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

Paul Castillo† Alberto Humala‡ Vicente Tuesta§

August 2006

Abstract

The link between inflation, inflation uncertainty, and inflation persistence is evaluated in a context of monetary policy regime shifts for the Peruvian economy. Several univariate time-series techniques are used to represent inflation dynamics. First, the paper follows the approach of Ball and Cecchetti (1990) to decompose inflation into its stochastic trend and its stationary part. Then, a Markov switching univariate model (MS-AR) is postulated to capture and identify regime shifting in inflation dynamics and to assess whether the link between uncertainty and inflation is conditional on the regime. Lastly, a model that simultaneously deals with both regime shifts and transitory and permanent components of inflation, in line with Kim and Nelson (1999), is estimated and analysed. Main findings can be summarized as follows: (a) Periods of high (low) inflation mean were accompanied by periods of high (low) short- and long-run uncertainty and persistence in inflation. (b) Three clearly differentiated regimes were identified. First, a period of price stability, then a high-inflation high-volatility regime, and finally a hyperinflation period, (c) Inflation and money growth rates share the same permanent component and similar regime shifts. As a result, it could be argued that monetary policy regimes can explain changes in inflation uncertainty and persistence in Peru.

JEL Classification: C22, E31, E42, E52

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

---

*We thank participants in the Economic Research Workshop at the Central Reserve Bank of Peru and Gabriel Rodríguez for useful 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 Bank.

†Central Reserve Bank of Peru and LSE.

‡Central Reserve Bank of Peru, Corresponding author: alberto.humala@bcrp.gob.pe, Jr. Miró Quesada 441, Lima-Perú. (511)-613 2785.

§Central Reserve Bank of Peru.

1 Introduction

There is some evidence that high inflation rates impose considerable costs to the economy, since those rates enlarge both inflation uncertainty and inflation persistence[1]. Higher inflation uncertainty induces larger stabilization costs because it makes more difficult to forecast inflation. Besides, as Milton Friedman (1977) pointed out in his Nobel prize lecture in 1976, higher inflation uncertainty generates larger relative price distortions.

In addition, Ball (1992) suggests that when inflation uncertainty is high, uncertainty about future monetary policy actions increases. More recently, Lansing (2006) shows, using a simple model of inflation determination, how high uncertainty about inflation renders inflation more persistent and, thus, increases stabilization costs. All in all, establishing empirically the relevance of this link between mean, uncertainty and persistence of inflation and assessing what role monetary policy plays in it, is crucial to understand inflation and the costs of stabilization programmes.

In this paper, then, several univariate time-series techniques are used to assess extensively the existence and relevance of the aforementioned link.[2] A relative long-span set of inflation data (quarterly observations for 1949.1 to 2006.2) is looked at to identify the relationship between average inflation, inflation uncertainty, and inflation persistence, across different monetary policy regimes. 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 as to whether or not monetary policy shifts have induced indeed changes in inflation dynamics. In particular, there are not many attempts to assess the impact in inflation uncertainty and persistence of the recently-adopted inflation-targeting regime that put in perspective policy regime switches from the pre-hyperinflation period. This paper intends to fill this gap focusing on the feasible regime switching nature of the inflation rate over a sample that spans almost six decades.

In the literature, the link between inflation and inflation uncertainty has been addressed mainly for the case of the US. For instance, Ball and Cecchetti (1990) 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

[1] We define inflation uncertainty as the variance of the forecast error of inflation.

[2] Future research agenda considers extending the analysis to a multivariate approach.

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 might become 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.[3]

Methodologically, this paper proceeds as follow. First, this paper explores whether 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, it considers possible regime shifts in inflation trend and short-term volatility by estimating a first-order Markov switching autoregressive model of inflation (MS-AR).[4] This type of inflation modelling allows assessing the extent to which (short- run) uncertainty about future regime shifts enhances inflation uncertainty and persistence. Thereafter, the study focuses on a Markov switching heteroskedasticity model of inflation, whereby inflation dynamics is decomposed into a stochastic trend and a stationary part both subject to regime switching in their disturbances, as in Kim and Nelson (1999). Finally, the possibility of permanent inflation and money growth rates sharing their permanent component is studied. In a recent theoretical approach, Sargent, Williams, and Zha (2006) impose economic structure into a model to link "normal" inflation dynamics to money-financed fiscal deficits, and "extraordinary" dynamics to destabilizing expectations. Our empirical results are consistent with their structural approach and results.

Main empirical results in this study indicate that there exists a link between inflation and inflation uncertainty at both short and long-horizons in Peru. Moreover, periods of high long-run inflation uncertainty are associated to higher persistence in inflation dynamics. Supporting the association of the regime switching nature of inflation dynamics and of monetary policy, three regimes are clearly identified in both inflation and money growth. A low-inflation stable regime 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. Interestingly, inflation persistence in the high-inflation high-

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

[4]This representation is also evaluated for a money aggregate (M2) to assess similarities in price's and money's regime switching behaviour and dynamics.

volatility regime is twice the level of persistence during the low-inflation stable regime. Finally, empirical evidence supporting the hypothesis of a common permanent component for inflation and money growth is found. So far, these results suggest that monetary policy shifts from low- to high-inflation regimes, first, and from high- to low-inflation regimes later on, explain the rise in inflation volatility and persistence from 1949-1975 to 1976-1994, and the opposite movement from 1994 on, respectively. 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, attributes the shifting in inflation regimes to stochastic switches between rational expectations (normal inflation) and adaptive expectations (high- and hyper-inflation periods).

