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Working Papers

N° 2006-11

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November 2006

**Documento de Investigación**
2006-11

**Working Paper**
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# Inflation Dynamics in Latin America[1]

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**Abstract**

We analyze inflation's persistence in the 1980-2006 period for the ten largest Latin American economies using univariate time-series techniques. Although the estimated degree of inflation persistence appears to be different across countries, for the region as a whole the persistence seems to be very high. However, the estimated degree of persistence falls in all countries once we permit structural breaks in the mean of inflation. The timing of these breaks coincides with shifts in the monetary policy regimes and is similar across countries. Regardless of the changes in the mean, the degree of persistence appears to be decreasing in the region, even though for some countries persistence does not seem to be changing.
**Keywords**: Inflation, Inflation Persistence, Latin America, Monetary Policy, Multiple Breaks, Time Series.
**JEL Classification**: E31, E42, C22

**Resumen**

Analizamos la persistencia de la inflación en el periodo 1980-2006 para los diez países más grandes de América Latina utilizando técnicas de series de tiempo univariadas. A pesar de que el grado de persistencia estimado parece ser diferente entre países, para la región en general la persistencia parece ser muy alta. Sin embargo, el grado de persistencia estimado disminuye en todos los países si controlamos por cambios estructurales en la media de la inflación. Las fechas de estos cambios estructurales coinciden con cambios en los regímenes de política monetaria y son similares entre países. Sin importar los cambios en la media, el grado de persistencia parece estar disminuyendo en la región, a pesar de que en algunos países la persistencia no parece estar cambiando.
**Palabras Clave**: Inflación, Persistencia de la Inflación, América Latina, Política Monetaria, Múltiples cambios Estructurales, Series de Tiempo.

[1]We thank Daniel Chiquiar, Alejandro Díaz de León, Gabriel López Moctezuma, Antonio E. Noriega, Alberto Torres, and many seminar participants for helpful comments. Andrea San Martín provided outstanding research assistance.
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# 1. Introduction

In recent times inflation has been low worldwide and Latin America is not the exception. In the ten largest Latin American countries (LAC), consumer price inflation went from a monthly average of 6.6 percent in the 1980s to 3.8 percent in the 1990s and to 0.7 percent in the first five and a half years of the 2000s.[4] As argued by Rogoff (2003), among others, low world inflation has been most likely the result of a combination of good policies and the effects of globalization.

In view of this evidence, one question that arises is whether the inflation process has changed fundamentally, in a manner that makes low inflation more likely in the future. One way to give an answer is by looking at the dynamics of inflation, in particular at its persistence. As is well known, the degree of inflation persistence in an economy is important to model the impact of monetary policy (i.e., to assess the response of inflation to monetary conditions) and to determine the short-run trade-off between inflation and real activity.

In order to analyze the persistence of inflation, a first step is to investigate what is driving the non-stationarity of the inflation process. The fact that the level of inflation is much smaller now than in the 1980s indicates that inflation has not been stationary in the past, but whether the non-stationarity comes from a unit root or from structural changes in the mean is essential to investigate inflation's persistence. If inflation has a unit root, any shock would have permanent effects on inflation and the persistence would be infinite, whereas if it does not the persistence could be either high (but the shocks would eventually die away), or low, with shocks dying away almost immediately.

In this context, we want to assess the level of inflation persistence in LAC and to investigate if, as is the case for the mean, the degree of inflation persistence has also declined through time. Various studies conducted for the United States (Clark, forthcoming; Fuhrer and Moore, 1995; Levin and Piger, 2004; and Pivetta and Reis, forthcoming), for the United Kingdom (Benatti, 2005), for some countries of the OECD (Gadea and Mayoral, 2006), and for the euro area (Batini, 2002), show that inflation persistence seems to be high.[5] However, other studies such as Cogley and Sargent (2005) and Taylor (2000) present evidence that inflation persistence in the U.S. has fallen recently.[6] In this sense, a question that the literature on inflation persistence has been asking is: Do we need theories that generate persistence in

[4] Inflation is measured as the monthly variation of the Consumer Price Index. The ten largest LAC, ranked by the dollar amount of their Gross Domestic Products, are Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Peru, Uruguay, and Venezuela.

[5] An excellent survey of the literature on inflation persistence in the euro area and the U.S. can be found in Altissimo et al. (2006).

[6] Although Stock (2001) and Pivetta and Reis (forthcoming) argue that persistence in the U.S. has remained unchanged.

inflation as structural or have the changes in monetary policy changed inflation persistence?[7]

Our evidence comes from a sample of ten Latin American countries: Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Peru, Uruguay, and Venezuela, the ten largest economies in Latin America measured by their 2005 GDPs in dollars. These countries have gone through several dramatic changes in their fiscal, monetary, and trade policies since the beginning of the eighties. To characterize these different regimes, our approach is to use knowledge of economic events that affected these countries and its similarities, along with statistical analysis. We make particular emphasis on looking for regimes that could be characterized as the same or very similar across countries. In this regard, we find a marked decreasing tendency in the mean of inflation, captured via structural changes, and that the region seems to have moved in tandem.

Using univariate techniques, in this paper we provide evidence that in some countries the degree of inflation persistence does not seem to change through time, but that it does in others. In particular, we find that the degree of persistence does not seem to have changed during the sample period analyzed in Chile, Colombia, and Venezuela, whereas it seems to have decreased in Argentina (although it remains high), Brazil, Ecuador, Mexico and Peru, and increased in Uruguay, with mixed results for Bolivia. For the countries whose persistence does not change, we find relatively high persistence in Venezuela, medium persistence in Colombia, and low persistence in Chile. When looking at the region as a whole, we find some evidence of a general decline in persistence in recent times.

Our results on the changes in persistence are robust to the shifts in the mean of inflation, but the degree of persistence seems to be overall lower when these are taken into account. This result is in line with what has been found for other countries. For example, Altissimo et al. (2006), Clark (forthcoming) and Levin and Piger (2004) find evidence of multiple break shifts in the mean of inflation and that allowing for these lowers the estimated degree of persistence.

The paper is organized as follows. Section 2 describes the data. Section 3 presents the analysis of the behavior of the mean of inflation through time in all the countries. In Section 4 we estimate different measures of inflation persistence per-country using both time-domain and frequency-domain approaches. Finally, Section 5 presents the conclusions.

[7] See Sargent (1999); Pivetta and Reis (forthcoming); and Sims and Zha (2006), among others.

## 2. Inflation Processes in Latin America

Our data is the monthly percentage change in consumer prices in each country from January 1980 to June 2006.[8] The data is presented in Figure 1.[9] The striking heterogeneity in the levels of the series immediately catches the eye. For instance, Peru reached an inflation around 420 % per-month in 1990, whereas Colombia's highest level was 4.5 %, reached in 1985.

Four features of inflation in these countries make it, in general, difficult to uncover the process that is generating each of these series, in particular during the eighties, as well as any co-movements between them. First, inflation seems to exhibit some non-stationarities. Second, the series present high volatility for some periods of time (i.e., "volatility clustering").[10] This effect is more clear in Chile, Argentina, and Peru. Third, the variance and the level of inflation seem to be positively related.[11] Finally, in most cases both the level and the volatility of inflation present a moderately inverted-U shape: the level and volatility of inflation are relatively low at the beginning of the sample, then increase around the late eighties and early nineties, to return to low values at the end of the sample. This inverted-U pattern, however, is not very smooth. For instance, after having inflation above 14 % in 1988, Mexico had achieved levels below 2 % in the early nineties, then rebounded to levels around 8 % immediately after the 1995 crisis, to later return to levels below 1 % at the end of the sample.

Table 1 shows the mean, the standard deviation and the first order autocorrelation of inflation per-decade in our sample. The relation between the mean and the variance of inflation is clear, as is clear the so-called "great moderation" reflected in the table as the very low mean and standard deviation in each country during the last six years. In all the countries, the mean from the 1990s (already low by historical standards) further decreases in the 2000s. The same can be said about the standard deviation, with the exception of Ecuador, where it increases (but from already low levels). The average of the autocorrelations, and most of the individual estimates, are above 0.5. These are a first measure of inflation persistence. This measure also seems to have declined in the region, but the order of magnitude of this

[8] We use data that is not seasonally adjusted. The estimation of the seasonal factors in these countries for a sample that includes the eighties is complicated by the fact that the seasonal factors change markedly through time, possibly following different inflation regimes. This renders the adjustment for seasonality a task that may induce changes in the stochastic process of inflation unrelated to the objectives of this paper.

