The Dynamics of Inflation: A Study of a Large Number of Countries

by

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Abstract

Over the last twenty years the statistical properties of inflation persistence has been the subject of intense investigation and debate without reaching a unanimous conclusion yet. In this paper we attempt to shed further light to this debate using a battery of econometric techniques in order to provide robust evidence on the degree of inflation persistence and whether this has changed during the period in which several countries have followed inflation-targeting regimes or new monetary regimes. We consider the inflation rates of thirty developed and emerging economies using quarterly data for the period 1958-2007 which include alternative monetary policy regimes. The coefficient of the inflation parameter is estimated by OLS, ARMA and ARFIMA models. Furthermore, the grid-bootstrap median unbiased estimator approach developed by Hansen (1999) is used to estimate the finite sample OLS estimates coupled with the 95% percent symmetric confidence interval. We also examine parameter stability of persistence coefficients by estimating a model with time-varying parameters and we provide evidence that the AR coefficient has remained, in most cases and for several periods, high, although there is a tendency for lower inflation persistence in the late 1990s and during the 2000s and this downturn may be the result of a shift in monetary policy. This finding is more evident for the case of the EMU countries, since the adoption of the euro.

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1. Introduction

Persistence defines the extent to which the effect of a shock persists both in terms of size and length of time. The statistical properties of inflation and its degree of persistence and stability over time is a subject of great interest and debate because the variable plays a central role in monetary policy. Furthermore, with the great moderation beginning in about 1984 an examination of the changing behavior of inflation pre- and post-great moderation is of interest. A further reason to examine the time series dynamics of inflation is that many countries have shifted towards an inflation targeting regime. There is a growing body of research suggesting that the monetary regime in place has an impact on the persistence properties of inflation, or in other words, inflation persistence is not an inherent characteristic of industrial countries. Kim and Nelson (1999), McConnell and Peres-Quiros (2000), Brainard and Perry (2000), Taylor (2000), Kim *et al*. (2004) and Carlstrom *et al*. (2009) report evidence that US inflation persistence during the Volcker-Greespan era was substantially lower than during the previous two decades. Ravenna (2000) shows a large post-1990 decline in Canadian inflation persistence; Stock and Watson (2002, 2003) demonstrated for several other countries that the Great Moderation was taken to be in the mid-1980s; Batini (2006) finds limited evidence in favour of shifts in inflation persistence for the Eurozone countries; and Benati (2008) demonstrates that the UK and US inflation had no persistence during the metallic-standard era (prior to 1914) whereas both economies exhibited very high persistence during the 1970s and very low during the last decade.

Nelson (2001) points out how monetary policy in the UK underwent several regime changes over the last 50 years; From a fixed exchange rate regime with foreign exchange controls until 1972; to free-floating with no domestic nominal anchor until 1978, followed by a system of monetary targeting until the mid-1980s; then back to exchange rate management, the period of 'shadowing' the Deutsche Mark, which culminated in the membership into the Exchange Rate Mechanism (ERM) from 1990-1992; since 1992, inflation targeting has been the official regime governing the UK monetary policy, with interest rate decisions made by the UK government in concert with the Bank of England up until May 1997 and after it by the Bank of England alone, under a new law mandating its procedures and targets. Nelson and Nikolov (2004) and Nelson (2007) examine these regime shifts. For the period as a whole there were large swings in both inflation and output growth. Inflation was in double digits during most of the 1970s. Nelson (2001) demonstrates UK output growth was lower than major trading partners in the 1960s, underwent a further slowdown after 1973, with partial recovery beginning in the early 1980s. There were recessions in 1972, 1974-75, 1979-81, and 1990-92.

There is yet to be a consensus on the most appropriate way to model the inflation rate. Two issues emerge in the macroeconomic debate. The first is how to measure the persistence of inflation accurately and the second has this persistence changed over time. The degree of inflation persistence is important in the monetary transmission mechanism and a determinant of the success of monetary policy and of obtaining a stable inflation and output simultaneously. Coenen (2006) and Angeloni *et al*. (2003) study optimal monetary policy conditional on the degree of persistence in inflation. These papers demonstrate how robust monetary policy could be designed when there is uncertainty about the correct persistence of inflation. They conclude that it would be optimal to design the monetary target assuming a high degree of inflation inertia. Finally, Taylor (1998) and Hall (1999) have pointed out that tests in the spirit of Solow (1968) and Tobin (1968) will tend to reject monetary neutrality if persistence estimates are revised downwards. Thus, an understanding of inflation dynamics has important policy implications.

During the last decade several papers have shown that there is evidence on the presence of multiple structural breaks in the mean of inflation over the post-WWII period for many countries. Thus, Kim and Nelson (1999), McConnell and Peres-Quiros (2000), Taylor (2000) Cogley and Sargent (2001), Stock and Watson (2002, 2003), Levin and Piger (2004), Corvoiser and Mojon (2005) and Rapach and Wohar (2005) argued that not controlling for shifts in the mean of a process spuriously increases its degree of persistence. Most often the Bai and Perron (1998, 2003) methodology is employed. We choose not to pursue this discrete structural shift in a series for a number of reasons. First, our unit root tests indicate that for most countries examined in our sample we cannot reject the unit root. The Bai and Perron procedure is not meant to be applied to an I(1) series. Second, Bai (1998) proves that if a series is truly I(1) and a structural break procedure is applied, one will find with certainty a structural break in the middle of the series. Third, Engle and Smith (1999) shows that a random walk has breaks every period. Fourth, our tests also indicate that inflation is highly persistence. Prodan (2008) has shown that if the Bai and Perron procedure is applied to persistent series one will find spurious breaks in the mean. We acknowledge that there is a change in persistence over time but we choose to employ an alternative approach. This approach is a continuous time varying parameter procedure as opposed to the Bai and Perron (1998, 2003) procedure that allows for only up to 5 discrete structural breaks in the mean of inflation and may in fact result in spurious breaks.[1] The advantage of using such an approach as opposed to break tests is reinforced in the cases when the true data generation process (DGP) is characterized by random-walk time variation, i.e. the notion of gradual change a stylized fact of inflationary process.

[1] One of the implications of the above analysis is that one can approximate the data generating process of inflation as either a persistent or I(1) series or a stationary series around mean shifts. If one approximates inflation as stationary around level shifts in the mean then one cannot use inflation in tests of the fisher effect where an I(1) inflation series is tested to be cointegrated, with cointegrating vector (1,-1), with an I(1) interest rate series.

The purpose of this paper is to provide insights in the ongoing debate regarding the degree of inflation persistence and its temporal stability. This task is accomplished by analyzing inflation rates for 30 different countries employing a number of different econometric techniques. Our analysis has a number of novel features and the testing procedure is conducted in a sequential order using quarterly inflation data for thirty developed and emerging economies over the period 1958-2007. First, we apply unit root tests in order to examine whether the series are either I(1) or I(0). This is of crucial importance since a stylized fact emerged from previous studies of U.S. and other major economies inflation dynamics have demonstrated that inflation is an I(1) variable over certain periods of time whereas is an I(0) process in other periods of time.[2] Therefore, we need to examine when inflation is an I(1) and when an I(0) process. Second, we estimate the AR and MA components of an ARMA for all cases. Third, given the concerns about the appropriateness of models that allow only for I(1) or I(0) processes with respect to medium- and long-term monetary policy implications we allow for a wider statistical framework and we consider long memory and fractional integration to describe the inflation dynamics. We estimate the value of the fractional d parameter with the recently developed exact local Whittle estimator (Shimotsu, 2007; Shimotsu and Phillips, 2005, 2006) which is shown to have good properties in Monte Carlo experiments. Fourth, we provide efficient estimates of the AR coefficients by applying the Hansen (1999) grid-bootstrap Median Unbiased estimator and Romano and Wolf (2001) subsampling procedures to compute the 95 percent symmetric confidence intervals for the autoregressive coefficient. Fifth, we test for changes in persistence with the application of nine

[2] Nobay *et al*. (2010) note that the unit root feature of inflation is now reflected in theoretical models of inflationary process. However, they argued that the assumption of a random walk in the inflation target implies a negative inflation target which is implausible given the zero lower bound on nominal interest. Recently, Buiter (2009) has challenged this notion and he demonstrated three ways to overcome the zero lower bound and therefore the possibility of negative nominal interest rates. In addition, Nobay *et al*. (2010) argued that the inflation behaves as a near unit root process for rates near the inflation target but it is mean reverting for large deviations.

test statistics recently developed by Harvey *et al*. (2006). Finally, we consider the case of time-varying coefficients in our model and we follow Beechey and Osterholm (2007, 2009) in order to examine the temporal stability of our results.[3]

The main findings of our analysis are summarized as follows. First, we show that for most cases the inflation series are I(1) processes with only a few cases exhibiting stationary behaviour. Second, we fit an ARMA model to the data and we found that for all cases the sum of the AR coefficient is close to one indicating a high degree of inflation persistence in inflation in both developed and emerging economies. Third, we consider the case of fractional integration and we found that the values of the estimated coefficient d is between 0.5 and 1 indicating that I(1) and I(0) processes may be rejected and therefore their consideration may lead to persistence overestimation. Germany, Greece, Malaysia, Netherlands, Philippines and Singapore are the only cases with an estimated value of d below 0.5. Fourth, we provide efficient estimates for the AR coefficient and compute 95 percent confidence intervals for ρ using the Hansen (1999) and Romano and Wolf (2001) methodologies. Of the 30 countries examined 17 have point estimates of the persistence measure greater or equal to 0.85 and confidence intervals that range between 0.75 and above 1.00. Fifth, when we looked at changes in persistence the application of the suggested testing procedure leads to the conclusion that we only receive clear-cut evidence of no change in inflation persistence for the case of Belgium, China, Denmark, France, Germany, Greece, Korea, Luxembourg, Malaysia and Thailand. Finally, we examine parameter stability of persistence coefficients by estimating a model with time-varying parameters and we provide evidence that the AR coefficient has remained in most cases and for several periods high

[3] We chose to work with alternative univariate models for two reasons. First, in studying individual features of a time series, like persistence or volatility, using more sophisticated specifications such as SVAR model or other type of structural models will not give us more information. Second, it may well be the case that the use of a more complicated model will most likely lead to the emergence of more questionable assumptions and therefore an increased risk associated with them. Therefore, we prefer to work with the minimum econometric framework.

although there is a tendency for lower inflation persistence in the late 1990s and during the 2000s and this continuous downturn may be the result of a shift in monetary policy leading to the possible existence of a structural break. This finding is more evident for the case of the EMU countries since the adoption of the euro.