The rest of the paper is organized as follows: in Section 2, data preliminary analysis is conducted as a first approximation to the link between the level of inflation and inflation uncertainty. An unobserved component model of inflation is estimated and presented. Section 3 evaluates in detail possible regime switches in inflation dynamics by estimating two Markov switching models of inflation. The first one is to identify shift dates and persistence evolution. The second model is to allow conditional and unconditional heteroskedasticity in shocks to permanent and transitory components of inflation. This section also assesses whether or not those regime switches in inflation are linked to monetary policy shifts. Section 4 discusses a theoretical framework that provides rationale to the relationship between inflation uncertainty and persistence. A last section concludes and outlines research agenda.

2 A Glance at Inflation Uncertainty

This section provides a *prima facie* evidence of the relationship between inflation and the variance of future inflation. It follows closely Ball and Cecchetti (1990) to assess this link.[5] 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.

First, correlation between sample mean and inflation variance, across non-overlapping periods of different length (from 1 to 10 years), is estimated (see Table 1 in the Appendix). That is, there are 50 data-points for one-year-based calculation, 25 data points for two-year-based calculation and so on.[6] The mean-variance correlation increases the larger the length period. Thus, for instance, correlation for the one-year period is 0.59 and it increases steadily up to

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

[6] For these calculations, observations from the period 1985-1991, which correspond to a large extent to high-level inflation and to the hyperinflation episode, are disregarded.

0.95 for the ten-year-based calculations. These results suggest an important relationship between inflation and inflation volatility (uncertainty).[7] Though this link is not trivial at short horizons, it becomes stronger and more important at longer horizons.

Next, correlation between inflation and inflation variability, for different quarters ahead, is evaluated (see Table 2 in the Appendix). Thus, the following correlation is calculated: $Corr\left(\pi_t, (\pi_{t+x} - \pi_t)^2\right)$ where π_t denotes inflation, x represents the number of periods ahead and $(\pi_{t+x} - \pi_t)^2$ denotes inflation variability. Figure 1 depicts this correlation for each x. Remarkably, the correlation increases from $x = 1$ until $x = 20$, approximately. Thereafter, it remains at a near-constant level with a light-downward trend from $x = 40$ on. Notice out the correlation is 0.26 for the first quarter and it more-than doubles at the end of the first year ($x = 4$). Hence, this pattern shows evidence of inflation levels having stronger links to uncertainty over the end of the year than over the first quarter. Yet, first quarter correlation is not at all negligible, but implies an important link between current inflation levels and future short-run uncertainty.[8] The correlation's pattern observed in Figure 1 suggests also that the current inflation level is very informative about inflation uncertainty in distant future times.[9] In fact, the correlation for $x = 41$ is 0.62 which is slightly above the one observed after the second year $x = 11$ (0.61).

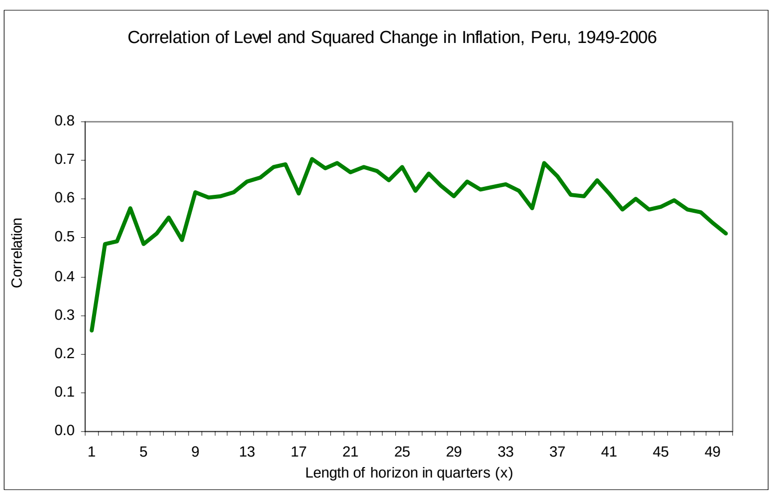

Figure 1

Overall, data analysis supports the existence of a link between inflation and inflation uncertainty (volatility). Furthermore, the link seems to be stronger for long-run uncertainty, but

[7] Variance or volatility does not necessarily mean uncertainty, since part or all of it might be expected. This distinction is stressed later on in Section 3, where shocks to permanent and transitory components of inflation are evaluated.

[8] In Ball and Cecchetti (1990), results for the U.S. economy show a weak link at short horizons.

[9] Instead, Ball and Cecchetti (1990) find, at very long-horizon, a weak correlation between current level of inflation and inflation variability.

in no sense negligible for short-run uncertainty. This evidence remains indicative of such a link, even if inflation variance is standardized.[10]

2.1 An unobserved component model for inflation

This section reviews a simple model for inflation dynamics. The model allows establishing the feasible link between inflation level and uncertainty at short and long horizons. In particular, following Ball and Cecchetti (1990), the inflation rate series is decomposed into its permanent and temporary (but persistent) parts. This decomposition facilitates obtaining different combinations between uncertainty and level of inflation at different horizons. The univariate benchmark model reads as follows:

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

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

where ε_t and η_t denote shocks to the permanent and transitory unobserved components of inflation, respectively. For simplicity, it is assumed that shocks are uncorrelated disturbances with mean zero and variances σ_ε^2 and σ_η^2. 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 a 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. Conveniently, this model of inflation dynamics can be re-written as an ARIMA model with a single shock. Moreover, it can also be expressed as a MA(1) model for the change in inflation rather than inflation itself:[11]

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

where

$$\begin{aligned} \sigma_\varepsilon^2 &= (1+\theta)\, \sigma_v^2 \\ \sigma_\eta^2 &= -\theta \sigma_v^2 \end{aligned} \tag{4}$$

In order to assess effects from trend inflation levels $\left(\pi_t^T\right)$ over short-run and long-run uncertainty $\left(\sigma_\eta^2, \sigma_\varepsilon^2\right)$, it is further assumed that variances of inflation are functions of the previous

[10] In this case, estimation controls for a possible scale effect in the covariance between level and variance of inflation.

