Inflationary Dynamics and Persistence in Costa Rica: Period 1953-2009

Carlos Chaverri Morales and Carlos Torres Gutiérrez*

1. INTRODUCTION

There are various definitions of *inflation persistence* in the economic literature, however, for the purposes of this paper we adopt the definition suggested by Marques (2004), who sees it as the speed with which inflation returns to its long term equilibrium value after a disturbance.[1] High values of the inflation rate suggest high values of persistence after a disturbance, thus because inflation converges more slowly towards its underlying or long-term equilibrium, while low values of this coefficient show a rapid convergence of the variable to its equilibrium level (Álvarez, Dorta and Guerra, 2000).[2] This is consistent with the statistical definition of persistence attributable to Fuhrer (1995), who conceived it as the tendency of a

* C. Chaverri Morales and C. Torres Gutiérrez, Departamento de Investigación Económica of the Banco Central de Costa Rica. An earlier version of this paper was presented at the XIV Meeting of the Researchers Network of Central Banks of the Americas (Salvador-Bahia, Brazil, November 2009). The authors thank Laura D'Amato, Manfred Esquivel and other participants of the workshop to discuss progress of the "Joint project on inflation dynamics and persistence," coordinated by the Center for Latin American Monetary Studies (CEMLA) and the Central Bank of the Republic Argentina (Brasilia, Brazil, July 2009). Email: ⟨chaverrimc@bccr.fi.cr⟩.

[1] He quotes definitions of Batini (2002), Batini and Nelson (2002) and Willis (2003), Andrews and Chen (1994) and also Pivetta and Reis (2001), pointed out that inflation persistence defined similarly. D'Amato, Sotes and Garegnani (2007) also embrace this definition.

[2] That is, there is high (low) inflation persistence when they are very long (short) deviations of inflation from its steady state level, After the occurrence of a shock (Céspedes et al., 2003)

variable to stay away from its average level over an extended period of time, once it is disturbed.[3]

According to that definition it is possible to indicate that the persistence is conditional on the long-term equilibrium level of inflation that is assumed. Marques (2004) argues that it is not always appropriate to assume that this equilibrium level is constant over time, although it is exogenous and unaffected by the inflationary shock.

It is important to study the phenomenon of inflation persistence with regard to the economic environment and particularly to the field of action of the central banking. For economic policy in general it is relevant to understand the rigidities that underlie the persistence of inflation, because it helps to shape and guide the process of structural reform aimed at improving efficiency and economic welfare. In fact, bringing about changes in the indexation of prices and wages in the economy, from retrospective to prospective factors, liberalizing regulated markets and promoting competition in key economic sectors can result in the reduction of inflation persistence (IMF, 2009).

For the conduction of the monetary policy it is also important to estimate the level (extent) and sources of inflation persistence in the economy, because that can help into to design appropriate responses to inflationary shocks and also because it allows to improve macroeconomic modeling and inflation forecasting. When inflation persistence is high, the response of the monetary policy should be gradual and the horizon of inflation targets relatively long. This because the higher inflation persistence is associated to a greater *sacrifice ratio* (in terms of cost in product stabilizing inflation in the short term) and it is more difficult to anchor inflation expectations to the inflation target of the central bank.

However, it is also argued that inflation persistence is affected by endogenous factors, including changes in the behavior of the price fixing process in the economy resulting from policy changes and the credibility

[3] It should be noted that inflation persistence, thus defined, is different from the concept of inflation inertia. According to Lendvai (2004), inertia is the slow response of one variable to unexpected changes in economic conditions, so that past levels of the variable (or past expectations about its current level) have a direct influence on their level today. He adds that the persistence of a variable can be generated by various sources, being the only one inertia. Also, inflation persistence is considered different to inflation uncertainty, which is difficult to predict inflation because of its variability (measured from the standard deviation of forecast errors of inflation, generated by an econometric model for a certain period). Thus, a greater variance of inflation forecast errors implies higher inflation uncertainty and vice versa (Solera, 2002).

of monetary policy. When the central bank attaches great importance to achieving its inflation target, it counteracts inflationary shocks, so they come to have less impact and lower inflation persistence. In the other hand, if the central bank has a more flexible inflation target, and it is also concerned with other factors (such as growth or employment) after an inflationary shock, then the central bank has to face greater persistence (Hansson et al., 2009).

In accordance with the importance of this issue, the main objective of this paper is to study the dynamics of inflation and persistence in the case of Costa Rica. We identified relevant facts of the inflationary process in Costa Rica, modeled changes in the average rate of inflation over time, depending on the major internal and external inflationary shocks and changes in monetary policy strategy and finally we estimated the level of inflation persistence.

It should be noted that it is not objective of this paper to study the sources or causes of inflation persistence commonly mentioned in literature, such as the volatility of inflation itself, the degree of indexation of prices and wages in the economy and changes in the credibility of economic policy, including the changing nature of inflation targeting central bank policies.[4]

The document is structured as follows: The second section contains the most important conceptual issues. The third section briefly describes the methodology used. The fourth section includes a summary description of the evolution of inflation in Costa Rica. The fifth section contains the main empirical results and section six concludes.

2. CONCEPTUAL ISSUES

The issue of inflation persistence is discussed in the literature under two different approaches (Marques, 2004): a simple univariate time series representation of inflation, and a structural econometric, multivariate, approach of the behavior of this variable.

Under the univariate approach, it is usual to assume an autoregressive model of inflation, where the shocks on the residual component are

[4] The empirical study of the sources of inflation persistence in the country exceeds the objectives of this research, although a brief description of these factors, from the theoretical point of view, is provided in Annex 1 and in Annex 12. It also presents some measurements of inflation persistence sources in Costa Rica.

designed as a summary measure of all shocks that affect inflation in a given period.

The multivariate approach assumes a causality relation between inflation and its determining variables (usually the components of a Phillips curve[5] or variables in a structural VAR model), in which the shocks are structural in the sense that they are susceptible to economic interpretation, such as monetary policy shocks.

Following D'Amato et al. (2008), this paper adopted an univariate inflation approach.[6] The starting point assumes that this variable follows a stationary autoregressive process of order *p*:[7]

$$\pi_t = \alpha + \sum_{i=1}^{p} \beta_i \pi_{t-i} + \eta_t \ , \tag{1}$$

where π_t is the contemporary inflation rate; α, a constant term; β_i, the autoregressive coefficient of the inflation rate; and η_t, the random disturbance or shock.

Inflation persistence (ρ) is reflected in the sum of the significant autoregressive coefficients of order *p* in equation (1):

$$\rho = \sum_{i=1}^{p} \beta_i \ ; \ \forall\, i = 1, 2, ..., p. \tag{2}$$

Since it is assumed that inflation is stationary ($0 < \rho < 1$),[8] a shock would have a transient effect on inflation. As mentioned in the introductory

[5] Inflation expectations, real imbalances between actual and potential, real depreciation, among others.

[6] According to Marques (2004), this approach and its close substitute, which he calls low-frequency autocovariance of a series (the spectrum at zero frequency), seem able to deliver the best estimate of inflation persistence, compared with alternative measures, such as the *half-life* inflationary impact and the maximum autoregressive root. However, the univariate approach of the sum of the autoregressive coefficients is also point to certain limitations (Vladova and Pachedjiev, 2008): the likely bias of underestimation of the parameters of persistence when the latter is close to the unit, the choice of appropriate lag length, the likely overestimation of the persistence if not properly controlled or if structural changes are not considered variable in half time and the possibility of changes in persistence over time, especially over long periods.

[7] That is, it is assumed that inflation is a stationary series around its mean. Thus, a stationary representation of the behavior of inflation can be understood as the true inflation persistence (Fuhrer, 1995).

[8] Indeed $|\rho| < 1$, but according to Marques, to be a problem of interest it is suppose that ρ is not negative (if negative, a shock would have a contractionary transient impact on inflation).

section, if the effect of the shock takes a long time to disappear, inflation would be highly persistent, but if its effect is short-lived then it would exhibit low persistence.[9]

An important feature of a stationary time series is the property of reversion to the long term mean value after a shock. That is, if in the previous period the shock led to the series to be over (under) its average, in the current period it should decrease (increase) to converge to its mean. This is the basis for claiming that when evaluating the persistence of inflation, what really matters is the persistence of deviations of inflation from its mean, so, based on Marques (2004), D'Amato et al. rewrite the equation (1) as a correction mechanism to balance, as deviations of inflation from its mean value:

$$\pi_t - \mu = \sum_{i=1}^{p-1} \varphi_i \Delta(\pi_{t-i} - \mu) + \rho(\pi_{t-1} - \mu) + \eta_t \,, \tag{3}$$

where $\mu = \alpha/(1-\rho)$ is the unconditional inflation mean. Thus, under the stationary autoregressive inflation approach, the mean of the series plays the role of the long run equilibrium value of inflation to which this variable returns after a shock (Marques, 2004).

Since the inflation mean varies over time, assuming it is constant may erroneously lead to conclude that inflation is highly persistent when it is not and vice versa.[10] As we mention in the following methodological section, we deal with this problem by identifying structural breaks in the level of the series.

It is also important to note that the empirical use of the coefficient ρ as a measure of inflation persistence has the advantage of being simple to estimate and to test statistically, but according to Marques (2004),[11] it is easy to argue that this is also an abstract measure, which has difficulties of interpretation and comparison on a practical level.

An alternative measure of persistence that is easier to interpret, communicate and compare can be estimated using as a reference the value of

[9] According to Marques, it is possible to think that an integrated process has persistence unit (if the series is integrated of order one, $\rho = 1$ in that case a shock over the inflation would have permanent effects).

[10] Assuming a very flexible mean, it can be concluded that inflation is not persistent when in fact it is highly persistent.

[11] Unfortunately, according to the empirical literature it is not possible to clearly establish ranges of variation for this ratio, in which inflation persistence can be classify as *low, medium* or *high*.

ρ. This is called the *half-life* of an unexpected inflationary shock of unit magnitude, which is defined as the number of periods over which the effect of the shock value remains above 0.5 (Marques, 2004). The greater the number of periods required for half of the adjustment to takes place, the greater the degree of inflation persistence and vice versa. The total time it takes to complete the shock also gives an idea of the magnitude of the problem, because the more widespread it is, the higher the inflation persistence and vice versa.

For an autoregressive process of order one, the *half life(h)* can be computed as (Marques, 2004):

$$h = \ln(1/2)/\ln(\rho), \tag{4}$$

where *ln* denotes the natural logarithm.[12] However, it is more complex to calculate the *half-life*, for an autoregressive process of order *p*, so equation (4) is only an approximation to the *true* estimate of the concept. In this case, the *half-life* is usually estimated by simulating the effect on inflation of a unit shock on itself, this method is called an impulse response function (IRF) and it computes the number of periods that it takes to adjust the half of this shock.[13]

3. METHODOLOGY

Under the univariate time series approach adopted to study inflation persistence, we assume an autoregressive model of order *p* for the monthly change T (1,1) of the consumer price index (CPI) of Costa Rica, using as base period July of 2006 (July 2006 = 100).[14] However, for comparative purposes we also estimated the inflation persistence by using the definition of *half-life* of a shock of unit magnitude.

