BANCO CENTRAL DE RESERVA DEL PERÚ

Monetary Policy, Regime Shifts, and Inflation Uncertainty in Peru (1949-2006)

Paul Castillo Alberto Humala y Vicente Tuesta*

* Banco Central de Reserva del Perú

DT. N° 2007-005
Serie de Documentos de Trabajo
Working Paper series
Marzo 2007

Monetary Policy, Regime Shifts, and Inflation Uncertainty in Peru (1949-2006)*

Paul Castillo   Alberto Humala†   Vicente Tuesta
Central Reserve Bank of Peru

October 2006

Abstract

This paper evaluates the link between inflation and inflation uncertainty in a context of monetary policy regime shifts for the Peruvian economy. We use a model of unobserved components subject to regime shifts to evaluate this link. We verify that periods of high (low) inflation mean were accompanied by periods of high (low) both short- and long-run uncertainty in inflation. Interestingly, unlike developed countries, short run uncertainty is important. These relationaships are consistent with the presence of three clearly differentiated regimes. First, a period of price stability, then a high-inflation high-volatility regime, and finally a hyperinflation period. We also verify that during a recent period of price stability, both permanent and transitory shocks to inflation have decreased in volatility. Finally, we find evidence that inflation and money growth rates share similar regime shifts.

JEL Classification: C22, E31, E42, E52

Keywords: inflation dynamics, monetary policy, Markov-switching models, unobserved component models, sthocastic trends

*We thank Martin Ellison, Carlos Montoro, Gabriel Rodriguez and participants in the Economic Research Workshop at the Central Reserve Bank of Peru, the Workshop on Monetary and Financial Implications of Globalization: Issues for Central Bank organized by the Bank of Mexico and the ECB, and the 2nd Monetary Policy Research Workshop in Latin America and the Caribbean organized by the Central Reserve Bank of Peru and the Centre for Central Banking Studies of the Bank of England, for useful comments and suggestions. We also thank Gerardo Tirado for valuable research assistance. Any remaining errors are the authors' own responsibility. The views expressed herein are those of the authors and do not necessarily reflect those of the Central Reserve Bank of Peru.

†Corresponding author: alberto.humala@bcrp.gob.pe, Jr. Miró Quesada 441, Lima-Perú. (511)-613 2785.

1 Introduction

The literature and the empirical evidence suggest that the costs of high inflation rates are considerably larger when inflation uncertainty is high. On the one hand, higher inflation uncertainty induces larger stabilization costs because it makes more difficult to forecast inflation and renders inflation more persistent. On the other hand, as Milton Friedman (1977) pointed out in his Nobel Prize lecture in 1976, higher inflation uncertainty generates larger relative price distortions, increasing inefficiencies in production. High and volatile inflation experience in emerging markets should reveal interesting evidence about this relationship and provide policy lessons for monetary authorities in recently acquired low-inflation scenarios.

The goal of this paper is, thus, to assess empirically the link between inflation and inflation uncertainty for the Peruvian economy and to assess what role monetary policy plays in it. A set of relative long-span inflation data (quarterly observations for 1949.1 to 2006.1) is used to identify this relationship. Moreover, in order to account for the different monetary policy regimes that were in place during this sample period, inflation dynamics is modelled subject to regime switching. Indeed, monetary policy in Peru has evolved from money growth management to interest rate management, under different macroeconomic scenarios of price (in)stability. Peru suffered a hyperinflation experience in the late 1980s; implemented successfully a stabilization programme in the early 1990s; and adopted a fully-fledged inflation-targeting regime in 2002. Yet, there is no empirical assessment of the link between inflation and uncertinty against money growth dynamics (as an indicator of monetary policy stance). There is no consideration of regime shifts in monetary policy that might have shaped this link either.

Methodologically, this paper proceeds as follow. First, it explores whether or not there is a systematic link between inflation and (long- and short-run) inflation uncertainty by decomposing inflation dynamics between its stochastic trend and its stationary part, in line with Ball and Cecchetti (1990). Next, the study focuses on a Markov switching heteroskedasticity model of inflation, whereby inflation dynamics is decomposed into a stochastic trend and a mean-reverting (stationary) part both subject to regime switching in their disturbances, as in Kim and Nelson (1999).[1] Following a similar univariate approach, the paper studies also the regime-switching dynamics of money growth in order to assess feasible links to inflation dynamics.

Main empirical results in this study indicate that there indeed exists a link between inflation and inflation uncertainty (stronger for long-term uncertainty but also relevant for short-term volatility) in Peru and that, furthermore, this link has been subject to regime shifts. Three

[1] An atheoretical Markov switching autoregressive model (MS-AR) is first estimated to infer regime classification.

regimes are clearly identified in inflation dynamics. A low-inflation stable regime for inflation spans periods 1949 - 1975 and 1994 - 2006 and includes the recent inflation targeting experience. A high-inflation, high-volatility regime spans periods 1975 - 1987 (accelerating inflation) and 1991 - 1994 (disinflation). Lastly, an outlier-type hyperinflation regime prevails over the period 1988 to 1990. High-variance states of permanent shocks to trend inflation and of transitory shocks explain regime shifts towards higher inflation mean.

A by-product result from the auxiliary estimation of the Markov switching autoregressive model (to identify the regime shifts) is that periods of high long-run inflation uncertainty are also associated to higher persistence in inflation dynamics.[2] Notwithstanding the assumption of a random walk for the true generating data process for trend inflation, this atheoretical AR model is consistent with agents' perception of a persistent inflation dynamics as in Lansing (2006).

Supporting the association of the regime switching nature of inflation dynamics to monetary policy, three similar regime dating shifts are found for money growth. These results suggest that monetary policy shifts, from low- to high-inflation regimes and from high- to low-inflation regimes, explain the rise in inflation volatility and persistence from 1949-1975 to 1976-1994 and the opposite movement from 1994 on, respectively[3].

The rest of the paper is organized as follows: in Section 2, an unobserved component model of inflation is presented and estimated as a first approximation to the link between the level of inflation and inflation uncertainty. It is, then, extended to allow conditional and unconditional heteroskedasticity in shocks to permanent and transitory inflation components. Section 3 assesses whether or not those regime switches in inflation are linked to monetary policy shifts. Section 4 briefly discusses a theoretical framework that provides rationale to the relationship between inflation uncertainty and persistence. A last section concludes and outlines research agenda.

2 Inflation and Inflation Uncertainty

Average inflation and its volatility in the Peruvian economy have drastically changed in the last six decades or so. A simple look at the mean and volatility of the quarterly inflation rate over the sample 1949 - 2006, and over ten-year sub-samples, shows that the magnitude of those changes are far from being negligible (see Table 1 in Appendix A). Quarterly average inflation

[2] Interestingly, inflation persistence in the high-inflation high-volatility regime is twice the level of persistence during the low-inflation stable regime.

[3] The regime shifting nature of inflation dynamics is key to understand links to monetary policy changes and people's expectations. Sargent, Williams, and Zha (2006), for instance, attribute the shifting in inflation regimes to stochastic switches between rational expectations (normal inflation) and adaptive expectations (high- and hyper-inflation periods) associated, in turn, with fiscal deficit stances.

rate increased from around 0.7% during the 1960s to 2.4% during the 1970s, accompanied with an increase in inflation volatility from 0.97 to 2.4 across decades. During the 1980s, average inflation and volatility reached their highest levels, 16.4% and 38.2, respectively, whereas, during the 1990s found their lowest levels, 0.2 and 0.4, respectively.

A link between inflation mean and volatility emerges neatly from those basic statistics.[4] Likewise, those statistics suggest some structural breaks in inflation dynamics. In fact, monetary policy, the main long-run determinant of inflation, has evolved from money-aggregate targeting (with restricted independence before the 1990s) to an inflation-intolerant regime (after 1994), suggesting feasible regime switches on the permanent component of inflation[5]. Although no distinction of short and long run volatility is crystal from those indicators, these unobserved component could be estimated from inflation data. Hence, to properly account for the link between inflation and inflation uncertainty in the Peruvian economy, it is necessary to use a framework that simultaneously deals with regime switching and unobserved components.

In this paper we follow Kim and Nelson (1999), by using a model where both components, the stochastic trend and the stationary (autoregressive) part, are subject to regime switching. Yet, before moving to the regime switching estimation, we evaluate the link between inflation and inflation uncertainty using an unobserved component model in the line of Ball and Cecchetti (1990). They find a positive relationship between inflation and inflation uncertainty at long horizons by decomposing inflation into its stochastic trend and its stationary (autoregressive) part. However, as Gordon (1990) points out, their result is valid only in a situation in which the policy maker decides to disinflate the economy but not in any other regime and, thus, their empirical work is subject to the Lucas critique. In other words, empirical measures of inflation uncertainty at any horizon may be misleading if the econometric specification does not properly capture regime switching in monetary policy and in inflation dynamics. On this regard, Kim (1993) extends Ball and Cecchetti´s study assuming regime switching might be a key source of inflation uncertainty. Kim (1993) finds that high uncertainty about long-run inflation is associated with a positive shift in inflation levels and, as a result, monetary policy becomes unstable. Moreover, high uncertainty about short-term inflation is linked to a negative shift in inflation levels (and, therefore, a less-stable short-run monetary policy). This evidence shows that there are indeed costs of high-level inflation rates in terms of long-term uncertainty.[6]

[4] Even after adjusting for scale factors.

[5] The Central Reserve Bank of Peru started to announce annual inflation targets since 1994 and establish in 2002 a fully-fledged inflation-targeting regime.

[6] Evans and Wachtel (1993) develops a model of inflation from which they can derive measures of inflation uncertainty associated to different regimes.

2.1 A First Glance at Inflation Uncertainty

This section provides a *prima facie* evidence of the relationship between inflation and inflation uncertainty. It follows closely Ball and Cecchetti (1990) to assess this link.[7] Data for Peruvian inflation spans the period 1949 - 2006 and corresponds to the Consumer Price Index (CPI) inflation. The inflation time series has been seasonally adjusted at quarterly frequencies.

The inflation rate series is decomposed into its permanent and temporary (but persistent) parts and, thus, measures of short and long term inflation uncertainty are easily obtained. We define inflation uncertainty as the variance of the forecast error of inflation. The following unobserved component model for inflation is postulated:

$$\pi_t = \pi_t^T + \eta_t \tag{1}$$

$$\pi_t^T = \pi_{t-1}^T + \varepsilon_t \tag{2}$$

where ε_t and η_t denote shocks to the permanent and transitory unobserved components of inflation, respectively. π_t denotes the level of current inflation and π_t^T denote trend inflation, which follows a random walk. Trend is assumed a non-observable component of inflation. Equations (1) and (2) characterize inflation dynamics.