The rest of the paper is organized as follows. Section 2 provides a literature review. In section 3 we present and discuss the econometric methodology applied. Section 4 reports our empirical results with section 5 providing our conclusions and policy implications.

2. Literature Review

For univariate time series there is no consensus as to what measure one should use for persistence (see, Phillips, 1991; Andrews, 1993; Andrews and Chen, 1994; Marques, 2004; Murray and Papell, 2002; and Pivetta and Reis, 2007). If one assumes an AR process then measures of persistence most frequently used are the sum of the AR coefficients, the largest root, the half-life (i.e. the number of periods that inflation remains above 0.5 for a unit shock). For an ARMA process, the first two measures mentioned above are not appropriate as they ignore the MA coefficients. One other measure is the impulse response function. There is a large amount of evidence that the postwar inflation exhibits high persistence in developed countries. Some examples include, Pivetta and Reis (2007) for the US and O'Reilly and Whelan (2005), Batini (2006) and Beechey and Osterholm (2007) for the euro zone. However, the aforementioned result may be sensitive to the econometric techniques employed. The observed persistence may be due to the existence of structural change that are not accounted for, possibly from changes in the inflation targets of central banks, different exchange rate regimes, or shocks to key prices (see

Levin and Piger, 2004).[4]

There are a number of explanations for inflation persistence. One argument is that persistence is the result of aggregating prices from heterogeneous firms in their price adjustment costs. Buiter and Jewitt (1981), Fuhrer and Moore (1995), Fuhrer (2000), Calvo *et al*. (2003) and Christiano *et al*. (2005), all assume that high inflation persistence results from the structure of nominal contracts. Rotemberg and Woodford (1997), Dittmar *et al*. (2005), and Ireland (2004) generate persistence through the data generating process of the structural shocks hitting the economy. Zeira (1989) argues that inflationary inertia is associated (a) with positive autocorrelation across time; (b) that following monetary disinflations the decline of the inflation rate is gradual; and (c) supply shocks have a lasting effect on the increase of price level. Given these stylized facts then inflationary inertia is demonstrated to exist in a wage-price spiral framework. Dixon and Kara (2010) high inflation persistence can be explained with the use of microdata on the behaviour of prices for both the U.S. and the Eurozone within a model of nominal wage or price in which the specific distribution of the length of contracts can be explicitly modeled. An alternative view is that the degree` of inflation persistence is not an inherent structural characteristic of industrialized economies, but instead is a function of the monetary policy regime (see West, 1988).

Some authors have found evidence of a decrease in the persistence of inflation; (Taylor, 2000; Cogley and Sargent 2001; Kim *et al*., 2004). Other authors employing different econometric techniques found inflation persistence to be unchanged (Stock, 2001; Batini, 2006; Levin and Piger, 2004; O'Reilly and Whelan, 2005; Pivetta and Reis, 2007). Gadea and Mayoral (2006) estimate fractionally integrated models and compare them with other specification such as

[4] The existence of regimes that are not explicitly taken into account may lead to spurious persistence (see Perron, 1989).

ARMA and ARIMA models[5]. They employ both classical and Bayesian techniques. They find that if ARIMA models are used to measure persistence they will overstate persistence. The main results of their paper are that once fractional integration is allowed for, I(0) and I(1) specifications are rejected. They find that for most of the countries the fractionally integration model is preferred to the I(0) specification.[6] Inflation rates are estimated with different techniques and they show that inflation is best characterized as a fractionally integrated proceses with a memory parameter between 0.6 and 0.8. This implies that inflation rates are very persistent nonstationary; however, as opposed to the I(1) variable, shocks have a non-permanent effect, and thus, the series is mean reverting. They find important differences across countries. According to the half-life (HL) measure US inflation is most persistent. Inflation persistence in Central and Nordic European countries have the lowest degree of persistence. They find no change in persistence for the majority of the countries they examine.

3. Methodology

3.1. Unit Root Tests

The standard practice is to subject variables to a battery of unit root tests. There are many different tests for a unit root in the autoregressive (AR) polynomial of a univariate process that have been proposed, but the most common is the augmented Dickey-Fuller (ADF) test proposed by Said

[5] Fractional integration can appear in inflation as a result of aggregating prices from heterogeneous firms in their price adjustment costs.

[6] Fractionally integrated processes and I(0) process with structural breaks look very similar. Testing for the difference is difficult.

and Dickey (1983). It is based on the AR approximation of a general ARIMA process and is given in (1).[7]

$$\Delta y_t = \alpha + \beta t + (1-\rho) y_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + e_t \tag{1}$$

The null hypothesis of a unit root can be tested by estimating (1) using OLS and then using a t-type test statistic to test the hypothesis $(1 - \rho) = 0$. The choice of the lag truncation parameter k is important for the small sample properties of the test because when the number of lags is greater than the true number of lags there is a decrease in the power of the test, while two few lags leads to under sized tests. There are some potential problems with unit root testing using (1), however.

The first problem is low power of the test relative to local alternatives. Elliot *et al*. (1996) (ERS) proposed an estimator that increases the power of the unit root test substantially by using a GLS detrending procedure. One can motivate the unit root tests using the DGP in (2)

$$y_t = d_t + u_t, \qquad u_t = \rho u_{t-1} + v_t \tag{2}$$

where $v_t = \varphi(L) e_t = \sum_{j=0}^{\infty} \varphi_j e_{t-j}, d_t = \zeta' z_t = \sum_{i=0}^{p} \zeta_i t^i \text{ for } p = 0, 1.$ When estimating equation (1) the parameters of the deterministic components are estimated via OLS and are treated as nuisance parameters in the distribution of the unit root tests. By estimating these nuisance parameters using OLS the power of the test statistics is diminished. ERS propose a weighted least squares or GLS method to estimate these parameters and then detrend the data prior to testing for a unit root. For series $\{x_t\}_{t=0}^{T}$ define $(x_0^{\overline{\alpha}}, x_t^{\overline{\alpha}}) = (x_0, (1 - \overline{\alpha} L) x_t)$ for some value $\overline{\alpha} = 1 + \overline{c}/T$. The GLS detrended series is then defined as $\widetilde{y}_t \equiv y_t - \widehat{\zeta}' z_t$ where $\widehat{\zeta}$

[7] An ARIMA or autoregressive integrated moving average process assumes that a time series can be modeled in the time domain as a function of lagged values of itself and current and lagged values of the innovation or error to the process. An ARIMA(p,d,q) takes the general form $\phi(L)\Delta^d y_t + \mu = \theta(L)\varepsilon_t$ where the autoregressive lag polynomial $\phi(L) = 1 + \phi_1 L + \phi_2 L^2 + \ldots + \phi_p L^p$ is of order p, the order of integration is given by the differencing parameter d and the moving average polynomial $\theta(L) = 1 - \theta_1 L - \theta_2 L^2 - \ldots - \theta_q L^q$ if of order q.

minimizes $S(\overline{\alpha}, \zeta) = (y^{\overline{\alpha}} - \zeta' z_t^{\overline{\alpha}})'(y^{\overline{\alpha}} - \zeta' z_t^{\overline{\alpha}})$. ERS suggest imposing $\overline{c} = -7.0$ for $p = 0$ and $\overline{c} = -13.5$ for $p = 1$.[8] Testing for a unit root can then be done by estimating equation (3) using OLS and calculating a t-type test statistic as in (1), which is refereed to as the DF-GLS$^{\mu}$ statistic when p = 0 and DF-GLS$^{\tau}$ when p = 1.

$$\Delta \widetilde{y}_t = (1 - \rho)\widetilde{y}_{t-1} + \sum_{j=1}^{k} \gamma_j \widetilde{y}_{t-j} + e_{tk} \tag{3}$$

Although low power is always a problem for unit root tests, another concern is that size distortions in the tests may be a problem because of the properties of the underlying data generating process (DGP). One source of size distortion is the presence of large and negative moving average (MA) parameters in the DGP. Schwert (1987) was one of the first to point out that standard unit root tests like the ADF are severely oversized when there are large negative MA terms in the DGP. He suggests increasing the value of *k*, the lag truncation parameter in (1) and (3), to more accurately allow the AR process in (1) to approximate the MA components in the ARIMA. We estimate ARIMA models for each of the series of interest in this study in order to gauge how serious this source of size distortion may be in our application. Table 1 displays estimation results for inflation.[9]

Two features of many economic time series tend to affect the size and power of usual unit root tests. In particular, a large negative moving average root may induce size distortions, while a large autoregressive root may result in low power. When this is the case it is preferred to apply the MZ_{α}, MZ_t, MSB and the MPT tests due to Ng and Perron (2001), which are precisely designed to overcome both size distortion and low power problems when the data are characterized by these features. These tests are extensions of the *M* tests of Perron and Ng (1996)

[8] Cook (2006) finds that the power of the tests in finite samples under alternative DGPs can be increased with alternative values for $\overline{c}$.

[9] We used the Box-Jenkins procedure to identify several candidate models for each series and then chose the best fitting model based on residual serial correlation tests, significance of the parameter estimates and R^2.

that use Generalized Least Squares (GLS) detrending of the data, together with a modified information criterion for the selection of the truncation lag parameter.