[12] Since $\rho^h = 1/2$ denotes the number of periods required for a unit shock to be reduced by half, applying the logarithmic transformation on both sides we have: $h\ln(\rho) = \ln(1/2)$. Solving for h it is possible to obtain the expression (4).

[13] It should be noted, however, that this alternative measure of persistence also has some disadvantages (Marques, 2004): may underestimate the persistence if the IRF declines on a swing; but monotonically decaying IRF still might not be appropriate to compare different series if one exhibits a more rapid initial decrease and a subsequent decrease slower than the other series and difficult to distinguish changes in persistence over time.

[14] Hereafter, the CPI monthly variation rate will be referenced as the monthly inflation rate; although, with the CPI high frequency data might be more appropriate named as price variation rate.

The total period of analysis covers from February 1953 to December 2009 (1953m02-2009m12). At the moment we started this, this period was the longest for which CPI official data were available from the National Institute of Statistics and Census (INEC) of Costa Rica.

Although the CPI monthly rates of change are more erratic, or have more *noise* that the inter-T(1,12), this are more informative about short-term movements in inflation, which are of interest under the approach chosen in this research. By contrast, inflation rate variations over 12 months (inflation) are more suited to analyze movements in the inflation trend, which is not the objective of this work, plus it can lead to significant overestimation of the persistence, as stated in Pincheira (2008).

3.1 Autoregressive Model of Order *p*

The methodology used in this research closely follows the research line of D'Amato et al. (2008). As a prerequisite it is important to verify that the monthly inflation rate series is stationary throughout the period of study. After that, it is necessary to cover the following stages:

i) First, determine the main structural breaks in the mean and in the autoregressive coefficient of monthly inflation. For this a preliminary qualitative analysis is carried out, which reviews the major economic events that shaped the history of inflation in Costa Rica since the early 50s. This is then complemented with a technical analysis to identify structural breaks in the inflation rate, through the study of the residuals coefficients of a recursive regression and then applying the Bai and Perron (1998 and 2003) multiple structural changes test.

The recursive analysis consists in the sequential estimation of model (1) for different sample sizes. In general, if the number of parameters of the model is $k+1$, the first sample used is of that size and then the remaining observations are added one by one until including the entire sample. In each aggregation of variables and model estimation, the predictions of the endogenous variable and the associated prediction error are calculated for the next period. This sequence of values was used to generate the coefficients and the recursive residuals. In general, if there is no structural change, the estimated parameters are kept constant and the residuals will not deviate significantly from zero when the sample increases gradually (Carrascal et al., 2001).

The Bai-Perron test allows to formalize hypotheses tests about the presence of structural breaks in stationary variables over the selected sample. The procedure offers a series of tests: the $SupF_T(K)$ test considers the null hypothesis (H_o) of no structural breaks versus the alternative hypothesis (H_A) of k breaks. The $SupF_T(l+1 / l)$ test the existence of l breaks (with l = 0.1 ...., and H_o) against the alternative of l+1 changes. The *UDmax* and *WDmax* tests prove the H_o absence of structural breaks against the alternative hypotheses H_A of the existence of an unknown number of breaks at a significance level of 1%. Both tests evaluate a F statistical for 1 to 5 breaks where the break points are selected by maximizing the total of the squared sum of residuals.[15] The selection and interpretation of the breaks are determined according to the criteria suggested by Bai and Perron: *sequential procedure supF, Bayesian information criterion* (BIC) and *Liu, Wu and Zidek* (LWZ).[16] Once the breaks are detected confidence intervals can be formulated that allow data and the errors have different distributions between the segments in which the test separates the sample, enabling to empirically compare the coincidence between the breaks detected and the actual behavior of a series.

ii) The second stage estimates the average non-constant rate of inflation (μ_t),using a model that takes into account the different subperiods in which structural changes were detected in the average or autoregressive component of this variable.

iii) The third stage calculated the deviations of the monthly inflation rate for the non-constant average value (estimated in the previous stage) in order to define a variable, $z_t = (\pi_t - \mu_t)$, that is used in estimates of inflation persistence (ρ), according to the following equation of correction of deviations of inflation (3).

(5) $$z_t = \sum_{i=1}^{p-1} \varphi_i \Delta z_{t-i} + \rho z_{t-1} + \eta_t \, .$$

3.2 *Half-life* of an Inflationary Shock

The methodology for calculating the *half-life* of an inflationary shock

[15] The UDmax test equally weighted the five F statistics, while the WDmax statistical weights F so that the marginal p-values are equal across the number of breaks.

[16] The test program is implemented in GAUSS and was obtained from the website of Pierre Perron, Department of Economics, Boston University (http://people.bu.edu/perron/).

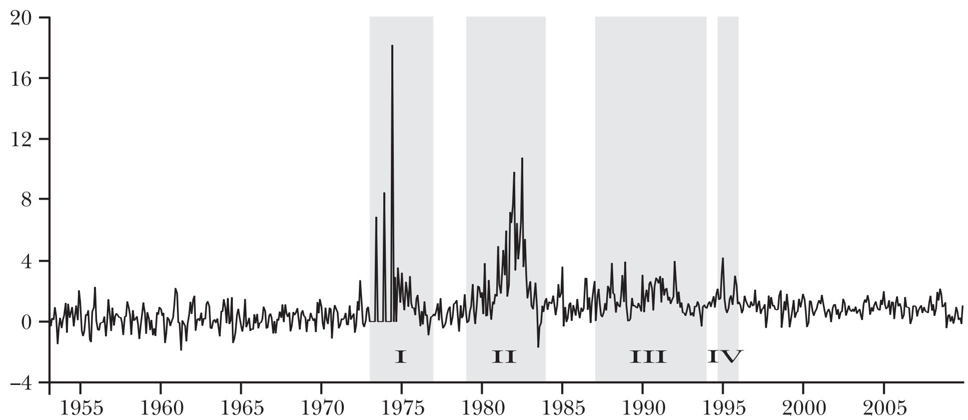
can also be performed by evaluating the stationarity of the monthly inflation rate in the period studied. Then an autoregressive model is estimated for this variable, whose initial optimal lag is determined with the FPE, AIC, LR and SC[17] criteria. This lag was subsequently adjusted according to the exclusion lag test applied sequentially to determine only those delays that are statistically significant.

Finally, after verification of the model stability and suitability of the residual of the regression, an impulse response function was estimated (IRF), by simulating the effect of a unitary shock on inflation itself and we computed the number of periods required for half of the adjustment to take place and for the rest of the adjustment to complete.

4. BRIEF HISTORY OF INFLATION IN COSTA RICA

During the 1953-2009 period inflation in Costa Rica recorded a monthly average of 0.9%. However, its behavior was highly volatile in some periods, as shown in Figure 1 and Table 1.

FIGURE 1. COSTA RICA: MONTHLY INFLATION RATE. PERIOD 1953M02–2009M12



SOURCE: Authors.

From 1953 to 1971 the monthly inflation rate was relatively low and stable, remaining at about 0.2% on average. This period was characterized by the absence of a formal process to define monetary policy rules by the

[17] Final prediction error (FPE), Akaike information criterion (AIC), sequential modified LR test statistic (LR) and Schwarz (SC).

TABLE 1. COSTA RICA: MONTHLY INFLATION RATES BY SUBPERIODS (PERCENTAGES)

Selected periods	*Mean*	*Standard deviation*
1953m02-1972m12	0.20	0.72
1977m01-1978m12	0.54	0.64
1984m01-1986m12	1.14	0.85
1994m01-1994m08	1.26	0.35
1996m01-2009m12	0.87	0.52
Periods to control		
I 1973m01-1976m12	1.39	3.02
II 1979m01-1983m12	2.54	2.51
III 1987m01-1993m12	1.42	0.91
IV 1994m09-1995m12	1.80	0.96

SOURCE: Authors.

monetary authority. Costa Rica's economy faced an external environment governed by new international monetary system and the adoption of the new exchange rate scheme called gold-dollar-standard. The exchange rate during this period was fixed and could be considered that served as anchor for monetary policy.

This relative inflation stability suffered its first major impact during the 1972-1982 period, when the monthly inflation average becomes higher and more variable. This period was characterized by global instability due to the first international oil prices crises in the mid 1970s, along with a break in the international monetary order. During these years Costa Rica experiencied an increase of capital inflows, which caused the typical problems related to the *Dutch disease:* appreciation of the currency affecting the competitiveness of exports, coupled with persistent fiscal imbalances. It is also necessary to highlight the effects as a result of the severe external debt problem that spanned most Latin American countries in the early 1980s, when average monthly inflation was 2.5% (Table 1).

During 1983-1990 inflationary problems were accentuated, dued to the exhaustion of the economic model in place since the sixties (Solera, 2002). This period is often categorized as the most inflationary one in the recent years, mainly due to adverse weather conditions, the abandonment of a fixed exchange rate (adopting a crawling peg exchange rate regime in 1983) when there was an acceleration of devaluation and higher interest rates. Additionally, the unfavorable behavior of international prices deepened the trade deficit. The average monthly inflation rate during this period was 1.3%, but with less variability than in the period described above.

Between 1991 and 2000 monthly inflation remained at 1.2% on average, a low level compared with that of previous periods, but high relative to international standards.[18] The fiscal deficit, associated to high public spending, was one of the main causes of inflation over this period, which led to an adjustment on the prices of public services, but also the accelerating pace of devaluation as well as the oil crises associated to the military conflict in the Persian Gulf and the rapid growth in money aggregates on the second half of 1994, due the intervention and posterior closure of the Banco Anglo Costarricense (see Solera, 2002) explain the relatively high inflation rates observed over this period.