Since ε_t captures permanent (stochastic) shocks to trend inflation, it is the source of long-run uncertainty. On the other hand, η_t represents transitory deviations of inflation from its trend and, therefore, it is associated to short-run uncertainty. For simplicity, it is assumed that shocks are uncorrelated disturbances with mean zero and variances σ_ε^2 and σ_η^2, respectively. We use estimations of σ_ε^2 and σ_η^2 as measures of short and long run uncertainty.

A simple way to estimated σ_ε^2 and σ_η^2 is to use the indirect approach of Ball and Cecchetti (1990), which uses the fact that the model of inflation, given by equations (1) and (2), is observationally equivalent to an ARIMA model with a single shock, which can be characterized as follows[8],

$$\triangle\, \pi_t = v_t + \theta v_{t-1} \tag{3}$$

where, $v_t \sim iid(0, \sigma_v^2)$ and $0 > \theta > -1$. Using the estimated parameters of this equivalent representation, $\widehat{\sigma}_v^2$ and $\widehat{\theta}$, the short and long run measures of inflation uncertainty, σ_ε^2 and σ_η^2

[7] For a recent survey of inflation dynamics modelling, see Rudd and Whelan (2005).

[8] The MA coefficient, θ, lies between 0 and -1 capturing the fact that temporary shocks eventually die out. See Ball and Cecchetti (1990) for details.

, can be obtained from the following identities,

$$\begin{aligned} \widehat{\sigma}_{\varepsilon}^{2} &= \left(1+\widehat{\theta}\right)^{2} \widehat{\sigma}_{v}^{2} \\ \widehat{\sigma}_{\eta}^{2} &= -\widehat{\theta}\widehat{\sigma}_{v}^{2} \end{aligned} \tag{4}$$

Equation (3) is estimated for five-year sub-samples. In order to mitigate the effect of the hyperinflation episode, data observations from the sub-sample 1985.01 - 1994.04 are disregarded. For a given period, equation (3) is estimated and its parameter $\widehat{\theta}$ and variances $\widehat{\sigma}_{v}^{2}$ are recovered and saved. Estimation results are reported in Table 2 in Appendix A. Thereafter, equation (4) is used to construct estimations of both $\left(\sigma_{\varepsilon}^{2}\right)$ and $\left(\sigma_{\eta)}^{2}\right.$. Then, these estimations are used to assess the link inflation and inflation uncertainty.

For that purpose, we calculate the correlation between our measures of short and long run inflation uncertainty and the average inflation rate , $\overline{\pi}_{t-1}^{T}$, calculated using the same span period employed to estimated $\left(\widehat{\sigma}_{\eta}^{2}, \widehat{\sigma}_{\varepsilon}^{2}\right)$. A simple way to calculated these correlations is estimated the following equations:

$$\sigma_{\varepsilon}^{2}(t) = \beta_0 + \beta_1 \overline{\pi}_{t-1}^{T} \tag{5}$$

$$\sigma_{\eta}^{2}(t) = \delta_0 + \delta_1 \overline{\pi}_{t-1}^{T} \tag{6}$$

If β_1 is large and δ_1 is small, then trend inflation $\left(\pi_{t-1}^{T}\right)$ has a larger effect over uncertainty at long horizons (a result that was pointed out by Ball and Cecchetti). Figure 2 plots average inflation for the nine five-year periods *vis-à-vis* implied measures of long-run $\left(\sigma_{\varepsilon}^{2}\right)$ and short run $\left(\sigma_{\eta}^{2}\right)$ inflation uncertainty. Both types of shocks seem to co-move positively with the average level of inflation, although it can not be concluded, by plotting inspection, which shocks link stronger to inflation.

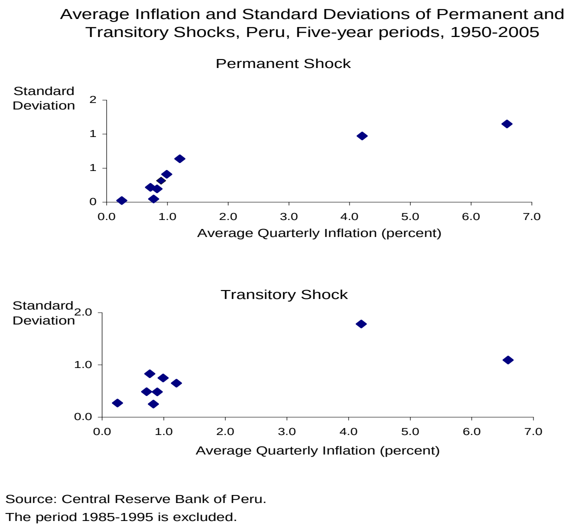

Figure 2

Table 3 in Appendix A reports results from estimating equations (5) and (6). Estimation confirms previous graphic inspection. Effects from average inflation on standard deviation of the permanent shock, $\left(\sigma_{\varepsilon}^{2}\right)$, is positive $(\beta_1 = 0.173)$ and significant $\left(R^2 = 0.84\right)$. There is also evidence supporting a possible link between the level of inflation and short-run uncertainty, although not as strong as for the long-run horizon. The parameter, in this latter case, is positive $(\delta_1 = 0.163)$ and significant, though the R^2 is considerably smaller than the one obtained in the permanent shock estimation.

The above preliminary evidence suggests the inflation process in Peru has been affected by both permanent and transitory shocks. A crucial assumption behind these estimations is that the economy is not affected by regime shifts. However, these switches might arise from changes towards inflation-fighting monetary policies or from changes in the way private agents learn the state of the economy. An inflation-intolerant regime should reduce uncertainty about economy fluctuations. By controlling any form of regime switches, a model of inflation will allow to verify whether or not short-run uncertainty is important, conditional on the regime that is in place in the economy.

2.2 Inflation and Inflation Uncertainty: A Markov Switching Model

In this section, we formally test the link between inflation and inflation uncertainty with shifts in regime using a Markov switching model that allows for conditional and unconditional heteroskedasticity. Although, before moving to this model, we first test for the existence of regime switching by simply using a Markov switching autoregressive (MS-AR) model, which does not allow for unobserved components[9].

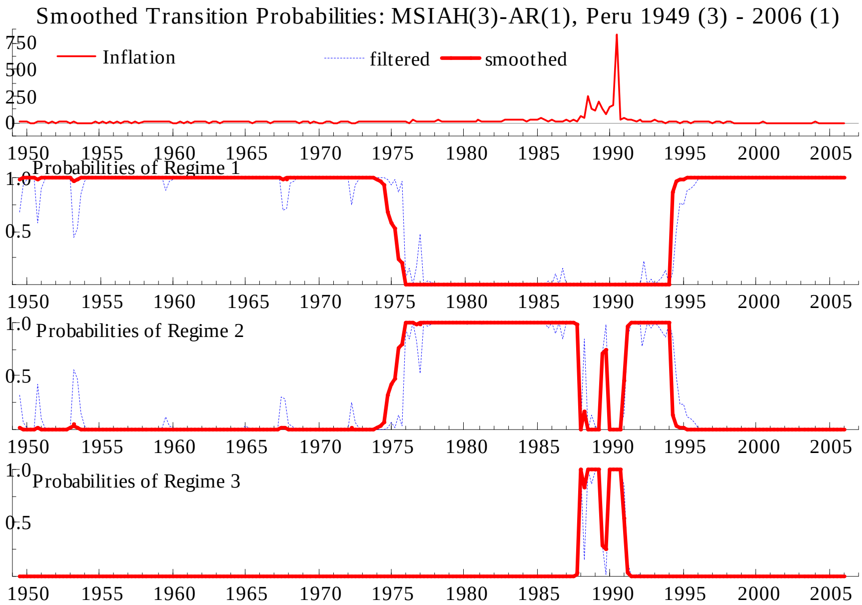

Figure 3

Estimation results show the presence of three clearly differentiated regimes over the entire sample (see Table 4 in Appendix A for parameter estimates).[10] Visual inspection of smoothed probabilities highlights regime's sequence (Figure 3). The first regime corresponds to a low-level inflation rate (an intercept of 1.3), low volatility (standard deviation of 1.7), and relative low persistence (0.29).[11] The periods 1949:3 - 1975:2 and 1994:2 - 2006:1 are classified into this first regime. The second regime refers to high-level inflation and highly volatile scenarios (intercept, 5.8; autoregressive parameter, 0.6; and standard deviation, 7.9).[12] Periods considered into this second inflation regime are 1975:3 - 1987:4, and 1991:2 - 1994:1. The third regime appears for an extremely volatile and outlier-type hyperinflation period in Peru (1988:1 - 1991:1).[13]

[9] See Appendix B for a detailed discussion of the model estimated and results.

[10] Rodríguez (2004) presents a time series analysis of the inflation rate series for various South American economies. His findings of nonstationarity of Peruvian inflation for a shorter sample are not in conflict with the presence of regime switching in this time series.

[11] All coefficients are significant at usual levels.

[12] All parameters are, again, statistically significant at usual levels.

[13] In the work of Sargent, Williams, and Zha (2006), the hyperinflation experience in Peru is modelled as the "extraordinary dynamics" inflation follows after a certain threshold level is reached and destabilizing expectations divorce inflation from its permanent path (even if fiscal deficit is zero).

2.2.1 Model description

We borrow from Kim and Nelson (1993). In their model, both components, the stochastic trend and the stationary (autoregressive) part, are subject to regime switching. A key feature of the model is, then, that it allows for conditional and unconditional heteroskedasticity. A much more illustrative association of the changing dynamics of inflation to regime switching in monetary policy is revealed by this approach. The equations for this model are:

$$\pi_t = \pi_t^T + \mu_2 S_{1,t} + \mu_3 S_{2,t} + \mu_4 S_{1,t} S_{2,t} + (h_0 + h_1 S_{2,t})\eta_t \tag{7}$$

$$\pi_t^T = \pi_{t-1}^T + (Q_0 + Q_1 S_{1,t})\varepsilon_t \tag{8}$$

where $\eta_t \sim N(0,1)$ is the shock to the transitory autoregressive component and $\varepsilon_t \sim N(0,1)$ is the shock to the stochastic trend component of the inflation series, both as in Ball and Cecchetti (1990). The stochastic component is subject to regime switching and $S_{1,t}$ is the unobserved state variable that represents this regime shifting. Similarly, the transitory component is also subject to switches in regime and $S_{2,t}$ captures the states for it. Both $S_{1,t}$ and $S_{2,t}$ are assumed to evolve according to two independent (of each other) first-order two-state Markov chains. Each state variable defines a low-variance state for the shocks, for which it takes on the value 0, and a high-variance regime for which it takes on the value 1. These discrete Markov processes are represented by the transition probabilities:

$$\Pr[S_{1,t} = 0/S_{1,t-1} = 0] = p_{00}, \qquad \Pr[S_{1,t} = 1/S_{1,t-1} = 1] = p_{11}, \tag{9}$$

$$\Pr[S_{2,t} = 0/S_{2,t-1} = 0] = q_{00}, \qquad \Pr[S_{2,t} = 1/S_{2,t-1} = 1] = q_{11}, \tag{10}$$

Shocks to the permanent (transitory) component take on the value Q_0 (h_0) if they are in a low-volatility fashion and Q_0 (h_0) + Q_1 (h_1) otherwise. This model of inflation involves, thus, the existence of up to four different economic states resembling possible combinations of regime occurrence at time t.[14] Regime 1 corresponds to a low-variance state for both Markov chains ($S_{1,t} = 0$ and $S_{2,t} = 0$), with Q_0 and h_0; regime 2 stands for a low Q_0 and a high h_1 ($S_{1,t} = 0$ and $S_{2,t} = 1$); regime 3 is for a high Q_1 and a low h_0 ($S_{1,t} = 1$ and $S_{2,t} = 0$); and, finally, regime 4 represents a high Q_1 and a high h_1 ($S_{1,t} = 1$ and $S_{2,t} = 1$). High-variance states of the shocks to the stochastic and transitory components of inflation affect inflation mean through the parameters μ_2 (if permanent shocks are highly volatile), μ_3 (if transitory shocks are highly volatile), and μ_4 (if both shocks are in a high-variance state, i.e. regime 4).