[11] 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.

level of trend inflation, as in:

$$\sigma_{\varepsilon}^{2}(t) = \beta_0 + \beta_1 \pi_{t-1}^{T} \quad (5)$$

$$\sigma_{\eta}^{2}(t) = \delta_0 + \delta_1 \pi_{t-1}^{T} \quad (6)$$

The evidence reported in the previous section supports the idea that lag-trend inflation has a stronger impact over long-run inflation uncertainty $\sigma_{\varepsilon}^{2}(t)$ than over short-run uncertainty $\sigma_{\eta}^{2}(t)$. To sum up, 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). Next subsection provides parameter estimates for these coefficients.

2.2 Permanent and transitory shocks

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 θ and variances σ_{v}^{2} are recovered and saved. Estimation results are reported in Table 3 in the Appendix. Thereafter, equation (4) is used to construct variances of both permanent $\left(\sigma_{\varepsilon}^{2}\right)$ and transitory $\left(\sigma_{\eta}^{2}\right)$ shocks. Figure 2 plots average inflation for the nine five-year periods *vis-à-vis* implied values of permanent $\left(\sigma_{\varepsilon}^{2}\right)$ and transitory $\left(\sigma_{\eta}^{2}\right)$ shocks. Both type 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.

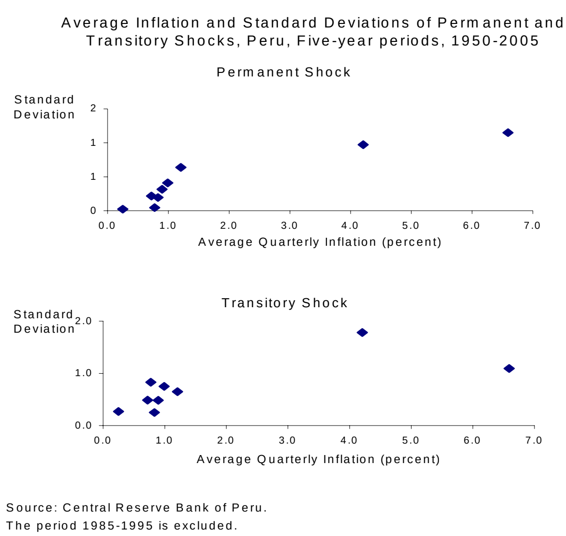

Figure 2

Table 4 in the Appendix 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.

3 Regime Shifts in Inflation Uncertainty

In the Peruvian economy, both inflation trend and short-term deviations from it, have probably changed in the last six decades or son. Those changes seem far from being negligible. First, monetary policy targets have moved from money aggregates to inflation itself and, second, sources of supply and demand shocks have shifted in time. Importantly, the nature and extensions of shocks to both permanent and transitory components of inflation might have also changed during such a long span of time. In order to assess whether or not inflation dynamics has been altered by those shifts, this section takes a regime switching approach that focuses on shifts on the data generating process and on the economic rationale behind those shifts. Plain vanilla evidence of such regime shifting behaviour in inflation dynamics could be gathered by looking at mean and volatility statistics for the inflation rate over several (five-year) sub-samples of the entire period of analysis (see Table 5 in the Appendix). A link between inflation mean and volatility clearly emerges from those basic statistics, though no regime switching behaviour could be formally inferred from them. In what follows, regime switching in inflation dynamics is evaluated, documented, and decomposed into its trend and autoregressive parts. Initially, a Markov switching autoregressive model for inflation shows the likely timing of shifting in regimes. Thereafter, an extension of the unobserved component model of inflation, which allows for stochastic trend and conditional heteroskedasticity, is used very much in line with Kim and Nelson (1999).

3.1 Markov switching autoregressive model of inflation

This sub-section uses a Markov switching autoregressive (MS-AR) model to characterize inflation dynamics in Peru. Several specifications are estimated and evaluated. Fixed transition probabilities are considered in all of them.[12] Suitability of regime switching representations of inflation dynamic is assessed and confronted with economic rationale.

The class of MS-AR models is part of the non-linear representations, where the non-linearity comes from the existence of switching regimes that account for time-varying parameters. If those changes are known to be happening under certain values of a deterministic model of the switching regimes, we have the kind of threshold autoregressive models. Nevertheless, if the process governing the switch in regimes is rather stochastic and is assumed to follow a Markov chain then we deal with the class of Markov switching AR models.[13]

3.1.1 Model description

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

$$\pi_t = c(s_t) + \sum_{j=1}^{p} \beta_j(s_t)\pi_{t-j} + \eta_t \tag{7}$$

where η_t is assumed to be a Gaussian innovation process, conditional on the regime s_t, and $\eta_t \sim NID(0, \sigma(s_t))$.[14] 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 $\sum\limits_{j=1}^{M} p_{ij} = 1 \forall i, j \in \{1, ..., M\}$. These probabilities could also be represented in the transition matrix for an irreducible ergodic M state Markov process (s_t):[15]

[12] A seminal application of regime switching models to dynamic macroeconomics is the Markov switching model from Hamilton (1989). An appealing extension is the case with time-varying transition probabilities in which the feasibility of being in a particular regime varies along with some information variables. See Diebold, Lee, and Weinbach (1994) for a detailed technical treatment of this case. Kim and Nelson (1999) made an empirical application of Markov switching models to inflation dynamics. 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).

