and Ramos- Francia (2009)	Econometric estimation of inflation persistence for the 10 largest Latin-American economies is done using univariate time-series methodologies and monthly data from 1980 to 2006. Estimation is also carried by sub-samples, selected by episodes of change in monetary policy regime. Reported inflation is the sum of autoregressive coefficients obtained using 2000-2006 sub-sample.	0.67	2000- 2006:06
Echavarria et al. (2010)	Estimation of inflation persistence in Colombia is carried using monthly data form CPI inflation between 1990 and 2010 and several econometric methods including markov-switching models and state-space models. Reported persistence is the estimate from the MSIAH model.	0.34	1990:1- 2010-6
Echavarría et al. (2010)	An unobserved components model with regime-switching is used to estimate persistence, as a second order autoregressive process, and structural breaks for several inflation indexes using quarterly sample ranging from 1979 to 2010. Reported persistence is the mean of the sum of autoregressive coefficients for the 1989-1999 and 1999-2010 sub-samples.	0.31	1979-2010

Indeed, Colombia used to be a country of high and volatile inflation. At the beginning of the nineties the newly independent central bank (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.<sup>1</sup> 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

<sup>1</sup>A crawling band was implemented during the first years of the 90s.

around the world, the Banco de la República established its first quantitative inflation target of 22%.

The disinflation process was long and gradual. The bank missed its inflation target for 6 years in a row. It was until 1999, in the midst of a mayor financial and economic crisis, that inflation reached a single digit. Later, inflation declined steadily (but slowly) from about 9% in 1999 to about 5% in 2005. Nowadays, after the 2007-2008 global financial crisis, Colombian inflation is under control and within a 2-4% inflation target range. With this record at hand in such a prolonged period, one can easily guess that credibility has not been one of the main assets of the central bank (see table 2). Thus, it is natural to think that lack of credibility may be one factor behind the persistence of inflation.

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	18.0	-0.32	-0.77	0.45	
1998	16.70	17.95	16.0	0.70	-1.25	1.95	
1999	9.23	15.789	15.0	-5.77	-6.56	0.79	
2000	8.75	9.89	10.0	-1.25	-1.14	-0.11	33.0
2001	7.65	8.85	8.0	-0.35	-1.20	0.85	46.9
2002	6.99	6.95	6.0	0.99	0.04	0.95	69.1
2003	6.49	6.58	5.5	0.99	-0.09	1.08	16.1
2004	5.50	6.13	5.5	0.00	-0.63	0.63	70.4
2005	4.86	5.78	5.0	-0.14	-0.92	0.78	79.0
2006	4.48	5.23	4.5	-0.02	-0.75	0.73	90.1
2007	5.69	4.50	4.0	1.69	1.19	0.50	25.9
2008	7.67	4.14	4.0	3.67	3.53	0.14	6.3
2009	2.00	4.65	5.0	-3.00	-2.65	-0.35	53.1
2010	3.17	5.22	3.0	0.17	-2.05	2.22	93.8

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 inflation target would be met that year.

Source: Banco de la República de Colombia.

The hypothesis that lack of credibility on monetary policy is a source of inflation persistence which, in turn, determines the sacrifice ratio is not new. For instance, [Ball \(1994\)](#), seminal contribution, was to showed that imperfect credibility can raise the output costs of disinflation. Also, [Sargent \(1999\)](#) argued 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. [Erceg and Levin \(2003\)](#), in another influential paper, 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. The paper calibrates a standard staggered contracts Neokeynesian model to the US economy and find that the cost of the Volcker disinflation was

1.6 percentage points for each percentage point reduction in the inflation rate.<sup>2</sup> 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 adaptive expectations. Later, [Kozicki and Tinsley \(2005\)](#) examined the financial market implications of shifts in the inflation target. Using a time-series model of the term structure they showed that failure to account for imperfect policy credibility may explain empirical rejections of the expectations hypothesis of the term structure of interest rates.

This paper asks whether lack of credibility on the central bank's inflation target could have played an important role in explaining the persistence of inflation observed in the Colombian data at business cycle frequencies. To that end we define the concept of persistence and use an econometric model to measure it. The econometric model is able to capture the low frequency fluctuations of the inflation rate and confirms the results found by [Echavarría, López, and Misas \(2010\)](#) and [Echavarría, Rodríguez, and Rojas \(2010\)](#) that the conventional measure of inflation persistence has not changed significantly in the last decade. Despite this, our results allows us to infer (heuristically) that the importance of persistent shocks to inflation relative to transitory shocks has diminished.

Our main contribution is that we are able to identify, through an economic model, the sources behind inflation persistence. We follow [Erceg and Levin \(2003\)](#) and use their imperfect information model to understand them. We compare it against the conventional Neokeynesian model with inflation indexation. Following [Schorfheide \(2000\)](#), we use Bayesian analysis to discriminate between them. We compute the posterior odds and find that the Colombian data supports the lack of credibility model.

We also use the model to estimate the speed at which agents learn about the ultimate intentions of the Colombian central bank. We find that credibility has been higher after the central implemented the inflation targeting strategy by the end of 1999. 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. To our knowledge, this joint estimation of the monetary policy rule under imperfect credibility is another new result for Colombia.<sup>3</sup>

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<sup>2</sup>Typically staggered contracts models have been criticized for not being able to reproduce the observed inflation persistence present in the data. Many modelers use tricks to induce persistence, like adding lags. One of the main implications of Erceg and Levin 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.

<sup>3</sup>[Erceg and Levin \(2003\)](#), in an influential paper, studied the episode of the Volcker disinflation in the US using a model in which agents have 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 1.7 percentage

In light of these results, we proceed to calculate the sacrifice ratio implied by the imperfect credibility model. The estimated sacrifice ratio (0.83%) for the full sample is in line with those estimated previously in the literature by [Gómez and Julio \(2000\)](#), [Reyes \(2003\)](#), [Sarmiento and Ramírez \(2005\)](#), but higher than those obtained by [Hamann, Julio, Restrepo, and Riascos \(2005\)](#) for Colombia, and [Hofstetter \(2007\)](#) for the average of 18 Latin American countries in a 30-years sample.