[9] The data for Ecuador starts in February 1980 instead of January since we could only get the data for the price index since January 1980.

[10] Engle's first application of ARCH was to UK's monthly inflation, precisely because of this feature of inflation data (Engle, 1982).

[11] There is an extensive literature that documents this feature of inflation for different countries. For evidence about Latin America see Daal, Naka, and Sánchez (2005).

decline is nowhere near those of the mean and the standard deviation.

# 3. Evidence on Changes in the Mean of Inflation

Concerning the apparent non-stationarity of the series, in Figure 1 we see that inflation does seem to have time-varying first and second moments, but from the graphs it is not easy to distinguish if the non-stationarity comes from a unit root or from a process that has structural breaks. Although, during some periods, it does seem that some of the processes could have had a unit root, for instance, Brazil in the late eighties and the beginning of the nineties, or Mexico during the eighties. Nonetheless, it is very difficult to think of the whole period as a unit root process in each country, as this would imply that the monetary authorities had allowed any shock to inflation to permanently affect it, and this is not in accordance with the fact that in all these countries there have been several programs to contain inflation (e.g., stabilization plans at the end of the eighties and beginning of the nineties).[12]

To investigate the possibility that each inflation series could be modelled as a series with breaks, we apply Bai-Perron tests for multiple breaks in the mean to the ten countries (Bai and Perron, 1998; 2001). The Bai-Perron methodology allows for changes in the variance through regimes, and for very general stochastic processes. The strategy we follow for each country has two steps. In the first step we look for up to four breaks in the whole sample. In the second step we look for one more break in the sub-sample starting with the last break. We do the last step in order to prevent the trimming of the sample involved in the testing procedure and the very different levels and volatilities of inflation through time to interfere with finding breaks near the end of the sample. Due to the low power of the tests, we feel that this is a sensible strategy that allows us to detect candidate break-dates that, although not necessarily statistically significant, may be useful to characterize different inflation regimes.[13]

The results are reported in Table 2, where we show all the candidate break-dates estimated by the program.[14] In bold we report those dates that are significant according to the BIC criterion.[15] With the exception of Colombia, the last break for each country corresponds to the second step in our implementation of the Bai-Perron methodology. We can see that there

[12] We do formal unit root test in the next section. The results are reported in Table 3.

[13] We are looking for a total of up to five breaks because this number coincides with the number of mayor events that may have affected inflation in the region during the period we analyze: crises in LAC or elsewhere around 1982, 1987, 1995, 1998, and 2001.

[14] One candidate break-date is not reported: April 2003 for Colombia, estimated in the second round but not significant, similar to April 2002 estimated in the first round –not significant either.

[15] Bai and Perron (2001) p. 15 indicate that BIC performs reasonably well to estimate the number of changes in the mean of a series when breaks are present.

is at least one significant break per-country, evidence that inflation is indeed better modelled as a series with different regimes. The way we arranged the dates is meant to suggest that the changes are not isolated events within one country, but that these Latin American economies seem to move through regimes in tandem.[16]

Table 2 also presents information on the sign of the change in the mean of inflation after the corresponding break. We can see a general pattern for inflation: it accelerates in the eighties and decreases, although with some jumps, from the early nineties onwards. In general, the high inflation in the region during the eighties can be seen as a consequence of fiscal dominance in a context of high external indebtness.[17] With high external debt, the repayment needed to maintain the access to external financing was a priority. The transfers to external creditors were made mainly through real depreciations brought about by inducing nominal depreciations. But, as is well known, once this mechanism was in place, the countries implementing it had to be more aggressive each time in order for the nominal depreciations to translate into real ones. In this situation of fiscal dominance, it is clear that monetary policy had objectives that were not consistent with the maintenance of low and stable inflation. This period can be called the "debt overhang" period and is characterized by large fiscal deficits.

By the end of the eighties and beginning of the nineties, all the countries in our sample renegotiated their debts, which, coupled with large revenues from the privatization of public firms in some cases, opened the door to ending fiscal dominance.[18] This was a very important first step in bringing inflation down, and indeed led to a period of lower inflation, albeit with spikes. At the same time, these countries focused on stabilization programs to bring inflation down, generally based on exchange-rate anchors.[19] The breaks associated with the beginning of the stabilization programs and debt renegotiations (third column with results in Table 2) are related to reductions in inflation in most countries, except Venezuela, where inflation increased as a result of a banking crisis in 1993.

By the end of the first half of the nineties all of these countries had achieved relatively

[16] One could even think of the breaks as common features in the series, with co-breaking combinations highly likely (Hendry, 1996).

[17] All our countries belong to the seventeen middle income developing countries that came to be recognized by the IMF as heavily indebted, and all received assistance from the IMF in some form during the eighties (Boughton, 2001).

[18] Mexico started the debt renegotiations with a Brady Plan on July 1989. Argentina, Brazil, Chile, Colombia, Ecuador, Uruguay, and Venezuela also had Brady plans since 1989. Peru reached a Brady style debt reduction agreement in 1995, whereas Bolivia re-purchased its debt.

[19] The countries that had stabilization programs based on exchange rate objectives are Argentina (1991-2001), Bolivia (1985), Brazil (1994-1998), Ecuador (1992-1998), Mexico (1988-1994), Uruguay, and Venezuela (1989-2002). Chile, Colombia, and Peru focused on an inflation target as the nominal anchor, but their arrangements involved monitoring other variables, in particular monetary aggregates and the exchange rate (Singh et al., 2005).

low inflation. But rigidities in the economies, in particular rigid exchange rates, coupled with other structural problems at least in some countries, made them very vulnerable to external shocks –e.g., sudden stops of external financing– which brought about financial crises in the region and in other emerging markets in the second part of the nineties. However, in contrast to the eighties, the movements were less synchronized across countries. In Table 2 we can see the breaks around Mexico's Tequila crisis in December 1994 (the fourth column with results).[20] In most countries inflation decreased, probably as a consequence of the stabilization programs, except in Mexico where inflation increased as a result of the crisis, and Ecuador. The breaks in the next to the last column, associated with the turbulence brought about by the crises in Asia and Russia, have a negative sign next to them, indicating that inflation decreased even further. This likely happened because the increase in inflation in some LAC brought about by those crises were not large enough or with enough duration to be detected by the structural break tests, and the tests may be detecting the reduction of inflation brought about by the added efforts to stabilize the affected LAC. The only country with a positive sign next to it in this next to the last column is Argentina, as a consequence of its crisis. It is in this context that some countries in the region turned to inflation targeting (IT).[21] The last column of Table 2 presents the breaks possibly associated with globalization and China entering the World Trade Organization (end of 2001).[22] The effects of these last breaks on inflation are to reduce inflation in most countries with the exception of Brazil, that experienced an increase in inflation (although it returns to relatively low levels after six months -the beginning of 2003), and Bolivia.

After a decade of a decreasing trend in inflation, at the end of the nineties periods of fiscal dominance seem to have been over and inflation was low in all countries by historical standards, helped in part by the benefits brought about by the latest globalization wave.

[20] December 1994 is a candidate break-date for Mexico, although it is not significant. This is a result of the way the algorithm for finding multiple breaks works, and the trimming necessary to implement it. Since in the first round the algorithm finds one break in February 1999, the algorithm cannot determine if December 1994 is significant because it is close to the end of the sub-sample between March 1988 and February 1999 after trimming the sample. We see this as supporting our decision of reporting candidate break-dates even when they are not significant, and for running a second round of the algorithm to find breaks at the end of the sample.

[21] Chile and Peru are the economies with the longest experience with IT, implementing schemes very similar to IT as early as 1991 (Chile) and 1994 (Peru). The rest of the economies implemented IT around 2000: Colombia and Brazil adopted IT in 1999, while Mexico did a first step towards IT in 1999, and consolidated it in 2001. Argentina, Bolivia, Ecuador, Uruguay and Venezuela have not implemented inflation targeting. See Corbo and Schmidt-Hebbel (2001) and Schmidt-Hebbel and Werner (2002) for excellent accounts of inflation targeting experiences in LAC.

