

STABILIZATION, PERSISTENCE, AND INFLATIONARY CONVERGENCE: A COMPARATIVE ANALYSIS

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The role of inflationary persistence has become a recurrent theme in discussions on stabilization programs (Fischer, 1986). This has particularly been the case in recent debates on the merits of exchange-rate-based stabilization plans. Some authors claim that anti-inflationary programs based on predetermined nominal exchange rates are effective in reducing inflation. Others argue, however, that in the context of stubborn inflationary inertia these type of programs are bound to generate real exchange rate overvaluation and, in some cases, can even result in major crises.¹

Discussions on the selection of exchange rate regimes in developing countries and transitional economies have also dealt with persistence issues. For example, critics of exchange rate systems based on a crawling peg—including backward-looking exchange rate bands, such as the one operating in Chile since 1984—argue that this type of system tends to generate a very high degree of inflationary inertia, which can become highly destabilizing in the presence of shocks to fundamentals. Those that favor a fixed exchange rate system, on the other hand, argue that one of the most important merits of this regime is

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1. See, for example, Bruno (1993); Dornbusch and Warner (1994); Edwards (1993).

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that it tends to generate a rapid (almost instantaneous) elimination of inflationary inertia; this helps stabilize prices at a relatively low cost in terms of output.

In his work on chronic inflationary experiences in four Latin American countries, Pazos (1990) points out that backward-looking wage indexation is bound to generate a high degree of inertia, serious destabilizing forces, and a costly and protracted inflationary process.²

More recently, there has been increasing interest in analyzing inflationary persistence in the context of staggered-contracts models of the U.S. economy. Fuhrer and Moore (1995), for example, argue that standard Taylor-type models (Taylor, 1980) fail to generate the degree of inflationary persistence observed in the U.S. data. They propose an alternative model in which wage setters care about relative real wages over the length of the contract. They show that in this case inflation demonstrates considerable persistence over time. Their empirical analysis—based on autocorrelations and impulse response functions—supports their model: in the period 1965-91 inflation in the United States did indeed experience a considerable degree of inertia. Like most studies on the subject, however, Fuhrer and Moore's paper assumes that the degree of inflationary inertia in the United States has been invariant throughout the period under study. There is no reason, of course, for this to be the case. Indeed, one would expect that as the average length of contracts and the expectations formation process change, so will the degree of inertia. A changing degree of inflationary persistence through time would indeed have important policy implications, especially regarding the output costs of disinflationary programs.³

The purpose of this paper is to analyze empirically the importance of inflationary persistence in a group of sixteen advanced and developing countries. We do this by examining the time series properties of inflation, using both univariate and multivariate techniques, on the basis of monthly data from 1970 through the end of 1996. We ask several key questions. What has been the degree of persistence, if any, in the inflationary processes? Has inflation exhibited a unit root in any of the countries in the sample? Has inertia been stable through time, or has it experienced significant changes? In those countries where inertia has varied, are the changes associated with specific policy actions, such as the implementation of exchange rate-based stabilization programs? Is the degree of inertia associated with the level of inflation?

2. For inertial models based on staggered contracts, see Calvo 1983; Fischer 1986; Taylor 1980.

3. On the costs of disinflation see, for example, Ball (1991).

Is there preliminary evidence suggesting that the degree of inflationary inertia is associated with institutional arrangements governing the existence of indexation in the labor or financial markets?

The paper is organized as follows: the next section provides a brief discussion of alternative ways of measuring inertia. Section 2 presents a preliminary data analysis. We begin the analysis by providing a brief description of the data set and explain the criteria used for selecting the sample. We then present results from our analysis of inflationary persistence using the full sample (1970-94). In section 3 we provide a detailed comparative analysis of inflationary inertia in sixteen countries. We perform a battery of tests, including rolling augmented Dickey-Fuller tests, sequential Perron tests to test the unit root hypothesis in the presence of structural breaks, and sequential Quandt likelihood ratio tests to analyze the instability of the autoregressive coefficients. We then turn to the estimation of rolling measures of persistence. For all countries we estimate rolling autoregression (AR[1]) coefficients and report rolling estimates of the dominant root from rolling AR(2) estimates. We then provide an analysis of the connection between our estimates of inertia and the levels of inflation in the countries in the sample. Section 4 provides a discussion of the evolution of inflationary inertia in three countries that at one point or another have relied on a nominal exchange rate anchor to achieve price stability: Chile, Israel, and Mexico. Our interest here is to investigate the extent to which these programs have reduced the degree of inflationary inertia. We close in section 5 with concluding remarks.

1. MEASURING INFLATIONARY PERSISTENCE: ANALYTICAL ISSUES

A central purpose of this paper is to provide simple, reliable measures of inflationary persistence for a set of countries.⁴ A large literature, both theoretical and applied, addresses the persistence displayed by economic time series. Applications include the analysis of real gross domestic product (GDP), commodity prices, and inflation in the United States. The measurement of inflationary persistence can be approached either from the perspective of a particular theoretical model or through a model-free strategy. While the model-based approach

4. In this paper, we use the terms inflationary inertia and inflationary persistence interchangeably. See the discussion later in this section for formal definitions.

has some clear advantages, including its elegance and the existence of sharp hypotheses, it also has some limitations. Obstfeld (1995), for example, has persuasively argued that model-based measures of inertia are overly restricted and are subject to misspecification problems that can result in serious biases. This paper follows Obstfeld's recommendations by analyzing the degree of inflationary inertia from a model-free perspective.⁵

In general, identifying the degree of persistence displayed by a particular time series takes place on two levels. First, we might be interested in determining whether a series is covariance stationary, given that in a stationary series the effect of a shock eventually dies out. In other words, the impulse response is zero if T is large enough. Nonstationary time series exhibit very high persistence, with all shocks qualifying as permanent ones. Second, within the class of stationary time series, some may exhibit higher persistence in the sense that the effect of the shock takes longer to wear off. Measures like the largest autoregressive root, the half-life, and the impulse-response function have been used elsewhere for this purpose. In this paper we consider two approaches to measure inertia.

There are several ways of finding out whether a series is nonstationary. Consider the following univariate process:

$$y_t = \mu + \alpha(L)\varepsilon_t \quad \text{and} \quad (1)$$

$$\alpha(L)\varepsilon_t = \varepsilon_t + \alpha_1\varepsilon_{t-1} + \alpha_2\varepsilon_{t-2} + \dots, \quad (2)$$

where

$$\sum_{j=0}^{\infty} |\alpha_j| < \infty,$$

the roots of $\alpha(z) = 0$ lie outside the unit circle, and $\{\varepsilon_t\}$ is a white-noise sequence with mean zero and variance σ^2 . The process described by equations 1 and 2 is a covariance-stationary process. In terms of the persistence displayed by the process, the most interesting property is that the dynamic multiplier (or impulse response function) goes to zero as time goes on:

$$\lim_{s \rightarrow \infty} \frac{\partial y_{t+s}}{\partial \varepsilon_t} = 0. \quad (3)$$

5. See Edwards (1996, 1998) for inflationary analyses based on credibility models.

That is, the effect of a shock eventually dies off. Consider, instead, the following unit root process:

$$(1-L)y_t = \delta + \beta(L)\varepsilon_t, \quad (4)$$

where it is required that $\beta(1) \neq 0$.⁶ In this case, an innovation in ε_t has a permanent effect on y_t . To see this, notice that

$$\frac{\partial y_{t+s}}{\partial \varepsilon_t} = 1 + \beta_1 + \beta_2 + \cdots + \beta_s,$$

and therefore

$$\lim_{s \rightarrow \infty} \frac{\partial y_{t+s}}{\partial \varepsilon_t} = \beta(1). \quad (5)$$

Three main approaches have been used to detect extreme persistence (nonstationarity) in economic time series. Campbell and Mankiw (1987a, b) use a parametric approach to analyze the time series properties of real gross national product (GNP). Their test makes use of the fact that a nonstationary process requires $\beta(1) \neq 0$, by calculating impulse-response functions obtained from the estimation of autoregression moving average (ARMA) representations of the first-differenced series. A second approach outlined in Cochrane (1988) provides a nonparametric assessment of the same question. Intuitively, Cochrane's test, known as the variance-ratio test, exploits the fact that if y_t is a random walk [$\beta(L) \neq 1$] then the variance of its s -differences grows linearly with s , while if it is stationary, this variance approaches a constant.⁷ Consequently, the variance ratio is simply

$$VR(s) = \frac{1}{s+1} \frac{9(y_{t+s+1} - y_t - \mu s)}{9(y_{t+1} - y_t - \mu)}. \quad (6)$$

6. To understand why it is that this condition is required for nonstationarity of y_t , notice that the stationary process in equation 1 can be first differenced and rewritten as

$$(1-L)y_t = (1-L)\alpha(L)\varepsilon_t = \beta(L)\varepsilon_t.$$

Because y_t is stationary, then Δy_t is also stationary. Although this process has the same form as equation 4, it is clear that necessarily $\beta(1) = 0$.

7. More generally, let y_t be a nonstationary process of the form indicated in equation 4. If we define $V(s) = E(y_{t+s} - y_t - \mu s)$, then

$$\lim_{s \rightarrow \infty} \frac{V(s)}{s} = \sigma^2 [\beta(1)]^2.$$

This ratio approaches one for a random walk and zero for a stationary time series. It is easy to show that in the general case of a nonstationary time series, the square root of $VR(s)$ would be a lower bound of $\beta(1)$.

A third approach to measuring persistence is known as the unobserved components approach. In this method, the series is modeled as the sum of two components. The best-known decomposition separates the series into a random walk and a stationary component:

$$\begin{aligned} y_t &= z_t + c_t, \\ z_t &= \mu + z_{t-1} + \eta_t, \text{ and} \\ F_t &= \gamma(L)\delta_t, \end{aligned} \tag{7}$$

where $E(\eta_t\delta_t)$ is arbitrary. This approach measures persistence as the variance of the random walk component relative to the total variance, $\sigma_{\Delta z}^2 / \sigma_{\Delta y}^2$.

These three measures of persistence are simply different ways of scaling the spectral density of Δy_t at zero frequency (Pesaran, Pierse, and Lee, 1993). This last function can be efficiently estimated using the Bartlett kernel,

$$\hat{s}_{\Delta y}(0) = \frac{1}{2\pi} \left[\hat{\gamma}_0 + 2 \sum_{j=1}^q \left(1 - \frac{j}{q+1} \right) \hat{\gamma}_j \right]. \tag{8}$$

These three approaches are subject to an important limitation, however.⁸ For any nonstationary general process there exists some stationary process, which is impossible to distinguish from the former for any finite sample of size T . A way of circumventing this problem and increasing the power in testing for unit root processes is to assume that the series is well represented by a low-order autoregressive process. Although it is generally not possible to refute the hypothesis that a general process is a unit root process, it is usually possible to test the less general hypothesis that a process is an $AR(p)$ process containing a unit root. The most popular unit root tests among applied economists, including the Dickey-Fuller and Phillips-Perron

8. See Hamilton (1994) for a discussion of the observational equivalence problem.

tests, belong to this family, in that they assume that the series is represented by a first-order autoregressive process.