[14] Actually, up to 16 possible combinations of outcomes from the two Markov chains representing permanent and transitory shocks.

2.2.2 Inflation regimes and estimation results

Considering the three clearly differentiated regimes in inflation dynamics (inferred from the MS-AR), parameter estimation for the Markov switching heteroskedasticity model should include data from the entire sample. However, sample observations during the hyperinflation period are, in general, far above levels (and variability) of inflation reported in the other two regimes. Thus, maximum likelihood estimation that involves three regimes in two Markov chains is not easy to implement. The presence of additive outliers during the hyperinflation regime further complicates the estimation effort.[15] Therefore, considering the regime shifts indicated by the MS-AR, the model is rather estimated by sub-samples that exclude the hyperinflation regime (1988 - 1990).[16] Thus, initial estimation of the model is for the sample 1949 - 1987, which includes observations from the first regime of price stability (during the 1950s and 1960s) and those from the high-inflation regime that started with the oil crisis in the mid-1970s. Two alternating regimes in each component (trend and transitory) are then defined for low-variance and high-variance of shocks. Notice out most of the volatile regime corresponds to a period of increasing inflation rate.

Parameter estimates (and standard deviations) are shown in Table 7 in Appendix A. Transition probabilities of remaining in low-variance regimes for both the permanent and transitory inflation components (p_{00} and q_{00}) are higher than those of remaining in the high-variance states (p_{11} and q_{11}). The shift on the variance of permanent shocks is quite remarkable (as the ratio Q_1/Q_0 indicates), not only because the variance of shocks is indeed high during the volatile regime but also because volatility of shocks in the calm regime are rather negligible.[17] The effects of high-variance states of shocks over inflation mean are both positive (parameters μ_2 and μ_3) and are further emphasized by their simultaneous occurrence (parameter μ_4).

An important outcome from this model estimation is the inference of regime probabilities at each sample observation. In particular, plots of the inflation rate and the probability of high variance regimes for permanent and transitory shocks are illustrative of the switching nature of shocks. The shift to a highly volatile environment for permanent shocks is clearly spotted by mid-1970s and reinforced continuously during the mid-1980s (see the first panel of Figure 4). Volatility of transitory shocks in the first regime of price stability, during the 1950s and 1960s, are sporadic and clearly associated to inflation peaks. However, they become frequent and

[15] Kim and Nelson (1999), for example, modify Hamilton's (1989) algorithm to estimate an univariate Markov switching model of output (where only the mean is time-varying) to include the possibility of a third regime by incorporating dummy variables. The task in hand here, not only involves having dummy variables into every single parameter, but also extending this treatment to two Markov chains.

[16] Kim and Nelson (1999), in their study of U.S. inflation (for 1950 - 1990), avoid estimating a three-state variance structure by excluding some initial sample observations.

[17] However, parameter Q_0 is not significant statistically.

more persistent during times of higher mean and variance of inflation (second panel of Figure 4). These results are consistent with the view that shifts in inflation trend are associated to shifts in trend money growth, which also started by mid-1970s (see next subsection), and regime switching in transitory shock is more associated to demand and supply shocks (which become more frequent in an uncertain environment). Once inflation rates reach escalating levels, both permanent and transitory shocks seem to feed each other back (as the parameter μ_4 indicates).

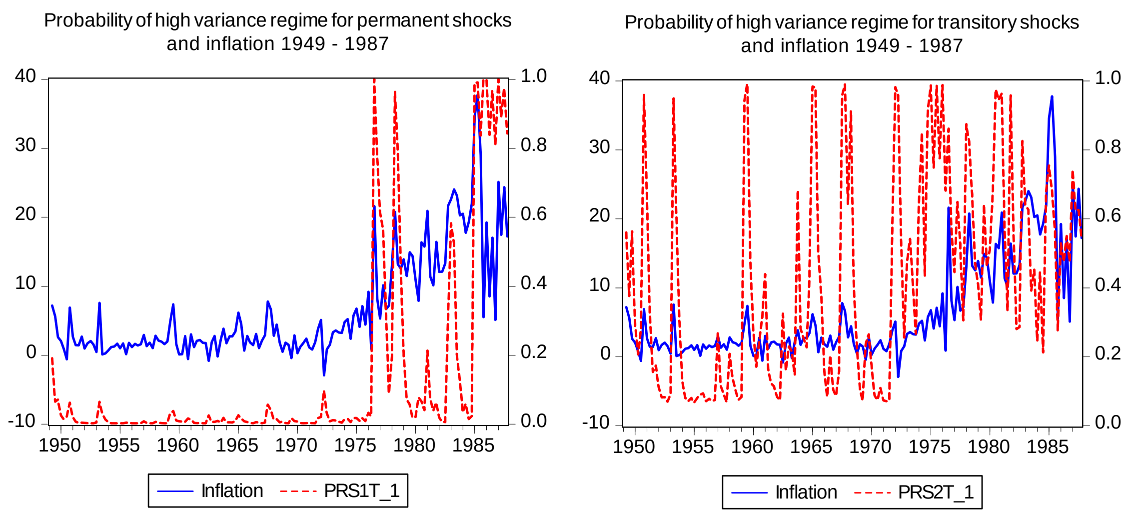

Figure 4

In order to fight hyperinflation, by the early 1990s, monetary and fiscal authorities adopted a stabilization programme to cut off money-based fuel to inflation and bring down inflation expectations. Therefore, the second sample for model estimation includes the period 1991 - 2006, so that again two regimes of low-inflation and high-inflation are included. This time, the high-inflation regime mainly corresponds to high but decreasing inflation rates (approximately during 1991 to 1993). Parameter estimates are shown in Table 7 in Appendix A (columns 3 and 4). Switches to the high-variance regimes in both trend and transitory components have greater effects on shocks' volatilities this time. In this case, although the increase in volatility of permanent shocks is larger (Q_1 is much higher), the ratio Q_1/Q_0 is lower than in the first sample estimation because the parameter Q_0 is not negligible. Still, this ratio shows the large increase in volatility once a shift in regime occurs. Graphs of the probability of high variance regime in permanent and transitory shocks show also important results (see Figure

5). First, the switch in trend occurs at the beginning of 1994 and a low-variance regime of permanent shocks follows thereafter. At this shift date, the Peruvian central bank started to pre-announce inflation objectives though still not committed to a fully-fledged inflation targeting scheme. Transitory shocks remain at a high-variance state for a while longer, but finally dies out at around 1999 and remains at a low-variance state after that. Contrary to the previous price-stability period of the 1950s and 1960s, the low-level and low-variance inflation regime in recent periods involves not only a stable trend but also a very stable sequence of transitory shocks. Importantly, this non-existence of shifts to high-volatility regimes, both in the permanent and transitory components of inflation, is not due exclusively to the adoption of the inflation targeting scheme of monetary policy (from 2002 onwards) but to the downward-expectations orientation of the monetary policy (from 1994 onwards) after successfully fighting hyperinflation. The merit of the inflation-targeting regime is to reinforce this orientation by smoothing transitory shocks around the inflation target.

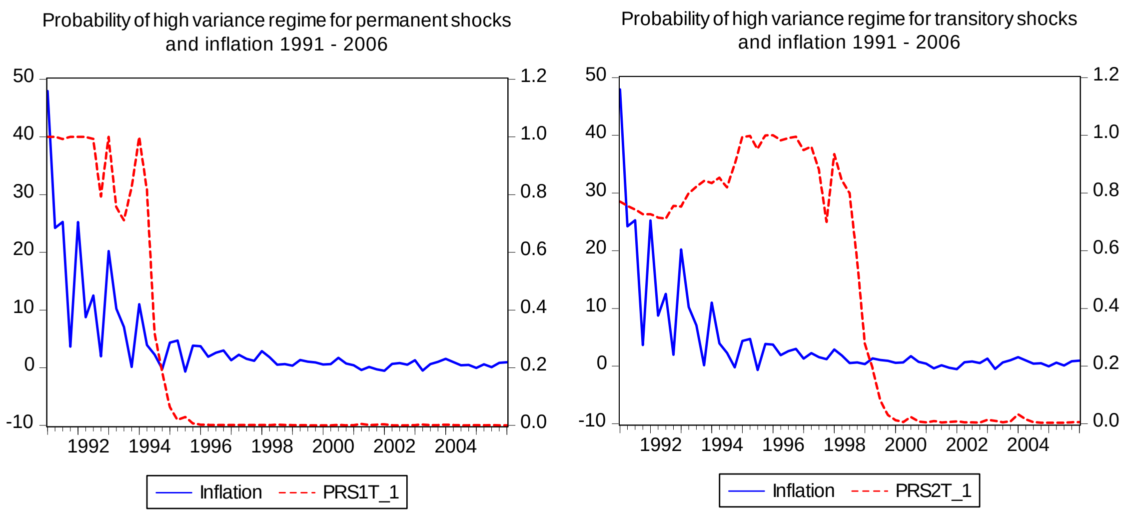

Figure 5

Summing up the results so far, the estimated Markov switching heteroskedasticity model of inflation is consistent with splitting up inflation dynamics between two regimes. More importantly, the heteroskedasticity models moves forward to infer that regimes switches occur in both (permanent and transitory) unobserved inflation components. Thus, for the permanent (transitory) shocks, a high-variance scenario is identified to alternate with a low-variance regime over a sample that spans almost six decades (though it excludes the hyperinflation regime).