Table 3: Sacrifice Ratios for Colombia and Latin America

Authors	Methodology	SR Estimate	Period
<a href="#">Gomez &amp; Julio (2000)</a>	Econometric estimation of a set of equations describing the transmission mechanisms of monetary policy for Colombia using quarterly data from 1990Q1-2000Q1. The sacrifice ratio is calculated as the accumulated loss in the output gap after a shock of permanent reduction of annual inflation in 1%.	0.79%	Simulated Scenario
<a href="#">Reyes (2003)</a>	Identification of different disinflation periods in Colombia and estimation of the sacrifice ratio during each one using the standard methodology of Ball (1994) and alternative methodologies of Zhang (2001) and Cecchetti and Rich (1999). SR estimate corresponds to the result obtained using Zhang's method for the latest period analyzed.	0.89%	1991-2001
<a href="#">Sarmiento &amp; Ramirez (2005)</a>	Estimation of a monetary SVAR with short-run restrictions as in Buiter and Miller (1982). Estimates of the SR are the average difference of the steady-state product gap with and without inflationary shocks for the period 1998-2003, obtained from the historical decomposition of estimated shocks.	0.88%	1998-2003
<a href="#">Hamann et al.(2005)</a>	Simulation of a small open economy DSGE calibrated for Colombia. Calculation of costs of reducing the steady state inflation rate from 5.5% to 3% from the transition dynamics simulated after this permanent shock.	0.04%	Simulated Scenario
<a href="#">Hofstetter (2007)</a>	Identification of disinflation episodes between 1973 and 2000 for 18 Latin-American Countries and estimation of average SR for each of the three decades composing the sample using three alternative methodologies. Results show a negative disinflation cost for the 1990-2000 period under all methodologies. Results reported are the average of this period estimate under the LL&L methodology.	-0.57%	1990-2001

\*What percent the current output has to fall from its long run level, due to a reduction of 1% in trend inflation

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 adaptive expectations

The rest of paper proceeds as follows: in the next section we describe the main facts about inflation persistence in Colombia for the period 1990-2010. Then, in section 3, we describe the model. In the fourth section we briefly present the estimation procedure. In the fifth section we report the results. In the sixth we compare the two models while in the seventh we compute the sacrifice ratios under imperfect credibility. The last section concludes.

## 2 Our measure of inflation persistence

Standard measures 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 constant mean.<sup>4</sup> [Marques \(2004\)](#) 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 \(2005\)](#) and [Ciccarelli and Mojon \(2010\)](#) 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. Furthermore, [Echavarría, López, and Misas \(2010\)](#) finds similar evidence for Colombia. In these papers inflation persistence is measured as sum of the auto regressive 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 \(2003\)](#).<sup>5</sup>

In line with this strand of the empirical literature, to characterize the evolving changes in the mean of inflation, we use the following state-space model:

$$\begin{aligned}
 y_t &= \mu_t + x_t & (1) \\
 x_t &= \rho x_{t-1} + \nu_{3t} \\
 \mu_t &= \beta_{t-1} + \mu_{t-1} + \nu_{1t} \\
 \beta_t &= \beta_{t-1} + \nu_{2t}
 \end{aligned}$$

with  $\nu_{it} \sim N(0, \sigma_{\nu_i}^2)$ ,  $i = 1, 2, 3$  and  $E[\nu_{it}\nu_{js}] = 0$  for  $i \neq j$  and  $t \neq s$ . Equation (1) decomposes the inflation

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<sup>4</sup>There are other approaches to measure inflation persistence, like the estimation of the autocorrelation function and unit root tests, to mention just two of them. Recently, in Colombia, [Echavarría, López, and Misas \(2010\)](#) compares different measurements of statistical persistence and estimates the evolution of inflation and inflation gap persistence in Colombia for the period 1990-2010 using a regime switching model and a Kalman filter.

<sup>5</sup>Cogley, Primiceri, and Sargent (2007) and Cogley and Sbordone (2009) stress the importance in econometric models of inflation of recognizing the low frequency movement of inflation. They called it “trend inflation”.



process into an 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  $\rho \in (-1, 1)$  forms a persistence measure. The trend component is given by  $\mu_t$  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 Harrison \(1999\)](#) and [Durbin and Koopman \(2001\)](#) for details. The advantage of (1) compared to other approaches is that trend and persistence are modeled simultaneously rather than sequentially<sup>6</sup>.

We use alternative indexes to measure inflation. 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:  $\pi_{\text{CPI-NF}}$  that excludes food from the CPI,  $\pi_{\text{CPI-T}}$  and  $\pi_{\text{CPI-NT}}$  that includes only traded goods and non-traded goods,  $\pi_{\text{CPI-R}}$  with only regulated goods and  $\pi_{\text{CPI-B}}$  that excludes food and regulated prices from the CPI. The sample consists of quarterly data for the period 1988:1 to 2010:4.

We report the results in [Table 4](#) and [Figures 7 to 8](#). These results confirm that most measures of inflation display an important amount of persistence at the business cycle frequency. The only exceptions being the inflation of regulated goods; for which the estimated trend component of inflation follows closely the observed inflation, so deviations from trend quickly revert to the mean, and non-traded sector, which displays a low persistence, compared with the other measures.<sup>7</sup> Notwithstanding this, most measures of inflation display high persistence.

Table 4: Estimated persistence for the different inflation rates.

	$\pi_{\text{CPI}}$	$\pi_{\text{CPI-T}}$	$\pi_{\text{CPI-NT}}$	$\pi_{\text{CPI-NF}}$	$\pi_{\text{CPI-R}}$	$\pi_{\text{CPI-B}}$
$\hat{\rho}$	0.33	0.60	0.24	0.66	-0.02	0.84

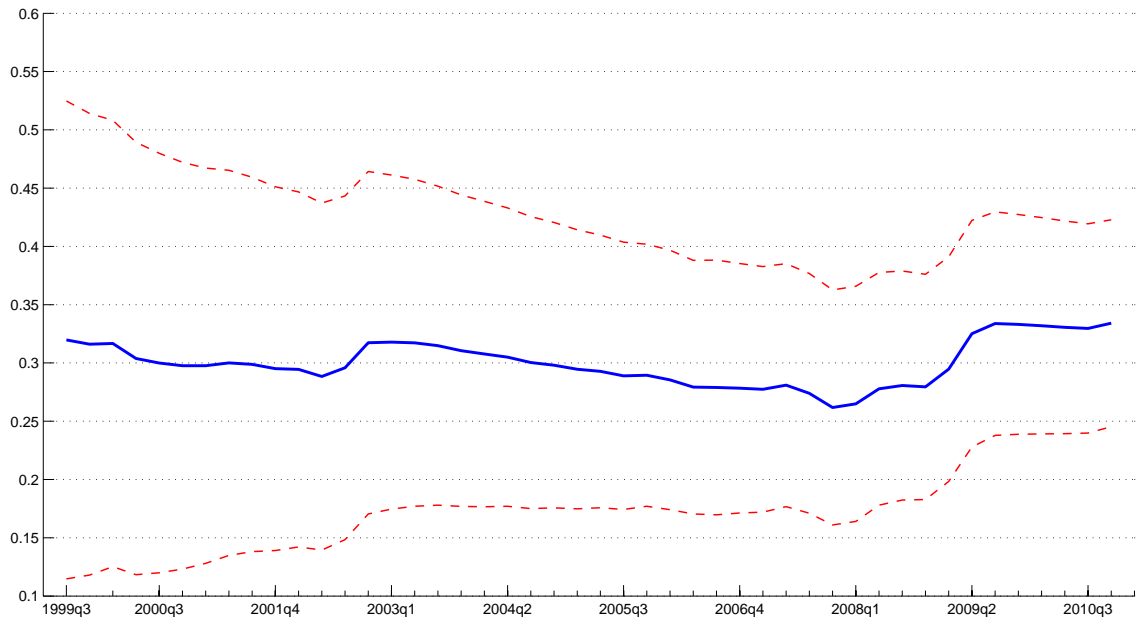
To have an indication of how much inflation persistence has changed in the last years, we perform a one-quarter rolling-window estimation of the parameter  $\rho$  in the stationary AR(1) process  $x_t = \rho x_{t-1} + \nu_{3t}$ , starting in the third quarter of 1999 and ending in the fourth quarter of 2010. This period covers the inflation targeting regime. [Figure 1](#) shows the result.