A different way of thinking about inertia is within the class of stationary processes. The stationarity of an ARMA(p, q) model depends on the roots of the autoregressive component lying outside the unit circle, regardless of the roots of the moving average component. This result provides a simple way of measuring the degree of persistence of a stationary series. In the long run, the dynamics of a series will be dominated by the largest root of its autoregressive component. Of course, if the series is stationary, with the modulus of all inverted roots being less than one, the effect of a random disturbance will eventually die off. In the meantime, however, the effect of the shock will last longer for a larger value of the inverted dominant root or eigenvalue. In this paper we use the inverted dominant root of the autoregressive component of the inflationary process of each country as a measure of the degree of inertia displayed by the series.

2. PRELIMINARY DATA ANALYSIS

This section presents a preliminary analysis of the time series properties of monthly inflation rates for sixteen countries. Our main interest is to analyze several measures of persistence using a full sample. In the subsequent section, we analyze the stability through time of our inertia measures and try to detect structural breaks in the dynamics of the different inflationary processes.

2.1 The Data

We use data from the International Financial Statistics (IFS) provided by the International Monetary Fund (IMF), on monthly consumer price index (CPI) inflation for the period 1970-94. The sample includes sixteen advanced and developing countries: Argentina, Brazil, Chile, Colombia, France, Germany, India, Israel, Italy, Japan, Mexico, Spain, Sweden, Switzerland, the United Kingdom, and the United States. In selecting these countries we made an effort to cover a wide variety of macroeconomic experiences and inflationary histories. Table 1 presents the average, coefficients of variation, and minimum and maximum of the monthly inflation rates. The diversity of the sample is considerable: while Argentina and Brazil had monthly average inflation rates of around 10 percent, Germany and Switzerland had monthly rates on the order of 0.3 percent.

Table 1. Inflation in Selected Countries, 1970.1-1996.8

Percent and coefficient of variation

<i>Country</i>	<i>Average inflation rate</i>	<i>Coefficient of variation rate</i>	<i>Minimum monthly inflation rate</i>	<i>Maximum monthly inflation rate</i>
Argentina	9.3	1.8	-6.0	196.6
Brazil	10.2	1.2	-0.3	80.7
Chile	3.7	1.8	-8.1	87.5
Colombia	1.7	0.6	-2.0	6.9
France	0.5	0.7	-0.3	1.9
Germany	0.3	1.1	-0.5	1.8
India	0.7	1.5	-2.9	4.3
Israel	3.5	1.3	-1.4	27.5
Italy	0.8	0.7	-0.4	3.1
Japan	0.4	2.0	-1.1	4.1
Mexico	2.5	1.0	0.0	15.5
Spain	0.8	1.1	-4.4	9.0
Sweden	0.6	1.1	-0.5	3.3
Switzerland	0.3	1.3	-0.7	2.1
United Kingdom	0.7	1.1	-0.9	4.3
United States	0.4	0.7	-0.5	1.8

Source: Authors' calculations.

2.2 Inflationary Persistence: Full Sample Measures

Table 2 presents nonparametric variance ratio tests, as well as Dickey-Fuller unit root tests for the sixteen countries in the sample.⁹ For all countries the variance ratios clearly decrease as k increases.

9. In the frequency domain, a nonstationary series will present a large spectrum at low frequency. The variance ratio is an asymptotically equivalent estimate of the spectral density at frequency zero, and we therefore provide estimates of this statistic. In the time-space representation a series will be nonstationary if it has a unit root in its autoregressive component. Unit root tests like the Dickey-Fuller, Phillips-Perron, and Sargan-Barghawa tests target that feature in a first-order, autoregressive, nonstationary time series. The analysis of the statistical properties of unit root tests is far from being the goal of this paper. However, it is important to keep in mind that most of the unit root tests are linear combinations of the same two statistics and that consequently there is no single, most powerful test of the unit root hypothesis. In this paper we use the Dickey-Fuller test to identify extreme persistence, as this is by far the most popular among applied economists.

We removed existing monthly additive, seasonal components by a standard regression method in all those countries for which the seasonal components are statistically significant. Jaeger and Kunst (1990) show that the method used to seasonally adjust series affects the measured persistence of a series. In particular, filtering the data through X-11 methods spuriously introduces persistence into the series. We therefore used the simpler method of regression on dummies.

The pace at which the ratio decreases is quite different across countries, however, suggesting that the degree of inflationary persistence may be higher in some of these nations than in others. Moreover, the ratio decreases very rapidly in some countries (Switzerland, Japan, and Spain, for example), indicating that inertia was not an issue in the period under study. This is confirmed by the variance ratio plots (which include confidence intervals) presented in figure 1.

It is particularly interesting to contrast Brazil and Mexico, on the one hand, with more developed, low inflation countries like Japan, Switzerland, and Sweden. In the former group, although the variance ratios decrease as k rises, they are still significantly different from zero even for k as large as thirty-six months. In contrast, in the second group of countries, the variance ratios are not significantly different from zero for k between eighteen and twenty-four months.

Table 2 also shows augmented Dickey-Fuller tests for the full sample. The table presents the t statistic for the autocorrelation coefficient obtained in an ordinary least squares (OLS) regression, including intercept, deterministic trend, and four lags of inflation first differences. Using the full sample, we are not able to reject the unit root hypothesis in six of the sixteen countries: France, Israel, and Mexico, at 95 percent confidence and Brazil, Chile, and Italy at 99 percent confidence.

Table 2. Full Sample Nonstationarity Tests

Country	Seasonal adjustment	Variance ratio test					ADF ^a
		$k = 3$	$k = 6$	$k = 12$	$k = 24$	$k = 36$	
Argentina	No	0.61	0.31	0.21	0.14	0.10	-4.80
Brazil	No	0.76	0.67	0.46	0.32	0.20	-3.89
Chile	No	0.25	0.16	0.08	0.06	0.05	-3.78
Colombia	Yes	0.25	0.15	0.08	0.04	0.03	-6.76
France	Yes	0.34	0.21	0.14	0.10	0.08	-3.40
Germany	Yes	0.28	0.17	0.09	0.06	0.05	-5.54
India	Yes	0.25	0.15	0.08	0.04	0.03	-4.16
Israel	No	0.25	0.15	0.08	0.04	0.03	-2.57
Italy	Yes	0.32	0.19	0.13	0.07	0.06	-3.87
Japan	Yes	0.25	0.15	0.09	0.05	0.04	-5.34
Mexico	No	0.56	0.44	0.28	0.19	0.14	-3.28
Spain	Yes	0.23	0.13	0.07	0.04	0.03	-6.13
Sweden	Yes	0.28	0.15	0.08	0.04	0.03	-6.85
Switzerland	Yes	0.34	0.16	0.08	0.05	0.04	-6.48
United Kingdom	No	0.33	0.16	0.06	0.04	0.03	-5.89
United States	Yes	0.36	0.23	0.14	0.11	0.08	-4.02

Source: Authors' calculations.

a. Critical values are -3.42 (5 percent) and -3.98 (1 percent).

Figure 1. Variance Ratio Test

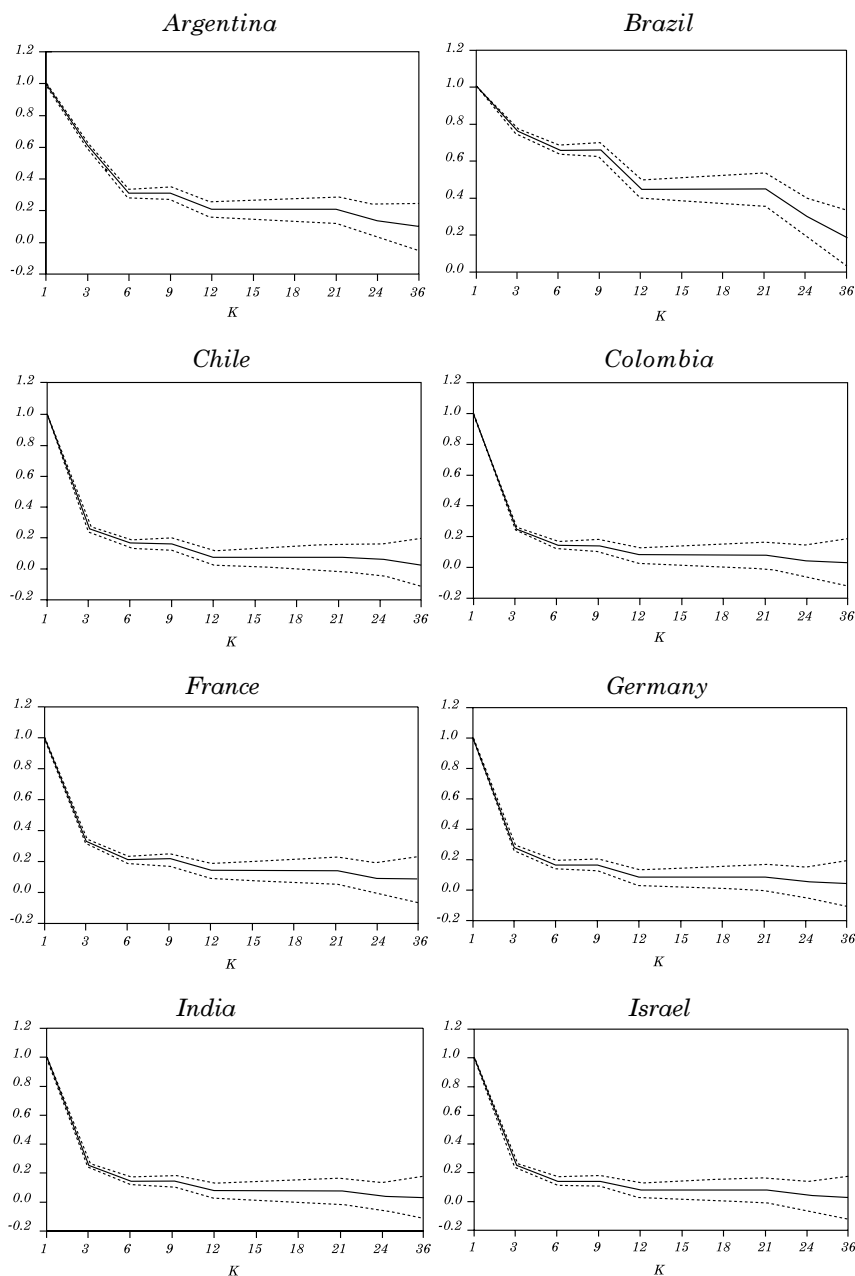
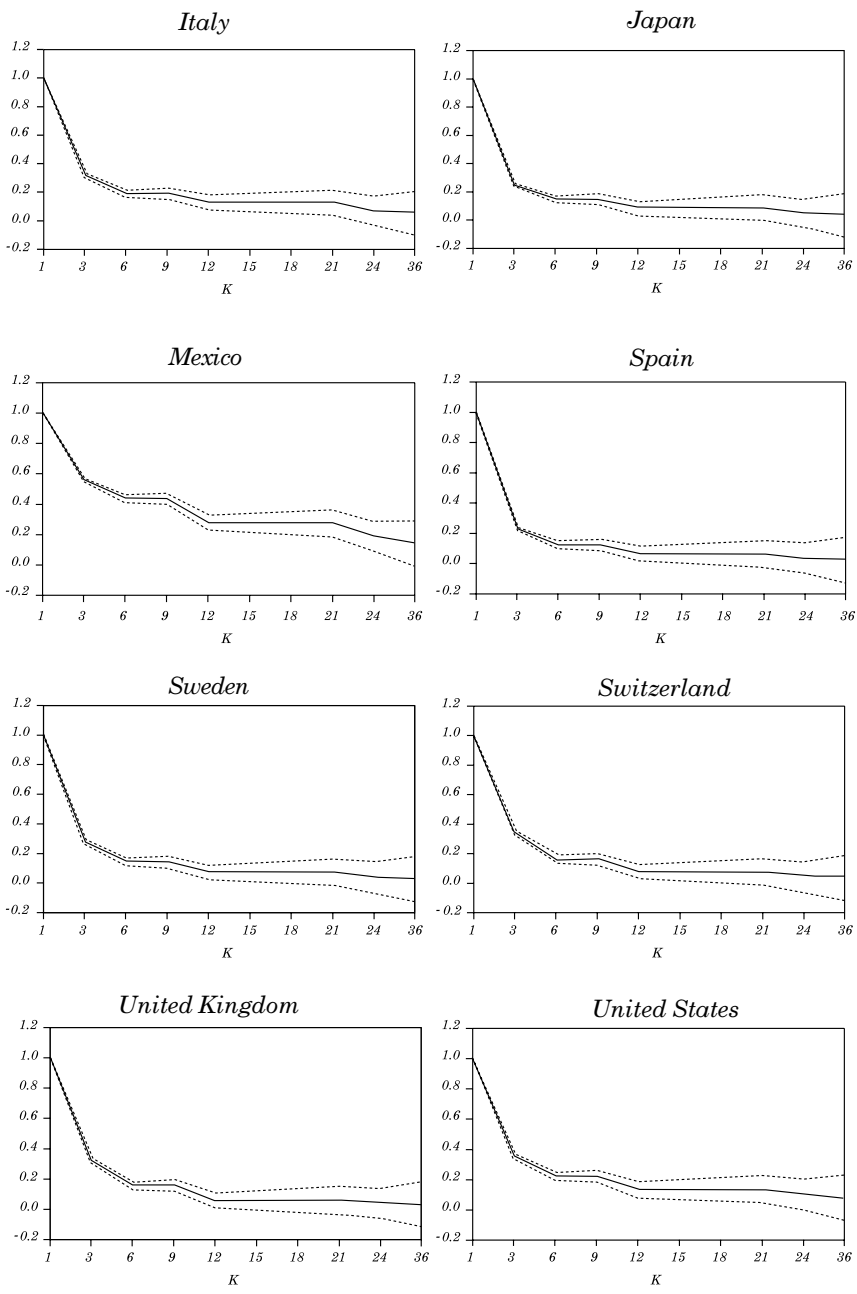