Inflation dynamics is subject to permanent and transitory shocks, which in turn are subject to switching between calm and volatile regimes. Figure 6 depicts the unobserved inflation components for the two sub-samples already presented here and Figure 7 shows the association of these components to the probability of being in a high-variance stance of the corresponding shock.[18]

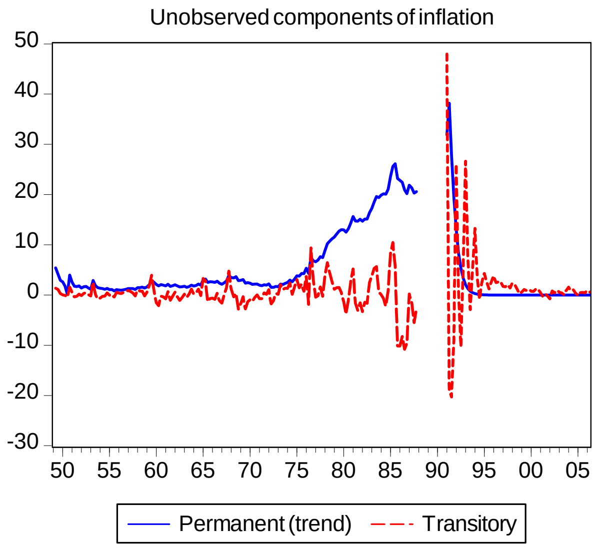

Figure 6

[18] Alternatively, the model has been estimated including both previous sub-samples but merged into a new time series of inflation rates (excluding hyperinflation observations). Though this is not a formal solution to the treatment of the third regime, estimated parameters confirm previous conclusions about the shifts in trend and transitory shocks. Results are available from the authors upon request.

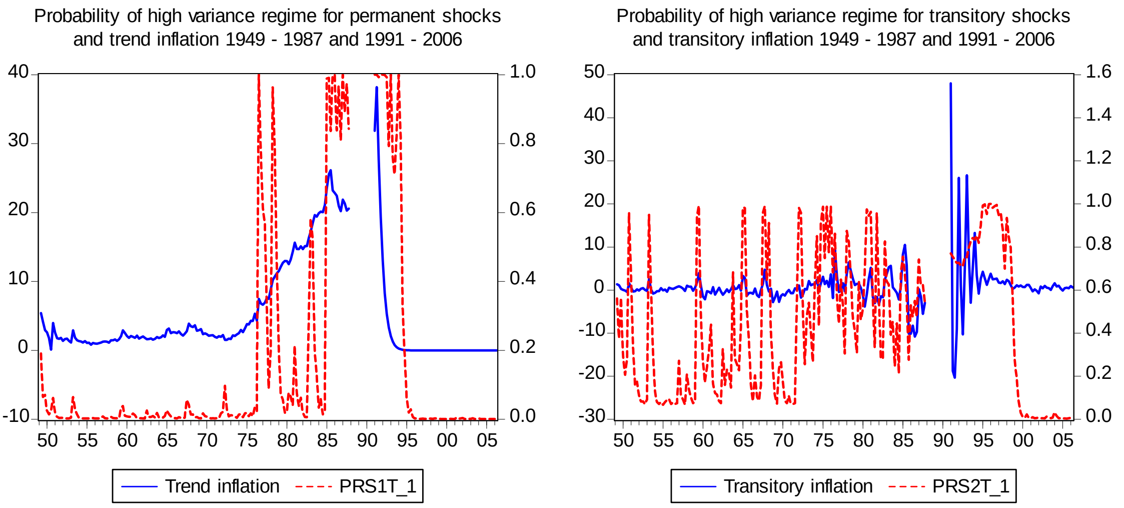

Figure 7

Despite capturing regime switches in inflation dynamics in these two sub-samples, so far estimation has dropped out observations from the hyperinflation period that are clearly differentiated as part of a third regime by the MS-AR approach. Therefore, in order to enhance understanding of inflation dynamics in switches from high-level and high-variance inflation to the hyperinflation regime, the sub-sample 1973 - 1993 is used for parameter estimation of the model.[19] Parameter estimates are shown in Table 7 in Appendix A (last two columns). Very important results and conclusions emerge from this sample. First, the increase in volatility of permanent shocks is very large in magnitude (larger than in the case of shifting between a low-inflation regime to a high-inflation regime). The coefficient Q_1 scales up to 43.1 from a Q_0 of 1.3. Furthermore, effects on inflation mean from the high-variance state in permanent and transitory shocks are considerably much larger too. Figure 8 shows plots of the inferred probabilities of high variance regimes for permanent and transitory shocks against the inflation rate series. The first panel strongly represents the hyperinflation regime as a shift in permanent shocks. Meanwhile, large volatility of transitory shocks span over three to four years before and after the hyperinflation period but decrease, in probability, somehow during the hyperinflation itself.

In a hyperinflation scenario, volatility of both type of shocks have strong effects on inflation

[19] Observations at 1988.3 and 1990.3 (September in each year) are treated as outliers even during the hyperinflation regime. They correspond to policy-adopted large price shocks (in attempts to drastically cut down price increases).

level and uncertainty. Actually, mean rising is larger as a response to transitory shocks (a parameter μ_3 of 23.3) than as a response to permanent shocks (a parameter μ_2 of 4.5).[20] In such a regime, inflation dynamics can only be switched out of its spiral by an explicit and drastic shift in trend money growth, as it actually happened by the early 1990s. As shown above, inflation persistence increases with level and uncertainty of inflation and, therefore, it becomes harder to abandon accelerating inflation scenarios unless the monetary authority commits itself to non-indulgent policy of inflation fighting.

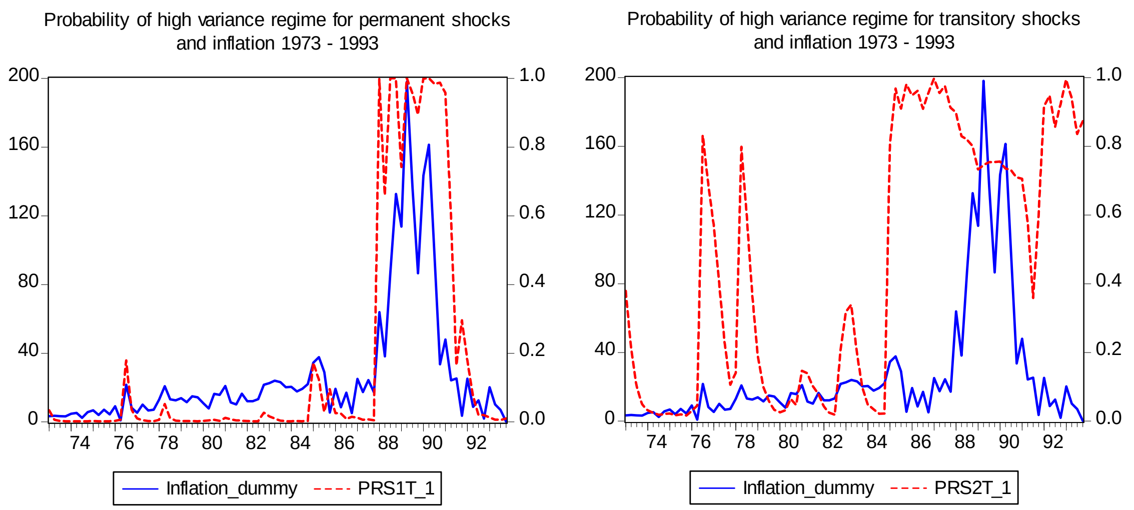

Figure 8

Conclusions for the conduct of monetary policy from the above discussion are of the most importance. A commitment to keep shocks to trend inflation to a minimum are certainly fruitful to bring inflation down, but a further commitment to anchor down inflation expectations reinforces low-variance scenarios in both permanent and transitory inflation. Insofar as an inflation targeting monetary scheme succeeds in keeping inflation under control, price stability scenarios feedbacks from its own dynamics. Of course, regime switches in the transitory component of inflation might occur for reasons other than local supply or demand management, but credibility in the central bank's commitment should help to keep inflation anchored at the chosen target. Importantly, once authorities (for whatever reason) start losing control, chances of rapidly changing into a high-variance regime increase.

[20] Recall that quarterly observations of percentage change in the CPI is being used for model estimation.

3 Inflation, Inflation Uncertainty and Monetary Policy

The institutional framework of monetary policy in Peru has radically changed during the sample period. Before the 1990s, the Central Bank of Peru was not entirely autonomous since the evolution of fiscal deficit partially conditioned monetary policy and, specially, money growth rates.[21] In contrast, during the 1990s, formal autonomy was granted to the central bank (by a new Peruvian Constitution and Central Bank Charter) and price stability was adopted as the unique objective of central bank's monetary policy. More recently, in 1994, the central bank took the first steps towards adopting an inflation targeting framework by pre-announcing inflation targets. In 2002, the bank decided to adopt a fully-fledged inflation-targeting regime.[22]

The previous section provided empirical evidence of regime shifting in both permanent and transitory components of inflation in Peru. Although no link to monetary variables or policy was pursued empirically, switching regimes in inflation trend are known to be associated to money market considerations. Therefore, this section addresses formally this issue by studying money growth dynamics subject to shifts in regime (given the many changes in monetary policy experienced during the sample).

To start with, a MS-AR model is estimated to test for the presence of regime switches in money growth (measured as M2, total liquidity in domestic currency).[23] Smoothed probabilities inferred at each observation are presented in Figure 9.[24] As in the case of inflation, three regimes are identified for the money growth rate. Interestingly, dates of regime shifts in money growth coincide, or are very similar, with those of inflation. The first period of low-level, low-volatility money growth goes up to 1978 (instead of 1975, as in the case of inflation) and prevails again from 1995:1 (instead of 1994:2) onwards. A high-mean and volatile regime comes for periods 1978:4 - 1988:2 and 1991:2 - 1994:4. Lastly, an explosive-rate regime of money growth goes for the period 1988:3 - 1991:1 (that coincides mostly with the hyperinflation episode). Thus, both inflation and money growth rates share fundamentally the same regime shifts over the sample 1949 to 2006. Indeed, cointegration analysis confirms that inflation and money growth rates share the same stochastic trend in Peru.[25]

With the regime classification obtained from the MS-AR estimation, it is then estimated a

[21] For a historical perspective of monetary policy in Peru, see Guevara (1999).

[22] For a detailed account of the monetary policy framework in Peru from 1991to 2001, see Quispe(2000) and De la Rocha (1999). For the inflation-targeting regime, see Armas and Grippa (2006) and Rossini (2000).

[23] Because of data availability, sample estimation is defined for1964 (not 1949, as in the case of inflation) onwards.

[24] Estimation results for money growth are available from the authors upon request.