[22] The increasing presence of China in international markets as a supplier with a comparative advantage of unskilled labor intensive manufacturing processes has increased the fall in prices of manufactured goods intensive in this type of labor and has also led to an increase in the relative prices of certain commodities. This has improved the terms of trade of commodities exporting countries.

# 4. Analysis of Inflation Persistence

It is possible, in principle, that the choice of economic policies, in particular monetary policy, could affect inflation persistence. For example, the marginal effect of inflation targeting on inflation could be to prevent shocks to affect inflation permanently. This seems to be the belief of Bernanke et al. (1999, chapter 10), who suggest, looking at the experiences of four IT countries and five non-IT countries, all of them industrialized, that IT contributes to the maintenance of price stability "... by preventing one-time shocks to inflation from permanently affecting the inflation rate.".

By estimating the degree of inflation persistence in a sample of Latin American countries we hope to provide some insights into what drives inflation and therefore into the question of whether it has changed fundamentally or not.

## 4.1. Assessing if there is a unit root in inflation

As an indication of the degree of inflation persistence we test for the presence of a unit root. In their classical analysis of inflation persistence, Fuhrer and Moore (1995), using augmented Dickey-Fuller tests, conclude that quarterly inflation in the U.S. has a unit root, and therefore that inflation is very persistent. In this subsection we investigate if this result holds for the inflation processes in the ten countries in our sample.

First we use the ADF-ERS test (Elliott, Rothenberg, and Stock, 1996), a modification of the augmented Dickey-Fuller test that has substantially improved power when an unknown mean or trend is present. We select the lags using the modified Akaike criterion (AIC) suggested by Ng and Perron (2001), a modification that further improves the size and power of the test when negative MA errors are present, as is typically the case with inflation.[23] For each country we selected an intercept or a trend or both at the 10 % significant level. The results are reported in Table 3. At 1 %, we find evidence to reject the null hypothesis of a unit root for Chile and Peru, at 5 % for Venezuela and at 10 % for Bolivia and Ecuador, although we do not find evidence (at 10 %) to reject the null for Argentina, Brazil, Colombia, Mexico and Uruguay.

It would seem that at least for five countries in our sample the apparent non-stationarity could be associated with a unit root. But the tests we have applied so far work under the

[23] We also applied the Phillips-Perron test (Phillips and Perron, 1988). We were able to reject the null hypothesis of a unit root for all the countries at 1 %. These results are interesting since the Phillips-Perron test allows for possibly heterogeneously distributed data, therefore making the results more robust to the presence of the volatility clustering. However, inflation typically exhibits a large negative moving average (MA) root. If MA errors are present in the variable of interest, the Phillips-Perron test suffers size distortions, and the authors recommend using procedures with long autoregressions, selected with one of their modified criteria.

assumption that there are no breaks in the stochastic process under consideration. The presence of breaks is not a problem if the null of a unit root is rejected, as is the case for Bolivia, Chile, Ecuador, Peru, and Venezuela, although in these countries there could still be breaks (as we already show in Section 2). The problematic case is when a test cannot reject a unit root, as this rejection could be driven by the presence of breaks. In order to see if this is the case, we apply Kapetanios' (2005) test of a unit root against the alternative of up to $m$ structural breaks to the five remaining countries.[24] The results are reported in the last column of Table 3. We reject the null of a unit root at 1 % for Argentina, Brazil, Colombia and Mexico, and at 10 % for Uruguay.[25] The conclusion from the battery of tests is that the apparent non-stationarity probably does not come from a unit root but from structural breaks, at least for the entire sample.[26] This conclusion is congruent with the fact that all the countries had dramatic changes in their economic policies throughout the sample.

## 4.2. Estimates of the persistence of inflation

Once we have discarded the possibility that the persistence of inflation in the full sample is driven by the presence of a unit root, we proceed to estimate direct measures of inflation persistence. If the series are stationary, a scalar value that can be used to measure the effect of one-time shocks on inflation is the sum of the autoregressive coefficients of an AR(p) model. There are several reasons of why this sum is a good measure to look at. First, it is positively related to the cumulative impulse response function (Hamilton, 1994). For an AR(p) stationary process, $\pi_t = \mu + \phi_1 \pi_{t-1} + ... + \phi_p \pi_{t-p} + \varepsilon_t$, with $\varepsilon_t \sim iid(0, \sigma^2)$, if we define the sum of the autoregressive coefficients as $\alpha = \sum_{i=1}^{p} \phi_i$, and if $\alpha \in (-1, 1)$, the cumulative effect of a one-time change in $\varepsilon_t$ on $\pi_t$, $\pi_{t+1,...}$ –the cumulative impulse response– is given by:

$$\sum_{j=0}^{\infty} \frac{\partial \pi_{t+j}}{\partial \varepsilon_t} = \frac{1}{1-\alpha}. \tag{1}$$

If $\alpha$ is close to one, the cumulative effect of a one-time shock is larger than unity, the shock having "second round" effects on $\pi$, whereas if $\alpha$ equals zero, one-time shocks have no "second round" effects, and the only effect on inflation is given by the the initial shock. We

[24] The methodology proposed by Kapetanios (2005) needs a unit root test and an algorithm to select the breaks. For the test we use the ADF, selecting the number of lags using the modified AIC, while for the algorithm we use Bai and Perron's (2001) procedure. We set $m = 1$ because one break was enough to reject the null of a unit root. Because $m = 1$, the test we are applying is almost identical to the one proposed by Zivot and Andrews (1992).

[25] The break-dates used in each case are: Argentina, April 1991; Brazil, June 1994; Colombia, May 1998; Mexico, March 1988; and Uruguay, December 1983.

[26] However, it is possible that the null hypothesis of unit-root could not be rejected for some sub-samples.

would expect countries with sensible monetary policies not to let one time shocks to have permanent effects on inflation.

Second, the sum of the autoregressive coefficients, $\alpha$, is also a scalar measure of the persistence of a series.[27] Since the long-run persistence properties of a series are displayed by the impulse response function –for a series with a unit-root this function never dies out, whereas it does for stationary series– the cumulative impulse response is also a measure of the long-run persistence. In this sense, series with $\alpha$ close to one are more persistent than series with $\alpha$ close to zero. We would expect countries with sensible monetary policies to have impulse response functions of inflation that die out, irrespective of the shape of the impulse response.[28]

Finally, $\alpha$ is a monotone transformation of the spectral density function at frequency zero: $S_y(0) = \frac{\sigma^2}{(1-\alpha)^2}$. The spectrum at zero frequency is a measure of the low-frequency autocovariance of the series and is proportional to the asymptotic variance of the sample mean of $\pi$. If for a particular country $\alpha$ is close to zero, then the long-run variance of inflation would be determined only by the variance of $\varepsilon$. We would expect countries with sensible monetary regimes to try to decrease the long-run variance of inflation to the minimum.[29]

To estimate $\alpha$ we first use the full sample and then estimate it using three sub-samples, one for each decade. The first covers the period from January 1980 to December 1989, the second from January 1990 to December 1999, and the third from January 2000 to June 2006. The partition is arbitrary, but we want to see how inflation's stochastic process changed from a period with fiscal dominance (roughly the 1980s) to a period without fiscal dominance and where half of the countries in our sample had adopted inflation targeting (the 2000s).

For the full sample and for each sub-sample per country we do an OLS estimation of a Dickey-Fuller re-parametrization of the AR(p) model that includes a constant and a trend:

$$\pi_t = \mu + \beta t + \alpha\pi_{t-1} + \psi_1\Delta\pi_{t-1} + ... + \psi_{p-1}\Delta\pi_{t-p+1} + \varepsilon_t, \tag{2}$$

where the coefficient attached to the first lag of inflation, $\alpha$, is the sum of the autoregressive coefficients (Dickey and Fuller, 1979). Since the data is monthly and is not seasonally adjusted we start with twelve lags and select the best model according to the modified AIC (Ng and

[27] Pivetta and Reis (forthcoming) applies this and other methods to estimate inflation persistence in the United States. They discuss the benefits and shortcomings of each measure.

[28] Although as pointed out by Andrews and Chen (1994), the cumulative impulse response is a useful measure to compare different series only if the impulse responses across the series are of the same basic shape.