Figure 1. (continued)



A key question from a policy perspective is whether the domestic inflation rate converges with world inflation, and, if so, how fast this convergence takes place. We can statistically distinguish three distinct cases of convergence: (1) inflation has a unit root and thus does not converge to a steady-state value; (2) inflation is white noise, and thus there is no inflationary inertia; and (3) there is some degree of inertia, with inflation converging slowly to its steady-state value. With regard to this latter case, we investigated whether different countries exhibit different speeds of inflationary convergence, proceeding as follows: we fitted MA(1), MA(2), AR(1), AR(2), ARMA(1,1), and ARMA(2,2) for all sixteen countries in the period January 1970 to August 1996. We also fitted higher-order autoregressive models in all those cases in which the correlograms and partial correlograms suggested that these were plausible processes. Table 3 summarizes the results. For each country we indicate whether the series was previously seasonally adjusted, the various models that were fitted to the series, and the Q statistic white-noise residuals.

The results suggest, first, that pure moving average representations generally perform poorly in terms of Q statistics. Second, low-order autoregressive processes seem to provide a good fit for the inflationary process in some countries, namely, Brazil, Colombia, India, Mexico, Sweden, Switzerland, the United Kingdom, and the United States. An AR(1) performs particularly well in Brazil, Colombia, India, and Mexico. The other four cases are well represented by an AR(2) process. Third, some of the more developed, relatively low inflation countries such as France, Germany, Italy, Japan, and Spain show low Q statistics in low-order ARMA representations.¹⁰ However, these models present the standard near root-cancellation problem. In all five cases, a very high inverted AR root has as its counterpart a very high MA inverted root; these results must therefore be interpreted carefully.¹¹ Consider the case of Germany. The AR(1) and AR(2) representations feature dominant inverted roots of 0.36 and 0.61, respectively. The MA(1), in turn, has a low negative root of -0.29. The ARMA(1,1) and ARMA(2,2) representations, however, both present autoregressive and moving average dominant roots close to unity. Finally, the inflation processes of Argentina, Chile, and Israel do not seem to be well represented by any of the models presented in the table. This situation may be due to the fact that we need to specify a higher-order model

10. Inflation in Germany and Italy could also be fairly well represented by an AR(2) process.

11. To understand this problem, note that the white-noise process $y_t = \varepsilon_t$ can be rewritten by multiplying both sides of the last equation by $(1-\rho L)\varepsilon_t$, such that $(1-\rho L)y_t = (1-\rho L)\varepsilon_t$ is also a valid representation of y_t for any value of p .

for inflation, or, as we discuss in the next section, it may result from the existence of a structural break in the inflation autoregressive process.

In summary, with few exceptions, a low-order autoregressive process seems to fit the inflationary process observed for most countries in the sample. The estimated dominant inverted roots differ dramatically across countries, ranging from 0.09 for Spain, 0.16 for Sweden, and 0.38 for Japan, to a high of 0.91 for Brazil.¹²

3. IS THE DEGREE OF PERSISTENCE STABLE THROUGH TIME?

The preliminary analysis presented in the preceding section assumes that inflationary persistence is stable through time. This need not be the case, however. Since inflationary inertia is clearly dependent on institutional price arrangements, including indexation rules, mechanisms determining staggered contracts, the degree of credibility of the exchange rate policies, and other economic variables, there are good reasons to believe that inflationary persistence will change through time. In particular, one would expect that a change in the exchange rate regime (from a crawling peg to a fixed rate, for instance) will result in a reduction in the degree of persistence.

In this section we investigate whether the data exhibit structural breaks in the degree of persistence. We do this by calculating unit rolling augmented Dickey-Fuller (ADF) tests and sequential Perron tests. The rolling ADF tests were constructed by rolling a thirty-six-month sample along the full period under analysis. For each of these subsamples, we performed an ADF test including a time trend. A full-sample interpretation of this test is that if the minimum t -ratio obtained is larger (in absolute value) than the critical value, then the unit root hypothesis can be rejected for the whole sample.¹³ The sequential Perron test provides an alternative to this procedure. We started by performing a sequence of full-sample ADF tests, allowing for a structural break in the intercept as well as in the time trend in each test. The break is sequentially moved one period ahead, until we reached the end of the sample. We then compared the minimum t -ratio obtained in the lagged inflation term

12. For this comparison, we consider the highest inverted root of those models with a relatively good Q statistic and with no indication of the root-cancellation problem.

13. Critical values for this test are provided in Banerjee, Lumsdaine, and Stock (1992). The critical value is typically larger in absolute value than the standard Dickey-Fuller critical value.

Table 3. Full Sample Measures of Persistence

Country	Seasonal adjustment	Model specification	Inverted AR roots				Inverted MA roots		Q statistic ^a	
									$p = 5^*$	$p = 10^{**}$
Argentina	No	AR(1)	0.68	...	...	...	...	...	15.6	62.4
		AR(2)	0.65	0.05	...	...	...	...	12.8	61.0
		MA(1)	...	...	...	...	-0.35	...	18.7	31.4
		MA(2)	...	...	...	...	-0.18 + 0.26i	-0.18 - 0.26i	13.3	23.1
		ARMA(1,1)	0.66	...	...	...	-0.04	...	13.4	61.3
		ARMA(2,2)	0.366 - 0.27i	0.36 + 0.27i	...	...	0.02 + 0.54i	0.02 - 0.54i	2.2	48.9
Brazil	No	AR(1)	0.91	...	...	...	...	...	9.4	13.3
		AR(2)	0.90	0.06	...	...	...	...	8.1	11.5
		MA(1)	...	...	...	...	-0.72	...	491.1	739.5
		MA(2)	...	...	...	...	-0.45 + 0.63i	-0.45 - 0.63i	259.6	397.8
		ARMA(1,1)	0.90	...	...	...	-0.05	...	8.3	11.8
		ARMA(2,2)	0.88	-0.48	...	...	-0.29 + 0.29i	-0.29 - 0.29i	2.3	6.2
Chile	No	AR(1)	0.46	...	...	...	...	...	49.1	92.1
		AR(2)	0.71	-0.37	...	...	...	...	39.3	69.4
		AR(3)	0.85	-0.29 + 0.49i	-0.29 - 0.49i	...	...	...	28.6	46.0
		AR(4)	0.92	-0.05 + 0.70i	-0.05 - 0.70i	-0.64	...	...	4.2	13.3
		MA(1)	...	...	...	...	-0.33	...	137.7	256.6
		MA(2)	...	...	...	...	-0.17 - 0.4i	-0.17 + 0.4i	82.4	159.1
		ARMA(1,1)	0.98	...	...	...	0.83	...	9.2	26.0
		ARMA(2,2)	0.98	-0.97	...	...	0.83	-0.99	9.3	25.2
Colombia	Yes	AR(1)	0.49	...	...	...	...	...	2.9	7.2
		AR(2)	0.54	-0.09	...	...	...	...	2.2	6.4
		MA(1)	...	...	...	...	-0.42	...	21.4	34.1
		MA(2)	...	...	...	...	-0.23 - 0.43i	-0.23 + 0.43i	8.7	18.0
		ARMA(1,1)	0.57	...	...	...	0.11	...	2.5	6.7
		ARMA(2,2)	0.52	-0.27	...	...	-0.01	-0.20	2.1	6.4

Table 3. (continued)