Ng and Perron (2001) have developed a modified information criterion that chooses *k* in (1) or (3) in a way that mitigates the size distortion in unit root tests. It is based upon an autoregressive estimate of the long-run variance of y_t, denoted s^2_{AR}. This estimate is calculated as

$$s^2_{AR} = \frac{\widehat{\sigma}^2_k}{[1-\gamma(1)]^2} \tag{4}$$

where $\gamma(1) = \sum_{i=1}^{k} \gamma_i$ and $\widehat{\sigma}^2_k = (T-k)^{-1}\sum_{t=k+1}^{T} \widehat{e}^2_{tk}$ and γ_i and $\{\widehat{e}_{tk}\}$. The parameters can all be estimated from equation (3) using OLS.[10] The modified information criteria (MIC) is given as

$$MIC(k) = \ln(\widehat{\sigma}^2_k) + \frac{C_T(\tau_T(k)+k)}{T-k_{\max}} \tag{5}$$

where $\tau_T(k) = (\widehat{\sigma}^2_k)^{-1}\widehat{\rho}\sum_{t=k_{\max}+1}^{T} \widetilde{y}^2_{t-1}$ and $k_{\max}$ is the largest lag truncation considered. When $C_T = \ln(T-k_{\max})$, equation (5) represents the modified Bayesian information criterion (MBIC) and when $C_T = 2$ it is the modified Akaike information criterion (MAIC).

Ng and Perron (2001) also suggest using three tests that have less size distortion in the presence of MA errors than standard tests. These tests are $MZ_\rho, MZ_t,$ andMSB, collectively referred to as the M-tests. The tests are calculated from estimates of (3) as follows:

$$MZ_\rho = (T^{-2}\widetilde{y}^2_t - s^2_{AR})(2T^{-2}\sum_{t=1}^{T} \widetilde{y}^2_{t-1})^{-1} \tag{6}$$

and

[10] Perron and Qu (2007) suggest that small sample power can be improved if the parameters used to construct the estimate of the long-run variance are estimated from equation (1) rather than (3).

$$MSB = \left[\frac{T^{-2} \sum_{t=1}^{T} \widetilde{y}_{t-1}^{2}}{s_{AR}^{2}} \right]^{\frac{1}{2}} \tag{7}$$

and $MZ_t = MZ_\rho \times MSB$.

Another potential source of size distortion in the unit root tests is the conditional heteroskedasticity that characterizes the variances of each series. Kim and Schmidt (1993) were the first to analyze the effects of conditional heteroskedasticity on unit root tests. More recently Cook (2006) conducts a Monte Carlo analysis that finds that even the modified unit root tests above are subject to some size bias when the errors in the unit root testing equation follow a GARCH(1,1) process. The size of the unit root test increases as the degree of volatility of the GARCH process increases. To get some idea of how much this GARCH effect has on the data one would estimated the ARIMA models for each series allowing for GARCH errors. The GARCH specification is given as,

$$h_t^2 = \phi + \alpha u_{t-1}^2 + \beta h_{t-1}^2 \tag{8}$$

where h_t^2 is the conditional variance of the errors u_t from the ARIMA specification. The degree of volatility in the variance is measured by the parameter α and the persistence in the conditional variance is measured by $\alpha+\beta$. Cook (2006) finds that for any significant size distortion to occur, the volatility parameter α must be larger than 0.25.

We find that the ARIMA parameters are not much affected by including the GARCH effects and the GARCH volatility is not high enough to induce serious size bias in our unit root tests of inflation.

3.2. Long-memory and Fractional Integration

The exact Whittle estimator was proposed in Phillips (1999) and its asymptotic theory was developed in Shimotsu and Phillips (2005, 2006). For a fractional process defined as:[11]

$$x_t = (1-L)^{-d} w_t = \sum_{k=0}^{t-1} \frac{\Gamma(d+k)}{\Gamma(d)\Gamma(k+1)} w_{t-k} \tag{9}$$

where $\{w_t\}$ is a weakly dependent process with continuous spectral density, exact Whittle estimation of the memory parameter d involves maximizing the following Whittle log-likelihood function:

$$Q_m(G,d) = \frac{1}{m} \sum_{j=1}^{m} \left(\log G\lambda_j^{-2d} + \frac{1}{G} I_{\Delta^d(x)}(\lambda_j) \right) \tag{10}$$

where G is a positive constant, $I_{\Delta^d(x)}(\lambda_j)$ is the periodogram of $(1 - L)^d x_t$ with the fractional filter $(1 - L)^d$ defined in the same way as in Robinson (1994) and Phillips (1999), and m is a bandwidth parameter satisfying $m/n + 1/m \rightarrow 0$ so that the band $\{\lambda_j = 2\pi j/n, j = 1,2,...,m\}$ concentrates on the zero frequency as the sample size $n \rightarrow \infty$.

Shimotsu and Phillips (2005, 2006) show that the exact Whittle estimator $(\hat{G}_{EW}, \hat{d}_{EW})$ is consistent and that $\hat{d}_{EW}$ has the following limiting distribution as $n \rightarrow \infty$:

$$\sqrt{m}(\hat{d}_{EW} - d) \xrightarrow{d} N(0, 1/4) \tag{11}$$

[11] Two main approaches have been used in the literature to define a fractional process x_t. The first, which is adopted in Hosking (1981), among others, defines a stationary fractional process as an infinite order moving average of innovations: $x_t = \sum_{k=0}^{\infty} \Gamma(d+k)/\Gamma(d)\Gamma(k+1) w_{t-k}$ and defines a nonstationary $I(d)$ process as the partial sum of an $I(d-1)$ process (Velasco, 1999a,b). The second, which is used in Robinson (1994) and Phillips (1999), truncates the fractional difference filter and defines $x_t = \sum_{k=0}^{t-1} \Gamma(d+k)/\Gamma(d)\Gamma(k+1) w_{t-k}$ for all values of d. For a more detailed discussion of the definitions and their implications, see Shimotsu and Phillips (2006).

for all values of *d*. The robustness of the asymptotic properties of $\hat{d}_{EW}$ is especially appealing for practical work when the domain of the true order of fractional integration is controversial. The EQ estimate also provides guidance on the order of the fractional difference that can render the data stationary.

We estimate the value of the fractional d parameter, *d*, for inflation with the Shimotsu (2007) semiparametric two-step feasible exact local Whittle estimator that allows for an unknown mean in the series. This estimator refines the Shimotsu and Phillips (2005, 2006) exact local Whittle estimator, and these authors show that such local Whittle estimators of *d* have good properties in Monte Carlo experiments. Shimotsu and Phillips (2005, 2006) propose the exact Whittle estimators of long memory parameters as a general purpose estimation procedure. As no short-run dynamics need to be specified, the semiparametric estimators are robust to its misspecification. They show that their exact local Whittle estimators are consistent and asymptotically normally distributed. It is called exact because the procedure is based on the transformation of the Whittle likelihood function with a purely algebraic manipulation that hold exactly for any value of d.

We employ the conditional sum of squares (CSS) estimator to compute d, p, and q jointly in the ARFIMA specification which is similar to the Whittle estimator. Some of the properties of the CSS estimator are discussed in Chung and Baillie (1993). They show that the effect of initial observations is asymptotically negligible so that the CSS estimator is asymptotically equivalent to Maximum Likelihood Estimation (MLE). Chung and Baillie (1993) also report simulation results which show that the CSS estimator performs well when estimating ARFIMA(p,d,q) models with p and q being less than 2 and -0.5< d <0.5 and for sample sizes greater than 100. The CSS estimator can be regarded as being a time domain alternative to the

Fox and Taqqu (1986) frequency domain estimator of the AFIMA model, which is known to also be approximately Maximum Likelihood under Normality.

3.3. Tests for a Change in Inflation Persistence[12]

Testing for the presence of a unit root is now routine practice among practitioners analyzing the stochastic properties of macroeconomic time series. This practice is oriented towards the classification of series as stationary or nonstationary. Establishing this distinction is meaningful for several reasons. For the purposes of the present article, the most important one is that it helps in understanding the effect of shocks to macro variables; while the impact of such shocks will be transitory for a stationary series, for a nonstationary one any random shock may have persistent effects. In other words, while an $I(0)$ time series will display mean-reverting behaviour, an $I(1)$ variable will be persistent, i.e. shocks to it will have long-lasting effects, thus preventing the series from returning to any defined level.

It has been observed in recent years, however, that macroeconomic variables - such as the inflation rate - may display both stationary and nonstationary features within a specific period. Indeed, it seems some series could be switching from $I(0)$ to $I(1)$ behaviour, or vice versa.

To test for changes in the degree of persistence, we apply nine test statistics recently developed by Harvey *et al.* (2006) (HLT henceforth), which follow the work of Kim (2000), Kim *et al.* (2004) and Busetti and Taylor (2004). The model underlying the test statistics proposed by HLT is the following:

$$y_t = x_t' \beta + u_t \tag{12}$$

[12] This section draws heavily from Harvey *et al*. (2006).

$$u_t = \rho_t u_{t-1} + \varepsilon_t, \quad t = 1,...,T$$

where y_t is the inflation rate vector x_t contains either a constant, or a constant and a linear trend, and ε_t, is mean zero satisfying assumptions by Phillips and Perron (1988).