[25] Using the cointegration technique from Johansen and Juselius (1990), and under different specifications, the test suggests the presence of one cointegrating vector between inflation and money growth rate.

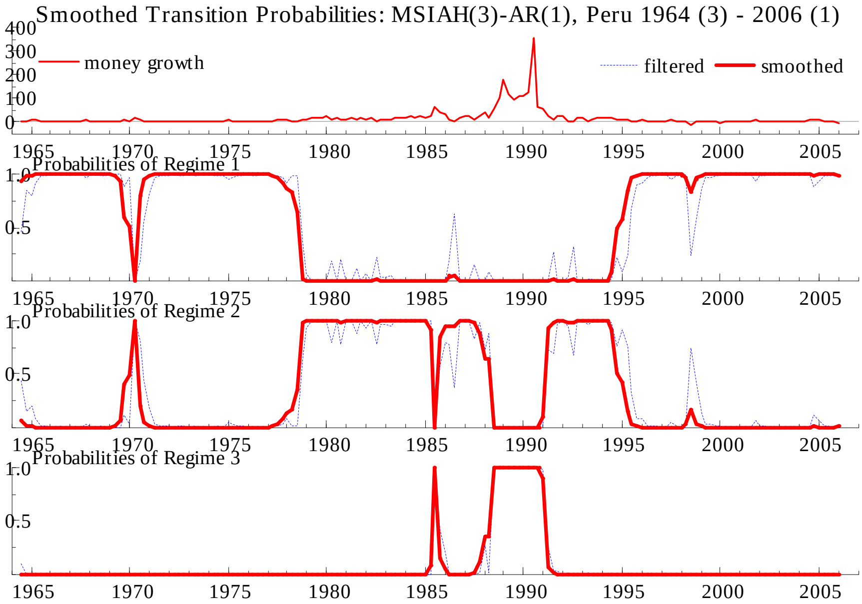

Figure 1: Figure 9

Markov switching heteroskedasticity model of money growth in a similar univariate fashion as in the case for inflation above.[26] Figure 10 shows the decomposition of the money growth rate into its permanent and transitory parts over the entire sample (1964 - 2006). Evolution from trend money growth shows long-term instability over most of the 1980‘s and the disinflation effort during the first half of the 1990‘s.

[26] Equations 7 and 8 are, thus, estimated for the money growth rate, with all the corresponding details equal to the model specification above.

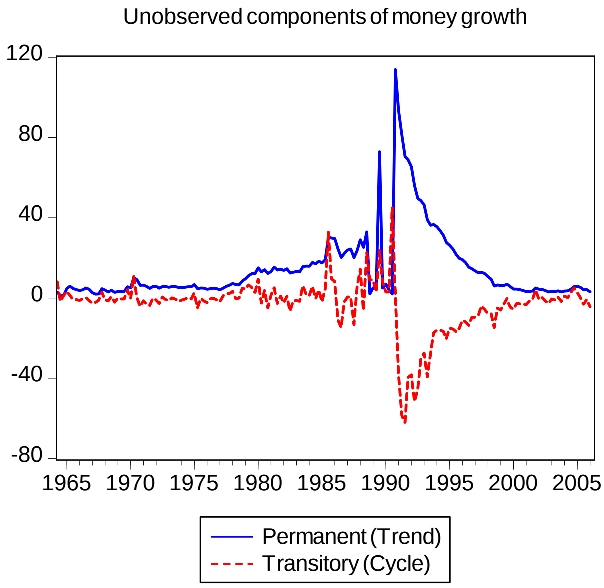

Figure 10

Sub-samples are defined for 1964 - 1988, 1991 - 2006, and 1978 - 1994, so that up to two regimes in each period are considered. The following figure depicts regime probabilities of the high-variance state for permanent and transitory shocks to the money growth rates for the first two sub-samples, thus excluding the hyperinflation episode. High-variance regime for long-term (permanent) shocks are most certain from 1985 to 1994 (with an interior gap for the hyperinflation regime, when these shocks actually shift to another regime). Probability of a high-variance state for the transitory shocks, in turn, shows a shift to this state earlier on, from around 1978, but also ends before (in 1988, when it switches again to the third explosive-rate of money growth). Notice out that these patterns of shock to the money growth rate are similar to those of inflation and signal the close relationship between changes in the conduct of monetary policy and inflation dynamics. Furthermore, an interesting empirical result is that high uncertainty in the short-term for money growth seems to have been more important in the building up of high-inflation periods (from 1978 to 1985). Nevertheless, high variability on permanent shocks is vital to keep inflation at those high levels (and uncertainty) and to raising it up to the hyperinflation stage.

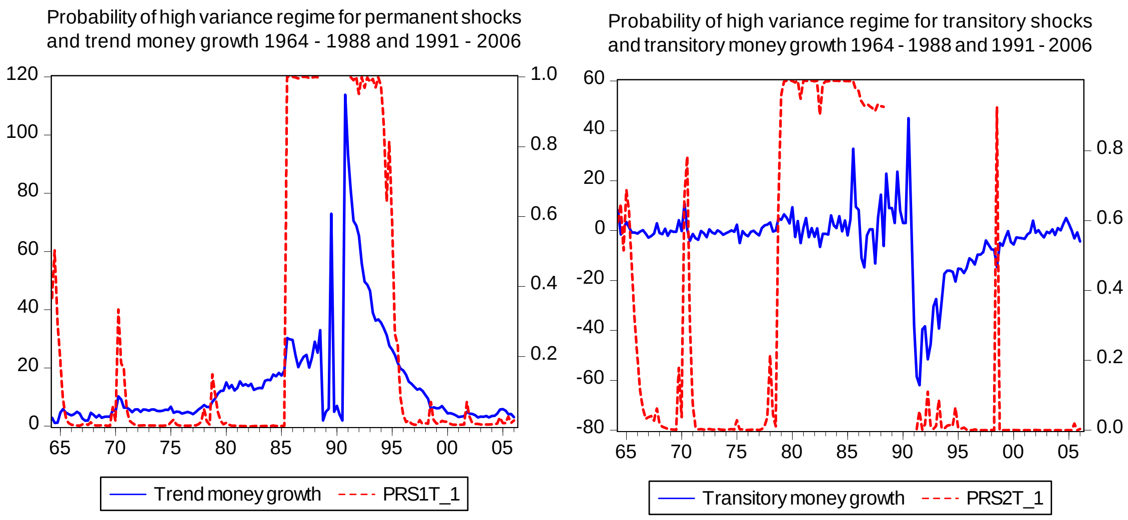

Figure 11

Considering the third sub-sample, 1978 to 1994, the shift from the high-mean and high-uncertainty regime to the explosive rate of money is shown as close-to-one probability of permanent shocks being in the high-variance state. It actually suggests the higher long-term uncertainty a few years before (1985) and after (1992) than the actual hyperinflation period (1988 to 1990).

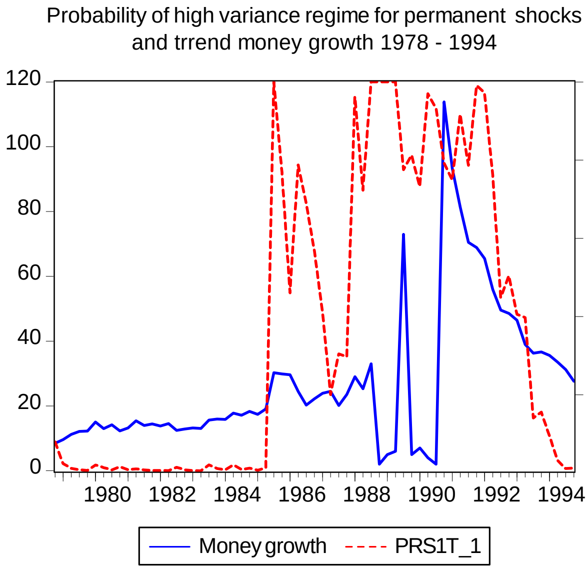

Figure 12

From a multivariate or multiequation perspective of inflation dynamics, Marcet and Nicolini (2005) introduce money growth subject to Markov switching as the exogenous driving force of permanent inflation dynamics (in a money demand function). Salomon (2001) studies the link to fiscal and monetary policies by modelling regime switching inflation with time-varying transition probabilities (depending upon policy stance). Furthermore, economic structure could be added on into the regime shifting feature in inflation dynamics by considering state variables subject to switches in regime for the fiscal deficit (rather than for inflation itself or money growth), as in Sargent, Williams, and Zha (2006).[27]

How much of the recent low-inflation and low-volatility trend is due to the inflation-targeting regime of monetary policy (or any other inflation-intolerant monetary policy) or to worldwide dragging-down effects on inflation is, however, an issue not directly studied here. Vega and Wilkeried (2005), in a international assessment of inflation targeting success, find that the adoption of inflation targeting delivers low mean inflation and low inflation volatility.[28] Complementary, Borio and Filardo (2006) argue that proxies for global economic slack add considerable explanatory power to inflation models, with inflation rates becoming less

[27] Mizuno, Takayasu, and Takayasu (2006) take a totally different (rather atheoretical) approach (borrowed from econophysics) to represent hyperinflations as double-exponential functions of time. Switches in regime, from non-hyperinflation to hyperinflation times, are set to reflect people's psychology.

[28] See Bratsiotis et al. (2002) for a study of the effects from the adoption of inflation targeting in inflation persistence.

sensitive to the domestic output gap.

4 Inflation Uncertainty and Persistence

In the MS-AR estimation three well-defined inflation regimes in Peru, over the sample 1949 - 2006, were identified. Model estimation reports a notorious change in the autoregressive parameter (ρ) between the price-stability regime (0.29) and the high-inflation regime (0.6).[29] This parameter can be interpreted as a reduced-form coefficient of inflation persistence. It is worth to mention that in a general equilibrium structure ρ will depend upon deep parameters and in the way in which agents form expectations. Remarkably, the significant change in the autoregressive coefficient coincides with regime shifts in both inflation and money growth rates and, therefore, this parameter might be associated to different episodes of long-run and short-run uncertainty.

This paper has not attempted an evaluation of the price-setting behaviour of economic agents at the micro level, but an assessment of how much inertia there is on inflation dynamics.[30] Thus far, evidence from inflation dynamics modelling suggests that inflation persistence diminishes with level and volatility of inflation. These results are consistent with the association of low persistence to predominant forward-looking inflation dynamics (as in the low inflation regime) and high persistence to predominant backward-looking inflation dynamics (as in the high- or hyper- inflation regimes). In terms of the Sargent, Williams, and Zha (2006)'s analysis of inflation dynamics, rational expectations support lesser inflation persistence than whenever there is a degree of learning (adaptive expectations) in the forming of inflation expectations. As Marcet and Nicolini (2005) state out, monetary supply shocks are incorporated more slowly into inflation expectations under learning than under rational expectations.