Country	Seasonal adjustment	Model specification	Inverted AR roots				Inverted MA roots		Q statistic ^a	
									$p = 5^*$	$p = 10^{**}$
France	Yes	AR(1)	0.81	...	...	...	...	...	27.8	44.0
		AR(2)	0.89	-0.24	...	...	...	...	24.6	35.4
		AR(3)	0.95	-0.19+0.52i	-0.19-0.52i	...	...	...	11.7	15.1
		MA(1)	...	...	...	...	-0.69	...	327.8	657.0
		MA(2)	...	...	...	...	-0.4-0.37i	-0.4+0.37i	182.5	373.3
		ARMA(1,1)	0.99	...	...	...	0.69	...	19.4	20.8
		ARMA(2,2)	0.99	-0.22	...	...	0.79	-0.53	4.0	5.8
Germany	Yes	AR(1)	0.36	...	...	...	...	...	20.0	43.1
		AR(2)	0.61	-0.30	...	...	...	...	8.0	18.9
		MA(1)	...	...	...	...	-0.29	...	47.8	93.7
		MA(2)	...	...	...	...	-0.14+0.38i	-0.14-0.38i	27.9	59.2
		ARMA(1,1)	0.97	...	...	...	0.85	...	3.9	6.1
		ARMA(2,2)	0.97	-0.41	...	...	0.86	-0.50	1.2	3.1
India	Yes	AR(1)	0.50	...	...	...	...	...	19.3	28.3
		AR(2)	0.63	-0.19	...	...	...	...	16.7	23.5
		MA(1)	...	...	...	...	-0.42	...	70.2	96.1
		MA(2)	...	...	...	...	-0.22-0.32i	-0.22-0.32i	43.1	61.1
		ARMA(1,1)	0.88	...	...	...	0.57	...	10.7	14.0
		ARMA(2,2)	0.90	-0.42	...	...	0.65	-0.57	3.5	6.1
Israel	No	AR(1)	0.75	...	...	...	...	...	45.6	71.6
		AR(2)	0.83	-0.21	...	...	...	...	36.5	57.5
		AR(3)	0.92	-0.18+0.57i	-0.18-0.57i	...	...	...	23.1	31.1
		MA(1)	...	...	...	...	-0.71	...	227.5	451.6
		MA(2)	...	...	...	...	0.39-0.06i	0.39+0.06i	158.1	302.6
		ARMA(1,1)	0.97	...	...	...	0.64	...	24.8	36.6
		ARMA(2,2)	0.98	-0.43	...	...	0.75	-0.75	99.3	188.2

Table 3. (continued)

Country	Seasonal adjustment	Model specification	Inverted AR roots				Inverted MA roots		<i>Q statistic</i> ^a	
									<i>p</i> = 5*	<i>p</i> = 10**
Italy	Yes	AR(1)	0.67	...	...	...	...	...	30.7	54.4
		AR(2)	0.85	-0.40	...	...	...	...	6.7	17.2
		MA(1)	...	...	...	...	-0.45	...	230.9	437.8
		MA(2)	...	...	...	...	0.25 - 0.54i	0.25 + 0.54i	102.6	220.1
		ARMA(1,1)	0.97	...	...	...	0.69	...	4.8	8.1
		ARMA(2,2)	0.97	-0.85	...	...	0.70	-0.88	49.4	8.1
Japan	Yes	AR(1)	0.38	...	...	...	...	...	38.5	86.5
		AR(2)	0.67	-0.38	...	...	...	...	12.5	35.3
		AR(3)	0.81	-0.29 + 0.43i	-0.29 - 0.43i	...	...	...	4.3	19.9
		MA(1)	...	...	...	...	-0.29	...	79.0	172.7
		MA(2)	...	...	...	...	-0.13 + 0.45i	-0.13 - 0.45i	46.4	112.7
		ARMA(1,1)	0.98	...	...	...	0.84	...	1.9	12.3
Mexico	No	ARMA(2,2)	0.97	0.16	...	...	0.82	0.18	2.0	12.2
		AR(1)	0.86	...	...	...	...	...	6.5	9.5
		AR(2)	0.87	-0.04	...	...	...	...	7.0	10.0
		MA(1)	...	...	...	...	-0.71	...	386.0	624.6
		MA(2)	...	...	...	...	-0.45 + 0.52i	-0.45 - 0.52i	170.5	280.6
		ARMA(1,1)	0.88	...	...	...	0.08	...	7.6	10.5
Spain	Yes	ARMA(2,2)	0.86	-0.96	...	...	0.04	-0.99	6.5	9.2
		AR(1)	0.09	...	...	...	...	...	44.8	111.1
		AR(2)	0.50	-0.44	...	...	...	...	18.0	49.5
		MA(1)	...	...	...	...	-0.06	...	50.5	124.5
		MA(2)	...	...	...	...	-0.02 + 0.39i	-0.02 - 0.39i	29.7	79.2
		ARMA(1,1)	0.99	...	...	...	0.94	...	13.3	20.0
		ARMA(2,2)	0.99	-0.07	...	...	0.91	0.13	4.8	13.8

Table 3. (continued)

Country	Seasonal adjustment	Model specification		Inverted AR roots			Inverted MA roots		Q statistic ^a	
									$p = 5^*$	$p = 10^{**}$
Sweden	Yes	AR(1)	0.16	...	...	...	...	...	10.4	39.1
		AR(2)	0.44	-0.30	...	...	...	...	2.3	24.7
		MA(1)	...	...	...	-0.13	...	...	13.7	47.8
		MA(2)	...	...	...	-0.07 + 0.36i	-0.07 - 0.36i	...	4.5	29.5
		ARMA(1,1)	0.99	...	...	0.93	...	...	4.0	14.6
		ARMA(2,2)	0.99	-0.99	...	0.93	-0.98	...	3.5	15.8
Switzerland	Yes	AR(1)	0.36	...	...	...	...	...	15.8	44.4
		AR(2)	0.62	-0.33	...	...	...	...	1.2	20.0
		MA(1)	...	...	...	-0.25	...	...	38.6	92.9
		MA(2)	...	...	...	-0.14 + 0.48i	-0.14 - 0.48i	...	5.9	38.4
		ARMA(1,1)	0.95	...	...	0.82	...	...	23.3	37.3
		ARMA(2,2)	0.67	-0.97	...	0.30	-0.93	...	0.9	15.2
United Kingdom	No	AR(1)	0.45	...	...	...	...	...	8.7	42.9
		AR(2)	0.62	-0.24	...	...	...	...	3.1	34.7
		MA(1)	...	...	...	-0.36	...	...	45.6	103.2
		MA(2)	...	...	...	-0.18 + 0.37i	-0.18 - 0.37i	...	23.4	73.4
		ARMA(1,1)	0.96	...	...	0.79	...	...	14.3	36.8
		ARMA(2,2)	0.98	0.15	...	0.89	-0.07	...	2.4	21.5
United States	Yes	AR(1)	0.67	...	...	...	...	...	23.7	45.7
		AR(2)	0.82	-0.33	...	...	...	...	5.4	18.1
		MA(1)	...	...	...	-0.48	...	...	183.3	350.6
		MA(2)	...	...	...	-0.23 + 0.55i	-0.23 - 0.55i	...	86.6	162.4
		ARMA(1,1)	0.94	...	...	0.58	...	...	14.6	24.6
		ARMA(2,2)	0.97	0.41	...	0.81	0.15	...	5.4	13.5

Source: Authors' calculations.

a. * Critical values at 95 percent confidence are 9.49 for the AR(1) and MA(1) models; 7.82 for the AR(2), MA(2), and ARMA(1,1) models; 5.99 for the AR(3) model; and 3.84 for the AR(4) and ARMA(2,2) models. ** Critical values at 90 percent confidence are 16.92 for the AR(1) and MA(1) models; 15.51 for the AR(2), MA(2), and ARMA(1,1) models; 14.07 for the AR(3) model; and 12.59 for the AR(4) and ARMA(2,2) models.

with the critical value. Table 4 reports the results obtained with these two tests. We reject the full-sample unit root hypothesis with 95 percent confidence in all cases. The implications are very clear: independently of the autoregressive roots presented in table 3, after allowing for structural breaks and parameter instability, the inflation rate seems to have been stationary for all countries in our sample.

To investigate whether the degree of inertia has been stable, we estimated sequential Quandt likelihood ratio (QLR) tests for each country. We constructed these tests by performing a sequence of full-sample QLR tests using an AR(1) model specification, in which we allowed for a change in the autoregressive coefficient. The break point was moved sequentially along the sample. Again, a full-sample interpretation of this long series of tests is that if the maximum QLR statistic obtained is larger than the 95 percent critical value, then we cannot reject the null hypothesis of an absence of a structural break in the sample.¹⁴ This procedure is not subject to the standard problem of choosing the break point date through data inspection.

Table 4. Unit Root and Trend Break Hypothesis Tests^a

<i>Country</i>	<i>Rolling ADF</i>	<i>Sequential Perron test</i>	<i>Sequential Quandt LR</i>
Critical value (95%)	-4.85	-4.39	27.87
Argentina	-16.89*	-6.91*	88.63**
Brazil	-5.62*	-4.45*	12.19
Chile	-8.45*	-5.52*	32.10**
Colombia	-7.56*	-9.82*	4.81
France	-6.44*	-6.16*	35.06**
Germany	-6.63*	-8.27*	6.79
India	-6.20*	-9.50*	4.62
Israel	-19.07*	-7.85*	47.20**
Italy	-6.38*	-5.95*	49.49**
Japan	-7.80*	-8.33*	10.55
Mexico	-9.26*	-7.12*	18.81
Spain	-8.39*	-7.58*	38.62**
Sweden	-6.65*	-8.37*	12.37
Switzerland	-6.27*	-8.68*	9.37
United Kingdom	-5.53*	-7.40*	17.81
United States	-5.08*	-6.95*	20.11

Source: Authors' calculations.

a. * Reject unit root hypothesis at 95 percent confidence. ** Reject nonstructural break hypothesis.

14. For a more detailed analysis of this testing procedure and of its asymptotic and empirical distributions, see Banerjee, Lumsdaine, and Stock (1992).

Column 3 of table 4 summarizes the sequential QLR test results. We strongly reject the hypothesis of an absence of a structural break in the cases of Argentina, Chile, France, Israel, Italy, and Mexico. These results may explain why we were unable to fit a unique ARMA representation for the inflationary processes of Argentina, Chile, and Israel, as reported in section 2. Unfortunately, we were unable to examine the possibility of a structural break following the implementation of the Real Plan in Brazil because the test requires at least fifty observations after an event.

Figure 2 presents the statistics obtained for each of these three procedures. In those cases in which the null hypothesis of an absence of a structural break is rejected, it is possible to determine the approximate date of the break. For Argentina, a structural break in the inflationary process seems to have happened in 1989 at the time of the Alfonsin hyperinflation. Structural breaks are detected in mid-1975 in Chile and in mid-1985 in Israel. Section 4 provides a detailed analysis of the latter two cases, together with Mexico.