The null hypothesis states that the inflation rate is stationary, i.e. y_t, is $I(0)$. In this setting, $\rho_t = \rho, |\rho| < 1, \quad t = 1,...,T$ in model (1). This hypothesis is denoted by H_0. In testing for a change in persistence, HLT allow for two different alternative hypotheses. The first corresponds to a change from 1(0) to I(1), denoted H_{01}, and the second to a change from $I(1)$ to $I(0)$, denoted H_{10}. Specifically,

$$\begin{aligned} H_{01}: \rho_t &= \rho, |\rho| < 1 \quad \text{for } t \le \lfloor \tau^* T \rfloor \text{and} \\ \rho_t &= 1 - \bar{\alpha}/T, \quad \text{for } t > \left[\tau^* T\right] \end{aligned}$$

$$\begin{aligned} H_{10}: \rho_t &= 1 - \bar{\alpha}/T, \quad \text{for } t > \lfloor \tau^* T \rfloor \text{and} \\ \rho_t &= \rho, |\rho| < 1 \quad \textit{for } t \le \left[\tau^* T\right] \end{aligned}$$

where $\bar{\alpha} \ge 0$ allows for a local to unit root, and τ^* denotes the unknown proportion of the sample size where the change in persistence occurs. τ^* is assumed to belong to the interval $\Lambda = [\tau_1, \tau_u] \in (0,1)$, where τ_1, τ_u stand for (arbitrary) lower and upper values for τ^*. Given the preliminary analysis presented above for our data set, the empirical applications below will concentrate in testing H_{01} against H_{10}.

The various tests to be applied are based on the following ratio introduced by Kim (2000), designed to test H_0 against H_{01}:

$$K[\tau T] = \frac{(T = [\tau T])^{-2} \sum_{t=[\tau T]+1}^{T} \left(\sum_{i=[\tau T]+1}^{t} \hat{u}_{i,\tau} \right)^2}{[\tau T]^{-2} \sum_{t=1}^{[\tau T]} \left(\sum_{i=1}^{t} \hat{u}_{i,\tau} \right)^2} \tag{13}$$

where $\hat{u}_{i,\tau}$ in the numerator (denominator) is the residual from applying Ordinary Least Squares (OLS) to Model (1) for $t = [\tau T]+1,...,T(t = 1,...,[\tau T])$. Note from Equation 13 that, under H_0, the sums in numerator and denominator should be equal. In order to test for a change in persistence (H_0 against H_{01}) Kim (2000), Kim *et al.* (2004) and Busetti and Taylor (2004) consider the following three statistics, all functions of the ratio defined above:

$$M(S) = T_*^{-1} \sum_{[t=\tau_1 T]}^{[\tau_u T]} K_t \tag{14}$$

$$M(E) = \ln T_*^{-1} \sum_{t=\tau_l T}^{\tau_u T} \exp(0.5 K_t) \tag{15}$$

$$M(X) = \max_{\tau \in \{[\tau_l T],...,[\tau_u T]\}} K_t \tag{16}$$

Where $T_* \equiv [\tau_u T] - [\tau_l T] + 1$. These authors derive the limiting distributions of the statistics as functionals of Brownian motion processes, and show that they are pivotal (free of nuisance parameters) under the null. Equations 14-16 correspond to Hansen's (1991) mean score statistic

(S), Andrews and Ploberger's (1994) mean exponential statistic (E), and Andrews' (1993) maximum statistic (*X*), respectively. This last statistic allows estimation of the true (and unknown) change point, over the interval Λ, and is the one used in the empirical applications below for estimating the date of change.

To test H_0 against H_{10} Busetti and Taylor (2004) proposed three other tests based on the reciprocals of K_t, namely

$$M(S)^R = T_*^{-1} \sum_{[t=t_l T]}^{[\tau_u T]} K_t^{-1} \tag{17}$$

$$M(E)^R = ln\ T_*^{-1} \sum_{t=\tau_l T}^{\tau_u T} \exp\left(0.5 K_t^{-1}\right) \tag{18}$$

$$M(X)^R = \max_{\tau \in \{[\tau_l T],\ldots,[\tau_u T]\}} K_t^{-1} \tag{19}$$

These tests are the analogous of Equations 14-16 with K_t replaced by K_t^{-1}, which we use in the empirical applications below. HLT propose six other tests, which are modified versions of Equations 9-11, with the modification being such that the critical values are precisely the same under the null and alternative hypotheses, and at the same time asymptotically equal to the unmodified statistics. These modified statistics are the following

$$M(Z)_m^R = \exp\left(-bJ_{1,T}\right) M(Z)^R \tag{20}$$

$$M(Z)^{R}_{m\,\min} = \exp\left(-bJ^{R}_{\min}\right)M(Z)^{R} \tag{21}$$

for $Z = S, E, X$. In the statistics implied by Equation 20 and Equation 21, b is a finite constant, chosen so that the modified tests are asymptotically correctly sized under H_0 (values for b are provided in Table 2 of HLT for all nine reciprocal statistics), and $J_{1,T}$ is T^{-1} times the Wald statistic (W) for testing the joint hypothesis $\gamma_{k+1} = \cdots = \gamma_9 = 0$ in the regression

$$y_t = x_t^{'}\beta + \sum_{i=k+1}^{9} \gamma_i t^i + \text{error} \tag{22}$$

For t=1,…, T. For the three statistics in Equation 21, $J^{R}_{\min} = \min_{\tau\in\Lambda} J_{[\tau T],T}$ and $J_{[\tau T],T}$ is $T^{1}W$ for testing $\gamma_{k+1} = \cdots = \gamma_9 = 0$ in Equation 22, for $t = [\tau T] + 1,..., T$. Critical values (both finite samples and asymptotic) for all nine statistics (Equations 14-19 and the 3 associated with Equation 20) to be applied in the next section for testing H_0 against H_{10}, are reported in Table 1 of HLT.

4. Data and Estimation Results

We conduct our analysis with the use of quarterly consumer price index data for the period from the first quarter of 1958 to the last quarter of 2007 for a group of thirty countries. The data have been obtained from the International Financial Statistics of the International Monetary Fund. The countries included in the study are Argentina (ARG), Australia (AU), Austria (AT), Belgium (BE), Brazil (BRA), Canada (CA), Chile (CHL), China (CHI), Denmark (DK), Finland (FI), France (FR), Germany (GE), Greece (GR), Ireland (IR), Italy (IT), Japan (JP), Korea (KO), Luxembourg (LX), Malaysia (MAL), Mexico (MEX), Netherlands (NL), New Zealand (NZ), Philippines (PHL), Portugal (PT), Singapore (SNG), Spain (SP), Sweden (SWE), Thailand (THI), United Kingdom (UK) and the United States (USA). All series are in natural logarithms. The inflation rates are computed as $\pi_t^i = \ln P_t^i - \ln P_{t-1}^i$, where π_t^i is annualized quarterly inflation. To construct the inflation rates we first seasonally adjusted the price series for each country using the X12 quarterly seasonal adjustment method of the U.S. Census Bureau.[13]

The testing procedures in this paper are carried out in a sequential order. First, unit root tests are applied to the data, in order to establish the apparent degree of integration of the series. Next we report AR(1) coefficients and ARMA models estimation results and we then estimate values of the long-memory parameter, d. The fractional integration parameter d is estimated alone and then jointly with AR and MA coefficients in an ARFIMA model. Moreover, in order to uncover any possible change in the degree of persistence, we implement the newly developed tests for a change in persistence. Finally, time-varying estimations of the basic model are provided in order to examine whether inflation persistence has changed overtime as a result of changes in the monetary regimes. This analysis is also coupled with further evidence provided by rolling regressions and recursive estimates.

Table 1 reports the results of the four unit root tests discussed in the methodology section applied to inflation rates for thirty countries. Eight of the thirty countries examined have two out

[13] Most of the previous studies on inflation dynamics and inflation persistence have been conducted for the case of the U.S. economy and several other industrialized countries. They used the GDP deflator or Harmonized CPI for the case of the EU countries since these price series are reported seasonally adjusted as well. However, in our case there is no consistent dataset that provides the CPI or GDP deflator for all the countries of our sample as seasonally adjusted. Therefore, we are forced to apply the X12 seasonal adjustment method of the U.S. Census Bureau which is widely accepted.

of four tests rejecting the null of unit root. These countries are Argentina, Australia, Chile, Finland, Germany, Netherlands, Thailand and the United States. Thus, in the majority of cases inflation rates are found to be nonstationary I(1) series.

We next estimate ARMA(p,q) models. We select ARMA(1,1) models.[14] We find that 22 countries have AR coefficients greater than 0.9. Of those 22 countries, 9 have MA coefficients greater than 0.7, suggesting that there may be size distortions in the unit root tests for these countries. The last column of Table 2 reports the value of the fractional d parameter using the exact local Whittle estimator. Recall, that a value of d less than 0.5 indicates a stationary series. A value of d greater than 0.5 indicates nonstationarity but mean reversion. Be clear that the value of d is estimated separately from AR and MA coefficients. There is not a one-to-one correspondence between the value of d and the value of the AR coefficient. That is, there are times when a large AR coefficient is associated with a small value of d. The only countries that have values of d below 0.5 are Germany, Greece, Malaysia, Netherlands, Philippines, and Singapore.

Table 3 reports estimates of the AR(1) coefficient from an OLS regression, and estimates of AR, MA, and d, from estimating an ARFIMA model. The estimator is a constrained sum of squares (CSS) estimator which is a time domain estimator and is asymptotically equivalent to MLE. The CSS estimator is sometimes referred to as the conditional maximum likelihood estimator (CMLE). We first turn our attention to the low values of the AR(1) coefficient in Table 3. Countries with the value of AR(1) coefficient less than 0.2 include Austria, Germany, Greece and the Netherlands. Of those countries, only Germany and the Netherlands were found

[14]Based on the Akaike and Swarch information criteria we selected an ARMA (1, 1) specification for all countries although for some cases a higher order ARMA (p,q) specification could also be selected. However the qualitative results derived by these alternative specifications were shown to be very similar to the ones obtained by the estimated ARMA (1, 1) models.

to be stationary according to the unit root tests of Table 1. What can explain why the other series are found to be a unit root? If we look back at the ARMA models in Table 2 some information can reveal a possible answer. In the case of Austria, Germany, Greece, and the Netherlands all the AR and MA coefficients are almost equal and opposite in sign. They thus cancel each other out to some degree and when only an AR(1) is estimated, the coefficient is very small. Looking at the ARFIMA model results and comparing them to the value of d in Table 2, we can see that these values are within one standard deviation of each other. It is not clear which is the preferred method to estimate the fractional parameter d, alone, or in within a ARFIMA model framework.