[13] More generally, a MS-VAR model, for multivariate time series systems.

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

[15] If the probabilities were independent of the previous occurring regime, then the model would be a simple (rather than Markov) switching model. That is, there would not be persistence in the states. See Hansen (1992).

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

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

The time series vector π_t contains the inflation rate, as measured by the change in CPI. In this general representation, following Krolzig (1997)'s notation, a MSIAH(m)-AR(p) would indicate all parameters are switching between regimes: intercept (I), autoregressive (A) and variance (H, for heteroskedasticity). For realizations of the data, there would also be a time series realization of the regimes prevailing at each possible observation, although this variable will not be directly observable. All that can be done is to infer the probability that a particular regime has occurred at each observation period. The regime-switching model describes the laws governing the transition from one regime prevailing at any particular time to a regime occurring next period through the so-called transition probabilities. Once a regime changes, if there were some degree of persistence in the resulting regime, then the transition probability depends on past values of itself (which regime occurred before). A plausible description of such a pattern is thus the assumed Markov chain process for the unobserved regime variable s_t. In other words, a Markov switching regime assumes that the transition probability at any time t is related to the past only through the most recent realization of regime (at time $t-1$).

As Hamilton (1994) argues, a Markov switching representation is plausible only if economic rationale supports statistical findings of regime shifting. For inflation dynamics modelling, a shifting trend (or mean), different degrees of inflation persistence, or time-variant volatility would suggest the presence of regime switches. Next, regime shifts and transition probabilities are inferred since they are indeed subject to economic interpretation. A by-product result from the parameter estimation is the evaluation of inflation persistence over the sample 1949 - 2006[16].

3.1.2 Regime shifts

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.[17] Insightful gains from this representation are that it

[16] The autoregressive coefficient of the MS-AR stands out as a measure of inflation persistence.

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

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 empirical estimation of this model 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 (original sample selection was 1992-2006). The inflation rate is calculated as: $\left(\frac{\pi_t}{\pi_{t-1}} - 1\right) * 100$, where π_t is the CPI (in levels).[18]

Estimation results show the presence of three clearly differentiated regimes over the entire sample (see Table 6 in the Appendix for parameter estimates).[19] 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).[20] The periods 1949:3 - 1975:2 and 1994:2 - 2006:1 are classified into this first regime. Notice out the inflation-targeting period (since the beginning of 2002) is considered part of this low-inflation regime, though the entire recent experience of price stability (it goes back to the beginning of 1994) is also included in it. Changes in the design and conduct of monetary policy (i.e., larger use of central bank's bonds for monetary operations and improved policy transparency) seem to have guided inflation expectations downwards. In historical perspective, one-digit inflation rates prevailed even during the outburst of several international financial crises that affected the Peruvian economy to different degrees during the 1990s.

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

[19] 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.

[20] All coefficients are significant at usual levels.

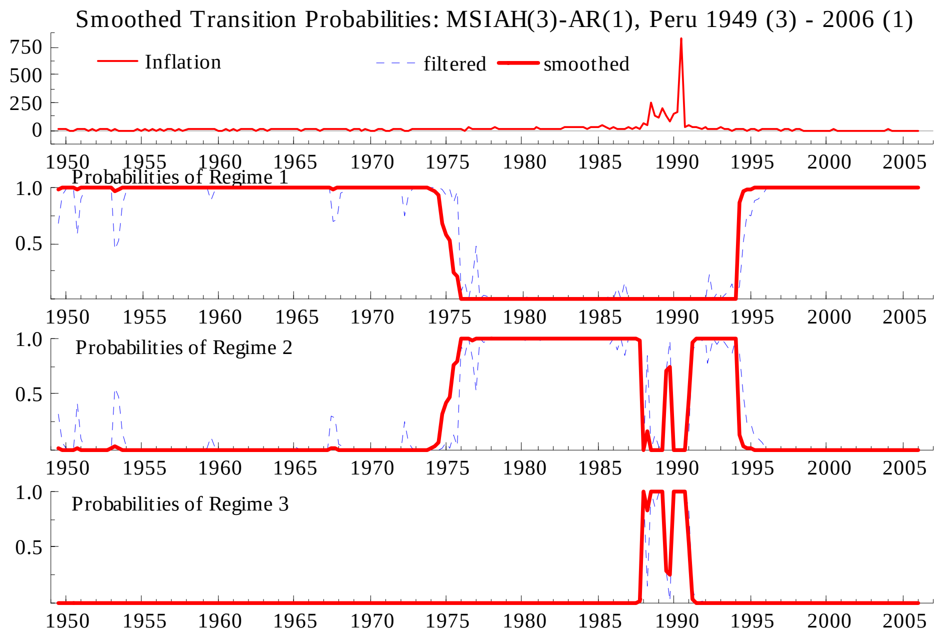

Figure 3

The second regime refers to high-level inflation and highly volatile scenarios (intercept, 5.8; autoregressive parameter, 0.6; and standard deviation, 7.9).[21] Periods considered into this second inflation regime are 1975:3 - 1987:4, in which inflation started mounting to new higher levels (with probably worsening fiscal accounts and relaxation of monetary conditions), and 1991:2 - 1994:1, in which disinflation policies took place. These 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). Although no actual fiscal deficit data has been used here, the link to fiscal policy stance is thoroughly discussed in a theoretical model by Sargent, Williams, and Zao (2006), from which they infer similar regime switching features in inflation dynamics in Peru.