Figure 2. Unit Root and Trend Break: Rolling ADF, Sequential Perron, and Sequential QLR

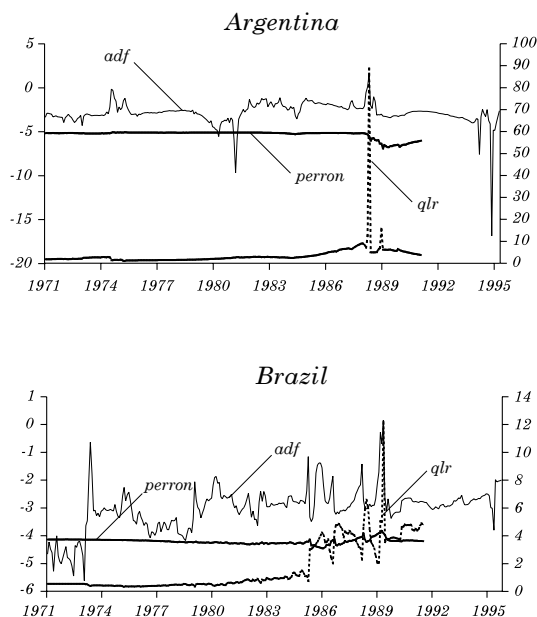


Figure 2. (continued)

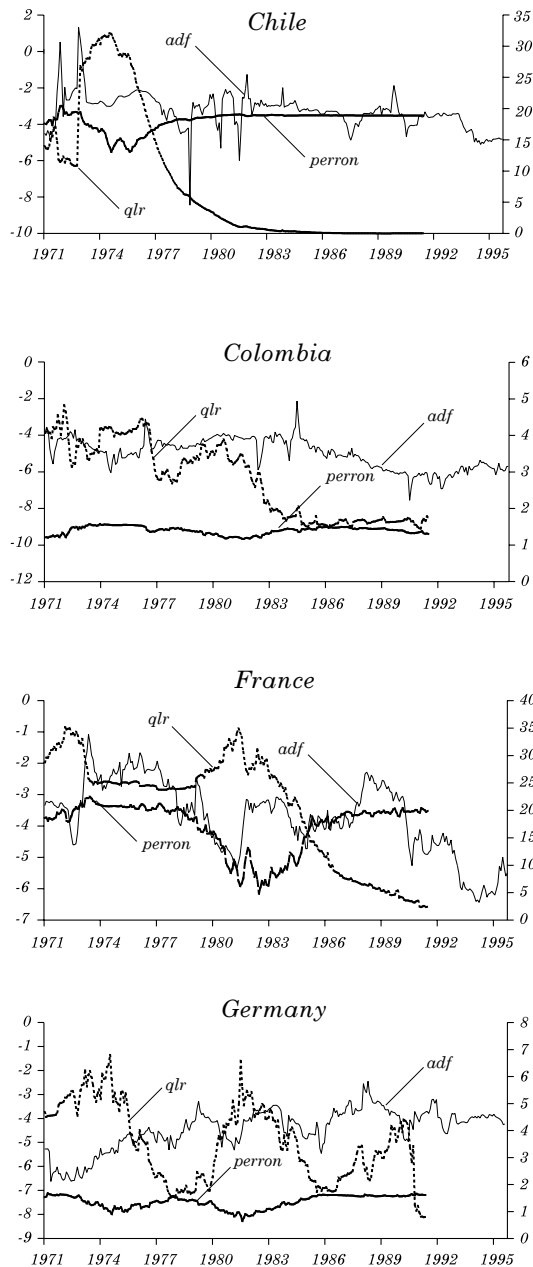


Figure 2. (continued)

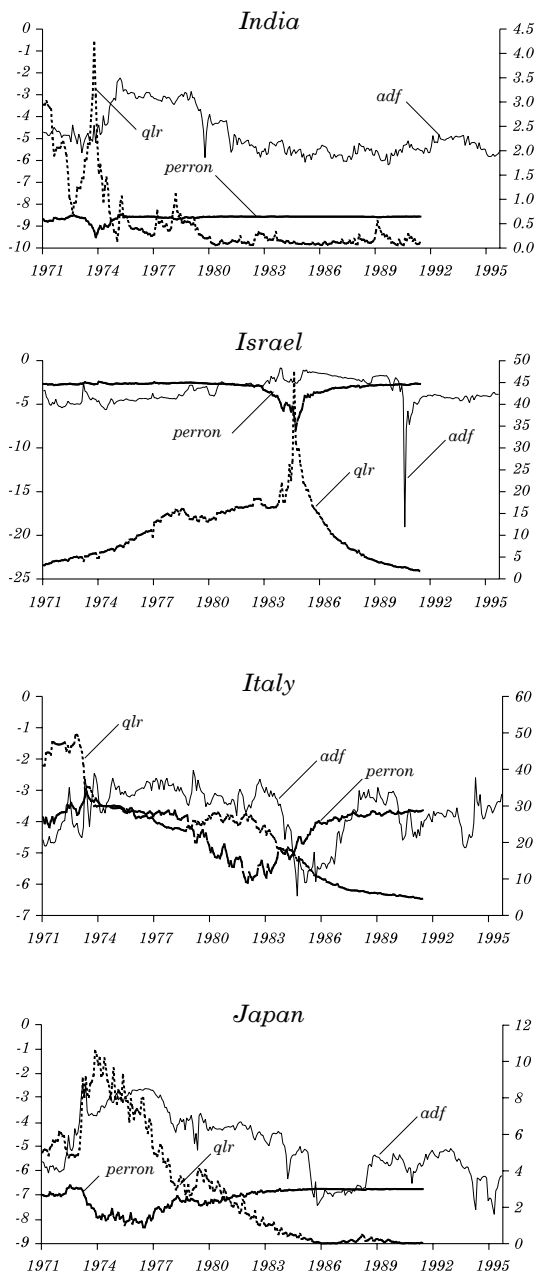


Figure 2. (continued)

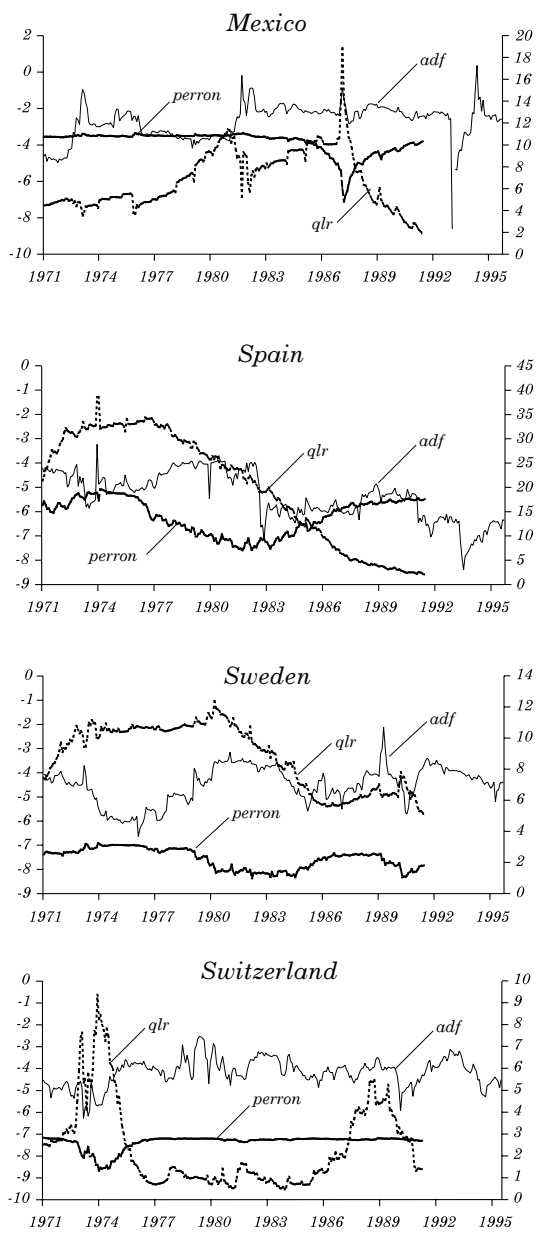
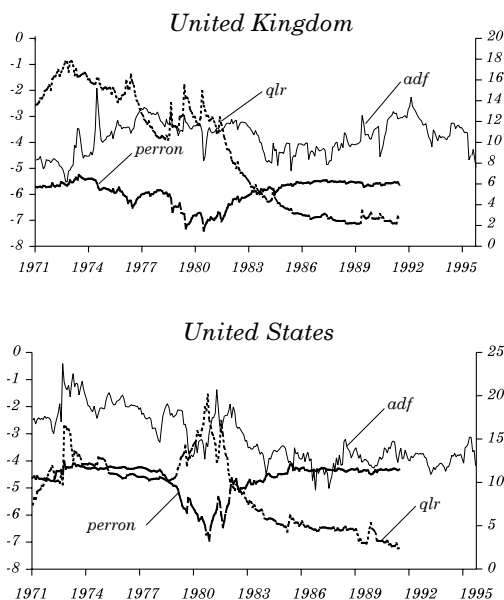


Figure 2. (continued)

3.1 Univariate Rolling Estimates of Inflationary Inertia

For each country we estimated AR(1) and AR(2) rolling regressions to further analyze the changing degree of inertia.¹⁵ Each regression was estimated for a rolling thirty-six-month sample. Figure 3 plots the rolling autoregressive coefficients obtained in the AR(1) regressions, with two standard deviation bands.¹⁶ The figure is quite revealing. First, the rolling AR(1) estimates show an important diversity of country experiences. While the degree of inertia is low and stable in some

15. Two alternative methods for addressing the parameter instability problem are recursive regressions and a Kalman filter. In both methodologies, the estimate at each date considers all the information contained in the sample up to that point. The rolling regression approach is more sensitive to changes in the inertia regime than is the Kalman filter because all the information used for the estimation is renewed at short intervals.

16. In the figure, our convention is to date the autoregressive coefficient obtained in each rolling regression in the last month included in the sample. Monte Carlo exercises showed that when using rolling regressions with thirty-six observations, a structural break is statistically detected when at least twenty months of the new regime are included in the sample. Hence, changes in the rolling coefficients should be associated with events happening one and a half years before the date in the figure.

nations, such as Sweden and Switzerland, in others it is highly volatile during the period under analysis (Israel, for example). Second, AR(1) estimates clearly suggest that in many countries, the degree of inflationary inertia varies significantly through time. This is the case of Argentina, for example, which exhibits no significant inflationary persistence in the early 1970s, a significant increase in inertia in the 1980s, and a marked decline in the 1990s. This latter development is related to the adoption of the convertibility stabilization plan in 1991. The estimates for Italy also suggest a rapid significant change in the degree of inertia. This is consistent with the results on the Quandt rolling structural break test reported above. Third, a number of countries seem to have lost their nominal macroeconomic anchor, at least temporarily. In this case, the AR coefficient is not significantly different from unity at conventional levels. This appears to be the case in Argentina in 1989 and in Brazil in the late 1980s and early 1990s. Fourth, some countries appear to have been able to defeat inflationary persistence, with the estimated coefficients of inertia collapsing from a positive—and in some cases very high—value to a value that is not significantly different from zero. An important question is whether those countries that have succeeded in eliminating inertia have followed similar policy courses. Section 4 addresses this issue with regard to nominal exchange rate policy and analyzes in greater detail the evolution of inertia in three countries that have implemented stabilization programs based on a nominal exchange rate anchor in the last ten years or so.