We confirm the findings in [Echavarría, López, and Misas \(2010\)](#) and [Echavarría, Rodríguez, and Rojas \(2010\)](#), which show that inflation persistence, measured as the component defined by the autoregressive

<sup>6</sup>A similar model has been used by [Clark \(1987, 1989\)](#) to decompose the US real GDP between trend and cycle

<sup>7</sup>One explanation could be that price formation in the non-traded sector could be better anchored to the inflation target. The non-traded 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 non-traded goods). We speculate that this event may have caused a price setting behavior that puts an important weight on inflation expectations. This is an open question.

Figure 1: CPI Inflation Persistence and volatility: Rolling



coefficient has remained roughly constant. However, it is worth taking a look at the right panels of Figures 7 to 8. A visual inspection of these graphs show that the relative importance of permanent shocks to CPI inflation has declined (relative to transitory shocks), in particular for the period 2001-2007. Measuring this relative volatility accurately is very difficult. Thus, if we take this graphical analysis as an heuristic measure of persistence, we can cautiously say that inflation persistence has somewhat declined.

In the next section we use a model in which lack of credibility plays a key role in explaining inflation dynamics at the business cycle frequency.

### 3 The model

Following Rabanal and Rubio-Ramírez (2005), 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 through equations (2) to (4):

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

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

$$i_t = \gamma_i i_{t-1} + (1 - \gamma_i) [\gamma_\pi (\pi_t - \bar{\pi}_t) + \gamma_y y_t] + z_t, \quad (4)$$

where  $\pi_t$  is the inflation rate,  $\bar{\pi}_t$  is the inflation target,  $mc_t$  is the real marginal cost and  $i_t$  is the nominal 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  $u_t$ , which represents the present value of private agent's error when forecasting future inflation and distinguishes this model from a standard Neoclassical model. 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  $\alpha$  is the share of labor factor in production,  $\theta$  the probability of keeping prices fixed during the period,  $\epsilon$  the elasticity of substitution between slightly differentiated types of goods and  $\beta \in (0, 1]$  the discount factor.  $\lambda$  measures the slope of the Phillips curve.  $\sigma > 0$  is the elasticity of intertemporal substitution and  $\gamma_\pi$ ,  $\gamma_y$  and  $\gamma_r$  measure the degree of responsiveness of the monetary authority to deviations from target, the output gap and past interest rate, respectively.

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

$$y_t = a_t + (1 - \alpha)n_t \quad (5)$$

$$mc_t = w_t + n_t - y_t \quad (6)$$

$$mrs_t = \frac{1}{\sigma} y_t + \eta n_t - g_t \quad (7)$$

$$w_t = mrs_t \quad (8)$$

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. Finally,  $\eta$  is the inverse of the elasticity of labor supply to the real wage.

The exogenous variables of the model evolve according to the following stochastic processes:

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

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

$$z_t = \varepsilon_t^i$$

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

When information about the inflation target is perfect,  $\bar{\pi}_t$  is known for all  $t$ . The interesting case is when information is not perfect. As proposed in [Erceg and Levin \(2003\)](#),  $\bar{\pi}_t$  varies over time due to a combination of a white noise shock,  $\varepsilon_t^q$ , and a shock  $\varepsilon_t^p$  with persistent effects on the inflation target. The central bank's reaction function is observable to agents, but the underlying components of the inflation target are not. Therefore, private agents must solve a signal extraction problem to infer the components of  $\bar{\pi}$ .

In this model, the central bank's inflation target is the sum of a constant steady-state of inflation  $\bar{\pi}$  and two zero-mean stochastic autoregressive components,  $\pi_t^p$  and  $\pi_t^q$ . The former is assumed to have an autoregressive root close to unity while the latter is assumed to have a much smaller autoregressive root. That is, the inflation target evolves according to the following process:

$$\bar{\pi}_t - \bar{\pi} = (\pi_t^p - \bar{\pi}) + \pi_t^q \tag{9}$$

where the time-varying components follow the first order vector autoregression:

$$\begin{bmatrix} \pi_t^p - \bar{\pi} \\ \pi_t^q \end{bmatrix} = \begin{bmatrix} \rho_p & 0 \\ 0 & \rho_q \end{bmatrix} \begin{bmatrix} \pi_{t-1}^p - \bar{\pi} \\ \pi_{t-1}^q \end{bmatrix} + \begin{bmatrix} \varepsilon_t^p \\ \varepsilon_t^q \end{bmatrix}. \tag{10}$$

We can write equations (9) and (10) in state-space form defining  $Z_t = [\pi_t^p - \bar{\pi}, \pi_t^q]'$ ,  $F = \text{diag}(\rho_p, \rho_q)$  and  $H = [1, 1]$ . In particular, the state equation

$$Z_t = FZ_{t-1} + \varepsilon_t$$

represents equation (10).