[29] For example, a strict inflation targeter would try to reduce the variance of inflation to the minimum, which is given by the variance of $\varepsilon$. But a flexible inflation targeter (with considerations for output in the loss function) would try to be as close as possible to this minimum, in accordance to the considerations for output.

Perron, 2001), with the restriction that all the models should have at least $\psi_1$. Table 4 reports the estimated value of $\alpha$ for each country and sub-sample, as well as 90 % confidence intervals, calculated using Hansen's (1999) Grid-Bootstrap, using 200 gridpoints and 1000 bootstrap replications at each gridpoint.[30]

For the full sample, with the exception of Chile, all the countries seem to have high persistence, as reflect by the fact that the lower bounds of the confidence intervals (CIs) are above 0.5. If we invert these CIs, they can be thought of as a test for a unit root at the 10 % significance level, with CIs including one providing evidence of a unit root. For five countries, Argentina, Bolivia, Colombia, Ecuador and Uruguay a unit root can not be rejected, as the upper bounds for these countries are above 1, whereas four countries are very close to having a unit root (Brazil, Mexico, Peru, and Venezuela), as their upper bounds are above 0.85. These results seem to contrast with our previous conclusion, but there are two important caveats. First, these are not formal unit root tests, and the fact that the CIs include one should only be taken as evidence of very high persistence. Second, so far we have not corrected for the presence of breaks in the mean when applying this methodology.

Looking at the results from the sub-samples, the persistence seems to be decreasing through time. During the 1980s, for six countries one is inside their CIs, while this only happens for 2 countries in the 1990s and for 1 country in the 2000s. With respect to the information in the lower bounds, only 2 are below 0.5 in the 1980s and 2 are very close to 0.5, whereas 4 are below that level in the 1990s and 5 in the 2000s.[31]

For the results per-country we look at the CIs across sub-samples to see if they intersect each other. If they do not, this can be taken as evidence that a change in persistence has taken place. If this is the case we report the direction of the change. Otherwise, this can be taken as evidence that the persistence has not changed. Although we are not formally testing for a change in the persistence of inflation, looking at the information in the CIs for different samples is informative about the behavior of persistence through time.

According to this set of results, inflation persistence does not seem to change in four countries. For Argentina and Venezuela, persistence seems to be high for the whole period (0.83 is a number that can be included in all the per-decade CIs, yet there seems to be some evidence that persistence in Argentina may have decreased a little after the 1980s). Chile seems to have had low persistence throughout, around 0.25.[32] Colombia seems to have had

[30] The OLS estimates of $\alpha$ are biased (Andrews and Chen, 1994), but the confidence intervals have correct first-order asymptotic coverage as shown by Hansen (1999).

[31] These results partially support the evidence of Bernanke et al. (1999) and Levin et al. (2004) that the long-run persistence of inflation appears to decrease under inflation targeting.

[32] Chile seems to be ahead on their monetary policy with respect to the rest of the countries analyzed. In this sense, a major change in persistence may have happened between the 1970s and the 1980s.

medium persistence throughout. For four countries we find that the persistence seems to have decreased. The more dramatic changes seem to have occurred in Mexico and Peru, that went from possibly having a unit root in the eighties to numbers below 0.5 in recent times.[33] Ecuador also seems to have had a reduction in the persistence of its inflation, also close to a unit root in the eighties, but the recent number seems to be around 0.5. Brazil presents some evidence of a decline in the persistence, despite the fact that 0.7 is a number that can be included in all the per-decade CIs, the last interval is very wide and the point estimate is close to zero. There is only one country where inflation persistence seems to have increased, Uruguay, where our measure of persistence went from below 0.5 in the eighties to above 0.5 in recent times. Finally, for Bolivia we have mixed results. Bolivia stars very close to a unit root, then goes below 0.5, then up again to high levels.

The measure of persistence that we are using is very sensitive to the modelling of the deterministic component. The inclusion or not of a trend or an intercept sometimes can change the estimates of persistence dramatically. In Section 2 we reported results from structural break tests where it is clear that the mean of inflation in the ten countries has changed through time. To see how robust our results are to these changes, we estimate the persistence of inflation using the series of inflation controlling for their time-varying means. To de-mean each series we subtract the mean of each sample between structural breaks from the original series.[34] In this way we are treating the breaks as deterministic and unrelated to the persistence, in contrast to the exercise above, where these changes are treated as related to the persistence. The results for inflation without breaks are reported in Table 5. In general the measures of persistence are below those reported in Table 4, for instance no country seems to have a unit root. The qualitative results do not change for six countries, Chile, Colombia, Mexico, Peru, Uruguay and Venezuela. However, Argentina and Bolivia now seem to have a reduction in the persistence, while Brazil and Ecuador now seem to have an unchanged persistence.

One potential criticism to the results presented so far is that the division of the subsamples is arbitrary. We agree, but we also think that these results reflect what is actually happening to inflation processes in Latin America. To show that our results are not sample dependent, we estimated our measure of inflation persistence for each country using a 72 months rolling window and inflation without adjusting for the breaks.[35] We want to see how

[33] Evidence on the reduction in the persistence of inflation in Mexico can be found in Ramos-Francia and Torres (2006).

[34] This is equivalent to estimate equation (2) using dummy variables to take into account the changes in the level of inflation.

[35] For each country we estimate the Dickey Fuller re-parametrization in a 72 months rolling window using always 12 lags, a constant, and no trend. We generate a time series with the coefficients corresponding to the lagged inflation for each country.

the time series of the sums evolved over time. This strategy is in between the strategy of not controlling for the breaks at all (results presented in Table 4) and the strategy of fully controlling for the breaks (results presented in Table 5). The results are shown in Figure 2. What we found is that the point estimates are very volatile, which supports our previous decision of basing our analysis on the confidence intervals and not on the point estimates. Despite the volatility, the patterns are in general consistent with the results obtained with the estimations per-decade as shown in Table 5.

## 4.3. Estimates of the inflation spectra

One more exercise is informative about the persistence of inflation in each country: to look at estimates of the spectrum of each series not only at frequency zero, but at all frequencies. In fact, the spectrum has more information about the persistence of a series than what is reflected at frequency zero. As we have seen, the zero frequency is related to the cumulative effect of shocks on inflation (our measure of persistence), but the other frequencies give us an idea of how important is the long run effect with respect to high frequency movements. As shown in Granger (1966), the typical spectral shape of an economic variable has a hump shape, with the long run effects dominating the short run ones (i.e., most of the power is at the lower frequencies). This is indeed the case for inflation in the ten countries for the full sample if the spectra are estimated using inflation without adjusting for the structural changes in the mean (results not reported). However, as can be seen in Figure 3, when we use the data adjusted for the changes in the mean, the typical shape is not present anymore in some countries.[36] These countries are Chile, where the short-run variance seems relatively more important, Colombia and Peru, that seem to have a relatively low crossing at frequency zero but also present relatively high variance at low frequencies, and Uruguay, where the variance at all but the higher frequencies seems important.

In the same way that sensible monetary policies may try to reduce the mean, the variance, and the persistence of inflation, it is sensible to think that good policies, or other changes that have a positive effect on inflation, may affect the shape of the inflation spectrum. In particular, sensible monetary policies means that the variance of inflation is not dominated by low frequency but by high frequency components. A lower variance affects the height of the spectrum, but not necessarily its shape. A lower persistence, as we saw, implies a lower

[36] In Figures 3 and 4 the x-axis indicates the frequency (or inverted horizon) in fractions of $\pi$. For example, the zero frequency means the horizon for the spectrum is infinite periods, whereas frequency 0.5 means the horizon for the spectrum is 4 periods, frequency 0.2 means 20 periods and so on. Since the data are monthly, one period is a month. The y-axis is the spectrum estimated for each frequency, and indicates the relative importance of the components in terms of their contribution to the overall variance (Granger (1966)). The spectra are estimated nonparametrically using a tent window with a width of 25.

intercept (at zero frequency), but these again does not necessarily mean a change in the shape of the density. To see if a change in the shape has occurred we estimated the spectrum for the first ten and the last ten years in our sample (using the data without the time-varying mean) and plot the resulting graphs in the same figure for each country (Figure 4).[37]

Figure 4 shows that the estimated spectral density is lower in the second sub-sample for all countries except Ecuador. This merely reflects what we have already seen in Table 1 about the dramatic reductions in the standard deviations (except for Ecuador). However, interestingly enough, the shape has changed in some countries, in particular Argentina, Bolivia, Chile, Peru and Uruguay. Most of these changes go from the typical spectral shape to a shape that has power concentrated at higher frequencies (reducing the relative importance of the long run), except in the case of Uruguay, that shows the opposite. This evidence points to the fact that all the economic changes documented in Section 2, that had most of their effects on the mean of inflation, also changed the relative importance of different frequencies on the variance of inflation, at least on some countries.