The OLS estimates of the AR parameter measuring inflation persistence although informative are not efficient estimates in finite sample. In finite sample it is well known that the asymptotic distribution for OLS estimations as well as the t-statistics in autoregressive models become systematically less accurate approximations of the true finite sample distributions as the true value of the AR coefficient increases in value. To correct this problem we apply the Hansen (1999) grid-bootstrap Median Unbiased (MUB) estimator of the sum of the autoregressive coefficients and Romano and Wolf (2001) subsampling procedures to compute the 95 percent symmetric confidence intervals for the autoregressive coefficient which have correct first-order asymptotic coverage.

Table 4 presents our estimates for our sample of 30 countries and compute 95 percent confidence intervals for ρ. Of the 30 countries examined 17 have point estimates of the persistence measure greater or equal to 0.85 and confidence intervals that range between 0.75 and above 1.00. Specifically we observe that the Euro area countries have point estimates which are substantially high with the exception of Germany and the Netherlands for which the AR coefficient is 0.46 and 0.17 (the lowest value in the dataset) respectively. The corresponding 95%

upper bound of the estimated confidence intervals is shown to be greater or equal to one the exception being again Germany and the Netherlands and this implies that for most of the European economies included in the sample inflation exhibits high persistence.[15] The evidence of low inflation persistence in Germany and the Netherlands may be the outcome of the adoption of low inflation target by the monetary authorities of the countries over the period under examination. For the case of the UK, Sweden, the US, Canada, Japan, Australia and New Zealand the median-unbiased point estimates range from 0.80 to 0.86 with the upper bound of the confidence interval higher that one in this case as well, which implies that the median-unbiased representation of the inflation in these major economies is a unit root without drift. Looking into the Southeast Asia we find that the inflation persistence in Korea, Malaysia, Philippines and Singapore and Thailand is moderate although the value of the upper end of the confidence interval ranges from 0.67 to 0.90. In contrast China exhibits high inflation persistence with a point estimate of ρ equal to 0.90 with the upper and lower bounds of the corresponding confidence interval consistent with a high degree of persistence as well. Finally, the Latin America economies of Argentina, Brazil, Chile and Mexico exhibit high degree of inflation persistence as the large values of the AR coefficient and the corresponding confidence intervals indicate.

We now turn to tests for changes in persistence of the inflation series. The null hypothesis is no change in persistence (it could be I(0) through the sample or I(1) throughout the sample). These results are reported in Table 5. The first three columns of Table 5 have test statistics to test the null of no change in persistence against the alternative of a change from I(0) to I(1). (The first column provides a date of the break point of change in persistence). The second set of three

[15] Our sample does not include the recent members of the eurozone, namely, Cyprus, Malta, Slovakia and Slovenia.

columns have test statistics to test the null no change in persistence against the alternative of a change from I(1) to I(0). The third set of three columns have test statistics to test the null of no change in persistence against the alternative of either I(0) to I(1) or vice versa. For the countries of Argentina, Australia, Austria we cannot say anything about a change in persistence as all tests statistics are significant and the null is rejected in all 9 cases. However, for Belgium we find that we cannot reject the null of no change in persistence. The second set of test statistics are significant and we reject the null hypothesis of no change in persistence against the alternative of change from I(1) to I(0). We cannot say anything about Brazil, Canada, Finland, Ireland, Italy, Japan, Mexico, Netherlands, New Zealand, Philippines, Portugal, Singapore, Spain, Sweden, UK, and US as all test statistics are significant and reject the null. China, Denmark, France, Germany, Greece, Korea, Luxembourg Malaysia, Thailand, have test statistics that cannot reject the null of no change in persistence. They also reject the no change in persistence null against the alternative of a change from I(1) to I(0).

The tests we employed so far rely on the null hypothesis of the autoregressive parameter stability but not for any time varying behavior. The issue of parameter stability is an important one for the designing of optimal monetary policy since structural breaks and switching regimes in the mean and/or the variance must be taken into consideration by the monetary authorities. What is important to note is that even in the case that the application of formal tests for structural breaks leads to the inability to reject the null of no existence structural breaks this evidence does not necessarily implies that such a regime switch has not occurred. As we have already discussed if there is a decrease in inflation persistence then tests of the natural rate hypothesis of the type suggested by Solow (1968) and Tobin (1968) may lead to the rejection of money neutrality because of this decrease. Furthermore, as Levin and Piger (2004), Benati (2008) and Cogley and

Sbordone (2008) argue that if we consider an indexation parameter of inflation persistence into macroeconomic models such as the type of New Keynesian Phillips curve then if persistence is not constant overtime this will be consistent with Lucas' critique.

In order to take account the issue of time-varying inflation persistence we adopt the recently developed approach by Beechey and Österholm (2007, 2009), and we estimate an AR(1) model allowing the autoregressive parameter to follow a random walk process.[16] An important element of this approach is that time-variation may be due to changes in the structural characteristics of the economy or to changes in monetary regimes.[17] We estimate our model with maximum likelihood using the Kalman filter approach, with the autoregressive term treated as the unknown state variable.[18] The results of estimating this model are shown in Figures 1.1-1.5, where the time-varying autoregressive parameter is plotted together with two root mean squared error bands. These plots reveal that there is substantial time-variation in the autoregressive coefficient.[19] Moreover, while the coefficient is generally large and suggestive of persistent behavior it is typically in a range below one although it can be observed that the parameter does

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[16] Based on the Akaike and Swarch information criteria we selected an AR (1) specification for all countries although for some cases a higher order AR (p) specification could also be selected. However the qualitative results derived by these alternative specifications were shown to be very similar to the ones obtained by the estimated AR (1) models.

[17] An alternative approach to study time-variation in parameters is the use of split samples and rolling regressions. Following O'Reilly and Whelan (2005), Batini (2006) and Pivetta and Reis (2007) we also applied rolling regressions to examine the time-varying behaviour of the coefficient of inflation persistence. To save space our results are available upon request. In addition we also applied recursive estimation of the AR coefficient and we found that the overall results are consistent with those obtain from rolling regressions. These results are also available upon request.

[18] Stock and Watson (1996, 1998) proposed an alternative estimation method the Time Varying Parameter-Median Unbiased Estimation which corrects for possible distortion in estimates. However, in our case no distortion is estimates was detected.

[19]The estimated equation suggested by Beechey and Österholm (2007, 2009) features homoskedastic disturbances. There are some studies, Cogley and Sargent (2005) and Sims and Zha (2006) which found evidence of heteroskedasticity in the U.S. inflation rates. However, our testing for conditional heteroskedasticity provided no evidence at the 5 percent critical value for the presence of conditional heteroskedasticity for all cases.

approach the value of one and sometimes exceeds the value of unity.[20]

Looking into the plots for the case of the Euro area countries we observe that the time-varying pattern of the coefficient of inflation persistence is mixed before the third stage of EMU. Specifically we observe that in most countries the AR parameter was substantially large with values exceeding one for most of the period since the early 1970s until the mid-1990s although there periods in which the coefficient of inflation persistence was well below one possibly as the result of the adoption of deflationary monetary policy since the mid-1980s (see Austria, Belgium Denmark, Luxemburg and Spain). For Germany we note that the AR coefficient has large values well above one in the early 1990s and this may be attributed to the reunification of Germany that resulted in budget deficits since 1993 though the inflation-targeting monetary policy has led to moderate inflation persistence. A common feature for all eurozone countries is that a substantial decrease in inflation persistence is observed since the mid 1999 marking the adoption of the euro and the adoption of the common monetary policy with an inflation target below but close to 2 percent.

Turning now our attention to the case of the UK it appears that inflation persistence remained high and above one in several periods, although there was a marked decline in inflation and inflation persistence during the 1980s following the monetary policy adopted by the Thatcher administration. Inflation persistence has returned to a substantially high level in the recent years. In Sweden inflation persistence has remained high through our sample. For the US we note that during the 60s and 70s when Martin, Burns and Miller where sequentially Chairmen of the Fed inflation persistence and possibly inflation target remained high whereas during the chairmanship

[20] Canova *et al*. (2007) and Gambetti *et al*. (2008) argued that the use of a SVAR model is a more appropriate specification to study for the individual features inflation such as persistence and volatility. They argued that in the presence of conditional heteroskedasticity, persistence and volatility should not depend on a single source of variation. Although this argument may be true, in our case this is not valid as we examine only one feature of inflation that of persistence.

of Volker and Greenspan we observe a dramatic drop in inflation persistence confirming the results by Beechey and Osterholm (2007). Inflation persistence in Canada, Japan, Australia and New Zealand followed to some extend similar patterns with the AR coefficient taking values near or above one in the late 1970s and during the 1980s followed by a substantial decrease in the 1990s and during the recent years.

Inflation persistence has remained above one or close to one for most of the sample period for the case of the Latin America economies which is not surprising given the long period of hyperinflation they experienced in the 1970s and 1980s. A fall in inflation persistence is marked in early 1990s which is due to the disinflationary policies and financial openness measures adopted during that period although in Argentina and Chile an upward trend has been observed in the recent years. Finally, the picture that emerges with respect the Southeast Asia countries show that inflation persistence remains moderate in all five economies of our sample although it was substantially high in the 1970s and early 1980s.[21]

5. Conclusions and Policy Implications

In this paper we investigate the issue of inflation persistence for thirty developed and emerging economies using quarterly data over the period 1958-2007. Inflation is a key macroeconomic variable and it has received a great deal of attention over the last twenty years. However, controversy remains about its stochastic properties and especially persistence. Although the degree of inflation persistence is an important issue equally significant is the investigation whether persistence has changed over time providing evidence of lower inflation in the future. Knowledge of the degree of inflation persistence as well as shifts in persistence has

[21] For the case of China the algorithm did not converge because of the small sample.

important implications for the conduct of monetary policy. Specifically, it is well documented in the relevant literature that in order to evaluate the response of inflation to monetary conditions and to determine the short-run trade-off between inflation and output the degree of inflation persistence is significant variable. Our analysis offers the opportunity to provide further insights to the effectiveness of changes in monetary policy to affect inflation persistence. Furthermore, the use of a large set of developed and emerging economies allow us to make cross-country comparisons.