Households and firms are assumed to use optimal filtering to solve this signal-extraction problem; this requires the inflation-target innovations  $\varepsilon_t^p$  and  $\varepsilon_t^q$  to be mutually uncorrelated, with variances  $\nu_p$  and  $\nu_q$  respectively, and to be uncorrelated with any other shocks in the economy. With these assumptions, the Kalman filter can be used by private agents of the economy to obtain optimal estimates of the unobserved

components through the following recursion:

$$E_t Z_t = F E_{t-1} Z_{t-1} + L_{gain}(\bar{\pi}_t - H F E_{t-1} Z_{t-1}) \quad t = 0, \dots, t$$

where  $K_{gain} = F L_{gain}$  is the Kalman gain matrix and  $L_{gain}$  determines how agents respond to the forecast error,  $\bar{\pi}_t - H F E_{t-1} Z_{t-1}$ , by updating their estimates of the underlying components of the inflation target. Therefore, given the current estimate  $E_t Z_t$  of these components, the optimal forecast of the inflation target  $j$  periods ahead is given by:

$$E_t \bar{\pi}_{t+j} = \bar{\pi} + H F^j E_t Z_t$$

For simplicity, [Erceg and Levin \(2003\)](#) assume  $\rho_q = 0$  and  $\rho_p = 1$ . Thus, the households and firms expectations of the future inflation target depend only on a constant  $\bar{\pi}$  and the expectation of the highly persistent component of the target. Specifically, the persistent component of the inflation target evolves according to:

$$E_t(\pi_t^p - \bar{\pi}) = E_{t-1}(\pi_{t-1}^p - \bar{\pi}) + \kappa(\bar{\pi}_t - E_{t-1}\bar{\pi}_t). \quad (11)$$

So, agents update their assessment of the persistent component of inflation target by the product of the forecast error innovation and the Kalman gain parameter,  $\kappa$ .

Returning to Equation (2),  $u_t$  is the present value of the forecast error of private agents in the prediction of future inflation, that is:

$$u_t = \beta (\hat{\pi}_{t+1} - E_t \pi_{t+1}) \quad (12)$$

where  $\hat{\pi}_{t+1}$  is the rational forecast given all information available to private agents at time  $t$ , obtained by the optimal filtering process given by Equation (11). By replacing (12) 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 of the inflation target.

We can use Equations (12) and (2)-(4) to derive the following expression for the evolution of  $u_t$ :

$$u_t = (1 - \kappa)u_{t-1} + (1 - \kappa)\varepsilon_t^c$$

where  $\kappa$  is the Kalman gain parameter from Equation (11) that determines the speed at which agents learn to distinguish between the two components of the inflation target, and  $\varepsilon_t^c$  is a normally distributed zero mean shock with standard deviation  $\sigma_c$ , associated with the shocks to the permanent component of the inflation target. Notice that in the case of full information,  $\kappa = 1$ , agents learn at the highest possible rate, implying  $u_t = 0$  and therefore  $E_t\pi_{t+1} = \hat{\pi}_{t+1}$ .

Of particular interest is the value of the learning parameter  $\kappa$ . 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 Neoknesian model augmented by learning about the inflation target. Unlike [Erceg and Levin \(2003\)](#), where parameter  $\kappa$  is calibrated from survey data, we estimate it using this simplified version of their model and Bayesian methods. The advantages of estimating models using a Bayesian approach are discussed formally in [Lubik and Schorfheide \(2003\)](#). A review of the Bayesian tools for macro economists is presented in [An and Schorfheide \(2007\)](#).

## 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 \(2000\)](#) 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 measurement 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 or some point nearby with a high probability density, 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 1990:1 to 2010:4. 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 measured using the manufacturing industry real wage index from Banco de la República.

## 4.2 Priors

Let  $d_t = (\pi_t, y_t, i_t, w_t)$  denote the observed data and define the vector of parameters to be estimated as  $\Phi = \left(\frac{1}{\sigma}, \eta, \theta, \kappa, \gamma_i, \gamma_\pi, \gamma_y, \rho_a, \rho_g, \sigma_a, \sigma_g, \sigma_i, \sigma_c\right)$ . We impose strong priors on  $\beta$ ,  $\alpha$  and  $\epsilon$ . We set  $\beta = 0.98$ , which implies a real annual return close to 4%. To replicate the labor factor compensation share in real GDP we set  $\alpha = 0.36$ . The elasticity of substitution of the good produced by intermediate firms,  $\epsilon$ , is set to 6, which is a standard value in the literature<sup>8</sup>.

As explained in [Rabanal and Rubio-Ramírez \(2005\)](#), there is an identification problem in the model between the probability of adjusting prices and the elasticity of substitution. That is,  $\theta$  and  $\epsilon$  cannot be identified separately. In order to circumvent the identification problem we choose to estimate  $\theta$  for a given markup, since the estimated parameter tells us about the implicit frequency at which firms adjust prices in Colombia. Several studies, including [Misas, López, and Parra \(2009\)](#) and [Bonaldi, González, , and Rodríguez \(2010\)](#) , suggest that Colombian firms set prices every one or two quarters; we choose a beta-prior distribution for  $\theta$  with mean 0.36, which is the value estimated in [Bonaldi, González, , and Rodríguez \(2010\)](#) baseline model and implies that firms set prices every 4 or 5 months, on average.

The prior distributions for the rest of the parameters in vector  $\Phi$  are reported in [Table 5](#). The inverse of the elasticity of intertemporal substitution 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. For the Taylor rule coefficients we use the priors that are commonly used in the literature:  $\gamma_\pi = 1.5$  and  $\gamma_y = 0.125$ . 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; we

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<sup>8</sup>We use strong priors for  $\beta$  and  $\alpha$  because the model does not have capital and so the likelihood does not have information for their estimation.

Table 5: Prior and Posterior Distributions: Imperfect and Full Credibility

Parameter	Prior Distribution			Posterior Distribution			
	Distribution	Mean	Std Dev	Imperfect Credibility Model		Full Credibility Model	
				Mean	HPD-90 Interval	Mean	HPD-90 Interval
$\frac{1}{\sigma}$	Gamma	2.3	0.5	5.38	(4.27 - 6.44)	5.05	(3.96 - 6.17)
$\eta$	Gamma	3.2	0.5	3.40	(2.62 - 4.11)	4.15	(3.31 - 5.04)
$\theta$	Beta	0.36	0.02	0.29	(0.26 - 0.32)	0.37	(0.34 - 0.40)
$\kappa$	Uniform [0,1)	0.5	0.29	0.19	(0.07 - 0.32)	-	-
$\nu$	Uniform [0,1)	0.5	0.29	-	-	0.84	(0.66 - 1.00)
$\gamma_i$	Uniform [0,1)	0.5	0.29	0.11	(0.00 - 0.22)	0.17	(0.04 - 0.29)
$\gamma_\pi$	Normal	1.5	0.25	1.97	(1.68 - 2.24)	1.60	(1.31 - 1.89)
$\gamma_y$	Normal	0.13	0.13	0.14	(-0.07 - 0.34)	0.22	(0.02 - 0.42)
$\rho_a$	Uniform [0,1)	0.5	0.29	0.48	(0.38 - 0.59)	0.47	(0.33 - 0.62)
$\rho_g$	Uniform [0,1)	0.5	0.29	0.65	(0.58 - 0.74)	0.72	(0.64 - 0.80)
$\sigma_a$	Inv. Gamma	0.01	$\infty$	0.008	(0.006 - 0.009)	0.009	(0.007 - 0.01)
$\sigma_g$	Inv. Gamma	0.01	$\infty$	0.05	(0.04 - 0.06)	0.05	(0.04 - 0.06)
$\sigma_i$	Inv. Gamma	0.01	$\infty$	0.028	(0.02 - 0.03)	0.026	(0.02 - 0.03)
$\sigma_c$	Inv. Gamma	0.01	$\infty$	0.01	(0.008 - 0.01)	-	-
$\log(\hat{L})$					940.06		746.71