With a strengthened institutional framework with clearer objectives in terms of defining an inflation goal by monetary authorities, the period 2001-2009 recorded an average monthly inflation rate of 0.8%. Of note during this period were the financial crisis of 2001, the introduction of a crawling band regime for the exchange rate adopted in October 16, 2006 and the recent financial crisis in mid-2008, when world production and the international prices of raw materials fell (such as oil and food), leading to a substantial fall in domestic inflation in 2009.[19]

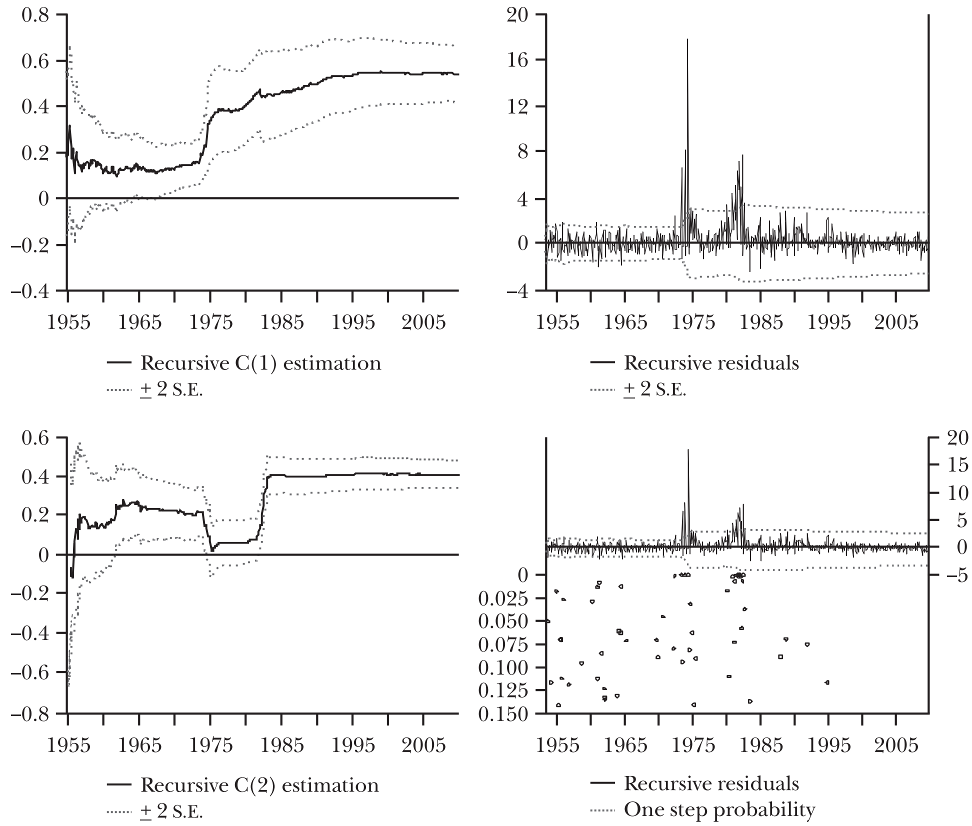
This brief characterization of the country's inflationary history helps to identify four subperiods (I to IV) of high inflation over the total period studied (Figure 2). So they must be taken into account to control the estimates of inflation persistence.

In short, these periods coincide with related shocks in international oil prices, the crisis of external debt and climate issues, and monetary and exchange rate issues typical of the Costa Rican economy. In general, they are periods that show a higher average monthly inflation and greater variability compared with other periods (Table 1).

[18] For example, the average monthly inflation rate of USA (main trade partner of Costa Rica) during this period was 0.21 per cent.

[19] The change in the CPI of December 2009 (4.0%) was 9.9 percentage points lower to that of December 2008 and the lowest since 1971. Many developed and developing economies also had disinflations because the slowdown or downturn and lower international prices of primary products (BCCR, 2010).

FIGURE 2. COSTA RICA: AUTOREGRESSIVE RECURSIVE ANALYSIS MODEL OF THE INFLATION RATE, PERIOD 1955–2009



SOURCE: Authors.

5. EMPIRICAL EVIDENCE

In this section we verify the stationarity of the monthly inflation rate and identify by technical means the main structural breaks in this variable, in order to estimate the non-constant average rate of inflation and the inflation persistence.

5.1 Degree of Integration Monthly Inflation

Table 2 contains the main results of the analysis of integration of the monthly inflation rate for the subperiods of mentioned in the previous section.

TABLE 2. COSTA RICA: ANALYSIS OF THE DEGREE OF INTEGRATION OF MONTHLY INFLATION, DICKEY FULLER F-STATISTIC, 1953-1997

Reference periods	*Constant*	*Trend*	*H_0: unitary root*
1953m02-1972m12	Not significant[a]	Not significant[a]	Rejected[a]
1977m01-1978m12	Not significant[a]	Not significant[a]	Rejected[b]
1984m01-1986m12	Significant[a]	Not significant[a]	Rejected[a]
1994m01-1994m08	Significant[a]	Significant[a]	Rejected[b]
1996m01-2009m12	Significant[a]	Not significant[a]	Rejected[a]

SOURCE: Authors.
[a] 1% significance. [b] 5% significance. [c] 10% significance.

It can be concluded that the monthly inflation rate has major breaks, but has no unit root, which is an important prerequisite to estimating inflation persistence; it ensures that the variable has no divergent or explosive behavior after the occurrence of a disruption, but has the property to revert to its mean.[20]

5.2 Recursive Analysis

To evaluate the presence of possible structural changes in the mean (constant) and in the autoregressive coefficient of monthly inflation, we estimate equation (1) recursively for the period 1953-2009.

As shown in the graphs on the left side of figure 2, there is evidence of instability of the estimated regression coefficients for both the coefficient associated with the constant, C(1), as for the autoregressive term, C(2). In the case of the constant, there is evidence of a break in the average inflation in the mid 1970s, while for the autoregressive coefficient breaks are seen in both the date as the early 1980s. Both structural changes in the monthly inflation coincide, in that order, with the first oil shock of the mid 1970s and the debt crisis of the early 1980s.

The analysis of the regression residuals stability in the first figure on the right side of Figure 2 shows that the errors are widely separated from zero and exceed the confidence band at 95% probability during the mid-1970s and early 1980s, which confirms the result mentioned above.

In the second figure on the right hand side, corresponding to one step ahead recursive residuals, the bottom shows the values of the probabilities

[20] According to the results of Dickey Fuller stationarity of monthly inflation is verified both in mean and variance.

for each of the points where the constancy of the coefficients can be rejected at various probability levels. It can be seen that the odds of record associated with the observations of the mid 1970s and early 1980s are zero.

In short, the recursive analysis concludes that no evidence of structural change in the monthly inflation series on those dates.

5.3 Evaluation of Multiple Breaks in the Inflation Rate (Bai-Perron test)

The above analysis is complemented with the study of the presence of multiple structural breaks in the monthly inflation rate, using the test developed by Bai and Perron (1998 and 2003). Two version of the test are used: breaks in the middle and breaks in the mean and the autoregressive coefficient for this variable.

As in D'Amato et al., we begin by testing the possibility of up to five structural breaks in mean inflation during the whole period, but only three are significant, according to the Bayesian information criterion (BIC), considered one of the most robust statistical tests of Bai-Perron. For this reason, applying the test again, restricting the number of breaks to three.

Both the *SupF* Sequential Procedure and again BIC confirm three significant structural breaks in mean inflation: May 1973, January 1983 and January 1997 (Table 3).

The dates of the first two structural breaks have been identified in the previous analysis (first oil shock and external debt crisis). The structural break January 1997 is a reflection of the sharp slowdown in inflation during the previous year (annual inflation in 1996 was nearly nine percentage points lower than in 1995), largely as a result of macroeconomic adjustment process and structural reforms in the Costa Rican economy in 1995-1996[21] (BCCR, 1997).

When testing breaks in the mean and in the autoregressive coefficients of monthly inflation rate, a Bai-Perron test for five breaks and then for three breaks suggest two significant structural changes at the discretion BIC: September 1974 and February 1996.

[21] In fact, restrictive economic policy pursued by the government in power in 1995-1996 resulted in a slower growth in domestic demand and lower inflationary pressures. Delgado (2000) states that heavily restrictive policies were the manifestation of political cycles in economic management.

TABLE 3. COSTA RICA: BAI-PERRON TEST FOR MULTIPLE STRUCTURAL BREAKS, 1953-2009

Change in mean (1953m02-2009m12)				
Specification				
z=1	q=1	p=0	h=100	m=3
Test				
SupFt(3)	UDmax	WDmax	$SupF_t(3/2)$	
35.71[a]	90.56[a]	90.56[a]	16.68[b]	
Number of selected breaks				
Sequencial	BIC	LWZ		
3	3	2		
T1	T2	T3		
1973m05	1983m01	1997m01		
Change in mean and recursive coefficients (1953m02-2009m12)				
Specifications				
z=2│3	q=2	p=0	h=136.6	m=3
Test				
$SupF_t(3)$	UDmax	WDmax	$SupF_t(3\|2)$	
44.58[a]	92.42[a]	92.42[a]	31.00[a]	
Number of selected breaks				
Sequencial	BIC	LWZ		
3	2	1		
T1	T2			
1974m09	1996m02			

SOURCE: Authors.
[a] Significant at 1%. [b] Significant at 2.5%.

The first of them had been previously detected by the test for changes in means only and corresponds to the effect of the oil shock of the mid-seventies.

The second break is near the breakdown previously identified by the test in January 1997 and may reflect delayed effects of high monetary expansion in late 1994,[22] due to the closure and liquidation of the country's oldest bank, which required funding from the Central Bank of Costa Rica (BCCR) in an amount close to 1.5% of nominal GDP in that year (Azofeifa and Rojas, 2000). The following events during 1996 also could have explained the break (BCCR, 1996): persistence of high fiscal

[22] Total liquidity increased during the year at an annual rate of 23.1 per cent.

deficits,[23] high level of imported inflation (5.8% in 1995),[24] increase in the pass-through of exchange rate to prices, increase by 10% to 15% sales tax (in September 1995) and high inflation expectations for the following periods.[25] Mention should also be made of an institutional change occurred in 1995, when in November of that year the new Organic Law of the BCCR was passed, and the beginning of a period of disinflation in the economy relatively widespread.

5.4 Estimation of the Average Non-constant Inflation in the Whole Period

In this section we estimate a non-constant mean for the monthly inflation rate for the total period (1953m02-2009m12), which includes major structural breaks occurred during the period. To do so, we combine elements of qualitative analysis of Costa Rican inflation mentioned in Section 4 and the technical results provided by the recursive analysis and the test for multiple breaks of Bai Perron in sections 5.2 and 5.3. This allows to define a number of subperiods in the evolution of the monthly inflation rate, starting with an initial period of subinflation (0.2%), which comprises 1953m02 to 1973m04 and in which prevailed a fixed exchange rate. It also identify five subsequent periods (denoted $d1$ to $d5$), each delimited by the previously discussed major inflationary shocks (Figure 3):

- $d1$: 1973m05-1974m06. First shock in international oil prices.
- $d2$: 1974m07-1980m12. Transition period.
- $d3$: 1981m01-1983m01. External debt crisis, high fiscal expansion, balance of payments problems, high pass-through of devaluation to prices.