# 5. Conclusions

The study of countries from Latin America allows us to control for the effects of large external shocks that affect the region as a whole. In this context, the changes in the mean and the persistence of inflation that we document give insights about the evolving structure of inflation in LAC.

Our main conclusion is that the persistence is high in the region and has only decreased in a few countries. In fact, we find that the main change in the inflation process over the past two and a half decades in LAC has been the decline in the mean of inflation, a result that supports the findings of Cecchetti and Debelle (2006) for a sample of industrialized countries.[38] However, the fact that the degree of persistence is, by historical standards, currently low in at least five of the ten countries analyzed contrasts with the common view that high inflation persistence is a structural part of an economy and favors the alternative that changes in monetary policy can affect inflation persistence.

In this respect, several changes in monetary policy have been used to explain changes in inflation. In particular, in Latin America the end of fiscal dominance and more recently inflation targeting have been highlighted in this regard. The results presented in this paper

[37] Since low frequencies are related to the very long run, the fact that we only have three and a half decades of data makes the estimation of the spectrum a difficult task. This has a larger effect for the spectra estimated with only ten years of data. Since $\alpha$ is a monotone transformation of the spectrum at zero frequency, this caveat also applies to our estimates of the persistence of inflation.

[38] Their sample also includes Chile.

suggest that the end of fiscal dominance is something that had profound effects on the dynamics of inflation, in particular lowering its mean and possibly its persistence. In contrast, the effects of inflation targeting are far from clear. First, because the mean and the variance of inflation have declined in all countries, and second because, although it appears that for inflation targeting countries in our sample the persistence of inflation does decline, the endogeneity involved prevents us from making strong conclusions about it. In addition, we have the case of Chile, for whom it seems that inflation persistence was already low before the adoption of inflation targeting, and the case of Ecuador, that apparently has achieved a low degree of inflation persistence without inflation targeting but with dollarization since 2000.

In this paper we have documented changes in the reduced forms of inflation for the ten largest Latin American economies. Lucas (1976) made us aware that, under rational expectations, reduced form coefficients are a function of so-called deep structural parameters. Future research should look at the implications for these deep structural parameters from the fact that reduced form equations have changed in several dimensions in these countries.

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# Appendix

We obtained the data for the consumer price index for each country from:

Argentina: Instituto Nacional de Estadística y Censos de la República Argentina.
Bolivia: Instituto Nacional de Estadística, Bolivia.
Brazil: Instituto Brasileiro de Geografia e Estatistica.
Chile: Banco Central de Chile.
Colombia: Banco de la República de Colombia.
Ecuador: Instituto Nacional de Estadística y Censos, Ecuador.
Mexico: Banco de México.
Peru: Banco Central de Reserva del Perú.
Uruguay: Instituto Nacional de Estadística, República Oriental de Uruguay.
Venezuela: Banco Central de Venezuela.

Table 1. Inflation Statistics (Monthly % Change)

| | 1980-2006 | | | 1980s | | | 1990s | | | 2000-2006:06 | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Mean | Standard Deviation | ρ | Mean | Standard Deviation | ρ | Mean | Standard Deviation | ρ | Mean | Standard Deviation | ρ |
| **Argentina** | 7.02 | 16.56 | 0.75 | 14.60 | 21.63 | 0.76 | 3.57 | 12.93 | 0.64 | 0.75 | 1.47 | 0.75 |
| **Bolivia** | 4.86 | 15.35 | 0.73 | 12.93 | 23.42 | 1.01 | 0.76 | 0.87 | 0.62 | 0.29 | 0.59 | 0.82 |
| **Brazil** | 9.48 | 14.86 | 0.67 | 11.76 | 9.41 | 0.67 | 13.10 | 20.77 | 0.64 | 0.59 | 0.54 | 0.13 |
| **Chile** | 0.99 | 1.53 | 0.31 | 1.59 | 2.10 | 0.21 | 0.89 | 1.01 | 0.58 | 0.25 | 0.37 | 0.58 |
| **Colombia** | 1.40 | 0.99 | 0.91 | 1.76 | 0.98 | 0.90 | 1.62 | 0.93 | 0.92 | 0.54 | 0.48 | 0.85 |
| **Ecuador** | 2.33 | 2.12 | 0.83 | 2.48 | 1.86 | 0.88 | 2.81 | 1.86 | 0.80 | 1.42 | 2.54 | 0.74 |
| **Mexico** | 2.29 | 2.47 | 0.93 | 4.31 | 2.79 | 0.93 | 1.51 | 1.16 | 0.92 | 0.42 | 0.33 | 0.75 |
| **Peru** | 7.10 | 25.77 | 0.30 | 10.84 | 13.83 | 0.79 | 7.98 | 39.27 | 0.20 | 0.17 | 0.34 | 0.44 |
| **Uruguay** | 2.73 | 2.38 | 0.85 | 3.83 | 2.29 | 0.83 | 2.95 | 2.32 | 0.89 | 0.74 | 0.94 | 0.78 |
| **Venezuela** | 2.20 | 2.08 | 0.85 | 1.72 | 2.52 | 0.72 | 3.14 | 1.76 | 0.95 | 1.49 | 1.00 | 0.76 |
| **Average** | 4.04 | 8.41 | 0.71 | 6.58 | 8.08 | 0.77 | 3.83 | 8.29 | 0.71 | 0.67 | 0.86 | 0.66 |

Table 2. Break Dates from Structural Change Tests [/1]

| Country | Break Dates | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Argentina** | | | **May-87** | + | **Apr-91** | - | Apr-95 | - | Dec-01 | + | **Jul-02** | - |
| **Bolivia** | | | **Jan-86** | - | Jan-91 | - | Feb-96 | - | Oct-00 | - | **Jun-02** | + |
| **Brazil** | Nov-83 | + | **May-88** | + | | | **Jun-94** | - | May-98 | - | Jun-02 | + |
| **Chile** | Mar-85 | + | | | Oct-90 | - | **Sep-94** | - | Oct-98 | - | Mar-03 | - |
| **Colombia** | | | Oct-87 | + | Jul-92 | - | | | **May-98** | - | Apr-02 | - |
| **Ecuador** | | | **Oct-87** | + | **Oct-92** | - | **Dic-96** | + | **Mar-01** | - | **Apr-03** | - |
| **Mexico** | **Dic-83** | + | **Mar-88** | - | | | Dic-94 | + | **Feb-99** | - | Jan-01 | - |
| **Peru** | | | **Sep-87** | + | **Aug-91** | - | Jul-95 | - | Jun-99 | - | Mar-01 | - |
| **Uruguay** | **Dec-83** | + | **May-88** | + | **May-92** | - | **Sep-96** | - | Feb-98 | - | | |
| **Venezuela** | | | **Jan-87** | + | **Jan-93** | + | **Dec-96** | - | Nov-00 | - | Jun-04 | - |

1/ Bai- Perron (1998) test for multiple structural breaks. The strategy that we follow for each country is to look for up to 4 breaks and to report all the candidate break-dates. Then we look for one more break for the sub-sample starting after the last candidate break-date and report the resulting candidate break-date. The trimming is always 15%. In bold are the dates selected by the BIC criterion. The last break dates for Argentina, Brazil, Bolivia, Chile, Ecuador, Mexico, Peru, Uruguay and Venezuela are obtained on the second round.