This section's goal is to provide a simple and preliminary explanation of how varying degrees of uncertainty (for long-run and short-run uncertainty) across regimes might explain changes in inflation persistence across those regimes.

A similar approach to Lansing (2006) is followed. He finds evidence that higher degrees of inflation uncertainty induce more inflation persistence. Similarly, based on the unobserved component model of inflation from Section 2, agents are assumed to perceive inflation evolution according to equations (1) and (2). Conveniently, a "signal-to-noise" ratio will be obtained from parameter estimates of the model.

It is further assumed that the Kalman filter implements an agent's optimal forecasting rule

[29] See Appendix B for details.

[30] Also called intrinsic persistence. See Angeloni et al. (2005) for a recent appraisal of new evidence on inflation persistence (in the Euro area) and a distinction of the main sources of it.

and that the error correction dynamics is given by the equation:

$$\begin{aligned} \widehat{E}_t \pi_{t+1} &= \widehat{E}_{t-1}\pi_t + K\left(\pi_t - \widehat{E}_{t-1}\pi_t\right),\ 0 < K < 1 \\ &= K\left[\pi_t + (1-K)\,\pi_{t-1} + (1-K)^2\,\pi_{t-2} + ...\right] \end{aligned} \tag{11}$$

The above equation implies the agent's forecast at time t is determined by an exponentially weighted moving average of past inflation rates. Hence, inflation dynamics could be represented as a function of both permanent and transitory shocks:

$$\pi_t - \pi_{t-1} = \varepsilon_t + \left(\eta_t - \eta_{t-1}\right) \tag{12}$$

Obtaining the unconditional moments:

$$Cov(\Delta\pi_t \Delta\pi_{t-1}) = E\left(\Delta\pi_t, \Delta\pi_{t-1}\right) = \left[\varepsilon_t + \left(\eta_t - \eta_{t-1}\right)\right]\left[\varepsilon_{t-1} + \left(\eta_{t-1} - \eta_{t-2}\right)\right] = -\sigma_\eta^2 \tag{13}$$

$$Var\left(\Delta\pi_t\right) = E\left(\Delta\pi_t^2\right) = E\left(\left[\varepsilon_t + \left(\eta_t - \eta_{t-1}\right)\right]\left[\varepsilon_t + \left(\eta_t - \eta_{t-1}\right)\right]\right) = \sigma_\varepsilon^2 + 2\sigma_\eta^2 \tag{14}$$

Equations (13) and (14) are used to obtain the signal-to-noise ratio, S. This ratio is defined as the relation between the variance of permanent shocks and the variance of transitory shocks to inflation $\left(\frac{\sigma_\varepsilon^2}{\sigma_\eta^2}\right)$:

$$\frac{\sigma_\varepsilon^2}{\sigma_\eta^2} = S = \frac{-1}{corr(\Delta\pi_t, \Delta\pi_{t-1})} - 2 \tag{15}$$

The solution for the optimal gain parameter in steady state is obtained from the error correction expression (11):[31]

$$K = \frac{-S + \sqrt{S^2 + 4S}}{2} \tag{16}$$

Then, the signal-to-noise ratio, S, and the implied optimal Kalman gain, K, are calculated for regime 1 (price stability) and regime 2 (high inflation) from the MS-AR estimation. Results are reported in Table 8 in Appendix A.

From equation (16), it is clear that there exists a positive link between the signal-to-noise ratio S and the Kalman gain K. Recall that, by definition, the signal-to-noise ratio measures long-run versus short-run uncertainty. In fact, a higher value of K implies that the representative agent is assigning more weight to recent inflation data since she perceives long-run uncertainty increases relative to short-run uncertainty (higher signal-to-noise ratio). Therefore, since agents put more weight to recent inflation, it induces larger persistence.[32]

[31] Equation (16) is in turn obtained as the solution to the signal extraction problem, where the objective is to minimize the mean squared forecast error. See Harvey (1996) for details.

[32] Lansing (2006) links the signal-to-noise ratio and the gain parameter to the structural parameters of inflation

Hence, a higher K or a larger S could be interpreted as if the central bank has become less credible in anchoring future expectations consistent with its target.

Calculations show that the signal-to-noise ratio is smaller in regime 1, (0.262) than in regime 2, (0.584), and consequently the parameter K is smaller in the first regime. Following previous intuition, in regime 1, agents assign less weight to past-observed values of inflation and hence we observe a lower degree of inflation persistence. The contrary occurs in regime 2. This simple evidence highlights the role of uncertainty at characterizing some features of inflation dynamics, in particular, inflation persistence.

Finally, the inverse of the parameter S can be interpreted as a measure of central bank credibility. Thus, some insights about people's expectations could be inferred by regime classification and S estimation. Both pre-IT low-inflation periods and the IT regime are considered into regime 1, for which a small signal-to-noise ratio (high inverse of S) is capturing credibility gains in the central bank's policy. It actually shows the inflation-intolerant position of the bank. Finally, a smaller degree of persistence is also associated with a more forward-looking behaviour within the economy, so that less-costly stabilization policies should be a feature of the recent price-stability regime.

5 Conclusions

This paper investigates the link between inflation, inflation uncertainty, and inflation persistence in the Peruvian economy, in a context in which monetary policy has been subject to regime switches. First, inflation time series is decomposed into its permanent and transitory components in order to establish the link between inflation and inflation uncertainty (both at long- and short-run). Second, regime switching behaviour in the variance of shocks to the permanent and transitory components of inflation is considered (into a Markov switching heteroskedasticity model of inflation) to disentangle this relationship. Lastly, influence from monetary policy changes is also assessed to associate them to switches in inflation dynamics.

Many novel results stand out from empirical estimations of these univariate models of inflation dynamics. To start with, it is found that inflation levels are associated to the variance of both permanent and transitory components. Yet, it seems that the link is stronger between inflation and long-term uncertainty (higher instability in trend inflation) than short-term variability. Given that trend inflation is explained by monetary policy actions, these results suggest that high-level inflation makes policy less stable and, hence, it implies rising stabilization costs.

and typical structural shocks. His model is able to generate time-varying inflation dynamics, in particular persistence, similar to those observed in long-run U.S. data. Castillo and Winkelried (2006) have used the same argument along agents' heterogeneity in order to explain why dollarization is so persistence even though inflation has declined to low levels.

Remarkably, short-run uncertainty is also important once we allow for regime switches in inflation dynamics. Indeed, there is evidence of three differentiated regimes over the entire sample. Sub-periods 1949:3-1975:2 and 1994:2-2006:1 are classified as low-level, low-volatility inflation regimes. The most recent period of price stability, that includes the inflation targeting experience in Peru, could be ascribed to shifting emphasis on monetary aggregates and/or on changes of policy makers' preference towards inflation-fighting policies. A particular important result from the analysis is that, before the recent price stability and inflation targeting regimes (1994 -2006), another low-uncertainty regime was in place from 1949 to 1975 but with a different pattern in its short-run uncertainty. The main difference comes out from the explicit inflation-intolerant monetary policy, reinforced by the adoption of the inflation-targeting scheme, in the most recent period. Not only this orientation might have contributed to achieve lower inflation levels than otherwise, but also might have helped to reduce considerably short-run volatility. This link between inflation levels and short-run uncertainty highlights the importance of the inflation-targeting scheme of monetary policy in curving down inflation expectations and shifting uncertainty to lower levels in the short-run. A third relevant finding is that inflation persistence increases with inflation and inflation variability.

Important conclusions arise for monetary policy's orientation. Keeping trend inflation under control and dragging inflation expectations down best reinforce credibility in a central bank's inflation-intolerant policy. Long-term, permanent shocks to inflation trend should be consistently and permanently avoided.[33] Once monetary authorities start losing control of trend inflation, chances of rapidly shifting to a high-level and high-variance inflation regime are not negligible at all and the danger of falling down into a hyperinflation spiral is latent. Domestic impulses to short-run transitory shocks are weakened if on top of a downward-trend-inflation management, inflation expectations are anchored towards low-level and low-variance inflation. Hence, inflation targeting regimes' contribution to monetary policy efficiency is best assessed under this perspective.

Overall, the empirical evaluation in the paper justifies studying inflation dynamics incorporating pre-hyperinflation observations to capture and distinguish regime shifts.[34] Recent experience of price stability reveals inflation-fighting policy's contributions to anchoring inflation expectations down, once the historical experience is set into perspective (and benefiting from the rich information contents in past inflation dynamics).

Using univariate modelling proved valuable for revealing inflation dynamics but it certainly reaches its limits when uncertainty about the sources of shocks is an issue. Further research will, then, be directed towards Markov switching structural multivariate models of inflation

[33] Something for which central bank autonomy should be granted and respected.

[34] Not a common approach, since inflation dynamics are highly distorted by hyperinflation periods.

dynamics, very much in the line of Sims and Zha (2005). Structural identification of the sources of regime switching is needed to assess if switching policy orientations or switching nature of volatility shocks are responsible for those inflation patterns studied here so far.

References

[1] Angeloni, Ignazio; Luc Aucremanne, Michael Ehrmann, Jordi Galí, Andrew Levin, Frank Smets (2005). New evidence on inflation persistence and price stickiness in the Euro area: implications for macro modelling. Mimeo, September.

[2] Armas, Adrian and Francisco Grippa (2005). Targeting inflation in a dollarized economy: the Peruvian experience. Seminar paper Inter-American Development Bank May.

[3] Ball, Laurence (1992). Why does high inflation raise inflation uncertainty?. *Journal of Monetary Economics*, Vol. 29, pp. 371-388.

[4] Ball, Laurence and Stephen Cecchetti (1990). Inflation uncertainty at short and long horizons. *Brooking Papers on Economic Activity,* Vol. 1990, No. 1, pp. 215-254.

[5] Bhar, Ramaprasad and Shigeyuki Hamori (2004). The link between inflation and inflation uncertainty: evidence from G7 countries. *Empirical Economics,* Vol. 29, No. 4, pp. 825-853.

[6] Borio, Claudio and Andrew Filardo (2006). Globalisation and inflation: new cross-country evidence on the global determinants of domestic inflation. Mimeo, Bank for International Settlements (March).

[7] Bratsiotis, George J.; Jakob Madsen, Christopher Martin (2002). *Inflation targeting and inflation persistence.* Mimeo, May.

[8] Bredin, Don and Stilianos Fountas (2006). Inflation, inflation uncertainty, and Markov regime switching heteroskedasticity: evidence from European countries. Mimeo, April.

[9] Castillo, Paul and Diego Winkelried (2006). Dollarization persistence and individual heterogeneity. Mimeo, January.

[10] De la Rocha, Javier (1999). The transmission mechanism of monetary policy in Peru. In, Bank for International Settlements, Policy papers, No. 3.