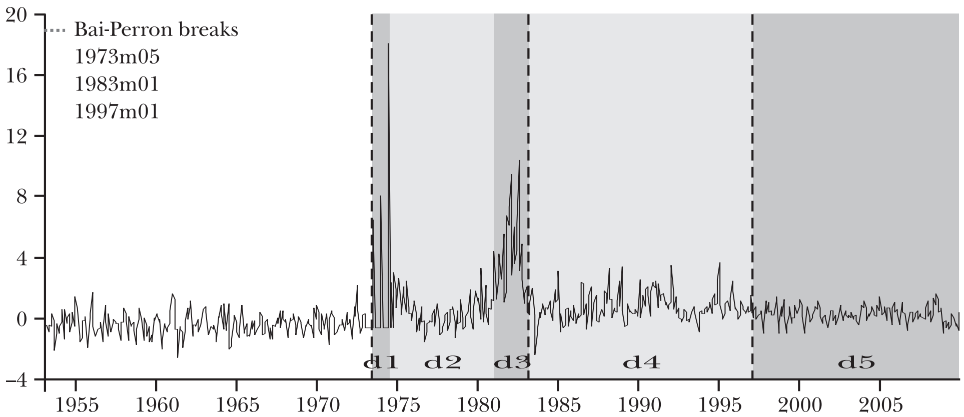
[23] For example, the fiscal gap in 1994 was estimated at 6.9% of GDP.

[24] In fact, in the period 1994-1996 several events were reported that affected the inflation rate in a group of Latin American countries, which were renegotiating their debts and stabilization programs focused on reducing inflation (Capistran and Ramos-Francia, 2007, pp. 7).

[25] Inflation expectations in 1995 were higher than 1994, mainly due to the following facts (BCCR, 1997): *i)* the year began with significant price increases, *ii)* the prospects of further devaluation, since the country had the resources of the PAE (Programa de Ajuste Estructural) III, *iii)* delay in the approval of the Tax Adjustment Act and the agreement with the International Monetary Fund; *iv)* persistence of high interest rates in the financial market.

- *d*4: 1983m02-1997m01. Introduction of a crawling peg regime, high pass-through, adverse weather conditions (Hurricane Jeanne), second shock in international oil prices (during the Persian Gulf War), the financing of losses due to the bankruptcy of BAC.
- *d*5: 1997m02-2009m12. Most recent period, which shows less variability in monthly inflation and includes amendments to the foreign exchange regime (crawling band was adopted in October 2006) and the international financial crisis contributed to lower domestic inflation 2009.

FIGURE 3. COSTA RICA: SUB PERIODS OF THE EVOLUTION OF THE MONTHLY INFLATION RATE, PERIOD 1953M02–2009M12



SOURCE: Authors.

Like Marques (2004) and D'Amato et al. variables dummy are considered that identify the average rate of inflation in each of the subperiods mentioned, which are incorporated in a regression whose dependent variable is the monthly inflation rate contemporary. The estimation of this regression is shown below.[26]

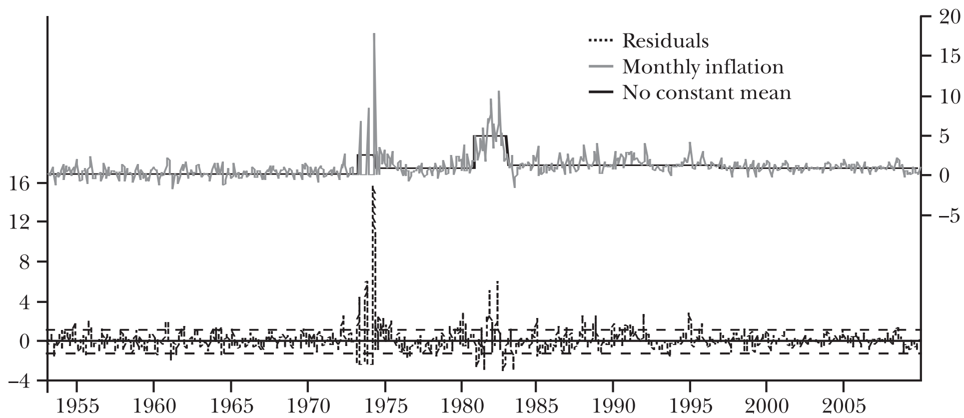
$$\underset{(HCSE)}{\pi_t} = \underset{(0.0449)}{0.2012} + \underset{(0.9828)}{2.1888}\,d_1 + \underset{(0.1765)}{0.7640}\,d_2 + \underset{(0.6320)}{4.4284}\,d_3 + \underset{(0.1090)}{1.1214}\,d_4 + \underset{(0.0680)}{0.6450}\,d_5 \tag{6}$$

[26] The model is estimated using ordinary least squares (OLS) with robust standard errors according to the covariance estimator of Newey and West (heteroskedasticity and autocorrelation consistent covariance-HAC-), which is consistent in the presence of heteroskedasticity and autocorrelation of unknown form (Annex 8.2).

The dummy variables (d 1 to d 5) correspond to the predetermined subperiod.[27.]

As expected, there is evidence of serial correlation in the residuals of the estimated equation for model (6). As in Marques (2004), lags of inflation were included in to overcome the problems of autocorrelation (Annex 3). Moreover, all coefficients maintained their significance and values did not change significantly, according to Wald test of coefficient restrictions (Annex 4). As stated by the author, these results can be seen as evidence that the time-varying (deterministic) mean implicit in (6) and represented by the stepped line at the top of Figure 4, is consistent with the data.

FIGURE 4. COSTA RICA: ESTIMATION OF THE AVERAGE NON-CONSTANT RATE OF MONTHLY INFLATION, PERIOD 1953M02–2009M12



SOURCE: Authors.

According to the model (6), the constant (0.20%) corresponds to the average monthly inflation in the early period, when fixed exchange rate and low inflation prevailed. However, the average inflation increases sharply during the first oil shock (2.39%),[28] the debt crises (4.63%) and the period in which significant internal and external shocks (adverse climatic factors, bankruptcy financing of BAC, high pass through and the Persian Gulf crisis, among others) were experienced (1.32%). During the transition period between the inflation shocks the average inflation reduces

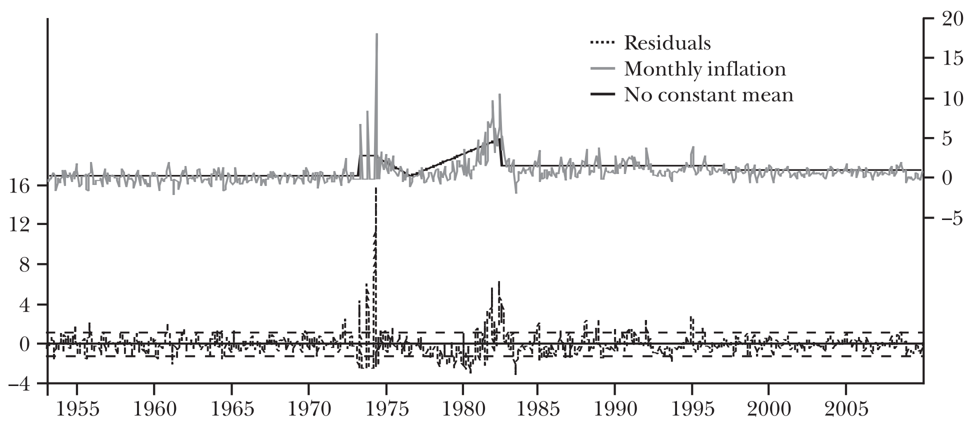
[27] Each dummy takes value one during the corresponding period and zero otherwise.

[28] The estimated mean value in each subperiod must be added the value of the finds of regression.

to 0.97%, while in the last period (1997 to 2008) the average monthly inflation was estimated at 0.85 per cent.[29]

Additionally, as suggested by the graphical analysis, we assessed the presence of deterministic trends in the transition period between high inflation of the mid-70s and the debt crisis in the early 80s. The first is a downward trend (t_1), covering the disinflation after the first oil shock (1974m07-1976m08) and the second is a growing trend (t_2),which accounts for the acceleration of inflation that led to the debt crisis of the early 80s (1976m09-1982m07) (Figure 5).

FIGURE 5. COSTA RICA: ESTIMATION OF THE NON-CONSTANT MEAN MONTHLY RATE OF INFLATION DETERMINISTIC TRENDS, PERIOD 1953M02–2009M12



SOURCE: Authors.

Lags of the inflation were again added to the calculation of variable mean to control for autocorrelation and the coefficients also maintained their significance, but Wald coefficients restriction tests revealed significant changes in mean values in most periods. Hence, we decided to keep results of the model (6) for purposes of calculating the time-varying average *(μ_t)*.

5.5 Estimated Inflation Persistence in the Total Period

As mentioned in Section 2, to evaluate the persistence of a time series what really matters is the persistence of deviations from the level of the se-

[29] For comparison this ratio amounts to 10.6% in annualized terms. However, it discourages the annualized monthly rates, to make the restrictive assumption that the rate will be maintained for a full year.

ries around their mean value, so we then calculated the deviations of the monthly inflation rate from μ_t, denoted as z, as shown in the model (5). In the econometric estimation of this model, the coefficient ρ is then the estimate of persistence when considering changes in average inflation (Table 4):[30]

TABLE 4. COSTA RICA: ESTIMATED INFLATION PERSISTENCE FOR CONSTANT MEAN AND CHANGES IN MEAN, PERIOD 1953M02-2009M12

	Constant mean	*Changes in mean*
Persistence (ρ)	0.78	0.18
NW-HAC[a]		(0.09)
Lags	1 a 2 y 6	1 y 3 a 5

SOURCE: Authors.
[a] Robust standard error as Newey y West covariance estimator (heteroskedasticity and autocorrelation consistent –HAC– covariances).

Under the assumption of a constant mean throughout the period, monthly inflation would be a highly persistent process (0.78),[31] with a significant lag up to six months. This result for Costa Rica is in line with international evidence for some Latin American countries (Annex 7). However, if it is recognized that there are changes in the mean of inflation, the estimated persistence is significantly reduced to 0.18, with a significant lag up to five months. This low rate of persistence may be, however, subject to the limitation mentioned in footnote 9, which refers to the possibility of changes in inflation persistence over time, especially when considering long periods. In particular, it can be influenced by the initial extended period of low inflation, in which no internal or external shocks of importance occurred and in which Costa Rica enjoyed great stability of prices, even higher than that observed in many developed and development (Delgado, 2000). For this reason, it is considered relevant to estimate inflation persistence over a more recent period that excludes the initial period and the subperiods following large external shocks and inflationary shocks of importance.