The third regime appears for an extremely volatile and outlier-type hyperinflation period in Peru (1988:1 - 1991:1).[22] Much shorter in length than the other two regimes, coefficient estimates for this 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).[23] Furthermore, the autoregressive parameter is no longer informative about inflation persistence (it becomes rather negative), as extremely-high volatility signals no information to people's expectations on infla-

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

[22] 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).

[23] With a short period, 1989:3 -1989:4, classified into the second regime.

tion. 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 correspond to policy-induced shocks).[24] 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.[25]

Interestingly, 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.[26] 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 7 in the Appendix for details on the transition matrix). Results suggest, then, it is unlikely a shift from the more recent period (1990s) of price stability to a hyperinflation regime unless there is a transition first to higher inflation level and uncertainty.

With this regime classification from the MS-AR estimation, summary statistics of mean and standard deviation of the inflation rate (see Table 8 in the Appendix) confirms the link between inflation rate levels and volatility and, more importantly, their association to regime shifts in inflation dynamics. Much more direct and relevant evidence of regime switching behaviour in the inflation's data generating process if thus revealed with the Markov switching approach and solid economic rationale lies behind it.[27]

3.1.3 Inflation persistence

This part 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 6 in the Appendix, 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

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

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

[26] 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.

[27] See below, an evaluation of the links between money growth and inflation.

high-inflation regime and 0.27 in the other regime.[28] The degree of inflation persistence is calculated in this case as the sum of all the autoregressive coefficients.[29] 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 outliers for 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. Considering a post-hyperinflation data sample (from 1992 to 2006), three regimes are single out, again, as representing inflation dynamics.[30] In this case, those regimes appear sequentially in time, without signalling obvious return to previous regimes (despite the low probabilities of such shifts, regimes are not entirely absorbent). Chronologically, the first regime corresponds to the period of disinflation (1992.4 to 1993.12) with large intercept coefficient (28.3) and standard errors (9.3). The second regime associates to a moderate inflation scenario (1994.1 to 1998.5) with an intermediate intercept and standard deviation (6.4 and 4.1, respectively). The more recent regime belongs to a much more stable economy (1998.6 to 2006.4) where both the intercept and standard deviation are the lowest over this sample size (1.7 and 3.2). Although the Markov regime switching AR(p) model allows a specification in which all parameters are regime dependent, the autoregressive coefficient does not seem to shift significantly among regimes in this sample size. It varies from 0.2 (in the highly volatile disinflation period) to 0.36 (in the moderate-inflation regime) and to 0.3 (in the stable-inflation regime). That is, inflation persistence does not seem to have changed significantly despite the presence of three regimes. However, taking into perspective the results from the larger sample size, these results are rather consistent with the finding that inflation persistence diminishes when inflation level and variability decrease. That is, inflation persistence diminishes from 0.36 in the intermediate regime to 0.3 in the price-stability regime,[31]

[28] 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. Results from the MSIAH(3)-AR(4) estimation are available from the authors upon request.

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

[30] Results are available from the authors upon request.

[31] Though the distinction between a "moderate-inflation" regime and a "low-inflation" regime is based here

while the lower persistence value for the first volatile regime might be due to the presence of some outliers observations in that period (as it was the case for the hyperinflation episode).

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.[32] 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.

3.2 Markov switching heteroskedasticity model of inflation

In order to allow the variance shocks to switch across regimes in the unobserved component model of inflation, the Markov switching heteroskedasticity model of Kim and Nelson (1999) is adopted in this section.[33] Thus, 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 monetary policy regime switching is revealed by this approach.

3.2.1 Model description

The equations for this model are:

$$\pi_t = \pi_t^T + \mu_2 S_{1,t} + \mu_3 S_{2,t} + \mu_4 S_{1,t} S_{2,t} + (h_0 + h_1 S_{2,t})\eta_t \tag{8}$$

$$\pi_t^T = \pi_{t-1}^T + (Q_0 + Q_1 S_{1,t})\varepsilon_t \tag{9}$$

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

on levels that are both classified as low-inflation in the original larger sample size.

[32] 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.

[33] Model estimations were carried on using Kim and Nelson (1999) Gauss codes from their webpage: www.econ.washington.edu/user/cnelson/ssmarkov.htm.

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:

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

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

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.[34] 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).

3.2.2 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.[35] Therefore, considering the regime shifts indicated by the MS-AR, the model is rather estimated by sub-samples that exclude the hyperinflation regime

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

[35] 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.

(1988 - 1990).[36] 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 9 in the Appendix. 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.[37] 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 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).

[36] 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.

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

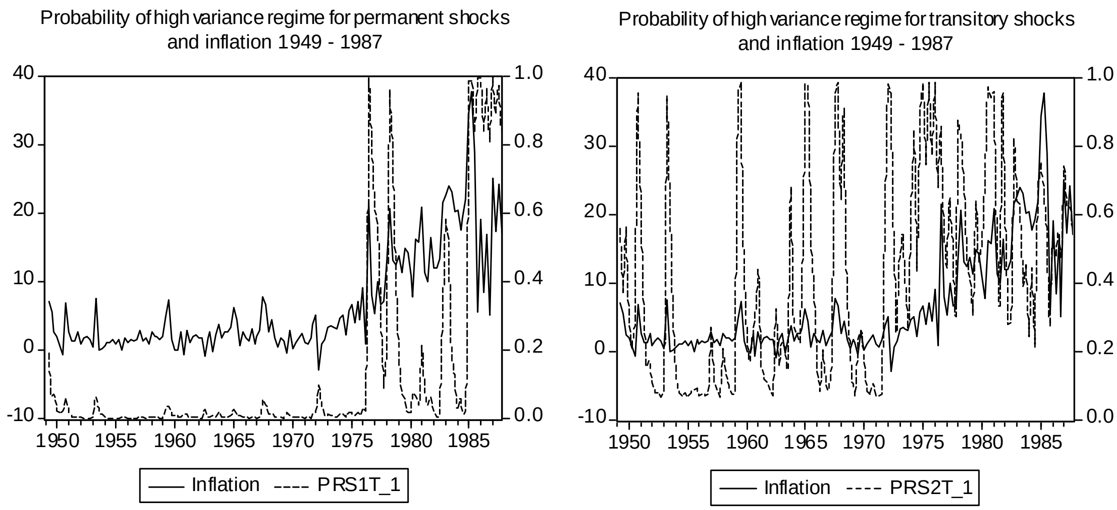

Figure 4

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 9 in the Appendix (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.