The analysis was conducted within a model-free framework and using a battery of econometric techniques we provide an assessment of inflation persistence, shifts in persistence and temporal stability of our estimates. Initially we analyzed the stochastic properties of inflation series and we found that they follow an I(1) process with the exception of Germany and the Netherlands in which cases we conclude that inflation rates are stationary. The next step of our analysis is the estimation of ARMA models and this procedure led to the identification of a value of the sum of the AR coefficients greater than 0.9 for most of the countries examined and this is taken as evidence of high inflation persistence. Furthermore, fitting ARFIMA models in the data provided us with more flexibility in studying the statistical properties of inflation and we show that in most cases the I(1) and/or I(0) polar cases are rejected in favor of a value of d higher than 0.5 but less than one indicating nonstationarity but mean reverting behavior. We then applied the Hansen (1999) grid-bootstrap Median Unbiased approach to correct for small sample inefficiency of the OLS estimates of the AR parameter measuring inflation persistence. We coupled the efficient OLS estimates with the calculation of the 95 percent symmetric confidence intervals for the autoregressive coefficient which have correct first-order asymptotic coverage as suggested by Romano and Wolf (2001).

The issue of the constancy over time of the parameters of the estimated AR models was then considered. Such an analysis was considered equally important since over a long period of time we expect that there will be changes in the structural characteristics of an economy or changes in the monetary regime leading to inflation persistence becoming time-varying. This issue was analyzed with four alternative approaches. First, formal tests for changes in persistence have shown that we obtain clear evidence for the null hypothesis of no change in persistence for Belgium, China, Denmark, France, Germany, Greece, Korea, Luxembourg, Malaysia and Thailand. Second, in order to take account the issue of time-varying inflation persistence we adopt the recently developed approach by Beechey and Österholm (2007, 2009), and we estimated an AR(1) model allowing the autoregressive parameter to follow a random walk process. The model was estimated with maximum likelihood using the Kalman filter approach, with the autoregressive term treated as the unknown state variable. The results reveal that there is substantial time-variation in the autoregressive coefficient. Moreover, while the coefficient is generally large and suggestive of persistent behavior it is typically in a range below one although it can be observed that the parameter does approach the value of one and sometimes exceeds the value of unity.

The overall evidence from our analysis suggest that for almost all the countries examined, inflation persistence is high and has not changed significantly over the years. These findings are in line with those of Levin and Piger (2004) O'Reilly and Wheelan (2005), Batini (2006), Pivetta and Reis (2007) and Beechey and Österholm (2007, 2009). In case of the European countries we find evidence that over the last forty years inflation persistence has changed very little and despite of the shifts in monetary policy during that period the persistence coefficients have been quite stable. This evidence may lead to arguments against the validity of the Lucas critique of reduced-form models. With respect to the four Latin American countries in our sample, we note that major monetary policy shifts at the end of 1980s, the end of fiscal dominance and the adoption of inflation targeting along with the financial liberalization process in the 1990s and 2000s possibly led to a lowering of inflation persistence and this finding could provide support to the Lucas critique as opposed to the Euro area and even more to the EMU inflation stance. Finally, looking into the case of the Southeast Asian countries and China we understand that inflation persistence has remained high and fairly stable although for the latter an increase in inflation has been detected possibly due to substantial increase in demand for goods and services.

References

Andrews, D.W.K., 1993, Tests for parameter instability and structural change with unknown change point, Econometrica, 61, 821-856.

Andrews, D.W.K and H.-Y. Chen, 1994, Approximately median-unbiased estimation of autoregressive models, Journal of Business and Economic Statistics, 12, 187-204.

Andrews, D.W.K. and W. Ploberger, 1994 Optimal test when a nuisance parameter is present only under the alternative, Econometrica, 62, 1383-1414

Angeloni, I., G. Coenen and F. Smetz, 2003, Persistence, the transmission mechanism and robust monetary policy, Scottish Journal of Economics, 50, 527-549.

Bai, J., 1998, A note on spurious breaks, Econometric Theory, 14, 663-669.

Bai, J. and P. Perron, 1998, Estimating and testing linear model with multiple structural changes, Econometrica, 66, 47-68.

Bai J. and P. Perron, 2003, Computation and analysis of multiple structural change models, Journal of Applied Econometrics, 18, 1-22.

Batini, N., 2006, Euro area inflation persistence, Empirical Economics, 31, 977-1002.

Beechey, M. and P. Österholm, 2007, The rise and fall of US inflation persistence, Finance and Economics Discussion Series 2007-26, Board of Governors of the Federal Reserve System.

Beechey, M. and P. Österholm, 2009, Time-varying inflation persistence in the Euro Area, Economic Modelling, 26, 532-535.

Benati, L., 2008, Investigating inflation persistence across monetary regimes, Quarterly Journal of Economics, 123, 1005-1060.

Braindard, W. and G. Perry, 2000, Making policy in a changing world, in G. Perry and J. Tobin (eds.) Economic Events, Ideas and Policies: The 1960s and After, Washington D.C., Brookings Institution Press.

Buiter, W., and I. Jewitt, 1981, Staggered wage setting with real wage relativities: Variations on a theme of Taylor, The Manchester School, 49, 211-228.

Buiter, W., 2009, Negative nominal interest rates: Three ways to overcome the zero lower bound, North American Journal of Economics and Finance, 20, 213-238.

Busetti, F. and A. M. R. Taylor, 2004, Tests of stationarity against a change in persistence, Journal of Econometrics, 123, 33-66.

Carlstrom, C.T., T.S. Fuerst and M. Paustian, 2009, Inflation persistence, monetary policy, and the Great Moderation, Journal of Money, Credit and Banking, 41, 767-786.

Calvo, G., O. Celasum, and M. Kumhof, 2003, A theory of rational inflationary inertia, in P. Aghion, R. Frydman, J. Stiglitz and M. Woodford, (eds.) Knowledge, Information, Expectations in Modern Macroeconomics: in Honor of Edmund S. Phelps, Princeton: Princeton University Press.

Canova, F., L. Gambetti and E. Pappa, 2007, The structural dynamics of output growth and inflation: some international evidence, Economic Journal, C167-C191.

Christiano, L., M. Eichenbaum, and C. Evans, 2005, Nominal rigidities and the dynamic effects of a shock to monetary policy, Journal of Political Economy, 113, 1-45.

Chung, C.F., and R.T. Baillie, 1993, Small sample bias in conditional sum of squares estimators of Fractionally Integrated ARMA models, Empirical Economics, 18, 791-806.

Coenen, G. 2006, Inflation persistence and robust monetary policy design, Journal of Economic Dynamics and Control, 31, 111-140.

Cogley, T. W. and A. M. Sbordone, 2008, Trend inflation, indexation and inflation persistence in the New Keynesian Phillips curve, American Economic Review, 98, 2101-2126.

Cogley, T. W. and T. J. Sargent, 2001, Evolving post-World War II US inflation dynamics, in NBER Macroeconomics Annual, 16, (eds.) Ben S. Bernanke and Kenneth S. Rogoff, Cambridge: MIT Press, 331-373.

Cogley, T. W. and T.J. Sargent, 2005, Drifts and volatilities: Monetary policies and outcomes in the post WWII U.S., Review of Economic Dynamics, 8, 262-302.

Cook, S., 2006, The robustness of modified unit root tests in the presence of GARCH, Quantitative Finance, 6, 359-363.

Corvoisier, S., and B. Mojon, 2005, Breaks in the mean of inflation: How do they happen and what to do with them, ECB Working Paper No. 451, March

Dittmar, R., W. Gavin, and F. Kydland, 2005, Inflation persistence and flexible prices, International Economic Review, 46, 245-261.

Dixon, H. and E. Kara, 2010, Can we explain inflation persistence in a way that is consistent with microevidence on nominal rigidity, Journal of Money, Credit and Banking, 42, 151-170.

Eliott, G., T. J. Rothenberg, and J.H. Stock, 1996, Efficient tests for an autoregressive unit root, Econometrica, 64, 813-836.

Engle, R.F. and A. D. Smith, 1999, Stochastic permanent breaks, Review of Economics and Statistics, 81, 553-574.

Fox, R. and M.S. Taqqu, 1986, Large sample properties of parametric estimates for strongly dependent stationary Gaussian time series, Annals of Statistics, 14, 517-532.

Fuhrer, J., 2000, Habit formation in consumption and its implications for monetary policy models, American Economic Review, 90, 367-390.

Fuhrer, J., and G. Moore, 1995, Inflation persistence, Quarterly Journal of Economics, 110, 127-160.

Gadea, M. D. and L. Mayoral, 2006, The persistence of inflation in OECD countries: A fractional integrated approach, International Journal of Central Banking, 2, 51-105.

Gambetti, L., E. Pappa, F. Canova, 2008, The structural dynamics of U.S. output and inflation: what explains the changes, Journal of Credit, Money and Banking, 40, 369-388.

Hall, R. E., 1999, Comment on "Rethinking the role of NAIRU in monetary policy: Implications of model formulation and uncertainty," by Arturo Estrella and Frederic Mishkin, in Monetary Policy Rules, John B. Taylor (ed.), Chicago: University of Chicago Press.