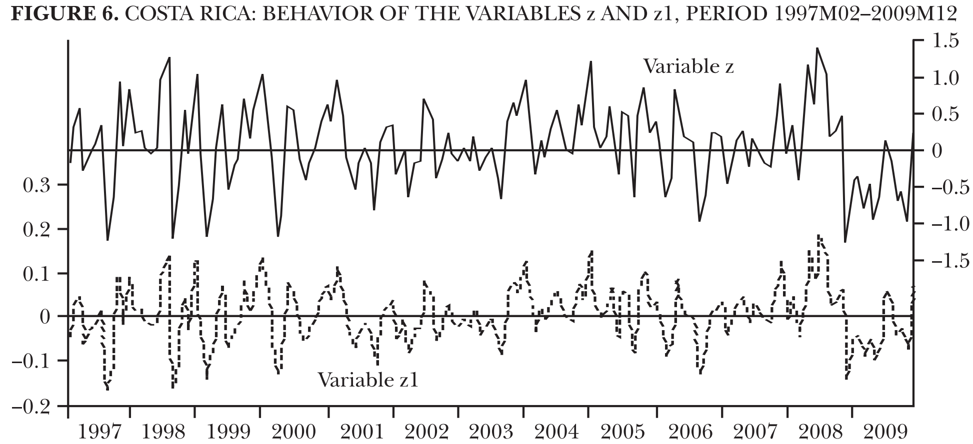
[30] The model was also estimated by OLS with robust standard errors as NW-HAC (Annex 5).
[31] To estimate the persistence failing to acknowledge changes in the average monthly inflation over the period, there are the autoregressive coefficients in the econometric estimation of equation (1) (Annex 6).

5.6 Estimation of Persistence in the Recent Subperiod

This section examines the persistence of inflation in the recent subperiod (1997m02-2009m12), both assuming that the average inflation rate (0.85%) represents the long-run equilibrium level of this variable in that period, such as using inflation targeting has been setting the BCCR at every opportunity as a characterization of the long-term level.[32]

As before, we begin by examining the requirement of stationarity of the deviations of the monthly inflation from its average level (variable *z*) and for the inflation of the BCCR from its average level (variable *z*1),[33] which is expected from the simple observation of the behavior of the series (Figure 6) and confirmed by the ADF test, Phillips-Perron and KPSS (Annex 8). This result is important because it confirms that the variables maintained the property of mean reversion.

FIGURE 6. COSTA RICA: BEHAVIOR OF THE VARIABLES z AND z1, PERIOD 1997M02–2009M12



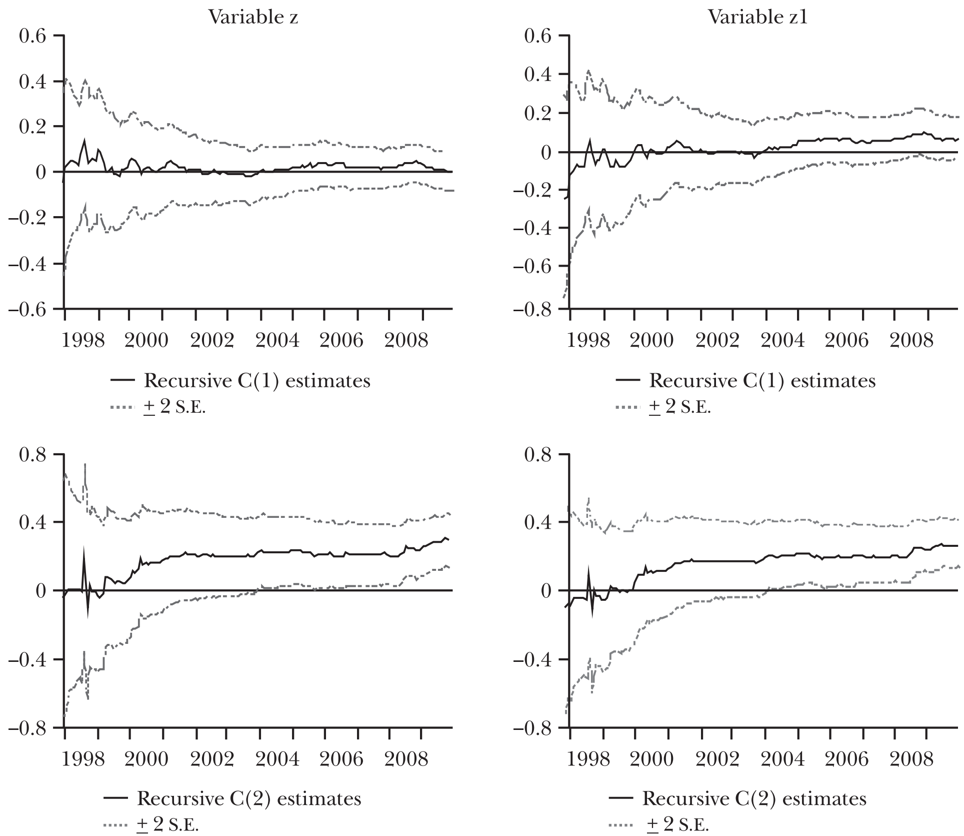
SOURCE: Authors.

[32] The use of inflation targeting central bank as a characterization of equilibrium level of long-term variable is a usual practice in inflation persistence estimates for countries with inflation targets. Although the BCCR does not operate under this monetary regime; in 2005 the BCCR authorities announced their will to adopt it in the future and for this has been preparing its monetary policy strategy, including the announcement and seeks forward-looking inflation targeting, consistent with the preparation of a medium-term macroeconomic program.

[33] To calculate the variable *z*1, monthly inflation rate is annualized and the inflation target has been established by the BCCR along this subperiod, which is expressed in annual price variations, is substracted. The monthly time series of the BCCR inflation target is taken from the work of Castrillo et al. (2008) and updated in this investigation.

According to the recursive analysis, to add data to estimate the coefficients does not produce significant variations and the confidence limits become narrower, ruling out the possibility of a structural change in the average or autoregressive coefficient of *z* and *z*1 in this subperiod (Figure 7).

FIGURE 7. COSTA RICA: ESTIMATED RECURSIVE COEFFICIENTS AND N-STEPS AHEAD PROBABILITY FOR THE VARIABLES z AND z1, PERIOD 1997M02–2009M12



SOURCE: Authors.

The main Bai-Perron statistical test suggests no breaks in the mean or the autoregressive coefficient of both variables when sequentially testing of up to five structural changes in these variables.

It is interesting to note that neither recursive coefficients nor the Bai-Perron test detected significant changes in the average or autoregressive component of *z* and *z*1 in connection with the adoption of the exchange rate band regime in October 2006. While after the adoption of this regime changes have been documented in the pass-through of the policy rate to other market interest rates (Durán and Esquivel, 2008) and changes in the pass-through of devaluation to prices (BCCR, 2009), these changes do not appear to have yet altered the data generating process of monthly inflation data as far the Bai-Perron test detected a significant structural break due to this new policy. The above evidence can be explained by the limited data available since the adoption of the new exchange rate system. Actually, there is structural change in the average inflation rate but not in the autoregressive coefficient, when the results are recalculated for the 2003m10-2009m12-subperiod, which takes into account the same number of observations before and after adoption of band exchange rate regime in October 2006 (Annex 9).

5.6.1 Autoregressive Model of Order p

Since there is no structural change in the average or autoregressive coefficient of variables (z) and ($z1$), the likelihood of spurious minimized results if one considers the constant mean of 0.85% implicit in the variable (z) is the inflation targets contemplated in $z1$ and the model (5) is directly used to estimate the persistence in this subperiod (table 5).[34]

The persistence of inflation in the recent subperiod is higher, estimated at between 0.31 and 0.42, with a maximum lag of 62 months in the case of the variable z and 19 months for the variable $z1$. Although this range of values is not considered high, the long lags show a slow rate of return of inflation to its average value or the value of long-run equilibrium

TABLE 5. COSTA RICA: ESTIMATION OF THE DEVIATIONS PERSISTENCE FROM DEVIATIONS OF INFLATION FROM ITS AVERAGE AND FROM THE INFLATION TARGET, PERIOD 1997M02-2009M12

	z variable	*z1 variable*
Persistence (ρ)	0,42	0.31
HCSE[a]	(0.08)	(0.08)
Lags	6,9,19,48,49,62	18,19

SOURCE: Authors.

[a] Robust standard error as Newey y West covariance estimator (heteroskedasticity and autocorrelation consistent –HAC– covariances).

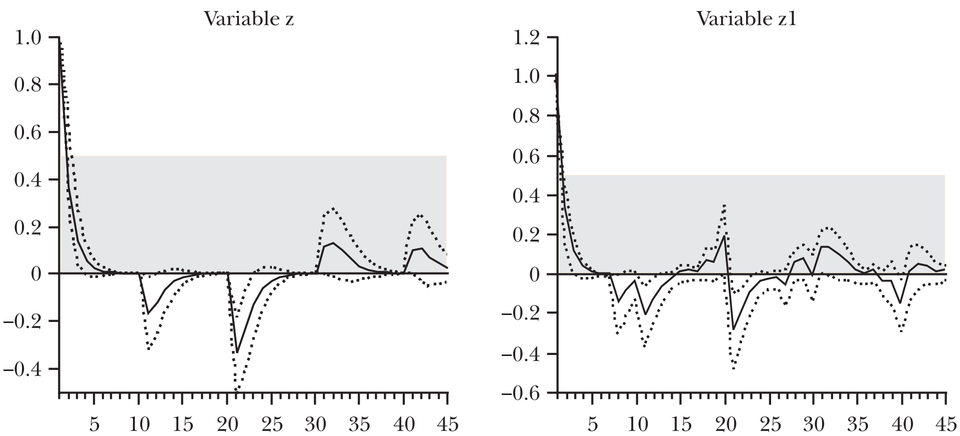
[34] OLS estimates with robust standard errors NW-SCC render have stationary, normal, homoscedasticity and not auto correlated. The models pass the specification test of Ramsey, recursive stability tests for the coefficient ρ and the test CUSUMQ (Annex 8.10).

after the occurrence of an inflationary shock, which points to inflation persistent in the last period analyzed.

5.6.2 Half-life *of an Inflationary Shock*

To complement the analysis and thus have stronger evidence, we also estimate inflation persistence according to the definition of *half-life* of an inflationary shock unit. This comes two autoregressive models: one for the variable *z* and one for the variable *z*1. According to the econometric tests, both models are stable and have normal, non-auto correlated errors.
The simulation of these unit shocks in the initial month ($t = 1$) shows the following impulse-response functions (figure 8):

FIGURE 8. COSTA RICA: IMPULSE-RESPONSE FUNCTION TO A UNIT INFLATIONARY SHOCK, PERIOD 1997M02–2009M12



SOURCE: Authors.

The *half-life* of both shocks is estimated at one month, as of the shock 50% is extinguished after that time. The total effect of the disruption is completed after 22 months for *z* and 40 months for *z*1 (table 6), affecting most persistent annualized monthly rates of change in this second case.