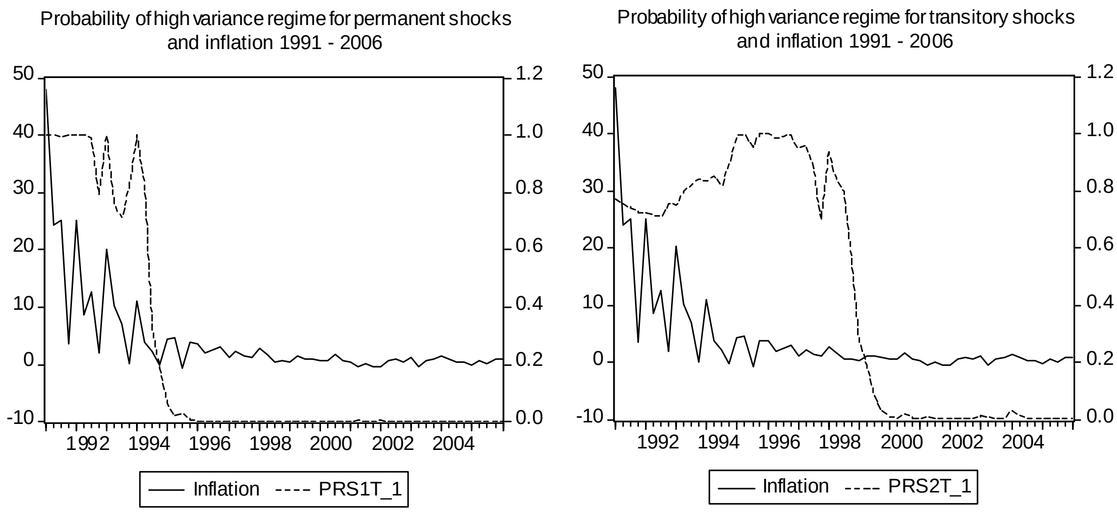

Figure 5

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.[38]

[38] 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.

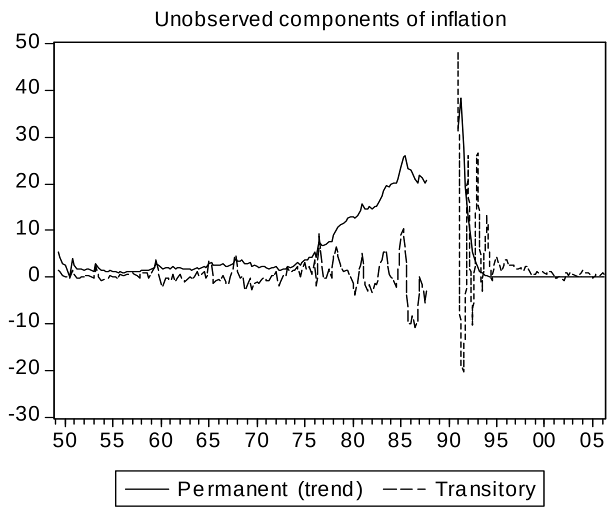

Figure 6

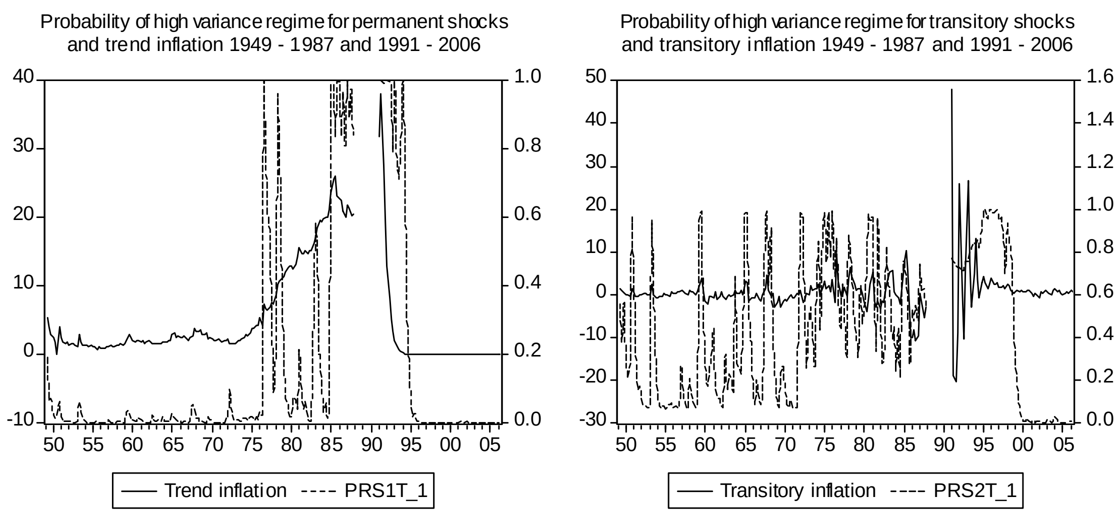

Figure 7

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.[39] Parameter estimates are shown in Table 9 in the Appendix (last two columns). Very

[39] 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).