use the estimates obtained in [Prada and Rojas \(2010\)](#) as priors for the estimation of the inverse of labor supply elasticity ( $\eta$ ) and the inverse of the intertemporal substitution elasticity ( $\sigma$ ). For all autoregressive parameters we use a uniform prior between [0,1) and, for all standard deviations of shocks, an inverse gamma distribution with mean 0.01.

As for the credibility parameters, results for the US suggest that  $\kappa$  is around 0.13. Nevertheless, we use an uniform which accounts for the lack of evidence regarding the value of these parameters in emerging countries and Colombia, particularly.

## 5 Results

We use a Random Walk Metropolis-Hastings algorithm to draw four chains of 200.000 draws from the posterior distribution of  $\Phi$  and construct the estimates for each parameter using half the draws of each chain. The acceptance rates for each chain were between 0.3 and 0.4. We use methods developed by [Brooks and Gelman \(1998\)](#) to monitor the convergence of the posterior draws estimates.<sup>9</sup> Estimation results are shown in [Table 5](#).

The data supports the idea of lack of credibility as  $\kappa$ , the speed at which agents learn in the economy,

<sup>9</sup>The results of this exercise are available upon request.



is different than one and closer to zero. So, we can reject the hypothesis of perfect credibility. Our findings show that the posterior mean of the speed of learning is  $\kappa = 0.19$ , while the estimated probability of keeping prices fixed during a quarter is 0.29, lower than the estimate obtained for the full-credibility model (0.37). We argue that this is due to imperfect credibility, which captures the persistence of inflation more closely than the standard model.

For the policy rule parameters, we find a the posterior mean response to inflation of  $\gamma_\pi = 1.97$ , to the output gap of  $\gamma_y = 0.14$  and a smoothing desire of  $\gamma_i = 0.11$ . The response to inflation differs from our prior, 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. However [Bernal \(2002\)](#), using a classical approach to estimate Taylor rules in a partial equilibrium model for Colombia, finds that  $\gamma_\pi = 1.34$ ,  $\gamma_y = 0.19$  and  $\gamma_i = 0.10$ . 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.

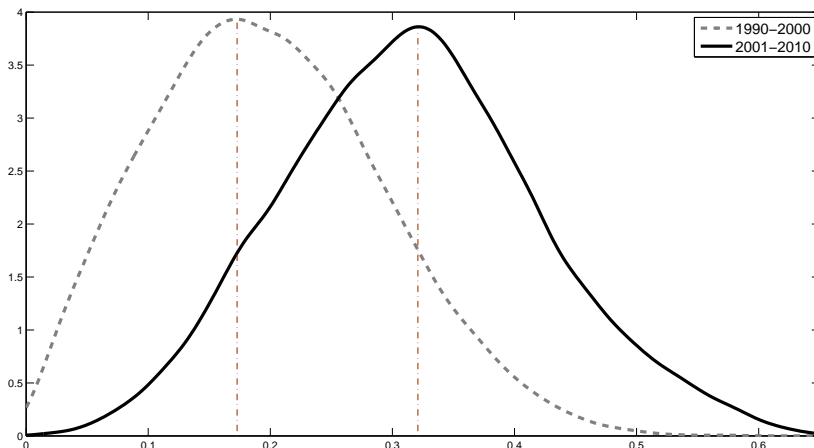
The 90% highest-posterior density interval (HPD) for the intertemporal rate of substitution coefficient ( $\sigma$ ) is of (4.27 – 6.44). The point estimate is 5.38, which lies in the upper end of the estimates obtained for the Emerging Market economies in the International Macroeconomics literature. The high values of the estimates of the coefficient of intertemporal rate of substitution reflect the higher variability of the macroeconomic time series typically found in Emerging Markets.

We also obtain an estimation of the labor supply elasticity. The posterior mean of  $\eta$  is 3.4, implying a labor supply elasticity of 0.29, which is in line with the results of [Prada and Rojas \(2010\)](#), who estimate the Frisch elasticity in Colombia for the period 2001-2006 obtaining a value of 0.31.

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 their autocorrelation coefficients are 0.48 and 0.65, respectively, with a standard deviation of 0.72 percent and 4.86 percent. The posterior mean of the volatility of the interest rate rule shock and the inflation target is 2.8 percent and one percent, respectively. 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.

We perform an additional exercise to assess the impact of inflation targeting on the estimated speed at which agents learn about the inflation target. We split the sample in two periods: one, from 1990 to 2000 and the second from 2001 to 2010. The break corresponds (approximately) to the date in which we think inflation targeting was implemented (by the end of 1999).

Figure 2: Estimated speed of learning  $\hat{\kappa}$  for pre- and post-inflation targeting periods



The graph shows a faster speed of learning for the post inflation targeting sample. Also notice that the full-sample estimated value of  $\kappa$  is closer to the value obtained for the first sample, the years before inflation targeting. Most of the information in the data appears to be concentrated in that part of the sample. This suggests that there are credibility gains in the last ten years. Nonetheless, the fact that inflation persistence has fallen, albeit not in a significantly, suggests that the impact of those gains has not been large.

Yet, the model reveals one important fact of the Colombian business cycle. A quick inspection of the smoothed shocks, shown in figure 6, obtained in our estimation shows that the variability of all shocks is lower since the end of 2000, close to the date in which the inflation targeting regime was implemented.<sup>10</sup>. In this sense we could argue that the implementation of IT in Colombia is associated with a higher degree of macroeconomic stability. This result is consistent with our finding of the previous section that the relative importance of permanent shocks has diminished in the last few years.

In the next section we compare the lack of credibility model to a standard Neoknesian model with *ad-hoc* inflation indexation.