Although the *half-life* of a shock dissipates quickly, the fact that the shock takes a considerable time to extinguish in both models also indicates persistent inflation in the latter subperiod.

When using the MMPT to estimate the *half-life* , the inflation persistence is higher: 10 to 11 months and the shock takes about 33 months to exhaust (Annex 11), which could be due to the use of annual price changes and feedback effects implicit in the functional interrelations that

TABLE 6. COSTA RICA: ESTIMATION OF INFLATION PERSISTENCE ACCORDING TO DEFINITION OF *HALF-LIFE* OF A UNIT INFLATIONARY SHOCK, PERIOD 1997M02-2009M12

	Variable z	*Variable z1*
Half life (*h*)	1 month	1 month
Total adjust	22 months	40 months

SOURCE: Authors.

NOTES: The optimal lags for both models initially set from 1 to 20 months, according to final prediction error statistics (EPE), Akaike information criterion (AIC) and sequential modified LR test statistic (LR). The statistically significant lags finally reduced to 1, 10 and 20, in the case of the *z* model and 1, 7, 10, 19 and 20 for the *z*1 model, according to the sequential lag exclusion test.

make up the model. In any case, the results are not entirely comparable, because of the multivariate approach, the different sample size (data from 1991) and the frequency of the data (quarterly observations).[35]

6. CONCLUDING REMARKS

The purpose of this study was to estimate inflation persistence, understood as the speed with which the monthly inflation rate returns to its long run equilibrium value of long term after a shock for Costa Rica in the period 1953-2009. Under a univariate approach of inflation and a stationary autoregressive process for this variable, it is assumed that the equilibrium value is the average of the variable throughout the period.

According to the empirical evidence, if the average rate of inflation does not change in the whole period studied, inflation is highly persistent (0.78), which is consistent with the evidence for some countries in Latin America. However, recognizing that this value breaks depending on structural changes and internal and external inflationary shocks faced by the economy, the estimated inflation persistence in the total period (0.18) reduces significantly, although this result can be due to the initial extended period of low inflation.

When studying the most recent period (1997-2009), the estimate of inflation persistence is greater (it ranges between 0.31 and 0.42) and there is no evidence of structural changes in the average monthly inflation when crawling band regime was adopted in October 2006.

[35] In the Annex 8.12 there are indirect measures of inflation persistence available in the BCCR.

It is only possible to detect a structural break due to lower average inflation rate from November 2008 when considering the 2003m10-2009m12 period, although the estimate of inflation persistence does not change in statistical terms (0.47). This break is located just over two years after adoption of the band, so the contractionary impact on the average monthly inflation rate might be picking up both the lagged effect of improving the independence of monetary policy that has been gained over time, due to lower BCCR intervention on the foreign exchange market, and the impact of lower imported inflation on domestic prices in 2009 due to the economic contraction and the reduction of commodity prices in the international markets, following the global financial crisis.

By supplementing the analysis with the calculation of the alternative definition of inflation persistence as the *half-life* of a unit shock in the recent period (1997-2009), we find that 50% of the shock is completed quickly (one month after it happened), but the total adjustment takes considerable time to exhaust (22 to 40 months, according to the definition used), which also points to persistent inflation. However, if this shock is simulated in 2003m10-2009m12 period, the *half-life* does not change but the overall effect of the adjustment is considerably shorter (8 to 11 months). On the other hand, if the *half life* is calculated using the BCCR Macroeconomic Model for Projection (MMPT) data from 1991, 50% of the disruption is consumed after 11 months and the total effect takes 33 months to extinguish. These greater persistence effects would be due to feedback effects present in a multivariate approach and the use of annual price changes.

When we resorted to other evidence not directly comparable of inflation persistence available in the BCCR for different periods and calculation of alternative definitions of this phenomenon (*intrinsic persistence* and autoregressive coefficients), we obtain estimates in a wider range (0.53 to 0.93) and long periods of adjustment to inflationary shocks (22 to 33 months lag). We do not exclude, however, that these indirect measurements overestimate the degree of inflation persistence, given the use of annual price changes. In any case, as stated by Marques (2004), the reliability of any estimate of inflation persistence ultimately depends on how realistic is the long-term path of inflation assumed. In this regard, with data from February 1997 to December 2009, the univariate approach adopted in this paper as an estimated proxy for the path of inflation rate of 0.85% monthly average (10.6% annualized) and inflation target average around 10%.

Clearly, the main implication for monetary policy is that, given the dependence of current inflation not only on long run and short run determinants, but also on past inflationary shocks, the greater the cost of the disinflation for the economy and, therefore the longer should be the horizon of inflation targets to be defined.

On the other hand, the control of inflation becomes a more complex problem than simply handling a short-term policy rate, gaining significance also control the main sources of inflation persistence cited in the literature, such as the volatility of inflation itself, the mechanisms of price and wage indexation in the economy and changes in the credibility of monetary policy, including changes in the inflation target.

Finally, in a broader sense , economic policy in general may also help reduce inflation persistence, through actions aimed at promoting competition in key economic sectors and through the reformulation of pricing policies (valuation cost model and price adjustment patterns or rates of regulated goods and services, among others) and deregulation of markets.

Annex 1

Of the Main Theoretical Causes of Inflation Persistence

Commonly mentioned sources or causes of inflation persistence in the literature are:

- The volatility of inflation itself: Pincheira (2008) argues that the volatility of inflation itself contributes to the persistence, to the extent that its trend component is stochastic and present continuous variations.
- The degree of indexation of prices and wages in the economy: Álvarez et al. (2000) argues that indexation mechanisms allows to readjust items such as wages or production costs of firms pegging them to the consumer price index or the exchange rate, in order to protect prices and real wages structure in economies characterized by moderate inflation. However, this practice hinders the adjustment of prices to real shocks (inflation persistence helps) and increases the costs of reducing inflation. Also, De Gregorio (1992 and 1995) explains that inflation persistence stems from a bias in inflation expectations of economic agents, caused by the indexation of prices in the economy.

- Changes in the credibility of economic policy: Álvarez et al. (2000) cites the work of Calvo and Végh (1994), in which an adjustment program based on the exchange rate anchor has credibility problems that slows down the convergence of inflation to the target inflation rate. For its part, Pincheira (2008) mentions that in Cukierman and Leviatan (1992) persistence is explained as a problem of credibility of economic agents facing the monetary authority. According to these authors, the credibility problem ends up generating higher inflation expectations, plus the fact that the monetary authority has no control over inflation, which generates a slow stabilization process that results in persistent inflation.
- Frequent changes in the inflation target by the central bank: As noted by Marques (2004), assuming that in the medium and long term inflation is determined by monetary policy, then the long-term level of inflation corresponds to the inflation target the central bank. Thus, movements of the inflation target can be a source of inflation persistence (if the central bank changes its target, it might take time to learn about the new target, so that inflation will take longer to converge to the target compared to a fixed target).

Annex 2

Equation (6) Estimation

Dependent variable: INFLAMEN
Method: Least squares
Date: 03/01/10
Sample (adjusted): 1953m02-2009m12
Included observations: 685 after adjustments
Newey-West HAC standard errors & covariance (lag truncation = 6)

	Coefficient	*Standard error*	*t-Statistic*	*Probability*
C	0.201152	0.044858	4.484156	0.0000
d1	2.188848	0.982819	2.227111	0.0263
d2	0.763976	0.176535	4.327618	0.0000
d3	4.428448	0.631953	7.007562	0.0000
d4	1.121407	0.108985	10.28953	0.0000
d5	0.645041	0.068049	9.479128	0.0000

R^2	0.371533	Mean dependent variable	0.917584
Adjusted R^2	0.366891	S. D. dependent variable	1.443632
S.E. of regression	1.148671	Akaike info criterion	3.123835
Sum squared residue	893.2650	Schwartz criterion	3.163599
Log likelihood	−1060.790	Hannan-Quinn criterion	3.139224
F statistic	80.04473	Durbin-Watson statistic	1.807464
Prob (F statistic)	0.000000		

Annex 3

Equation (6) Estimation with Inflation's Lags

Dependent variable: π
Method: Least squares
Date: 03/01/10
Sample (adjusted): 1955m02-2009m12
Included observations: 659 after adjustments
Newey-West HAC standard errors & covariance (lag truncation = 6)

	Coefficient	*Standard error*	*t-Statistic*	*Probability*
π (−2)	0.068775	0.039476	1.742208	0.0819
π (−6)	0.299682	0.150851	1.986602	0.0474
π (−15)	−0.072482	0.038847	−1.865833	0.0625
π (−16)	−0.095905	0.041172	−2.329400	0.0201
π (−21)	−0.082722	0.031977	−2.586934	0.0099
π (−24)	0.089776	0.040958	2.191939	0.0287
C	0.145428	0.058244	2.496875	0.0128
d1	1.877183	0.799906	2.346754	0.0192
d2	0.584961	0.183127	3.194284	0.0015
d3	3.313143	0.677504	4.890219	0.0000
d4	0.939556	0.205768	4.566103	0.0000
d5	0.529849	0.120288	4.404833	0.0000

R^2	0.465249	Mean dependent variable	0.939120
Adjusted R^2	0.456157	S. D. dependent variable	1.458533
S.E. of regression	1.075605	Akaike info criterion	3.001686
Sum squared residue	748.5314	Schwartz criterion	3.083459
Log likelihood	−977.0554	Hannan-Quinn criterion	3.033384
F statistic	51.17354	Durbin-Watson statistic	1.860015
Prob (F statistic)	0.000000		

Date: 03/01/10
Sample: 1955m2-2009m12
Included observations: 659

Autocorrelation	*Partial correlation*		*AC*	*PAC*	*Q-Statistic*	*Probability*
		1	0.069	0.069	3.1923	0.074
		2	0.019	0.014	3.4246	0.180
		3	–0.001	–0.004	3.4260	0.330
		4	–0.079	–0.079	7.5380	0.110
		5	–0.065	–0.055	10.338	0.066
		6	–0.024	–0.014	10.725	0.097
		7	–0.032	–0.028	11.426	0.121
		8	–0.027	–0.028	11.898	0.156
		9	–0.011	–0.016	11.974	0.215
		10	0.012	0.009	12.078	0.280
		11	–0.014	–0.022	12.219	0.347
		12	0.002	–0.004	12.222	0.428
		13	–0.041	–0.048	13.382	0.419
		14	–0.033	–0.030	14.114	0.441
		15	0.008	0.010	14.162	0.513
		16	0.008	0.004	14.204	0.584
		17	–0.062	–0.072	16.777	0.470
		18	–0.031	–0.035	17.448	0.493
		19	0.024	0.026	17.834	0.534
		20	–0.060	–0.066	20.291	0.440
		21	0.032	0.026	20.989	0.460
		22	–0.025	–0.044	21.405	0.486
		23	–0.031	–0.031	22.054	0.517
		24	–0.017	–0.026	22.257	0.564
		25	0.019	0.013	22.501	0.607
		26	0.022	0.012	22.847	0.642
		27	0.025	0.009	23.262	0.671
		28	0.030	0.017	23.902	0.687
		29	0.019	0.011	24.147	0.722
		30	–0.006	–0.012	24.173	0.764
		31	–0.007	–0.016	24.207	0.802
		32	–0.019	–0.010	24.461	0.827
		33	–0.047	–0.044	26.017	0.801
		34	–0.026	–0.023	26.480	0.818
		35	–0.014	–0.016	26.620	0.845
		36	0.032	0.032	27.347	0.850

Annex 4

Wald Test of the Equation (6)

Null Hypothesis	*Probability*	*Decision*
H0: C(7)=0.20	0.3387	Do not reject H0
H0: C(8)= 2.19	0.6968	Do not reject H0
H0: C(9)=0.76	0.3283	Do not reject H0
H0: C(10)=4.43	0.0997	Do not reject H0
H0: C(11)=1.12	0.3768	Do not reject H0
H0: C(12)=0.65	0.3382	Do not reject H0

SOURCE: Authors.
NOTE: Chi square test.