Hansen, B., E., 1991, Testing for structural change of unknown form in models with nonstationary regressors, Mimeo, Department of Economics, University of Rochester.

Hansen, B. E. 1999, The grid bootstrap and the autoregressive model, Review of Economics and Statistics, 81, 594-607.

Harvey, D. I., S.J. Leybourne, and A.M.R. Taylor, 2006, Modified tests for a change in persistence, Journal of Econometrics, 134, 441-469.

Hosking, J. R.M. 1981, Fractional differencing, Biometrika, 68, 165-76

Ireland, P., 2004, A method for taking models to the data, Journal of Economic Dynamics and Control, 28, 1205-1226.

Kim, C.J. and C. R. Nelson, 1999, Has the U.S. economy become more stable? A bayesian approach based on a Markov-Switching model of the business-cycle, Review of Economics and Statistics, 81, 608-616.

Kim, C.J., C. R. Nelson, and J. Piger, 2004, The less-volatile U.S. economy: A bayesian investigation of timing, breadth and potential explanations, Journal of Business and Economic Statistics, 22, 80-93.

Kim, J.-Y. 2000, Detection of change in persistence of a linear time series, Journal of Econometrics, 95, 161-82

Kim, K. and P. Schmidt, 1993, Unit root tests with conditional heteroskedasticity, Journal of Econometrics, 59, 267-300.

Levin, A. and J. M. Piger, 2004, Is inflation persistence intrinsic in industrial economies?, European Central Bank, Working Paper.

Marques, C., 2004, Inflation persistence: Facts of artifacts?, European Central Bank, Working Paper No. 371.

McConnell, M. and G. Perez-Quiros, 2000, Output fluctuations in the United States: what has changed since the early 1980s?, American Economic Review, 90, 1464-1476.

Murray, C. and D. Papell, 2002, The purchasing power parity persistence paradigm, Journal of International Economics, 56, 1-19.

Nelson, E. 2001, What does the UK's monetary policy and inflation experience tell us about the transmission mechanism?, CEPR Discussion Paper No. 3047.

Nelson, E., 2007, An overhaul of doctrine: The underpinning of UK inflation targeting, Federal Reserve Bank of St. Louis, Working Paper 2007-026B.

Nelson, E. and K. Nikolov, 2004, Monetary policy and stagflation in the UK, Journal of Money, Credit and Banking, 36, 293-318

Ng, S. and P. Perron, 2001, Lag length selection and the construction of unit root tests with good size and power, Econometrica, 69, 1519-1554.

Nobay, B., I. Paya and D. A. Peel, 2010, Inflation dynamics in the U.S.: global but not local mean reversion, Journal of Money, Credit and Banking, 42, 135-150.

O'Reilly, G. and K. Whelan, 2005, Has euro area inflation persistence changed over time?" Review of Economics and Statistics, 87, 709-720.

Perron, P., 1989, The great crash, the oil price shock and the unit root hypothesis," Econometrica, 58, 1361-1401.

Perron, P. and S. Ng, 1996, Useful modifications to some unit root tests with dependent errors and their local asymptotic properties, Review of Economic Studies, 63, 435-463

Perron, P. and Z. Qu, 2007, Estimating and testing structural changes in multivariate regressions, Econometrica, 75, 459-502.

Phillips, P.C.B., 1991, To criticize the critics: An objective bayesian analysis of stochastic trends, Journal of Applied Econometrics, 6, 333-364.

Phillips, P.C.B., 1999, Discrete Fourier Transforms of Fractional Processes, Cowles Foundation Discussion Paper No. 1243.

Phillips, P.C.B. and P. Perron, 1988, Testing for a unit root in time series regression, Biometrika, 75, 335-346.

Pivetta, F. and R. Reis, 2007, The persistence of inflation in the United States, Journal of Economic Dynamics and Control, 31, 459-502.

Prodan, R., 2008, Potential Pitfalls in Determining Multiple Structural Changes With an Application to Purchasing Power Parity, Journal of Business and Economics Statistics, 26, 50-65.

Rapach, D. E. and M. E. Wohar, 2005, Regime changes in international real interest rates: Are they a monetary phenomenon?, Journal of Money, Credit and Banking, 37, 887-906.

Ravenna, F., 2000, The impact of inflation targeting in Canada: A structural analysis, New York University, Working Paper.

Romano, J. P. and M. Wolf, 2001, Subsampling intervals in autoregressive models with linear time trends, Econometrica, 69, 1283-1314.

Robinson, P.M., 1994, Time series with strong dependence, in Advances in Econometrics, Sixth World Congress, C.A. Sims (ed.), Cambridge: Cambridge University Press, 47-96.

Rotemberg, J. and M. Woodford, 1997, An optimizing-based econometric framework for the evaluation of monetary policy, in B. Bernanke and J. Rotemberg (eds.), NBER Macro Economics Annual, Cambridge: The MIT Press.

Said, S.E., and D.A. Dickey, 1983, Testing for unit roots in autoregressive moving average models of unknown order, Biometrica, 71, 599-607.

Schwert G. W., 1987, Effects of model specification on tests for unit roots in macroeconomic data, Journal of Monetary Economics, 20, 73-103.

Shimotsu, K., 2007, Gaussian semiparametric estimation of multivariate fractionally integrated processes, Journal of Econometrics, 137, 277-310.

Shimotsu, K. and P.C.B. Phillips, 2005, Exact local Whittle estimation of fractional integration, Annals of Statistics, 33, 1890-1933.

Shimotsu, K., and P.C.B. Phillips, 2006, Local Whittle estimation of fractional integration and some of its variants, Journal of Econometrics, 130, 209-233.

Sims, C.A. and T. Zha, 2006, Were there regime switches in U.S. monetary policy?, American Economic Review, 96, 54-81.

Solow, R., 1968, Recent controversies on the theory of inflation: An eclectic view, in Inflation: Its Causes, Consequences and Control, Proceedings of a symposium, Stephen W. Rousseas (ed.), New York: New York University Press.

Stock, J., 2001, Comment on evolving post World War II U.S. inflation dynamics, in NBER Macroeconomics Annual, 16, (eds.) Ben S. Bernanke and Kenneth S. Rogoff, Cambridge: MIT Press, 373-378.

Stock, J. and M. Watson, 1996, Evidence of structural instability in macroeconomic time series relations, Journal of Business and Economic Statistics, 14, 11-30.

Stock, J. and M. Watson, 1998, Median-Unbiased estimation of coefficient variance in a time-varying parameter model, Journal of the American Statistical Association, 93, 349-358.

Stock, J. and M. Watson, 2002, Has the business cycle changed and why?, in B. Bernanke and K. Rogoff, eds. NBER Macroeconomics Annual, Cambridge: The MIT Press.

Stock, J. and M. Watson, 2003, Understanding changes in international business cycle dynamics, Journal of the European Economic Association, 3, 968-1006.

Taylor, John B., 1998, Monetary policy guidelines for unemployment and inflation stability, in Inflation, Unemployment and Monetary Policy, Robert H. Solow and John B. Taylor (eds.), Cambridge: MIT Press.

Taylor, J. B., 2000, Low inflation, pass-through, and the pricing power of firms, European Economic Review, 44, 1389-1408.

Tobin, J., 1968, Discussion, in Inflation: Its Causes, Consequences and Control, Proceedings of a symposium, Stephen W. Rousseas (ed.), New York: New York University Press.

Velasco, Carlos, 1999a, Semiparametric Gaussian estimation of non-stationary time series," Journal of Time Series Analysis, 20, 87-127.

Velasco, Carlos, 1999b, Non-stationary log-periodogram regresion, Journal of Econometrics, 91, 325-371.

West, K.D., 1988, On the interpretation of near random walk behavior in GDP, American Economic Review, 78, 202-209.

Zeira, J., 1989, Inflationary inertia in a wage-price spiral model, European Economic Review, 33, 1665-1683.

Table 1. Unit Roots Tests

Variables (MZ)	ADF	DF-GLS	MZ_a	MZ_t	MAIC lag
Argentina	-3.11**	-2.69***	-13.05**	-2.55**	2
Australia	-2.36	-2.02**	-7.15*	-1.89*	3
Austria	-2.35	-0.52	-0.29	-0.29	13
Belgium	-2.15	-1.48	-3.92	-1.33	5
Brazil	-1.83	-1.81*	-7.51*	-1.92*	5
Canada	-1.98	-1.68*	-4.91	-1.54	3
Chile	-2.69*	-2.57***	-18.44***	-3.03***	9
China	-1.86	-0.76	-1.47	-0.84	4
Denmark	-1.54	-0.93	-1.28	-0.76	11
Finland	-1.90	-1.82*	-6.05*	-1.70*	13
France	-1.64	0.06	0.32	0.49	3
Germany	-6.69***	-5.62***	-28.84***	-3.80***	0
Greece	-1.97	-1.63*	-2.35	-1.05	7
Ireland	-2.13	-2.09**	-5.23	-1.61	7
Italy	-1.60	-1.98**	-4.56	-1.51	10
Japan	-1.64	-0.89	-1.16	-0.72	11
Korea	-1.57	-1.56*	-3.78	-1.35	11
Luxembourg	-2.50	-0.87	-1.58	-0.74	14
Malaysia	-2.88**	-1.44	-3.49	-1.29	11
Mexico	-1.60	-1.57	-4.76	-1.53	11
Netherlands	-3.24**	-1.70*	-4.05	-1.41	12
New Zealand	-1.59	-1.03	-2.01	-0.98	10
Philippines	-2.82*	-1.22	-2.37	-1.08	11
Portugal	-1.60	-0.96	-1.84	-0.95	14
Singapore	-4.84***	-1.10	-1.84	-0.80	11
Spain	-1.53	-1.29	-3.21	-1.26	7
Sweden	-1.90	-1.48	-3.02	-1.23	7
Thailand	-3.65***	-2.85***	-15.31***	-2.75***	3
UK	-1.97	-1.41	-3.54	-1.31	11
US	-2.04	-2.05**	-6.44*	-1.78*	12

With the exception of the Augmented Dickey Fuller (ADF) test, all tests use GLS detrending. All tests include a constant but no trend as no series are trended.
The 10%, 5% and 1% critical values for the ADF test are -2.58, -2.89 and -3.50. The number in parentheses is the lag length used.
he 10%, 5%, and 1% critical values for the Elliot, Rotenberg, and Stock (1996) DF-GLS test are -1.62, -1.94, and -2.59.
The MZ_a and the MZ_t statistics are the Ng and Perron (2001) GLS versions of the Phillips-Perron tests.
The 10, 5% and 1% critical values for the MZ_a test -5.7, -8.1, and -13.8.
The 10, 5% and 1% critical values for the MZ_t test are -1.62, -1.98, -2.58. See Ng and Perron (2001, Table 1)
The last column indicates the number of lags included in the test, which was selected by modified AIC criterion.
*, **, *** indicate significance at the 10%, 5%, and 1% level respectively.