For most countries the results obtained from the estimation of rolling AR(2) (not reported here) tell a very similar story to those presented for AR(1). Some differences are appreciated in Colombia, France, Israel, Japan, Spain, and Sweden. As reported in table 3, however, the AR(2) representation appears significantly preferable to the AR(1) only for Japan, Spain, and Sweden.

The case of the United States deserves particular attention. Our results strongly suggest that the degree of inflationary persistence has not been stable through time. Between 1970 and 1975, the estimated AR coefficient was not significantly different from zero, suggesting that inflation was characterized by white noise and that there was an absence of inflationary inertia. Starting in 1976, however, the degree of persistence increased markedly. This peaked in the early 1980s, when the point estimate of the AR coefficient was 0.8—a figure not very different from that of the highly inflationary nations of South America. From then onward, inflationary inertia experienced a somewhat gradual decline, until virtually disappearing again in the mid-1990s. These results suggest that analyses that assume a stable degree of persistence, such as that of Fuhrer and Moore (1995), may result in highly misleading policy implications.

Figure 3. Rolling AR(1)

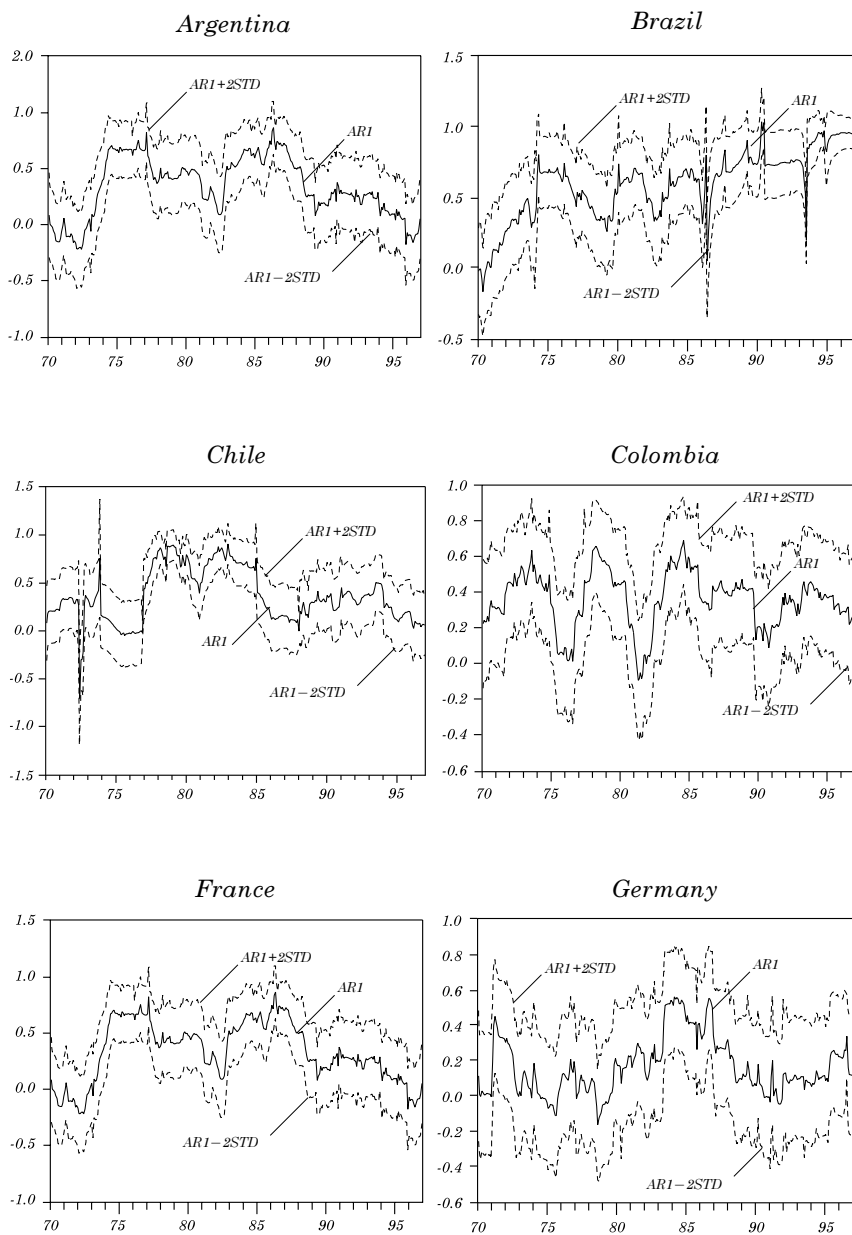


Figure 3. (continued)

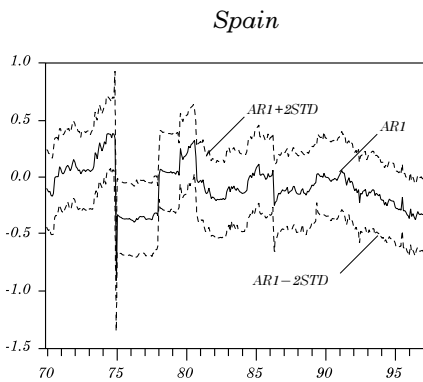
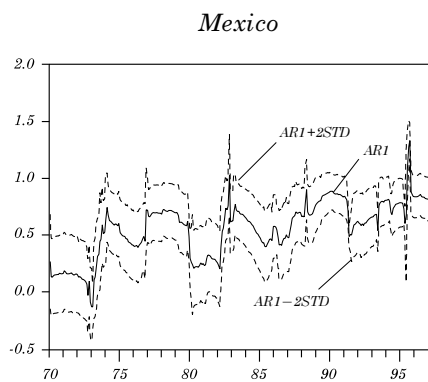
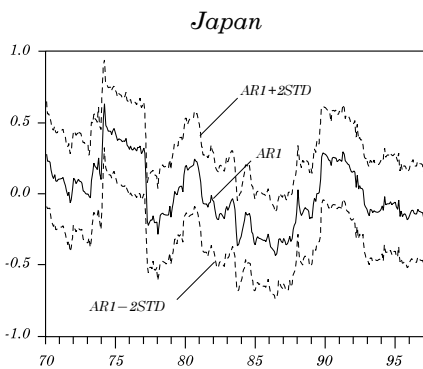
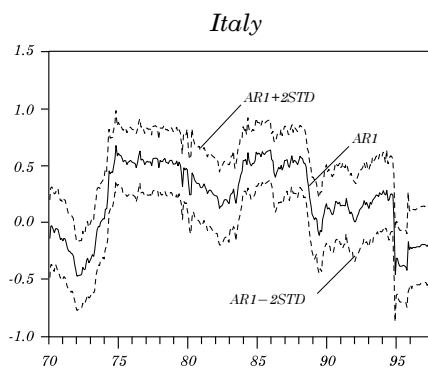
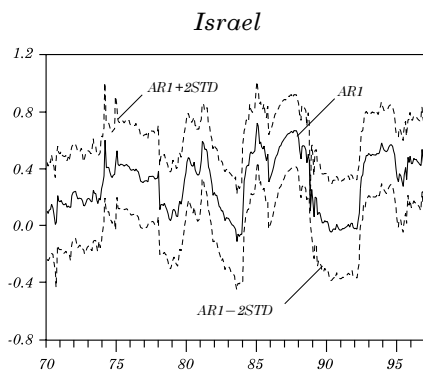
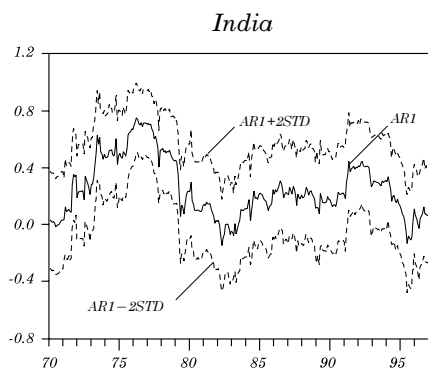
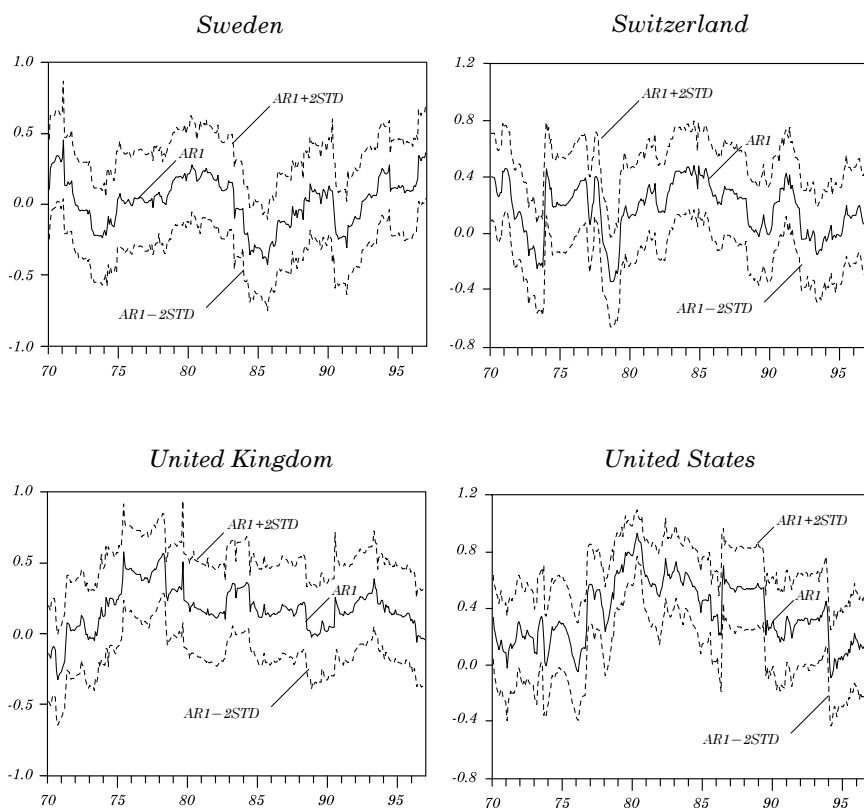


Figure 3. (continued)