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 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).[40] 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.

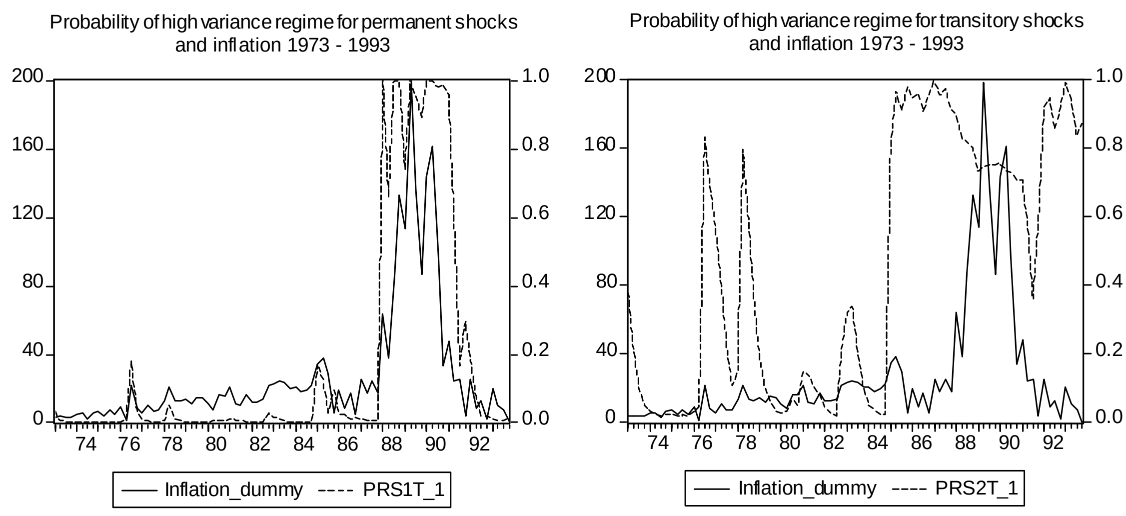

Figure 8

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

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

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.

3.3 Monetary policy and regime shifts

The previous sections provided empirical evidence of the existence of regime shifts in both mean and volatility of permanent and transitory components of inflation in Peru. However, no evidence of the determinants of these shifts was offered. This section addresses this issue. In particular, this section tests whether regime shifts in trend inflation have been related to changes in monetary policy design.

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.[41] 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.[42]

In order to evaluate the relationship between regime shifts in inflation and monetary policy, first, 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).[43] Smoothed probabilities inferred at each observation are presented in Figure 9.[44] As in the case of inflation, three regimes are identified for the money growth rate: low-rate, high-rate, and explosive-rate of money growth. 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, for example, goes up till 1978 (instead of 1975, as in the case of inflation) and prevails again from 1995:1 (instead of 1994:2) onwards. The high mean and volatile regime in the money growth rate is identified

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

[42] 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).

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

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

for periods 1978:4 - 1988:2 and 1991:2 - 1994:4. The explosive-rate regime of money growth goes for the period 1988:3 - 1991:1. Summing up, for Peru, both inflation and money growth rates share fundamentally the same regime shifts over the sample 1949 to 2006.

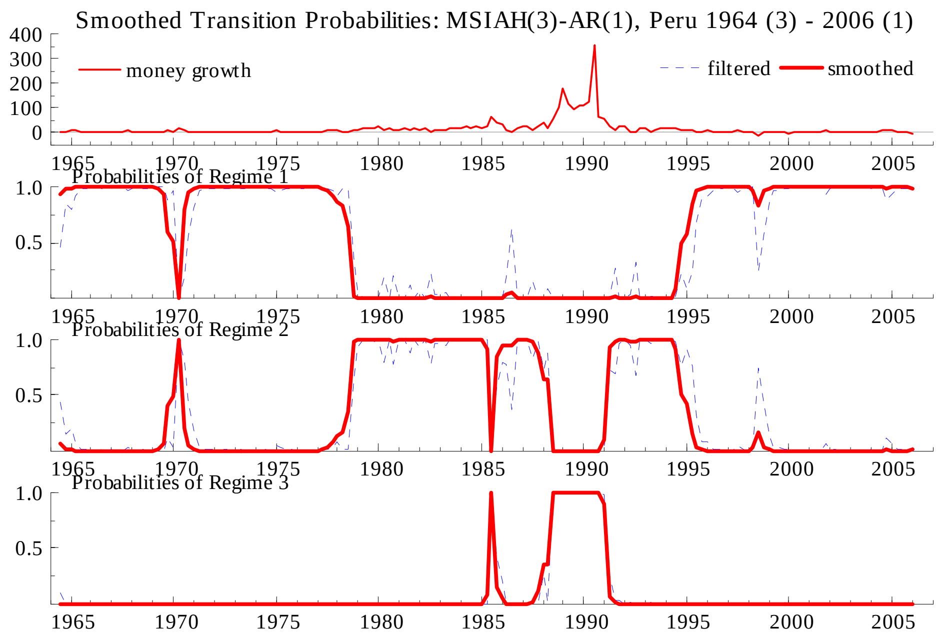

Figure 9

Second, using the cointegration technique from Johansen and Juselius (1990), it is tested whether or not inflation and money growth rate share the same stochastic trend. If so, it could be argued that main determinants of the inflation's regime shifting behaviour were observed changes in monetary policy. As Table 10 in the Appendix shows, under different specifications, the cointegration test indicates inflation and money growth rates share the same stochastic trend in the Peruvian economy. Actually, high-inflation and hyperinflation regimes could be associated to times in which money growth and monetary policy were closely linked to fiscal policy, while that price stability regimes are more associated to periods of more relative central bank autonomy.[45]

Moving on into multivariate or multiequation representation 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); and Salomon (2001) studies empirically the link to fiscal and monetary policies by modelling inflation dynamics subject to regime switches with time-varying transition probabilities (depending upon

[45] The role of money-financed fiscal deficits in accelerating inflation and inflation expectations is stated clearly in Sargent, Williams, and Zha (2006).