[11] Diebold, Francis X.; Joon-Haeng Lee; and Gretchen C. Weinbach (1994). Regime switching with time-varying transition probabilities. In *Hargreaves, C. (ed.) Nonstationary time series analysis and cointegration.* Oxford University Press.

[12] Evans, Martin (1991). Discovering the link between inflation rates and inflation uncertainty. *Journal of Money Credit and Banking,* Vol. 23, No. 2 pp. 169-184.

[13] Evans, Martin and Paul Wachtel (1993). Inflation regimes and the sources of inflation uncertainty. *Journal of Money Credit and Banking,* Vol. 25, No. 3, pp. 475-511.

[14] Franses, P.H. and N. Haldrup (1994). The effects of additive outliers on tests for unit roots and cointegration. *Journal of Business and Economic Statistics*, Vol. 12, pp. 471-478.

[15] Friedman, Milton (1977). Nobel lecture: inflation and unemployment. *Journal of Political Economy,* Vol. 85, pp. 451-472.

[16] Gordon, Robert J. (1990). Comments and discussion. In *Ball, Laurence and Stephen Cecchetti (1990): "Inflation uncertainty at short and long horizons". Brooking Papers on Economic Activity,* Vol. 1990, No. 1, pp. 215-254.

[17] Guevara, Guillermo (1999). La política monetaria del Banco Central del Perú, una perspectiva histórica. *Revista de Estudios Económicos N°5, Banco Central de Reserva del Perú.*

[18] Hamilton, James D. (1989). A new approach to the economic analysis of nonstationary time series and the business cycle. *Econometrica,* Vol. 57, No. 2, March, pp. 357-384.

[19] Hamilton, James D. (1994). *Times series analysis.* Princeton University Press, Princeton.

[20] Hamilton, James D. and Baldev Raj (Eds.) (2002). *Advances in Markov-switching models. Applications in business cycle research and finance.* Series: Studies in Empirical Economics. Springer Verlag.

[21] Harvey, Andrew C. (1993). *Time Series Models.* MIT press.

[22] Johansen, Soren and K. Juselius (1992). Maximum likelihood estimation and inference on cointegration with applications to the demand for money. *Oxford Bulletin of Economics and statistics*, Vol. 2, pp. 169-210.

[23] Kim, Chang-Jin (1993). Unobserved-component time series models with Markov-switching heteroskedasticity: changes in regime and the link between inflation rates and inflation uncertainty. *Journal of Business and Economic Statistics*, Vol 11, pp 341-349.

[24] Kim, Chang-Jin and Charles R. Nelson (1999). *State-space models with regime switching. Classical and Gibbs-sampling approaches with applications.* The MIT Press, Cambridge Massachusetts, London England.

[25] Krolzig, Hans-Martin (1997). Markov-switching vector autoregressions. Modelling, statistical inference, and application to business cycle analysis. Lecture notes in economics and mathematical systems, Springer.

[26] Lansing, Kevin (2006). Time-varying U.S. inflation dynamics and the new Keynesian Phillips curve. Mimeo, Federal Reserve Bank of San Francisco.

[27] Marcet, Albert and Juan Pablo Nicolini (2005). Money and prices in models of bounded rationality in high-inflation economies. *Review of Economic Dynamics,* Vol. 8, pp. 452-479.

[28] Mizuno, Takayuki; Misako Takayasu, and Hideki Takayasu (2006). The mechanism of double exponential growth in hyper inflation. Mimeo, Preprint submitted to Elsevier Preprint, August.

[29] Quispe, Zenón (2000). Monetary policy in a dollarised economy, the case of Peru. In *Monetary Policy Framework in a Global Context.* Edited by Lavan Mahadeva and Gabriel Sterne.

[30] Perron, Pierre (1994). Non-stationarities and non-linearities in Canadian inflation. In *Economic behaviour and policy choice under price stability.* Proceedings of a conference held at the Bank of Canada in October 1993. Bank of Canada, June.

[31] Robalo Marques, Carlos (2004). Inflation persistence: facts or artefacts?. Working Paper No. 371, European Central Bank, June.

[32] Rodríguez, Gabriel (2004). An empirical note about additive outliers and nonstationarity in Latin-American inflation series. *Empirical Economics,* Vol. 29, pp. 361-372.

[33] Rossini, Renzo (2001). Aspectos de la adopción de un regímen de metas de inflación en el Perú. *Revista de Estudios Económicos N°5, Banco Central de Reserva del Perú.*

[34] Rudd, Jeremy and Karl Whelan (2005). Modelling inflation dynamics: a critical survey of recent research. Mimeo for FRB/JMCB Conference "Quantitative evidence on price determination". September.

[35] Salomon, Marcelo F. (2001). The inflationary consequences of fiscal policy in Brazil: an empirical investigation with regime switches and time-varying probabilities. *Studies in Nonlinear Dynamics and Econometrics.* Vol. 5, No. 1, pp. 40 - 56.

[36] Sargent, Thomas, Noah Williams, and Tao Zha (2006). The conquest of South American inflation. Mimeo, in http://homepages.nyu.edu/$\sim ts43$/.

[37] Sims, Christopher A. and Tao Zha (2005). Were there regime switches in US monetary policy?. *American Economic Review*, forthcoming.

[38] Trabelsi, Abdelwahed and Maha Achour (2005). Markov switching and state space approaches for investigating the link between inflation level and inflation uncertainty. Mimeo, February.

[39] Vega, Marco and Diego Winkelried (2005). Inflation targeting and inflation behavior: a successful story?. *International Journal of Central Banking*, Vol. 1, No. 3, pp. 153-175.

[40] Vogelsang, Timothy J. (1999). Two simple procedures for testing for a unit root when there are additive outliers. *Journal of Time Series Analysis,* Vol. 20, No. 2.

A Tables

Table 1: Summary statistics

Period	Inflation Rate	
	Mean	Std. Dev.
1951 - 1960	0.60	0.99
1961 - 1970	0.78	0.97
1971 - 1980	2.41	2.40
1981 - 1990	16.40	38.28
1991 - 2000	1.88	2.58
2001 - 2006	0.16	0.36

Table 2: MA estimation of inflation change by sub-samples

Period	θ^*	σ_v^2
$1950-55$	-0.997	0.691
$1955-60$	-0.702	0.332
$1960-65$	-0.834	0.285
$1965-70$	-0.509	0.827
$1970-75$	-0.771	0.728
$1975-80$	-0.771	4.596
$1980-85$	-0.475	2.521
$1995-20$	-0.624	0.101
$2000-05$	-0.993	0.074

* Parameters are significant to the 95% confidence.

Table 3: Effects of average inflation on standard deviations of permanent and transitory shocks. Perú, five-year periods, 1950.01-2005:04

	Dependent Variable	Coefficient on Average Inflation	R^2	
	Permanent Shock (σ_ε^2)	0.173	0.84	
		(7.617)		
	Transitory Shock (σ_η^2)	0.163	0.52	
		(2.003)		

Numbers in parentheses are t-statistics. Iinformation for 1985-1995 is excluded.

Table 4: MSIAH(3)-AR(1) estimates for inflation in Peru 1949 - 2006

Parameters/ Regimes	Quarterly rate			
	Seasonally Adjusted	t-values	Seasonally Unadjusted	t-values
Regime 1				
Intercept	1.3405	6.6844	1.2970	5.9417
Autoregressive	0.2953	3.8531	0.3326	3.6913
Std. Dev.	1.6940		1.7376	
Regime 2				
Intercept	5.8310	4.3122	5.7279	6.0417
Autoregressive	0.6045	11.5191	0.6240	31.8427
Std. Dev.	7.9170		6.4160	
Regime 3				
Intercept	209.1280	1.6702	260.6374	2.1463
Autoregressive	-0.1315	-0.3676	-0.1754	-0.4853
Std. Dev.	211.7900		239.9200	

Table 5: Transition matrix for MSIAH(3)-AR(1)*
Quaterly inflation rate: 1949 -2006

	Regime 1	Regime 2	Regime 3
Regime 1	0.9926	0.0074	0.0000
Regime 2	0.0176	0.9523	0.0302
Regime 3	0.0002	0.1798	0.8200

*Seasonally adjusted series.

Table 6: Summary statistics by regimes

Switching regimes	Inflation Rate	
	Mean	Std. Dev.
Regime 1		
1949:01 - 1972:12	0.71	1.04
1994:01 - 2006:05	0.44	0.51
Total sample	0.61	0.91
Regime 2		
1973:01 - 1987:12	4.33	2.86
1991:01 - 1993:12	4.75	3.15
Total sample	4.40	2.91
Regime 3		
1988:01 - 1990:12	40.74	64.02

Table 7: Regime switching heteroskedasticity model of inflation in Peru 1949 - 2006

Parameters	Sample estimates*					
	1949 - 1987		1991 - 2006		1973 - 1993	
	Estimates	St. Dev.	Estimates	St. Dev.	Estimates	St. Dev.
p_{11}	0.9204	0.0792	0.9169	0.1376	0.9012	0.0805
p_{00}	0.9819	0.0106	0.9847	0.0159	0.9765	0.0205
q_{11}	0.7738	0.1040	0.9824	0.0235	0.9728	0.0410
q_{00}	0.8276	0.0648	0.9857	0.0169	0.9749	0.0307
Q_0	0.0001	0.1807	0.2841	0.1136	1.3040	0.5977
h_0	0.9621	0.1048	0.5597	0.0889	2.8742	0.4512
Q_1	8.4971	2.8329	12.7843	8.6773	43.0957	9.0696
h_1	2.0031	0.3975	5.9152	1.9081	5.9005	1.6237
μ_2	2.1537	0.6224	2.6994	2.3755	4.4619	5.4759
μ_3	2.2399	2.3368	1.2003	8.0934	23.3032	9.1136
μ_4	6.9749	4.5631	21.6089	14.6846	6.5318	12.6918
Q_1 / Q_0	106213.563		44.999		33.049	
h_1 / h_0	2.082		10.569		2.053	
Log likelihood	375.129		124.954		305.797	

Table 8: Signal to noise ratio and Kalman gain across regimes

	Regime 1	Regime 2
ρ	0.295	0.604
S	0.262	0.584
K	0.398	0.526

Regime 1 corresponds to the low-volatility period and Regime 2 to the high-volatility one.