Table 3. Unit Root Tests

| Country | ADF-ERS [2/] | | $\tau_{min}$ [3/] | |
|---|---|---|---|---|
| | $H_0$: Unit Root | | | |
| **Argentina** | -2.52 | | -5.85 | *** |
| **Bolivia** | -2.62 | * | | |
| **Brazil** | -1.26 | | -48.71 | *** |
| **Chile** | -3.68 | *** | | |
| **Colombia** | -2.08 | | -5.59 | *** |
| **Ecuador** | -1.67 | * | | |
| **Mexico** | -2.45 | | -9.12 | *** |
| **Peru** | -2.98 | *** | | |
| **Uruguay** | -1.77 | | -4.67 | * |
| **Venezuela** | -2.26 | ** | | |

* p < 0.10. ** p<0.05. *** p < 0.01.
1/ Sample: 1980:01- 2006:06.
2/ Augmented Dickey-Fuller, Elliott-Rothenberg-Stock; Includes trend and intercept, best model selected with modified Akaike criterion (Ng and Perron (2001)) using up to 12 lags.
3/ Kapetanios (2005) unit root test against the alternative of up to m structural breaks. For each country we selected an intercept or a trend or both at the 10% significant level. For the tests we use ADF, selecting the best model with the Modified Akaike criterion using up to 12 lags. We set m=1. To select the break we use Bai and Perron's (2001) procedure as described in the text.

Table 4. Sum of AR Coefficients [1/]

| Country | α | Confidence Interval [2/] Lower | Upper |
|---|---|---|---|
| | **Full Sample** | | |
| **Argentina** | 0.8542 | 0.7984 | 1.0099 |
| **Bolivia** | 0.8787 | 0.8216 | 1.0241 |
| **Brazil** | 0.8581 | 0.7700 | 0.9264 |
| **Chile** | 0.2899 | 0.0883 | 0.7391 |
| **Colombia** | 0.7874 | 0.7237 | 1.0566 |
| **Ecuador** | 0.8364 | 0.7602 | 1.0206 |
| **Mexico** | 0.8548 | 0.8117 | 0.9946 |
| **Peru** | 0.6657 | 0.5534 | 0.8602 |
| **Uruguay** | 0.8682 | 0.8312 | 1.0500 |
| **Venezuela** | 0.7666 | 0.7018 | 0.8612 |
| | **1980:01 - 1989:12** | | |
| **Argentina** | 1.0177 | 0.8081 | 1.2578 |
| **Bolivia** | 1.0499 | 1.0068 | 1.1038 |
| **Brazil** | 0.5803 | 0.4980 | 0.7332 |
| **Chile** | -0.1187 | -0.4726 | 0.4144 |
| **Colombia** | 0.5859 | 0.4787 | 0.7326 |
| **Ecuador** | 0.9453 | 0.8402 | 1.0329 |
| **Mexico** | 0.8305 | 0.7544 | 1.0168 |
| **Peru** | 0.9572 | 0.8653 | 1.0880 |
| **Uruguay** | 0.1380 | -0.1103 | 0.6362 |
| **Venezuela** | 0.8685 | 0.6651 | 1.0928 |
| | **1990:01 - 1999:12** | | |
| **Argentina** | 0.7150 | 0.6053 | 0.8333 |
| **Bolivia** | 0.2887 | 0.1748 | 0.4497 |
| **Brazil** | 0.8328 | 0.6592 | 0.9451 |
| **Chile** | -0.2146 | -0.5698 | 0.5845 |
| **Colombia** | 0.6155 | 0.5241 | 0.7747 |
| **Ecuador** | 0.9858 | 0.8971 | 1.1050 |
| **Mexico** | 0.9513 | 0.8828 | 1.0162 |
| **Peru** | 0.2530 | 0.0134 | 0.6903 |
| **Uruguay** | 0.2157 | 0.0190 | 0.5520 |
| **Venezuela** | 0.7899 | 0.7152 | 0.9266 |
| | **2000:01 - 2006:06** | | |
| **Argentina** | 0.7471 | 0.6004 | 0.8318 |
| **Bolivia** | 0.8211 | 0.6640 | 0.8731 |
| **Brazil** | -0.0560 | -0.5234 | 0.8144 |
| **Chile** | 0.1846 | 0.0078 | 0.4171 |
| **Colombia** | 0.6673 | 0.5745 | 0.8872 |
| **Ecuador** | 0.5943 | 0.3833 | 0.7221 |
| **Mexico** | 0.3091 | 0.1654 | 0.5399 |
| **Peru** | 0.2847 | 0.1226 | 0.4955 |
| **Uruguay** | 0.8520 | 0.6677 | 0.9703 |
| **Venezuela** | 0.9629 | 0.8292 | 1.1265 |

1/ OLS estimation of a Dickey-Fuller reparametrization of an AR(12) model that includes a constant and a trend:

$$\pi_t = \mu + \beta t + \alpha\pi_{t-1} + \psi_1\Delta\pi_{t-1} + ... + \psi_{11}\Delta\pi_{t-12+1} + \varepsilon_t$$

where α is the sum of the autoregressive coefficients.
We start with twelve lags and select the best model according to the modified AIC, with the restriction that all the models should have at least $\psi_1$

2/ 90% confidence intervals, calculated using Hansen's (1999) Grid-Bootstrap with 200 gridpoints and 1,000 replications at each gridpoint.

Table 5. Sum of AR Coefficients (Inflation without breaks) [1/]

| Country | α | Confidence Interval [2/] Lower | Upper |
|---|---|---|---|
| | **Full Sample** | | |
| **Argentina** | 0.5815 | 0.4630 | 0.7000 |
| **Bolivia** | 0.7569 | 0.6584 | 0.8641 |
| **Brazil** | 0.3634 | 0.1417 | 0.5970 |
| **Chile** | 0.0127 | -0.2250 | 0.2519 |
| **Colombia** | 0.5826 | 0.5106 | 0.6589 |
| **Ecuador** | 0.4668 | 0.2864 | 0.6502 |
| **Mexico** | 0.7631 | 0.6968 | 0.8354 |
| **Peru** | 0.0622 | -0.0276 | 0.1138 |
| **Uruguay** | 0.2387 | 0.0505 | 0.4264 |
| **Venezuela** | 0.5260 | 0.4178 | 0.6294 |
| | **1980:01 - 1989:12** | | |
| **Argentina** | 0.6007 | 0.2622 | 0.9965 |
| **Bolivia** | 0.7649 | 0.6105 | 0.9636 |
| **Brazil** | 0.7252 | 0.6025 | 0.8607 |
| **Chile** | 0.0370 | -0.3787 | 0.4378 |
| **Colombia** | 0.5795 | 0.4612 | 0.7180 |
| **Ecuador** | 0.5987 | 0.4164 | 0.7964 |
| **Mexico** | 0.7767 | 0.6679 | 0.9049 |
| **Peru** | 0.4855 | 0.1693 | 0.6659 |
| **Uruguay** | -0.0338 | -0.4448 | 0.3088 |
| **Venezuela** | 0.4373 | 0.2519 | 0.6189 |
| | **1990:01 - 1999:12** | | |
| **Argentina** | 0.6844 | 0.5840 | 0.7951 |
| **Bolivia** | 0.2699 | 0.1258 | 0.4099 |
| **Brazil** | 0.5326 | 0.1759 | 0.9722 |
| **Chile** | 0.3585 | -0.0539 | 0.8138 |
| **Colombia** | 0.5829 | 0.4652 | 0.7097 |
| **Ecuador** | 0.4731 | -0.0908 | 0.8690 |
| **Mexico** | 0.7584 | 0.6126 | 0.9064 |
| **Peru** | 0.0402 | -0.1276 | 0.1005 |
| **Uruguay** | 0.4355 | 0.2331 | 0.6526 |
| **Venezuela** | 0.6471 | 0.5247 | 0.7746 |
| | **2000:01 - 2006:06** | | |
| **Argentina** | 0.0408 | -0.1469 | 0.1917 |
| **Bolivia** | -0.1317 | -0.3202 | 0.0476 |
| **Brazil** | 0.5899 | 0.4318 | 0.7668 |
| **Chile** | 0.3888 | 0.2131 | 0.5713 |
| **Colombia** | 0.6750 | 0.5436 | 0.8153 |
| **Ecuador** | 0.4575 | 0.3414 | 0.6310 |
| **Mexico** | 0.2967 | 0.1160 | 0.4826 |
| **Peru** | 0.2721 | 0.0981 | 0.4618 |
| **Uruguay** | 0.6827 | 0.4930 | 0.9022 |
| **Venezuela** | 0.5758 | 0.2656 | 0.9298 |

Notes: See Table 4.