Annex 5

Equation (5) Estimation

Dependent variable: Z
Method: Least squares
Date: 03/02/10
Sample (adjusted): 1953m08-2009m12
Included observations: 677 after adjustments
Newey-West HAC standard errors & covariance (lag truncation = 6)

	Coefficient	*Standard error*	*t-Statistic*	*Probability*
ΔZ(−1)	−0.086894	0.035998	−2.413861	0.0161
ΔZ(−2)				
ΔZ(−3)	−0.071059	0.040377	−1.759892	0.0789
ΔZ(−4)	−0.184949	0.076282	−2.424534	0.0156
ΔZ(−5)	−0.257391	0.141215	−1.822691	0.0688
Z((−1)	0.177804	0.094053	1.890465	0.0591

R^2	0.098976	Mean dependent variable	−0.000403
Adjusted R^2	0.093613	S. D. dependent variable	1.149009
S.E. of regression	1.093906	Akaike info criterion	3.024746
Sum squared residue	804.1363	Schwartz criterion	3.058111
Log likelihood	−1018.876	Hannan-Quinn criterion	3.037663
Durbin-Watson statistic	1.941354		

Annex 6

Equation (1) Estimation

Dependent variable: π
Method: Least squares
Date: 03/02/10
Sample (adjusted): 1953m08-2009m12
Included observations: 677 after adjustments
Newey-West HAC standard errors & covariance (lag truncation = 6)

	Coefficient	*Standard error*	*t-Statistic*	*Probability*
C	0.201664	0.090765	2.221840	0.0266
π (−1)	0.198398	0.096122	2.064021	0.0394
π (−2)	0.175920	0.048674	3.614263	0.0003
π (−6)	0.408231	0.159779	2.554973	0.018
R^2	0.368892	Mean dependent variable		0.923530
Adjusted R^2	0.366079	S. D. dependent variable		1.448234
S.E. of regression	1.153072	Akaike info criterion		3.128628
Sum squared residue	894.8064	Schwartz criterion		3.155320
Log likelihood	−1055.041	Hannan-Quinn criterion		3.138962
F statistic	131.1261	Durbin-Watson statistic		2.003492
Prob (F statistic)	0.000000			

Annex 7

Inflation Persistence in Other Countries

Country	*Inflation persistence*[a]
Argentina	0.8542
Bolivia	0.8787
Brasil	0.8581
Chile	0.2899
Colombia	0.7874
Ecuador	0.8364
Mexico	0.8548
Peru	0.6657
Uruguay	0.8682
Venezuela	0.7666

SOURCE: Capistrán and Ramos-Francia (2007).
[a] Sum of the autoregressive coefficients (period 1980m01-2006m06).

Annex 8

Integration Analysis (recent periods)

INTEGRATION ANALYSIS OF THE MONTHLY INFLATION, PERIOD 1997M02-2009M12

Variable	*Option*	*ADF (P-value)*	*PP (P-value)*	*KPSS (LM-Stat)*
Z	CCCT	0.0000[a]	0.0000[a]	0.119230[b,c]
	CCST	0.0000[a]	0.0000[a]	0.140111[d]
	SCST	0.1152[a]	0.0001[a]	NA
D(Z)	CCCT	0.0000[a]	0.0001[a]	0.128210[b,c]
	CCST	0.0000[a]	0.0001[a]	0.225912[b]
	SCST	0.0000[a]	0.0000[a]	NA
Z1	CCCT	0.0000[a]	0.0000[a]	0.049261[c]
	CCST	0.0000[a]	0.0000[a]	0.149769[d]
	SCST	0.0000[a]	0.0000[a]	NA
D(Z1)	CCCT	0.0000[a]	0.0001[a]	0.126139[b,c]
	CCST	0.0000[a]	0.0001[a]	0.259787[d]
	SCST	0.0000[a]	0.0000[a]	NA

SOURCE: Authors.

NOTES: For ADF y PP, H_0: series has a unit root. For KPSS, H_0: series is stationary. NA stands for non-applicable.

[a] H_0 is rejected at the 1%, 5% and 10% significance level. [b] H_0 is rejected at the 10% significance level. [c] Critical values at 1%, 5%, 10%: 0.216; 0.146; 0.119, respectively. [d] Critical values at 1%, 5%, 10%: 0.739; 0.463; 0.347, respectively.

INTEGRATION ANALYSIS OF THE MONTHLY INFLATION, PERIOD 2003M10-2009M12

Variable	*Option*	*ADF (P-value)*	*PP (P-value)*	*KPSS (LM-Stat)*
Z	CCCT	0.0000[a]	0.0000[a]	0.096327[b,c]
	CCST	0.0000[a]	0.0000[a]	0.139136[d]
	SCST	0.0000[a]	0.0000[a]	NA
D(Z)	CCCT	0.0000[a]	0.0001[a]	0.027418[b,c]
	CCST	0.0000[a]	0.0001[a]	0.029775[b]
	SCST	0.0000[a]	0.0000[a]	NA
Z1	CCCT	0.0000[a]	0.0000[a]	0.076606[c]
	CCST	0.0000[a]	0.0000[a]	0.204467[d]
	SCST	0.0000[a]	0.0000[a]	NA
D(Z1)	CCCT	0.0000[a]	0.0001[a]	0.025414[b,c]
	CCST	0.0000[a]	0.0001[a]	0.26835[d]
	SCST	0.0000[a]	0.0000[a]	NA

SOURCE: Authors.

NOTES: For ADF y PP, H_0: series has a unit root. For KPSS, H_0: series is stationary. NA stands for non-applicable.

[a] H_0 is rejected at the 1%, 5% and 10% significance level. [b] H_0 is rejected at the 1%, 5% y 10% significance level. [c] H_0 is rejected at the 10% significance level. [d] Critical values at 1%, 5%, 10%: 0.216; 0.146; 0.119, respectively. [e] Critical values at 1%, 5%, 10%: 0.739; 0.463; 0.347, respectively.

Annex 9

Equation (5) Estimation on the Recent Period

VARIABLE Z

Dependent variable: Z
Method: Least squares
Date: 03/02/10
Sample (adjusted): 2002m05-2009m12
Included observations: 92 after adjustments
Newey-West HAC standard errors & covariance (lag truncation = 3)

	Coefficient	*Standard error*	*t-Statistic*	*Probability*
DZ(−6)	0.216157	0.065080	3.321420	0.013
DZ(−9)	0.145492	0.076912	1.891673	0.0619
DZ(−19)	0.230376	0.083233	2.767852	0.0069
DZ(−48)	0.169282	0.060851	2.781919	0.0067
DZ(−49)	0.102437	0.061974		0.1020
DZ(−62)	−0.126356	0.067888		0.0662
Z(−1)	0.418851	0.078728		0.0000

R^2	0.359942	Mean dependent variable	0.005219
Adjusted R^2	0.314761	S. D. dependent variable	0.497465
S.E. of regression	0.411797	Akaike info criterion	1.136466
Sum squared residue	14.41406	Schwartz criterion	1.328342
Log likelihood	−45.27745	Hannan-Quinn criterion	1.213909
Durbin-Watson statistic	2.082335		

VARIABLE Z1

Dependent variable: Z1
Method: Least squares
Date: 01/12/10
Sample (adjusted): 1998m09-2009m12
Included observations: 136 after adjustments
Newey-West HAC standard errors & covariance (lag truncation = 4)

	Coefficient	*Standard error*	*t-Statistic*	*Probability*
DZ1(−18)	0.175474	0.62916	2.789017	0.0061
DZ1(−19)	0.315224	0.056704	5.559140	0.0000
Z1(−1)	0.310072	0.080678	3.843335	0.0002

R^2	0.225359	Mean dependent variable	0.010541
Adjusted R^2	0.213710	S. D. dependent variable	0.068594
S.E. of regression	0.060824	Akaike info criterion	−2.739844
Sum squared residue	0.492046	Schwartz criterion	−2.675594
Log likelihood	189.3094	Hannan-Quinn criterion	−2.713734
Durbin-Watson statistic	1.848883		

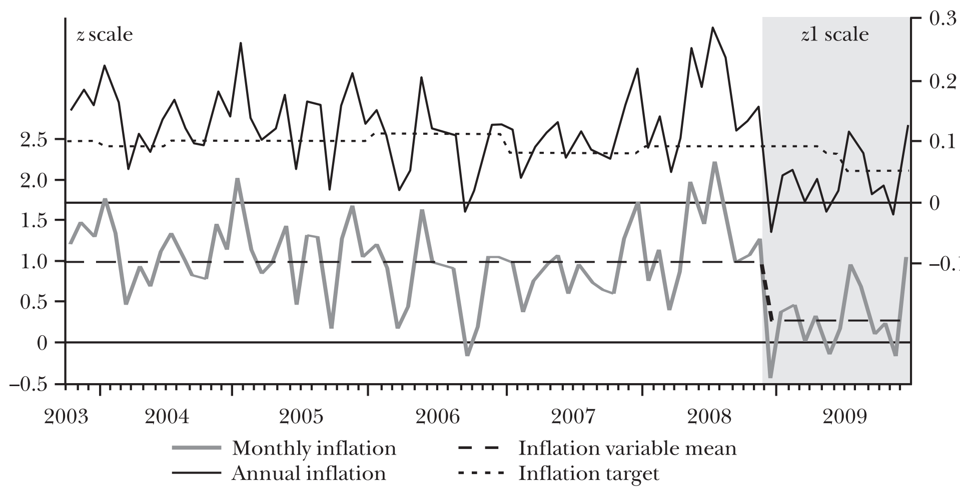
Annex 10

Inflationary Persistence in the Subperiod 2003m10-2009m12

In this annex we again test the hypothesis of structural change in the average monthly inflation after the adoption of the exchange rate band regime in October 2006. This takes into account the same number of observations before and after that date, so the new period of analysis is 2003m10-2009m12.