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).[46]

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.[47] Complementary, Borio and Filardo (2006) argue that proxies for global economic slack add considerable explanatory power to inflation models, with inflation rates becoming less 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). This parameter can be interpreted as a reduced-form coefficient of inflation persistence. 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 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

[46] 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.

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

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

$$\begin{aligned}\widehat{E}_t\pi_{t+1} &= \widehat{E}_{t-1}\pi_t + K\left(\pi_t - \widehat{E}_{t-1}\pi_t\right),\ 0 < K < 1 \\ &= K\left[\pi_t + (1-K)\,\pi_{t-1} + (1-K)^2\,\pi_{t-2} + ...\right]\end{aligned} \tag{12}$$

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:

$$\pi_t - \pi_{t-1} = \varepsilon_t + \left(\eta_t - \eta_{t-1}\right) \tag{13}$$

Obtaining the unconditional moments:

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

$$Var\left(\Delta\pi_t\right) = E\left(\Delta\pi_t^2\right) = E\left(\left[\varepsilon_t + \left(\eta_t - \eta_{t-1}\right)\right]\left[\varepsilon_t + \left(\eta_t - \eta_{t-1}\right)\right]\right) = \sigma_\varepsilon^2 + 2\sigma_\eta^2 \tag{15}$$

Equations (14) and (15) are used to obtain the signal-to-noise ratio, S. This ratio is defined here as the relation between the variance of permanent shocks and the variance of transitory shocks to inflation $\left(\frac{\sigma_\varepsilon^2}{\sigma_\eta^2}\right)$:

$$\frac{\sigma_\varepsilon^2}{\sigma_\eta^2} = S = \frac{-1}{corr(\Delta\pi_t, \Delta\pi_{t-1})} - 2 \tag{16}$$

The solution for the optimal gain parameter in steady state is obtained from the the error correction expression (12):[48]

$$K = \frac{-S + \sqrt{S^2 + 4S}}{2} \tag{17}$$

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 11 in the Appendix.

From equation (17), 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.[49]

[48] Equation (17) 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.

[49] 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, this link is re-assessed allowing for regime shifts in inflation dynamics. Lastly, 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 influence from monetary policy changes.

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´s 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.[50] 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. 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.[51] 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

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

[51] 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). *Dolarization persistence and individual heterogeneity*. Mimeo, January.

[10] De la Rocha, Javier (1999). The transmission mechanism of monetary policy in Peru. In "*The transmission mechanism of monetary policy in emerging markets*", 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] Hansen, Bruce E. (1992). The likelihood ratio test under non-standard conditions: testing the Markov switching model of GNP. *Journal of Applied Econometrics*, Vol. 7, Issue Supplement: Special issue on Nonlinear Dynamics and Econometrics, December S61-S82.

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

[23] 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.

[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, pp452-479.

[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.

Appendix

Table 1: Subperiod Correlations of Mean and Variance of Inflation, Peru, 1949-2006[a]

Length of subperiods in years	Correlation
1	0.5908
2	0.6092
3	0.6840
4	0.6664
5	0.7773
6	0.6510
7	0.6850
8	0.9936
9	0.7151
10	0.9449

[a] The period 1985-1991 is excluded.

Table 2: Correlations of Level and Squared Change in Inflation, Peru, 1949 - 2006

Length of horizon in quarters (x)	Correlation
1	0.260
2	0.483
3	0.490
4	0.577
5	0.484
6	0.513
7	0.554
8	0.496
9	0.620
10	0.603
11	0.609
12	0.617
13	0.647
14	0.655
15	0.682
16	0.689
17	0.615
18	0.705
19	0.681
20	0.694
21	0.669
22	0.684
23	0.674
24	0.648
25	0.685
26	0.623
27	0.666
28	0.635
29	0.609
30	0.647
31	0.623
32	0.630
33	0.640
34	0.623
35	0.577
36	0.695
37	0.658
38	0.611
39	0.609
40	0.649
41	0.616
42	0.573
43	0.602
44	0.575
45	0.580
46	0.598
47	0.574
48	0.565
49	0.538
50	0.510

[a] The period 1985-1991 is excluded.

Table 3: 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 4: 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 $\left(\sigma_\varepsilon^2\right)$	0.173	0.84	
		(7.617)		
	Transitory Shock $\left(\sigma_\eta^2\right)$	0.163	0.52	
		(2.003)		

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

Table 5: 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 6: 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 7: 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 8: 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 9: 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 10: Cointegration Johansen Test*
Quarterly Inflation and Money Growth Rate: 1964:1 - 2006:1

Data Trend: Test Type	None No Intercept No Trend	None Intercept No Trend	Linear Intercept No Trend	Linear Intercept Trend	Quadratic Intercept Trend
Trace	1	1	1	1	1
Max-Eig	1	1	1	1	1

*Critical values based on MacKinnon-Haug-Michelis (1999). Number of lags (13) chosen by the Akaike criterion. Selected number of cointegrating relations by model (0.05 level).

Table 11: 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.