Figure 1. CPI (Monthly % Change)

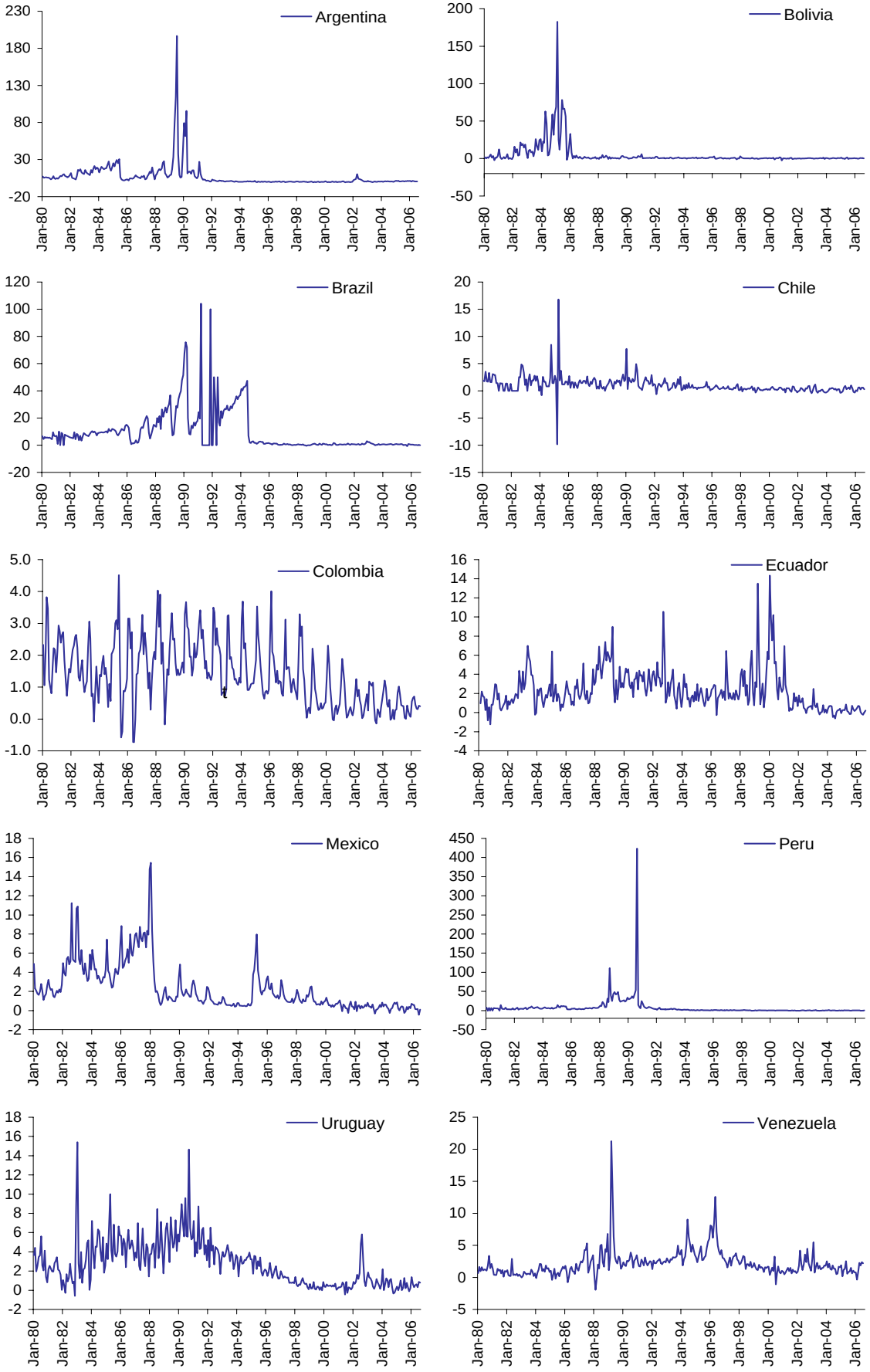
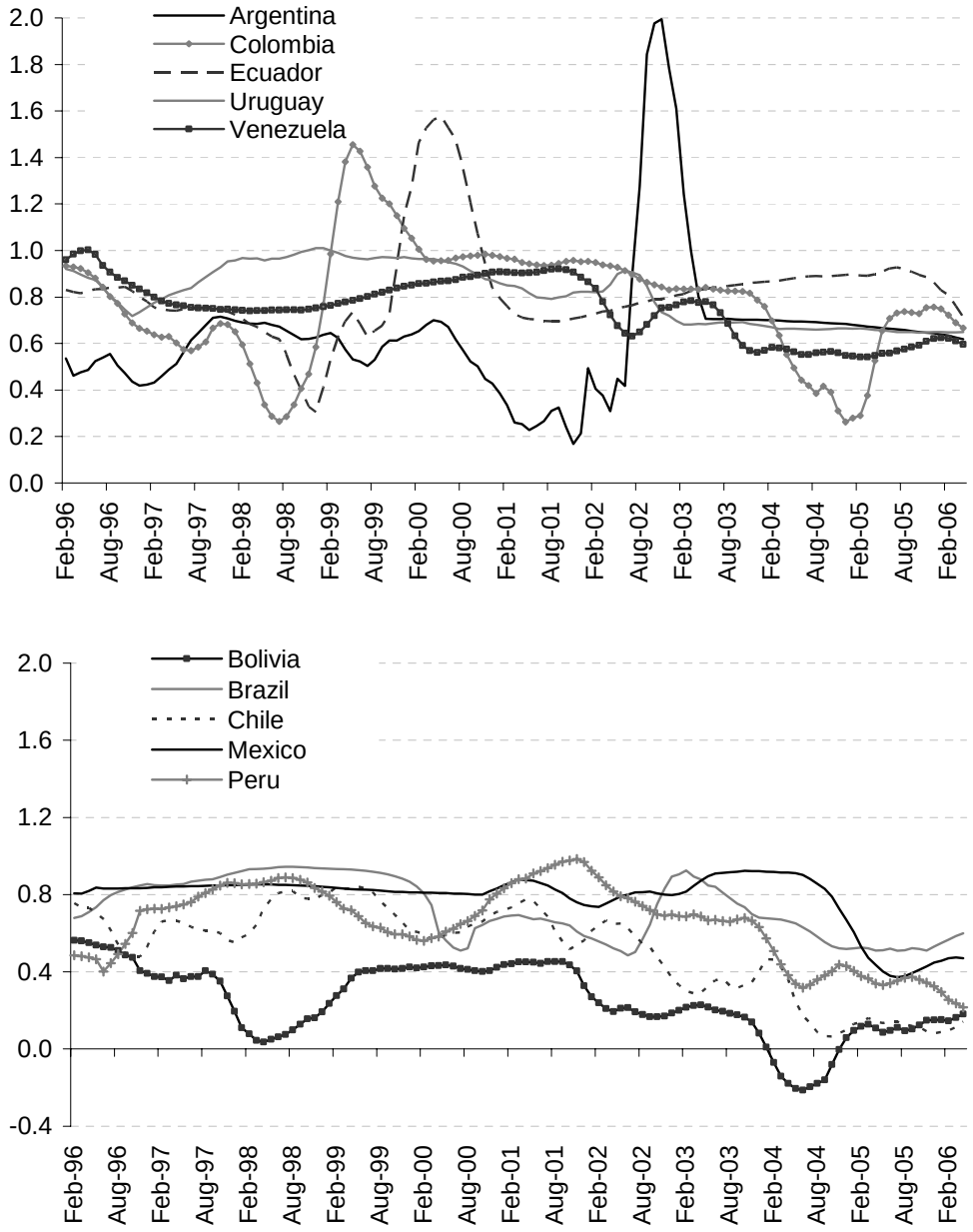
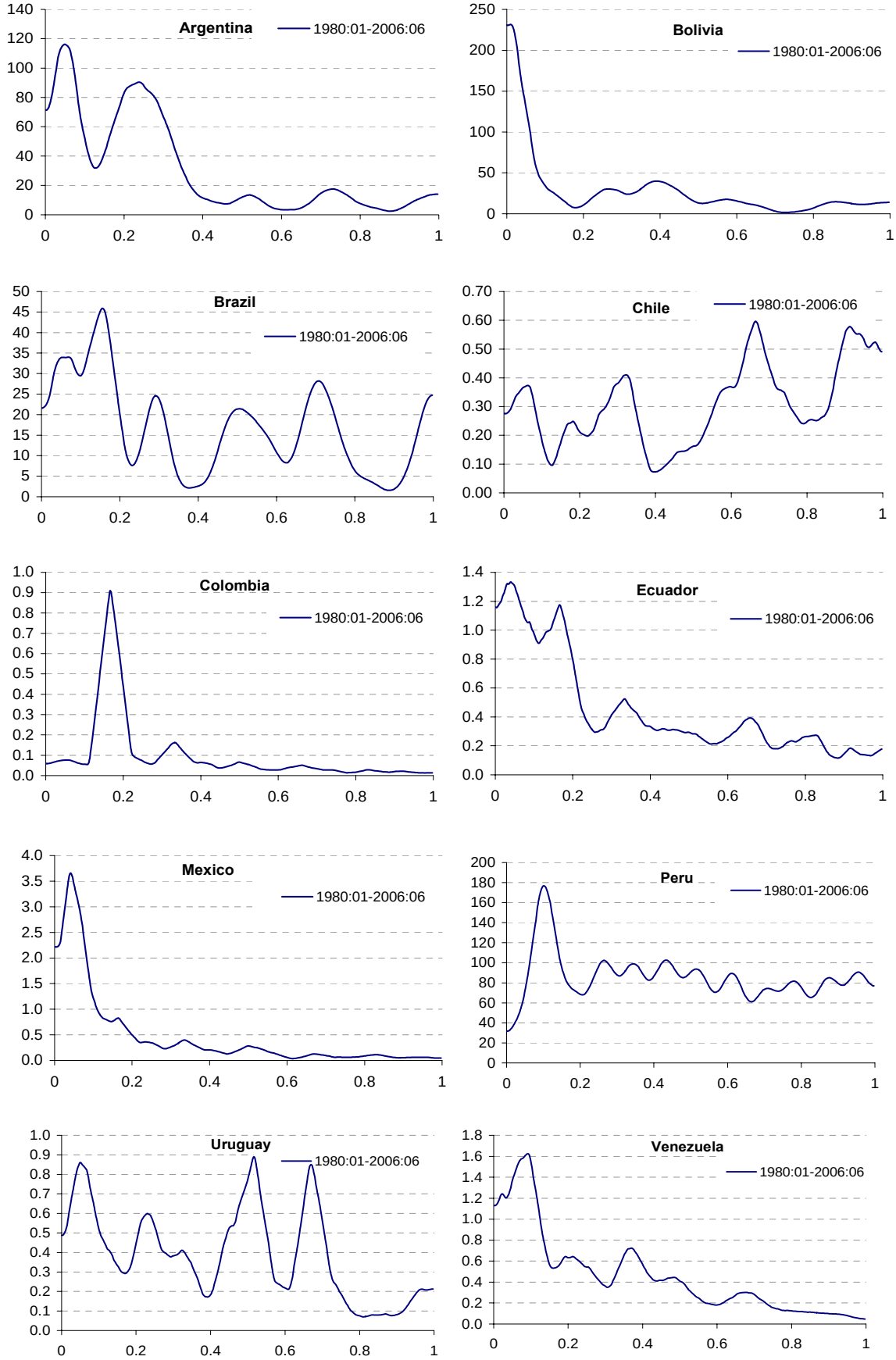