B Markov switching autoregressive model of inflation

This Appendix describes the Markov switching autoregressive (MS-AR) model that we preliminary use to characterize inflation dynamics in Peru. Regime switching explains time-varying parameters in the autoregressive representation of inflation. Several specifications are estimated and evaluated. Fixed transition probabilities are considered in all of them.[35]

B.1 Model description

Given the stochastic nature of regime shifts in inflation over the long-term, a MS-AR model of the inflation rate, π_t, is defined as a $AR(p)$ model conditional upon an unobservable regime $s_t \in \{1, ...M\}$ as in:

$$\pi_t = c(s_t) + \sum_{j=1}^{p} \beta_j(s_t)\pi_{t-j} + \eta_t \tag{17}$$

where η_t is assumed to be a Gaussian innovation process, conditional on the regime s_t, and $\eta_t \sim NID(0, \sigma(s_t))$.[36] The discrete random variable s_t describes the finite number of possible regimes, so that it could take on the values $1, 2, ..M$. This uniequational model is completed by assuming that the regime generating process is a discrete-state homogeneous Markov chain defined by the transition probabilities: $p_{ij} = \Pr(s_{t+1} = j/s_t = i)$ and the condition that $\sum_{j=1}^{M} p_{ij} = 1 \forall i, j \in \{1, ..., M\}$. These probabilities could also be represented in the transition matrix for an irreducible ergodic M state Markov process (s_t):

$$P = \begin{bmatrix} p_{11} & p_{12} & \cdots & p_{1M} \\ p_{21} & p_{22} & \cdots & p_{2M} \\ \cdots & \cdots & \cdots & \cdots \\ p_{M1} & p_{M2} & \cdots & p_{MM} \end{bmatrix}$$

where $p_{iM} = 1 - p_{i1} - ... - p_{i,M-1}$ for $i = 1, ...M$.

Following Krolzig (1997)'s notation, a MSIAH(m)-AR(p) is estimated. As Hamilton (1994) argues, a Markov switching representation is plausible only if economic rationale supports statistical findings of regime shifting. A by-product result from parameter estimation is the evaluation of inflation persistence over the sample.

[35] As in Hamilton (1989). For a detailed description of the case with time-varying transition probabilities see Diebold, Lee, and Weinbach (1994). For a recent survey of contributions in Markov switching modelling, see Hamilton and Balder (2002), and for an state-of-the-art update see Hamilton (2005).

[36] Notice out that this η_t term resembles directly the term η_t from Equation (1).

B.2 Inflation regimes

After careful experimentation with different representations and using the "bottom-up" procedure for model selection, as in Krolzig (1997), a MSIAH(3)-AR(1) has been selected to fit the inflation rate series for Peru.[37] Insightful gains from this representation are that it allows different trending in inflation (time-variant intercept), regime-dependent degree of inflation persistence (directly observable from the autoregressive parameter), and changing shock volatilities (standard deviation of errors). Data sample spans for almost six decades (from 1949 to 2006) and includes quarterly observations of percentage change in the seasonally adjusted CPI. Main conclusions from model estimation are robust to data frequency (using monthly observations with one-, three- and twelve-month percentage changes), variable adjustment (seasonally unadjusted series), lag selection (up to four lags), variable definition (in log differences), and sample length. The inflation rate is calculated as: $\left(\frac{\pi_t}{\pi_{t-1}} - 1\right) * 100$, where π_t is the CPI (in levels).[38]

Results suggest the presence of three regimes for the inflation dynamics in Peru. The two first regimes seem to mark inflation dynamics in the long run, with periods of price-level stability and episodes of unstable financial conditions (concentrated around the 1980s). A third regime corresponds to the hyperinflation experience in Peru by the end of the 1980s. Coefficient estimates for the latter regime are not statistically significant, but are well above historical levels for both the intercept (209.1) and the standard deviation of errors (211.8). The negativity and insignificance of the autoregressive parameter, from an econometric point of view, is most likely due to the presence of additive outliers in the hyperinflation period (observations at September 1988 and September 1990, which correspond to policy-induced shocks).[39] Indeed, additive outliers bias autoregressive parameters to zero and introduce moving average (MA) terms into the dynamics of the series. Thus, a rather artificial treatment of those two additive outliers renders the persistence parameter during the hyperinflation regime to a more interpretable 0.8 value.[40]

The transition matrix shows no probability of switching from the first (low-inflation) regime to the third (hyperinflation) regime. A near-zero transition probability denies such a scenario.[41]

[37] The model was estimated using the MSVAR application for Ox from Krolzig's webpage: www.kent-ac.uk/economics/staff/hmk.

[38] GDP deflator has not been used here because it is not readily available in Peru for such a long sample size.

[39] This issue was pointed out to us by Gabriel Rodríguez. See Franses and Haldrup (1994), Perron (1990), and Vogelsang (1999).

[40] Outlier observations were dropped out from the sample estimation. Estimation results are available from the authors upon request.

[41] This fact would support restricted estimation of the transition matrix, imposing a zero transition probability between regimes 1 and 3. It has not actually implemented here, but further structural estimation should consider such a restriction.

Staying in the quiet regime actually shows the highest probability of all possible transitions. Furthermore, chances of shifting from regime 2 (high-inflation) to regime 3 are clearly non-negligible and must signal indeed accelerating inflation risks once high levels of inflation are reached (see Table 5 in Appendix A for details on the transition matrix and Table 6 for summary statistics).

B.3 Inflation persistence

This section evaluates if there is a link between inflation and inflation persistence conditional to the regime in place. Remarkably, inflation persistence is positively linked to inflation level and variability. As reported in Table 4 in Appendix A, the degree of inflation persistence is around 0.6 in the high-inflation regime. It reduces to around half of it, (0.3), in the low-inflation scenario. These measures are robust to lag selection, data frequency, and variable definition. Thus, for example, estimating a MSIAH(3)-AR(4) renders a 0.56 persistence coefficient in the high-inflation regime and 0.27 in the other regime.[42] The degree of inflation persistence is calculated in this case as the sum of all the autoregressive coefficients.[43] Meanwhile, seasonally unadjusted series for the inflation rate produces persistence coefficients of 0.62 and 0.33, respectively. Changing to monthly data, a MSIAH(3)-AR(1) model suggests 0.64 and 0.28 values, respectively, for these parameters (0.69 and 0.3, when the inflation rate is calculated as log differences of the CPI). In all these cases, coefficient estimates are statistically significant. Empirical evidence clearly shows, then, that inflation persistence diminishes when inflation level and variability decrease. Furthermore, in all these model estimations, the third regime of hyperinflation does not support a statistically significant nor economic interpretable degree of inflation persistence. However, when the two additive outliers in the hyperinflation period are considered, estimation results support the increase in inflation persistence with level and volatility of inflation. These results on inflation persistence are robust even to a much shorter sample size, though not as decisively as shown above[44].

[42] This model outperforms marginally the one-lag model used here in terms of the simple log-likelihood ratio test criteria. However, the latter is preferred for the straightforward interpretation of the autoregressive coefficient as measuring inflation persistence.

[43] See Robalo (2004) for a discussion on measures of inflation persistence.

[44] Results for the sample 1992 - 2006 are available from the authors upon request.

Documentos de Trabajo publicados
Working Papers published

La serie de Documentos de Trabajo puede obtenerse de manera gratuita en formato pdf en la siguiente dirección electrónica:
http://www.bcrp.gob.pe/bcr/index.php?Itemid=213

The Working Paper series can be downloaded free of charge in pdf format from:
http://www.bcrp.gob.pe/bcr/ingles/index.php?Itemid=104

2007

Marzo \ March

DT N° 2007-004
Dollarization Persistence and Individual Heterogeneity
Paul Castillo y Diego Winkelried

DT N° 2007-003
Why Central Banks Smooth Interest Rates?: A Political Economy Explanation
Carlos Montoro

Febrero \ February

DT N° 2007-002
Comercio y crecimiento: Una revisión de la hipótesis “Aprendizaje por las Exportaciones”
Raymundo Chirinos Cabrejos

Enero \ January

DT N° 2007-001
Perú: Grado de inversión, un reto de corto plazo
Gladys Choy Chong

2006

Octubre \ October

DT N° 2006-010
Dolarización financiera, el enfoque de portafolio y expectativas:
Evidencia para América Latina (1995-2005)
Alan Sánchez

DT N° 2006-009
Pass–through del tipo de cambio y política monetaria:
Evidencia empírica de los países de la OECD
César Carrera, Mahir Binici

Agosto \ August

DT N° 2006-008
Efectos no lineales de choques de política monetaria y de tipo de cambio real en economías parcialmente dolarizadas: un análisis empírico para el Perú
Saki Bigio, Jorge Salas

Junio \ June

DT N° 2006-007
Corrupción e Indicadores de Desarrollo: Una Revisión Empírica
Saki Bigio, Nelson Ramírez-Rondán

DT N° 2006-006
Tipo de Cambio Real de Equilibrio en el Perú: modelos BEER y construcción de bandas de confianza
Jesús Ferreyra y Jorge Salas

DT N° 2006-005
Hechos Estilizados de la Economía Peruana
Paul Castillo, Carlos Montoro y Vicente Tuesta

DT N° 2006-004
El costo del crédito en el Perú, revisión de la evolución reciente
Gerencia de Estabilidad Financiera

DT N° 2006-003
Estimación de la tasa natural de interés para la economía peruana
Paul Castillo, Carlos Montoro y Vicente Tuesta

Mayo \ May

DT N° 2006-02
El Efecto Traspaso de la tasa de interés y la política monetaria en el Perú: 1995-2004
Alberto Humala

Marzo \ March

DT N° 2006-01
¿Cambia la Inflación Cuando los Países Adoptan Metas Explícitas de Inflación?
Marco Vega y Diego Winkelreid

2005

Diciembre \ December

DT N° 2005-008
El efecto traspaso de la tasa de interés y la política monetaria en el Perú 1995-2004
Erick Lahura

Noviembre \ November

DT N° 2005-007
Un Modelo de Proyección BVAR Para la Inflación Peruana
Gonzalo Llosa, Vicente Tuesta y Marco Vega

DT N° 2005-006
Proyecciones desagregadas de la variación del Índice de Precios al Consumidor (IPC), del Índice de Precios al Por Mayor (IPM) y del Crecimiento del Producto Real (PBI)
Carlos R. Barrera Chaupis

Marzo \ March

DT N° 2005-005
Crisis de Inflación y Productividad Total de los Factores en Latinoamérica
Nelson Ramírez Rondán y Juan Carlos Aquino.

DT N° 2005-004
Usando información adicional en la estimación de la brecha producto en el Perú: una aproximación multivariada de componentes no observados
Gonzalo Llosa y Shirley Miller.

DT N° 2005-003
Efectos del Salario Mínimo en el Mercado Laboral Peruano
Nikita R. Céspedes Reynaga

Enero \ January

DT N° 2005-002
Can Fluctuations in the Consumption-Wealth Ratio Help to Predict Exchange Rates?
Jorge Selaive y Vicente Tuesta

DT N° 2005-001
How does a Global disinflation drag inflation in small open economies?
Marco Vega y Diego Winkelreid