3.2 Multivariate-Based Measures of Inflationary Inertia

A possible limitation of the results reported above is that they are based on univariate time series techniques, without controlling for other variables. The omission of possibly important determinants of inflation may introduce some biases in the estimation. To address this potential source of bias, we estimated the same set of AR(1) rolling regressions for some of the countries in our sample including as an additional explanatory variable the growth rate of nominal money

(IFS line 34).¹⁷ The results for Argentina, Chile, Israel, and Mexico (which are not reported here) show that in all four cases, the differences with respect to the results of figure 3 are clearly unimportant. We went a step further in the cases of Chile and Mexico: we reestimated the AR(1) regressions controlling for estimated excess money supply. This latter variable was obtained, in turn, as residuals from error correction-based estimates of the demand for money. The broad results obtained show that our previous results are not affected by the inclusion of additional variables in the analysis.¹⁸

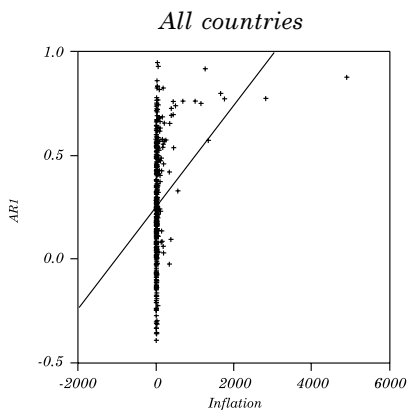
3.3 Inflation Levels and Persistence

The results summarized in the preceding two subsections present a simple measure of inflationary persistence for a heterogeneous set of countries. A key contribution of this paper is that we have allowed this measure of inertia to vary over time. This exercise provides a rich panel of data on inflationary persistence, which contains both cross-country and time series variation. An interesting question—and one that has important policy implications—is whether countries and periods of time featuring high inflation also tend to present high inertia. Figure 4 displays scatter diagrams of average inflationary inertia (measured as the average autoregressive coefficient obtained in the twelve rolling regressions for each year) and average annual inflation for different subsets of the countries in the sample. Interestingly, these diagrams show that there is a positive relationship between the average level of inflation and the measured degree of inertia. This relationship is present at different levels of average inflation. To demonstrate that this graphic evidence is also statistically significant, we performed panel regressions of inflation on persistence. We used the same information plotted in figure 4 to run panel regressions using pooled least squares, fixed effects, and random effects estimators. We also divided the sample into four different subsamples according to the average level of inflation. The results confirm those presented in figure 4 and discussed above (see table 5). All the regressions show a positive and statistically significant relationship between inflationary persistence and inflation levels. The results are robust to the estimation procedure and the subsample selected.

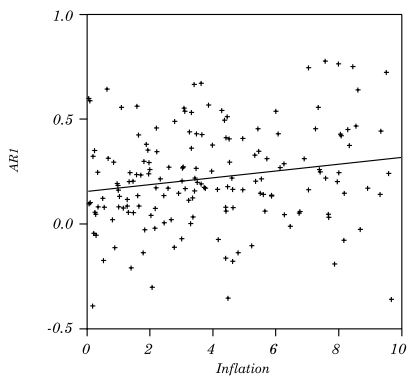
17. There is no monthly measure of money for several of the countries. For others, the series are too short or too incomplete.

18. We also estimated impulse response functions for alternative subsamples. Although we were unable to make fine timing distinctions, the estimates provided support for the results reported above.

Figure 4. Inflation and Persistence



All countries: Inflation between 0 and 10 percent



All countries: Inflation between 0 and 100 percent

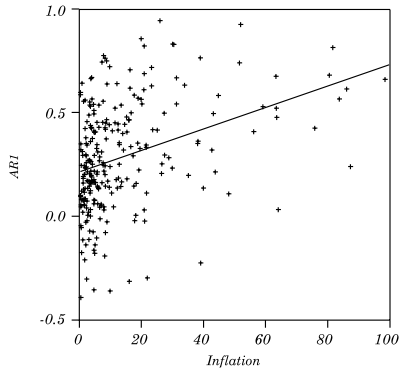
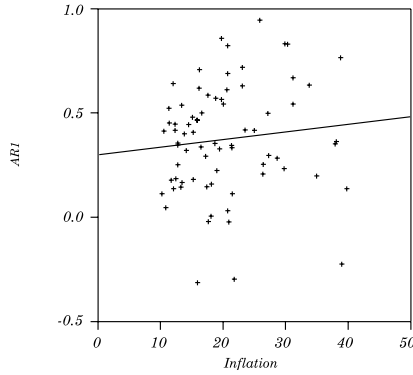
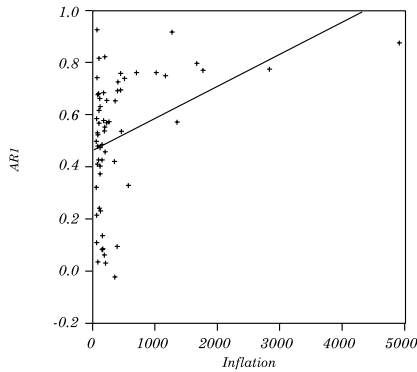


Figure 4. (continued)*All countries: Inflation between 10 and 40 percent**All countries: Inflation over 40 percent*

A possible explanation for this finding is that as the inflation level increases, societies develop institutional arrangements to protect themselves from the effects of inflation. Many of these protective measures—and, more specifically, indexation mechanisms—tend to increase the degree of inflationary inertia by adjusting prices according to some measure of past inflation. To find out whether countries with extensive indexation practices indeed exhibit higher inertia, we performed some preliminary nonparametric analyses. These results indicate that it is possible to reject the hypothesis that countries with high indexation and countries with no indexation come from the same sample. Moreover, average inflation rates in high-indexation countries are significantly higher than in countries where there are no generalized indexation practices.¹⁹

19. These preliminary results are available from the authors on request.

Table 5. Panel Regressions of Inflation on Persistence^a

<i>Regression method</i>	<i>Full sample</i>	<i>p < 0.1</i>	<i>p < 0.4</i>	<i>p < 1</i>	<i>p > 1</i>
Pooled least squares	3.371 (0.552)	0.029 (0.009)	0.112 (0.017)	0.223 (0.030)	15.766 (4.897)
Fixed effects	2.520 (0.695)		0.034 (0.015)		
Random effects	2.825 (0.639)		0.037 (0.014)		
No. observations	416	281	355	375	41

Source: Authors' calculations.

a. Standard errors in parentheses.

4. INERTIA AND INFLATION CONVERGENCE: THREE CASES

An important goal of most stabilization programs in developing (or emerging) economies is to reduce the domestic inflation rate to levels comparable to world inflation. In other words, stabilization programs are usually aimed at producing inflationary convergence. The speed at which inflation is reduced is important, however. Under fixed nominal exchange rates, a slow inflationary convergence will result in real exchange rate overvaluation and speculation; in the end, it can even generate a crisis. If inertia is eliminated or greatly reduced, inflationary convergence will be achieved rapidly, whereas if inertia is very high, convergence will tend to take a long time.²⁰ Historically many countries—including a number in our sample—have attempted to achieve inflation convergence by implementing stabilization programs based on the adoption of a rigid (and in some cases totally fixed) nominal exchange rate. In this section we briefly analyze the experiences of three countries in the sample that have adopted stabilization programs based on nominal exchange rate anchors within the last twenty years or so: Chile, Israel, and Mexico. Our main interest is to analyze, in the light of the battery of tests presented in the preceding sections, whether the adoption of this type of program indeed resulted in a reduction or elimination of the degree of inertia. We begin the discussion with a direct comparison of the Chilean and Mexican experiences of the 1970s and 1980s. We then move to the Israeli case.

20. Naturally, in making these statements we are assuming that the fundamental macroeconomic stance is conducive to inflationary reduction.

4.1 Chile (1978-82) and Mexico (1987-94)

The Chilean reform program was initiated in 1975, ten years prior to the launching of the Mexican reforms. Both programs, however, shared a number of features: drastic opening of the economy; ambitious privatization and deregulation; and a stabilization program based on a predetermined, nominal exchange rate anchor, supported by largely restrictive fiscal and monetary policies. Both programs were rooted in the notion that a predetermined rate of nominal devaluation, including a fully fixed nominal exchange rate, would constrain price increases, reduce inflationary expectations, and ensure that monetary and fiscal authorities would behave conservatively. In Chile, the exchange-rate-based stabilization program had two phases: from February 1978 through June 1979 the rate of devaluation was preannounced and deliberately set below the ongoing rate of inflation. In June of 1979 the nominal exchange rate was pegged at 39 pesos per dollar. The authorities announced that this new exchange rate would be in effect for the indefinite future.²¹

The Mexican program followed a somewhat different pattern. Between February and December 1988, the nominal exchange rate was fixed, thereby becoming the fundamental anchor of the anti-inflationary effort.²² Between January 1989 and November 1991, a system based on a preannounced rate of devaluation was in effect. As in Chile a decade earlier, during this period the actual rate of devaluation was deliberately set below the ongoing rate of inflation as a way of reducing expectations and price increases. The amount by which the peso was devalued was successively reduced from one peso per day in 1989 to 80 cents in 1990, 40 cents in 1991, and 20 cents in 1991. In November 1991, the authorities added some flexibility to the system when they adopted an exchange rate band with a sliding ceiling and a flat floor. This measure was justified on two grounds: it was supposed to discourage short-term capital inflows, and it was supposed to allow some real exchange rate corrections. Until October 1993—when the NAFTA controversy heated up in the United States—the actual peso-to-dollar rate was extremely stable, remaining in the lower half of the band. During 1994, however, and as a result of political and other developments, the

21. Parts of this comparison draw on Edwards (1995). See also Morely's review of Edwards in Morley (1997).

22. From October to December 1987 (the first months of the Pacto), nominal wages provided the anchor for the system. According to Vela (1993) the move to an exchange rate anchor in February 1988 was, in part, the result of labor union pressure.

exchange rate came under considerable pressure, moving towards the top of the band. As is well known, the attempt to widen the band in late December failed badly. It was too little, too late.

Early on, Chilean and Mexican authorities understood that consistent fiscal and monetary policies were a precondition for the success of the exchange-rate-based stabilization programs. In both countries, the primary balance of the public sector was under control even before the adoption of the exchange rate anchor, and monetary policy was largely restrained during most of the period.²³

The architects of both the Chilean and Mexican programs expected that a predetermined exchange rate would affect inflation through at least two channels. First, a fixed exchange rate in an open economy would impose a ceiling on tradable inflation; second, the new exchange rate policy would generate a major break in inflationary expectations and in inflationary inertia. Mexico's finance secretary, Pedro Aspe, put it the following way: "[Pegging the exchange rate] and... lowering... trade barriers in the tradables sector were indispensable for breaking down inertia. Consensus and a tendency toward purchasing power parity could reinforce each other to bring down inflation" (Aspe, 1993, p. 44).