Figure 2. Sums of AR Coefficients [1/]
(Centered 5 Months Moving Average)



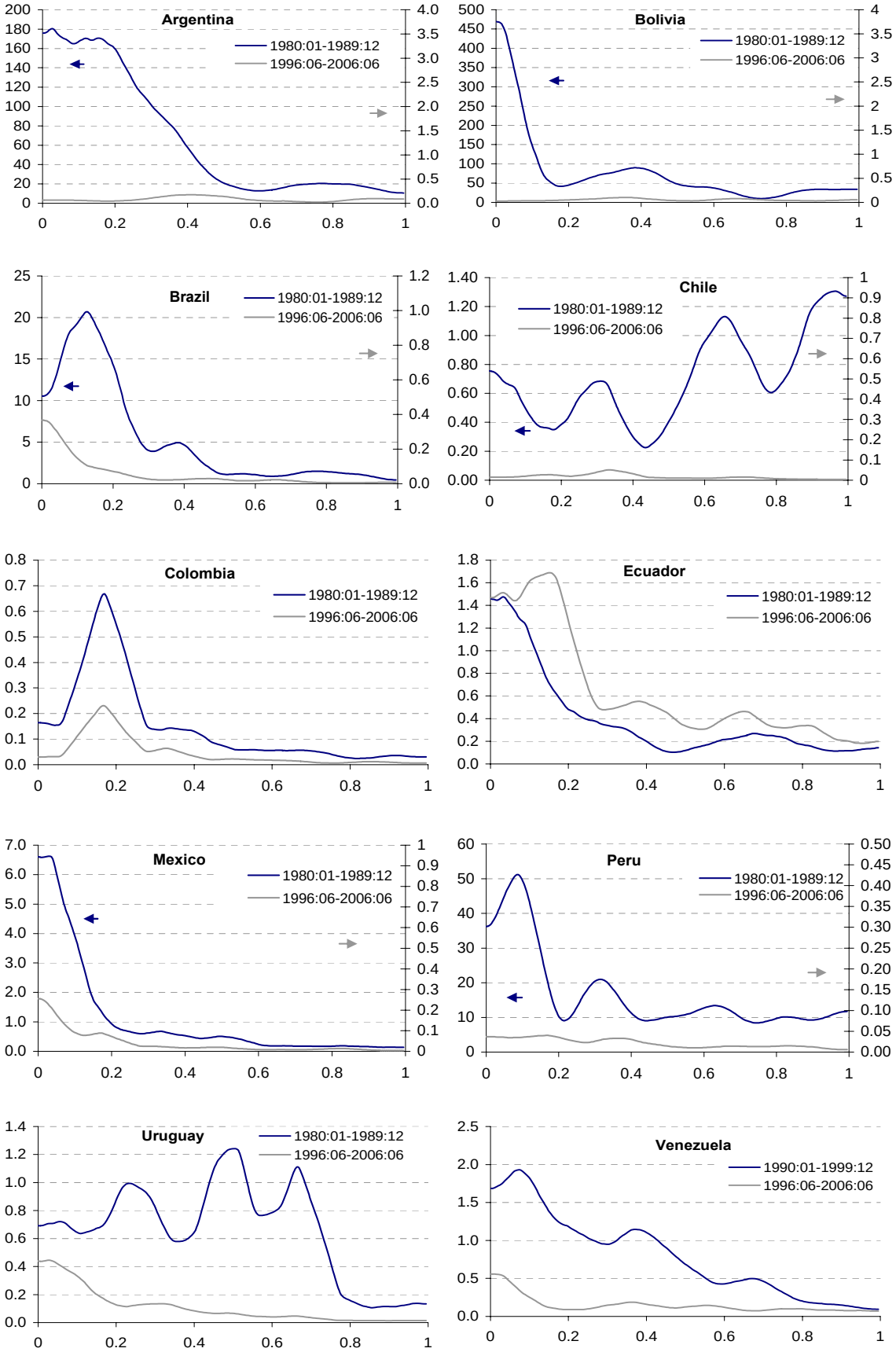

1/ Using the monthly percentage change in inflation, the strategy that we follow for each country is to estimate a Dickey Fuller re-parameterization of an AR(12) model in a 72 months rolling window. Then we generate a time-series with the coefficient corresponding to the lagged dependent variable for each country. The figure is the centered 5 months moving average of each time-series.

Figure 3. Inflation Spectrums [1/]
(Monthly % Change)




1/ Monthly inflation less mean according to all break dates. The X axis are fractions of pi.

Figure 4. Inflation Spectrum [1/]
(Monthly % Change, 1980:01-1989:12 / 1996:06-2006:06)




1/ Monthly inflation less mean according to all break date. The X axis are fractions of pi.