<sup>10</sup>See figure 6. The inflation targeting regime starts from position 50 in the graph

## 6 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. We estimate such a model by replacing equation 2 with:

$$\pi_t = \phi_1 \pi_{t-1} + \phi_2 E_t \pi_{t+1} + \phi_3 m c_t$$

Where  $\phi_1 = \frac{\nu}{1+\nu\beta}$ ,  $\phi_2 = \frac{\beta}{1+\nu\beta}$ ,  $\phi_3 = \frac{\lambda}{1+\nu\beta}$  and  $\nu$  is the degree of price indexation to past inflation <sup>11</sup>. We use the same priors for the estimated parameters in order to compare the performance of the imperfect credibility model against the conventional model. Table 5 reports the results of the estimation of the Neokeynesian model with price 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  $\theta_p$  and  $\theta_i$  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  $p(Y_T | \mathcal{M}_m)$ . We compute the marginal density of the data conditioned on the model:

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

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

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

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

$$\frac{p(\mathcal{M}_i | Y_T)}{p(\mathcal{M}_p | Y_T)} = \frac{p(\mathcal{M}_{pi}) p(Y_T | \mathcal{M}_i)}{p(\mathcal{M}_o) p(Y_T | \mathcal{M}_o)}.$$

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<sup>11</sup>We use a uninformative prior for the estimation of this parameter, as in [Rabanal and Rubio-Ramírez \(2005\)](#).

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

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

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

We find that  $p(Y_T | \mathcal{M}_i) = \exp(940.1)$  while  $p(Y_T | \mathcal{M}_p) = \exp(746.7)$  implying a log-Bayes factor around 193, so that the odds are strongly in favor of the imperfect credibility model.<sup>12</sup>

One 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, we can identify three sources of inflation persistence: extrinsic, intrinsic and expectations-driven. The first can be associated with the coefficient that accompanies the real marginal cost,  $\lambda$ , 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, while in an imperfect credibility model we can associate the degree of persistence with parameter  $\kappa$ .

To analyze the degree of extrinsic persistence the key parameter is the probability of keeping prices fixed during a quarter. Recall that in the imperfect information model this probability is 29 percent, implying price changes, in average, every 1.4 quarters. In the full information model with price indexation this parameter value raises to 1.5 percent implying price changes every 1.54 quarters. It seems that the assumption of price indexation coupled with full information tends to increase the degree of extrinsic persistence by lowering the responsiveness of inflation to the real marginal cost<sup>13</sup>.

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

Finally, the expectations-driven persistence on the full information model is closely linked to the intrinsic one through the discount factor, while in the imperfect information model is determined by parameter  $\kappa$ . Our results show that this is our main source of inflation persistence in Colombia. Agents learn at a relatively slow pace. Compared to the US disinflation period, our rate is more than one-fourth times the rate at which agents learn in the US. We could associate this result to the fact that the inflation

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<sup>12</sup>We use the Laplace approximation.

<sup>13</sup>To see this we draw from the distribution of parameter  $\lambda$  and find that its mode on the case of price indexation is 0.098, while in the case of imperfect credibility is 0.274.

target policy that the central bank follows in Colombia which has allowed anchoring inflation expectations to a long run inflation target.

The estimation of the policy rule is consistent with this idea. The responsiveness of the central bank to deviations of inflation from target is higher under imperfect credibility than under full information and price indexation. This means that the central bank has to exert more effort when it faces a lack of credibility problem than when it faces a price indexation problem. In the next section we use the model to estimate how costly has been to disinflate without anchored expectations.

## 7 Disinflation Costs

How large have been 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.

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.

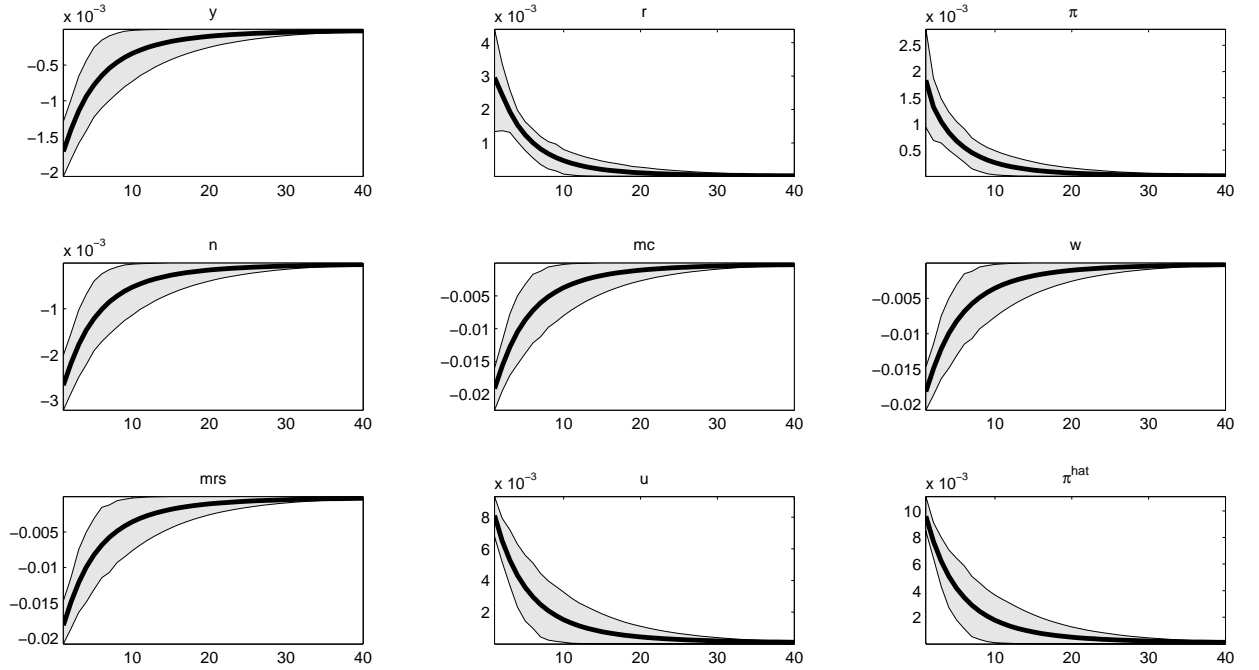
Given the estimated speed of learning, we compute the macroeconomic effects of a 100 bp disinflation focusing on the sacrifice ratio and the monetary policy effort. We measure the sacrifice ratio as the present value of the output gap during the disinflation period. To measure the monetary policy effort 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 to check the sensibility of our measurement of the sacrifice ratio to the discount factor used. Table 6 reports the sacrifice ratio and the monetary policy effort under alternative discount factors.