During this period π and $z1$ remain stationary (ADF test, PP and KPSS). Contrary to the results for other periods, recursive tests detected structural break in the mean and the autoregressive coefficient of π in mid-2008 and mid-2006 respectively, and in early 2006, both the mean and the coefficient autoregressive $z1$. Meanwhile, the BIC and statistical test LWZ Bai-Perron detect a break in mean inflation in November 2008 but not in the autoregressive coefficient (figure A.1).

FIGURE A.1. COSTA RICA: BEHAVIOR OF THE VARIABLES z=(π-μ) AND z1=(π_ANNUALIZED-TARGET), PERIOD 2003M10–2009M12



SOURCE: Authors.

This break is located a little after two years from the adoption of the crawling band regime, so that the contractionary impact on the monthly inflation rate could be picking up both the lagged effect of improving the independence of monetary policy that has been gained over time, due to lower BCCR intervention on the foreign exchange market,[36] as the impact of lower imported inflation on domestic prices in 2009 (due to problems with internal and external economic activity, following the international financial crisis), lower inflation expectations, nominal depreciation and lower pass-through to prices and improved supply of agricultural products (BCCR, 2010).

The variant monthly inflation average for the periods 2003m10-2008m11 and 2008m12-2009m12 is estimated at 1.0% and 0.27% respectively.[37] From this result, the variable z is redefined as inflation less this new average monthly variation in time. The variable $z1$ continues to define as before. With both variables, definitions of inflation persistence were recalculated (Table A.1).

The Wald coefficient restriction test, made on the basis of the evidence in table A.1, reveals that the estimated degree of persistence (coefficient

TABLE A.1. COSTA RICA: ESTIMATES OF INFLATION PERSISTENCE (PERIOD 2003M10-2009M12)

Variable z			*Variable z*[a]		
ρ[a]	*"Mean life"*[b]		ρ[a]	*"Mean life"*[b]	
	50% initial	*100% (total)*		*50% initial*	*100% (total)*
0.47 (0.12) 3/ Lag: 4,10,15,29,33	1 month	8 months	0,39 (0.10)[c] Lag: 18,19	1 month	11 months

SOURCE: Own elaboration.

[a] The OLS estimates, with robust standard errors NW-HAC, have stationary, normal, homoscedastic, and non-correlated residuals. Models passed the Ramsey specification test, recursive stability tests for the ρ coefficient and the CUSUMO test. [b] The autoregressive model for z is defined with 1, 4 and 6 significant lags, and for $z1$ with 1, 4, 6 and 8. Errors are distributed as a normal multivariate probability density, are homoscedastic and are not autocorrelated. [c] Robust standard errors according to the covariance Newey-West estimator (heteroskedasticity and autocorrelation consistent –HAC– covariances).

[36] In mid-November 2008 to mid-August 2009 the exchange rate was intermittently separating the *roof* of the exchange rate band and from that date keeps swinging freely inside without the intervention of the BCCR in market *wholesale* exchange (MONEX). This pointed to a lower variability of the money supply driven by the behavior of the exchange rate.

[37] The corresponding average inflation rates are annualized, in order, 12.8% and 3.3%. Estimates are as equivalent as in equation (6), using a variable dummy that takes a unit value during 2009m12-2008m12, and zero otherwise. As before, lags added to this equation is controlled for autocorrelation and coefficient restriction tests found no statistical changes in the estimated coefficients, which keeps its significance.

ρ) does not change in statistical terms between the subperiods studied (1997m02-2009m12 and 2003m10-2009m12) regardless of using the variable (z) or the variable ($z1$). Moreover, direct estimation of the *half-life* (50% of the initial shock) with (z) and ($z1$) indicates that the persistence does not change (remains invariant at 1 month) in the different subperiods, although the adjustment period for the remaining 50% of the shock to be extinguished reduces substantially in the 2003m10-2009m12.

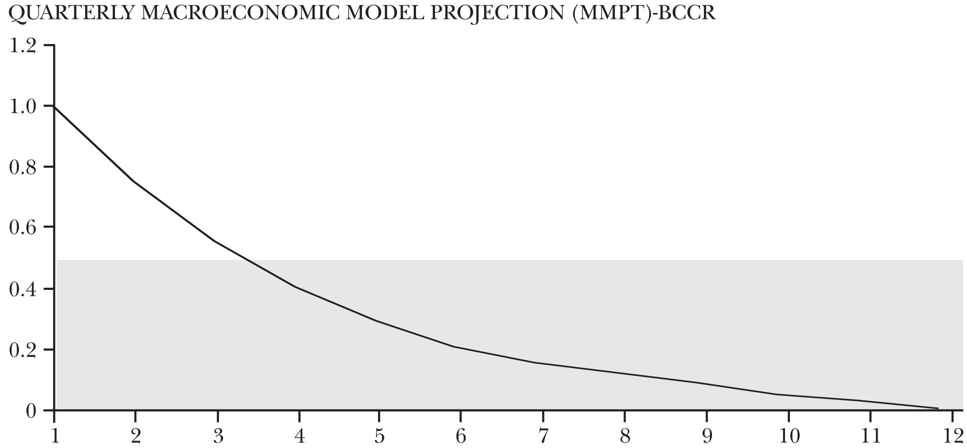
In short, structural change is detected in the average monthly inflation in November 2008, but the estimate of ρ does not change in statistical terms in this subperiod, although the adjustment period for the remaining 50% of the shock unit is extinguished is reduced, thereby reducing the persistence according to *half-life* definition.

Annex 11

Half-Life **of an Inflationary Shock Using the MMPT Model**

When the effect of an inflationary shock unit is simulated using quarterly data from 1991, covering all the functional relationships implicit in the Macroeconomic Model for Quarterly Projection (MMPT)-BCCR (Muñoz and Tenorio, 2008), *half-life* of the disruption is estimated between 10 and

FIGURE A.2. COSTA RICA: IMPULSE-RESPONSE FUNCTION OF INFLATION TO A SHOCK ITSELF. QUARTERLY MACROECONOMIC MODEL PROJECTION (MMPT)-BCCR



SOURCE: Estimation of the Departamento de Investigación Económica, BCCR.

11 months and the other half of the disruption shows that takes about 33 months to extinguish (figure A.2).[38]

Although these results are not directly comparable with the estimate of *half-life* unit shock generated with VAR models for the variables (z) and ($z1$), given the difference in methodologies, sample size and data frequency, it should be noted that most of the estimated inflation persistence is due to feedback effects implicit in the functional interrelations that make up the MMPT. Moreover, these results are consistent with the above mentioned of the annual rates of change in the CPI tend to overestimate the degree of persistence.

Annex 12

Indirect Measures of Inflation Persistence Available in the BCCR

In the international empirical literature is usual to use new hybrid models of inflation in the short term Keynesian Phillips curve type, to show three sources of inflation persistence (Whelan, 2004):

– *Intrinsic persistence* (essential): occurs when firms partially index prices of their products at prices that prevailed in previous periods. Thus, inflation in the current period is influenced by past inflation in the model of the Phillips curve.

– *Expectations-based persistence* takes place when expectations are not rational and economic agents do not have perfect information about the economy and the functional relations between key macroeconomic variables. Then, when an unexpected inflation shock occurs, they do not know if this is temporary or permanent, so that in its attempt to determine its nature and the best way they could react, they use the history of inflation to predict future inflation. This effect is reflected in the coefficient of expectations in the Phillips curve.

– *Extrinsic persistence* (not essential): arises when companies do not react at the same time by changing their product prices to changes in economic conditions. This lack of synchrony in pricing can lead to inflation becoming persistent. This effect is reflected in the coefficient

[38] The estimate of the effect the disruption took place at the Economic Research Department of the BCCR.

associated with the output gap or the real marginal cost in the Phillips curve.

To estimate these sources of persistence and other alternative means of this concept, using semi-structural multivariate models (Phillips curve implied by the MMPT) and partial (own estimate of the Phillips curve, model of pass through, and fiscal model) and as simple univariate models (ARMA models) previously estimated in the BCCR and own estimates (Table A.2).

TABLE A.2. COSTA RICA: INERTIA AND ALTERNATIVE MEASURES OF INFLATION PERSISTENCE INDIRECT APPROACHES. UNIVARIATE AND MULTIVARIATE APPROACHES TO INFLATION

Models	*Approach*	*Period*	*Inflation measurements*	*Alternative measuremets*	
New hybrid Phillips curve (MMPT)[a]	Semiestructural multivariate	1993Q1–2009Q3	12-month acumulated (quarterly data)	0.32	Intrinsic persistence
				0.57	Expectations-based persistence
				0.19	Extrinsic persistence
New hybrid Phillips curve (own estimation)[b]	Parcial multivariate	1997Q1–2008Q4	12-month acumulated (quarterly data)	0.53	Intrinsic persistence
				0.53	Expectations-based persistence
				0.89	Extrinsic persistence
Pass-through (Castrillo and Laverde, 2008)[b]	Parcial multivariate	1991m12–2007m12	Semester ahead	0.82	Inflationary inertia
Fiscal treasury bill model (Durán and Rojas, 2007)	Parcial multivariate	1996m09–2007m04	12-month acumulated (quarterly data)	0.91	Autoregressive coefficient sum
ARMA (1.1) (Muñoz, 2088)	Univariate	1996m01–2008m12	12-month acumulated (quarterly data)	0.87	Autoregressive coefficient
ARMA (2.0) (Rodríguez, 2009)	Univariate	1996m01–2009m09	12-month acumulated (quarterly data)	0.93	Autoregressive coefficient sum

SOURCE: Authors.
[a] Preliminary estimates by GMM estimates used in the recent inflation report. Explanatory variables: inflation leads and lags, inflation expectations one year lagged output gap and imported inflation. [b] Estimated by GMM own. Explanatory variables: lagged inflation, one year ahead inflation expectations and lagged output gap. Instruments: lags of the inflation and oil prices.

Again we see that in fact the coefficients are not comparable, but can give a rough idea about the maximum range of variation of these indirect measures of inflation persistence, since the rates of change of inflation used overestimate this effect. The *intrinsic persistence* is estimated at between 0.32 and 0.53 and inflationary inertia 0.82. The autoregressive coefficients, in turn, are around 0.9. Considering the current samples are observed between 0.53 and 0.93, which contrasts with the direct estimate of inflation persistence (ranging between 0.31 and 0.42) in the recent period.

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