Despite these similarities, the Chilean and Mexican programs had some important differences. Perhaps the most important referred to wage rate policy. Beginning in 1976, Chile adopted a formal backward-looking wage indexation scheme, through which all wages in the formal sector were periodically raised to compensate for the cumulative inflation rate of the previous period (Edwards and Edwards, 1991, chap. 6). In contrast, the Mexican *Pacto* encompassed a social and economic agreement among the government, the private sector, and labor unions, aimed at establishing guidelines for price, wage, and exchange rate changes.

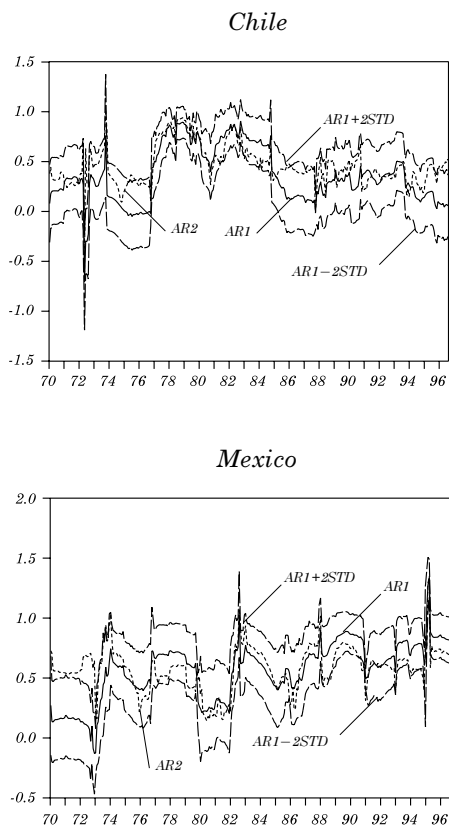
In both countries inflation was reduced, but at a significantly slower pace than what the authorities had anticipated. As their reforms proceeded and became consolidated, both countries were subject to very large capital inflows which helped finance increasingly large current account deficits and generated a very large real exchange rate overvaluation (Edwards, 1996). In Chile the current account deficit exceeded 12 percent of GDP in 1981, while in Mexico it surpassed 7 percent of

23. In the early years of the Mexican program, monetary policy was guided by a dual objective: on the one hand, the authorities were interested in reducing interest rates, which had reached extremely high real levels in 1987 and 1988, and on the other, they wanted to make sure that domestic credit policy would be consistent with the predetermined nominal exchange rate (Aspe, 1993, p. 37).

GDP in both 1992 and 1993. These very large deficits proved to be unsustainable, and both countries ended up facing major crises: the Chilean crisis erupted in June of 1982; the Mexican crisis in December 1994.

How successful were these two nominal exchange-rate-based stabilization programs in reducing (or eliminating) inflationary inertia? We can address this issue with the aid of the statistical analysis presented in this paper. In particular, we can use the different estimates of inertia developed in the preceding section to analyze whether the degree of inflationary persistence declined after the exchange rate anchor program was enacted, as the architects of the programs had intended. Figure 5 presents once again our estimates of inertia for these two countries using the AR(1) and AR(2) procedures.

Figure 5. Inflationary Inertia: Chile and Mexico



The results are quite interesting. Chile demonstrates a temporary decline in inertia by the end of 1979: both the AR(1) coefficient and the dominant root in the AR(2)-based measure decline temporarily to 0.33. From then on, however, measured inertia started to increase again, reaching more than 0.8 in the months immediately preceding the abandonment of the fixed exchange rate in June 1982. There are two possible explanations for this behavior. First, due to the mandated backward-looking wage indexation mechanism, structural inertia remained in the system despite the adoption of the anchor. Second, as Edwards (1998) argues, the program's credibility was eroded as real exchange rate overvaluation became more and more acute, which led firms to incorporate into their pricing decisions an expected change in the exchange rate regime and a resumption of the crawling peg policy.

Like Chile, Mexico experienced an initial decline in persistence. After a few months, however, measured inertia had once again increased. In mid-1994, seven years into the program, it was still in the range of 0.6 to 0.7. This high degree of inertia contributed to the creation of an acute real exchange overvaluation, which in turn contributed to the unleashing of the Mexican currency crisis of December 1994 (Dornbusch and Warner, 1994).

4.2 Israel

In 1983-84, Israeli inflation averaged almost 300 percent per year. This was the result of serious macroeconomic imbalances, including double digit fiscal deficits and an entrenched and perverse inflationary expectations process. In mid-1985 Israel launched an ambitious stabilization program based on three important elements: a control of the fiscal deficit; a social pact agreed among representatives of labor, government, and the private sector; and a fixed (but adjustable) nominal exchange rate. As in Chile and Mexico, the exchange rate anchor policy was aimed at breaking expectations, placing a ceiling on tradables inflation, and reducing inertia. The program was a success in many ways: in 1986 inflation was 20 percent, down from 185 percent the previous year, and by 1987 it was 16 percent.

As in Chile and Mexico, this program generated a process of real exchange rate appreciation, which hurt the degree of international competitiveness of Israeli exports. Between 1985 and 1988 this problem was handled through periodic nominal devaluations, which were followed by a new peg. In January 1989, however, a new exchange rate policy based on a crawling exchange rate band was implemented

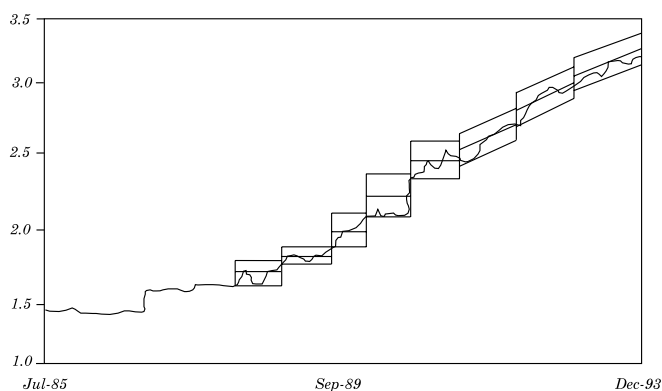
to combine the dual objectives of defeating inflation and avoiding an exchange rate crisis.

Figure 6 presents the evolution of Israel's nominal exchange rate from mid-1985 through December 1993. Initially this band allowed the exchange rate to fluctuate by ± 3 percent around a fixed central parity. The width of the band was subsequently increased to ± 5 percent, and in December 1991 the sliding band was replaced by a crawling band characterized by a moving central parity based on expected inflation rate differentials.

The move from the sliding to the forward-looking crawling band in late 1991 responded to two factors. First, the sliding band system was generating significant uncertainty. Every time the public expected that a realignment of the central parity was about to take place, speculation increased and domestic interest rates jumped. Second, by basing the rate of the crawl of the band on expected, rather than lagged, inflation, the authorities thought they would affect inflationary expectations and reduce—if not eliminate—inflationary inertia. With the exception of the first band that operated from January through June of 1989, the band width was maintained at ± 5 percent.

The adoption of the forward-looking band was successful in many respects. First, speculation on the realignment of the central parity was greatly reduced, and interest rates declined and became less volatile. Second, exchange rate overvaluation was avoided, and export performance improved significantly. Third, inflationary inertia was

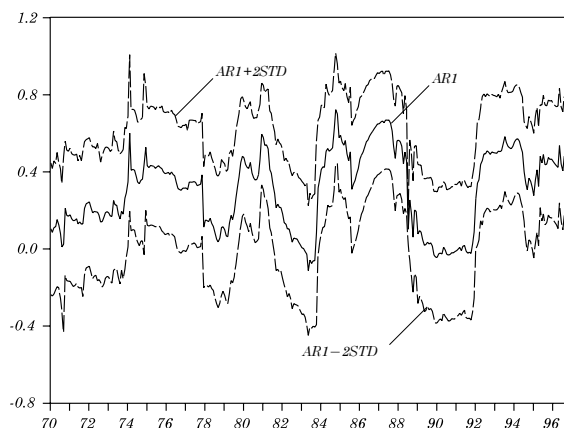
Figure 6. Israel's Exchange Rate Band



greatly reduced in the early 1990s. Fourth, as predicted by the program, inflation declined gradually through 1995. Finally, this band allowed the Israeli economy to accommodate in a nontraumatic fashion changes in fundamentals, including episodes of large capital outflows. In the last quarter of 1992, for example, capital outflows amounted to over 20 percent of international reserves.

From the perspective of this paper, an important question is how the Israeli exchange-rate-based program affected inflationary inertia. This can be addressed by reexamining the inertia measures computed above, which are reproduced in figure 7 for convenience. The figure is highly revealing and provides an interesting contrast against those for Chile and Mexico displayed above. The preferred AR(1) measure of inertia very clearly indicates that by late 1986 inflationary persistence in Israel was already experiencing a rapid collapse. In fact, it is not possible to reject, at conventional significance levels, the hypothesis of a complete absence of inertia between late 1988 and mid-1990. These estimates quite strongly suggest, however, that by 1991 the degree of inertia was once again on the rise, and by 1992-93 it was almost as high as it had been in 1983-86. The most likely explanation for this result is that with inflation in the 20 percent level for the sixth year in a row, it had become clear that the war against inflation had run to a standstill. The program suffered credibility problems, as old indexation practices slowly crept back into the system.

Figure 7. Inflationary Inertia: Israel



5. CONCLUDING REMARKS

This paper has empirically analyzed the extent of inflationary inertia in a set of countries. An important contribution of the paper has been to assume that the degree of inflationary inertia can change through time and to associate changes in the time-varying measures of inflationary inertia to specific economic events for a subset of countries. The paper has also discussed preliminary evidence regarding the effect of indexation practices and credibility on inflationary persistence.

We analyzed the time series properties of monthly inflation data from 1970 through the end of 1996. We performed a battery of tests: rolling augmented Dickey-Fuller tests; sequential Perron tests to test the unit root hypothesis in the presence of structural breaks; and sequential Quandt likelihood ratio tests to analyze the instability of the autoregressive coefficients. Results indicate that allowing for structural breaks, it is possible to reject the unit root hypothesis for all countries in the sample, and that it is also possible to reject the hypothesis of an absence of a structural break in the first-order autoregressive coefficient in Argentina, Chile, France, Israel, Italy, and Spain. In addition, the reported estimates of the rolling AR(1) coefficients and the dominant inverted roots of the rolling AR(2) estimates clearly show that inflationary inertia greatly varies both across countries and across time periods within a country. We demonstrated a clear connection between our estimates of inertia and the levels of inflation in the countries in the sample. Episodes (time and place) characterized by a high degree of inflationary persistence also feature a high level of inflation.

We ended the paper by discussing in more detail the evolution of inflationary inertia in three countries that, at one point or another, have relied on a nominal exchange rate anchor to achieve price stability: Chile, Israel, and Mexico. We have shown that the measure of inertia falls at the beginning of these stabilization programs, but then rises back as time goes by and the programs fail to cut inflation down to world levels. Unfortunately, we were not able to analyze the last Argentine stabilization program because of data constraints.

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