In the benchmark estimation, a 100 basis points shock to the inflation target requires the central bank

Table 6: Disinflation Costs and Monetary Policy Effort (basis points)

	$\beta$				
	0.95	0.96	0.97	0.98	0.99
Sacrifice Ratio	75	77	80	83	86
Policy Effort	134	138	144	149	155

Figure 3: Response to a decrease of the inflation target under imperfect credibility



to keep interest rates by nearly 150 bp above average generating a sacrifice of 83 bp in output. The effects of the disinflation shock takes about three to five years to dissipate. 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.

This sacrifice ratio is in line with those estimated previously in the literature by [Gómez and Julio \(2000\)](#), [Reyes \(2003\)](#), [Sarmiento and Ramírez \(2005\)](#), but higher than those obtained by [Hamann, Julio, Restrepo, and Riascos \(2005\)](#) for Colombia, and [Hofstetter \(2007\)](#) for the average of 18 Latin American countries in a 30-years sample.

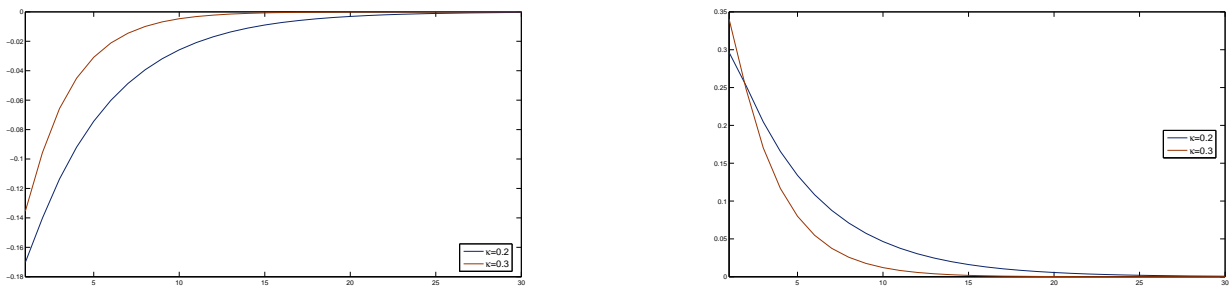
The differences with respect to these last two papers arise from the type of model used and from the type of estimation used. With respect to [Hamann, Julio, Restrepo, and Riascos \(2005\)](#), it is well known

that Neokeynesian models without strong *ad-hoc* persistence (as the one used in that paper) exhibit small sacrifice ratios. With respect to Hofstetter (2007), as the author explains, the negative sacrifice ratios arise from a unique set of conditions that occurred in Latin America during the period 1990-2000.

The sensitivity of this result depends on the discount factor used to compute the present value of the cost in terms of output. Using a discount factor of  $\beta = 0.95$ , the disinflation cost falls to 75 bp. Interest rates have to be 134 bp higher on average during the disinflation period.

More importantly, to gauge the impact of the degree of imperfect credibility on the sacrifice ratio, we can use the model to compute the disinflation cost under alternative values of  $\kappa$ , the speed of learning parameter. Notice how the larger the  $\kappa$  the smaller the disinflation cost and the lower the effort of the central bank.

Figure 4: Output and interest rate gaps after a decrease of 1% in the inflation target



## 8 Final Remarks

This paper asked whether lack of credibility on the central bank's inflation target could have played an important role in explaining the persistence of inflation observed in the Colombian data at business cycle frequencies. To answer that question we used a state-space model which decomposed the inflation process into a time-varying mean component and the cyclical movement around it. The model was estimated using the Kalman filter confirming the results found by Echavarría, López, and Misas (2010) and Echavarría, Rodríguez, and Rojas (2010) that inflation persistence has not changed significantly in the last decade. Despite this, we argue that our results allows us to infer (heuristically) that the importance of persistent shocks to inflation relative to transitory shocks has diminished. We take this as indirect evidence that inflation persistence has somewhat declined during the disinflation process.

More importantly, we were able to identify, using [Erceg and Levin \(2003\)](#) model, the sources behind inflation persistence. We estimated and tested that model against the conventional Neokeynesian model with inflation indexation using Bayesian analysis and showed that the posterior odds supports the lack of credibility model.

As a by-product of our estimation, we obtained estimates of the monetary policy rule of the central bank under imperfect credibility on the inflation target. We find that the central bank reacts more to inflation, less to output and smooths less the policy rate in comparison to the standard Neokeynesian model under full-credibility and ad-hoc inflation indexation.

We then obtained estimates of the speed at which agents learn about the ultimate intentions of the Colombian central bank. We found that credibility has been higher after the central implemented the inflation targeting strategy by the end of 1999. Although the standard measure of inflation persistence has remained roughly unchanged, the data and our econometric methodology allow us to infer that the importance of permanent shocks relative to transitory shocks has declined somewhat. Thus, we interpret these results as evidence that the credibility of the central bank has increased and has reduced inflation persistence but those gains have been modest. Why this happens is an important open question for future research.

Finally, we calculated the sacrifice ratio implied by the imperfect credibility model and found a sacrifice ratio (0.83%) in line with those previously estimated in the literature.

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# Appendix 1: Imperfect Credibility NKM-Estimation Results

Figure 5: Prior and Posterior Distributions:

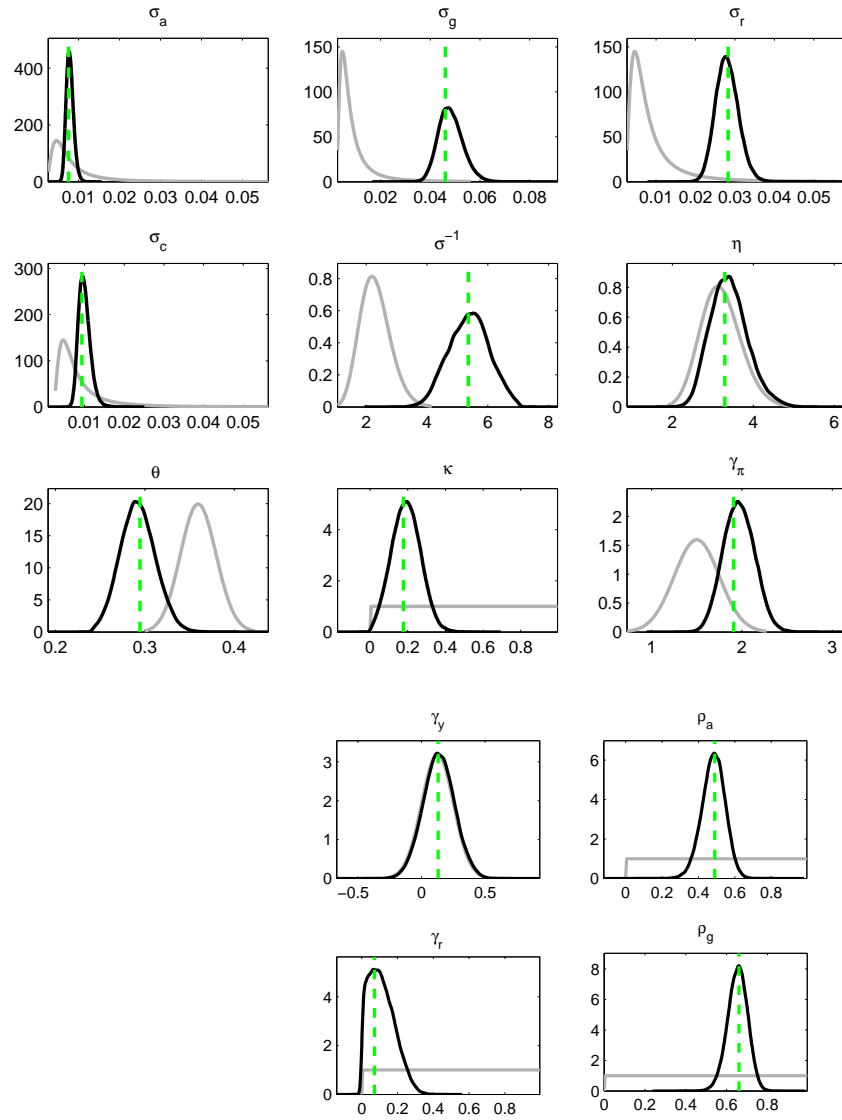
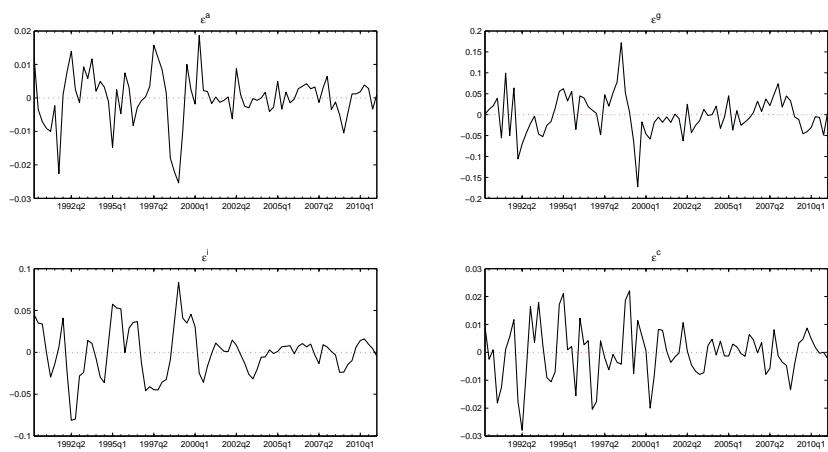


Figure 6: Imperfect Credibility Model: Smoothed Shocks



## Appendix 2: Inflation Persistence - Estimation Results

Table 7: Estimated  $\mu_t$  and  $x_t$ :  $IPC$  and  $IPC_t$

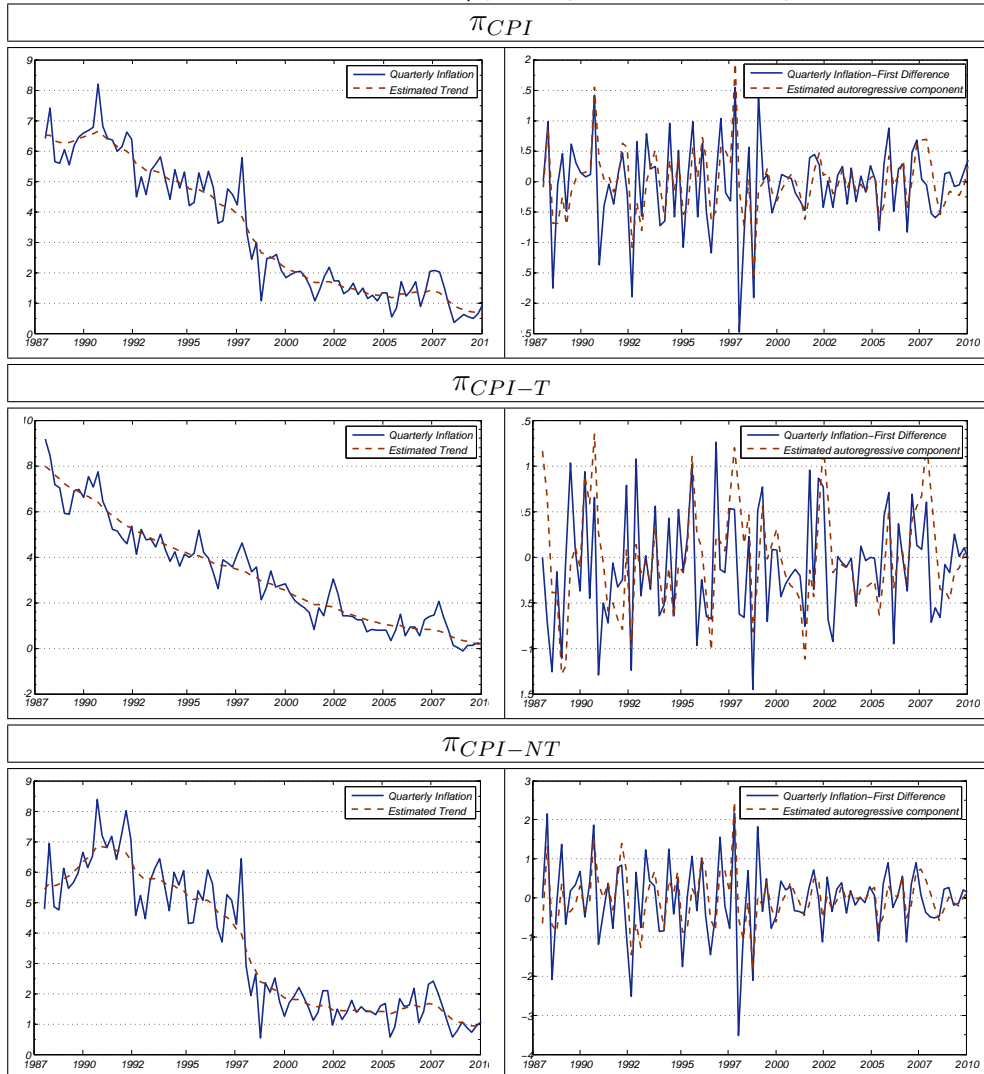


Table 8: Estimated  $\mu_t$  and  $x_t$ :  $IPC$  and  $IPC_t$