Table 2. ARMA(1,1) Estimates

Country	AR(1)	MA(1)	d
Argentina Inflation	0.89	-0.43	0.597
Australia Inflation	0.96	-0.64	0.719
Austria Inflation	0.97	-0.89	0.718
Belgium Inflation	0.95	-0.59	0.796
Brazil Inflation	0.82	-0.25	0.724
Canada Inflation	0.96	-0.66	0.879
Chile Inflation	0.94	-0.38	1.220
China Inflation	0.96	-0.52	0.685
Denmark Inflation	0.98	-0.83	0.571
Finland Inflation	0.94	-0.54	0.853
France Inflation	0.97	-0.57	0.713
Germany Inflation	0.96	-0.80	0.396
Greece Inflation	0.97	-0.86	0.447
Ireland Inflation	0.96	-0.73	0.747
Italian Inflation	0.95	-0.46	0.734
Japan Inflation	0.95	-0.68	0.657
Korea Inflation	0.88	-0.50	0.784
Luxembourg	0.95	-0.67	0.812
Malaysia Inflation	0.83	-0.47	0.376
Mexico Inflation	0.87		0.690
Netherlands Inflation	0.92	-0.88	0.129
New Zealand Inflation	0.93	-0.50	0.558
Philippines Inflation	0.75	-0.32	0.199
Portugal Inflation	0.98	-0.78	0.637
Singapore Inflation	0.78	-0.33	0.324
Spain Inflation	0.97	-0.71	0.568
Sweden Inflation	0.97	-0.78	0.584
Thailand Inflation	0.78	-0.24	0.500
UK Inflation	0.95	-0.67	0.638
US Inflation	0.93	-0.46	0.889

Note: The value of d is the exact Whittle estimator (Shimotsu, 2007; Shimotsu and Phillips, 2005, 2006).

Table 3. OLS and ARFIMA(1, d ,1) Estimates

Country	OLS AR(1)	ARFIMA			Country	OLS AR(1)	ARFIMA		
		AR(1)	d	MA(1)			AR(1)	d	MA(1)
Argentina	0.688	-0.1181	0.6696	-0.0968	Netherlands	0.04	-0.902	0.0756	0.8582
Australia	0.656	-0.1143	0.717	-0.3173	New Zealand	0.734	0.4379	0.7888	-0.7445
Austria	0.106	0.0277	0.5852	-0.6624	Philippines	0.501	-0.9891	0.5178	0.9327
Belgium	0.693	-0.0363	0.7704	-0.3862	Portugal	0.611	0.1417	0.8802	-0.7569
Brazil	0.88	0.2483	0.5046	0.2824	Singapore	0.552	-0.9916	0.5878	0.9354
Canada	0.728	0.1946	0.8326	-0.6321	Spain	0.592	-0.9879	0.4784	0.8764
Chile	0.856	0.1228	0.9298	-0.4395	Sweden	0.493	0.0687	0.6646	-0.5484
China	0.76	-0.3647	0.7806	-0.0026	Thailand	0.64	0.9054	-0.3306	-0.0359
Denmark	0.416	-0.0889	0.562	-0.349	United Kingdom	0.569	0.0015	0.6638	-0.4048
Finland	0.742	-0.3698	0.4878	0.4309	United States	0.773	-0.5282	0.48	0.8700
France	0.704	0.1610	0.8224	-0.6398					
Germany	0.178	0.0867	0.4826	-0.4635					
Greece	0.146	-0.9971	0.2886	0.9390					
Ireland	0.589	0.0922	0.7728	-0.6101					
Italy	0.846	0.3223	0.8534	-0.6071					
Japan	0.520	-0.1846	0.6168	-0.2432					
Korea	0.60	-0.3319	0.400	0.4166					
Luxembourg	0.575	-0.0786	0.7886	-0.4702					
Malaysia	0.489	-0.1171	0.3922	0.1038					
Mexico	0.873	0.0053	0.6222	0.1927					

Note: OLS estimates refer to full-sample univariate estimates of the AR(1) from an OLS regression.

Table 4. Median Unbiased Estimators and Sub-Sampling Confidence Intervals

Country	Grid Bootstrap	Sub Sampling	Country	Grid Bootstrap	Sub Sampling
Argentina	0.79 (0.69,0.95)	(0.64,0.94)	Japan	0.82 (0.70,1.02)	(0.63,1.00)
Australia	0.86 (0.77,1.03)	(0.74,0.99)	Korea	0.77 (0.66,1.02)	(0.57,0.99)
Austria	0.75 (0.60,1.04)	(0.47,1.02)	Luxembourg	0.85 (0.75,1.04)	(0.71,0.99)
Belgium	0.89 (0.80,1.04)	(0.76,1.01)	Malaysia	0.62 (0.47,0.83)	(0.40,0.83)
Brazil	0.91 (0.84,1.05)	(0.79,1.03)	Mexico	0.90 (0.84,1.01)	(0.80,1.00)
Canada	0.89 (0.81,1.04)	(0.70,1.08)	Netherlands	0.17 (-0.12,0.54)	(-0.25,0.60)
Chile	0.91 (0.85,1.02)	(0.79,1.03)	New Zealand	0.88 (0.81,1.03)	(0.77,1.00)
China	0.90 (0.82,1.05)	(0.77,1.03)	Philippines	0.51 (0.35,0.72)	(0.35,0.67)
Denmark	0.80 (0.67,1.04)	(0.57,1.02)	Portugal	0.86 (0.76,1.04)	(0.67,1.04)
Finland	0.88 (0.80,1.03)	(0.76,0.99)	Singapore	0.62 (0.47,0.82)	(0.45,0.78)
France	0.94 (0.89,1.04)	(0.85,1.03)	Spain	0.88 (0.80,1.04)	(0.74,1.02)
Germany	0.46 (0.03,1.08)	(0.14,0.61)	Sweden	0.80 (0.68,1.02)	(0.60,1.01)
Greece	0.77 (0.65,1.02)	(0.55,0.98)	Thailand	0.72 (0.58,0.92)	(0.53,0.90)
Ireland	0.86 (0.76,1.04)	(0.71,1.00)	UK	0.84 (0.75,1.01)	(0.67,1.01)
Italy	0.92 (0.86,1.04)	(0.82,1.02)	US	0.86 (0.79,1.00)	(0.74,0.99)

Note: Mean Unbiased Estimators of the AR coefficient are estimated using the Hansen (1999) grid-bootstrap method and the 95% percent symmetric confidence intervals for the autoregressive coefficient are calculated using the Romano and Wolf (2001) approach.

Table 5
Application of tests to Inflation series: Test of change in persistence: I(0) to I(1) or I(1) to I(0)

Series	T	MS MS_m *min 10%* MS_m *min 5%*	ME ME_m *min 10%* ME_m *min 5%*	MX MX_m *min 10%* MX_m *min 5%*	MS^R MS_m^R *min 10%* MS_m^R *min 5%*	ME^R ME_m^R *min 10%* ME_m^R *min 5%*	MX^R MX_m^R *min 10%* MX_m^R *min 5%*	MS^M MS_m^M *min 10%* MS_m^M *min 5%*	ME^M ME_m^M *min 10%* ME_m^M *min 5%*	MX^M MX_m^M *min 10%* MX_m^M *min 5%*
Argentina Inflation	200	36.88***	130.03***	269.48***	111.71***	588.63***	1186.85***	111.71***	588.63***	1186.85***
I(0)-I(1) estimated date 1974Q3		16.23***	36.87***	95.92***	75.90***	320.46 ***	730.43***	62.04***	247.38***	573.34***
I(1)-I(0) estimated date 1991Q3		11.03***	21.85***	60.24***	61.98***	247.24***	572.38***	48.93***	183.27***	430.57***
Australia Inflation	200	12.20***	48.23***	105.12***	21.63***	62.62***	134.82***	21.63***	62.62***	134.82***
I(0)-I(1) estimated date 1970Q2		4.41*	10.12***	29.23***	16.61***	41.34 ***	96.78***	14.48***	34.64***	82.03***
I(1)-I(0) estimated date 1976Q4		2.74	5.39**	16.43*	14.47***	34.63***	81.94***	12.31***	28.22***	67.46***
Austria Inflation	200	4.09*	11.40***	30.35***	6.44**	9.76**	26.46**	6.44**	11.40**	30.35**
I(0)-I(1) estimated date 1991Q2		3.60*	9.36**	25.83**	5.48**	7.58**	21.62**	5.31*	8.58**	23.92**
I(1)-I(0) estimated date 1967Q4		3.39	8.62**	24.02**	5.04**	6.80**	19.54**	4.91*	7.77**	21.77*
Belgium Inflation	200	2.14	3.61*	12.11	40.32***	92.48***	194.55***	40.32***	92.48***	194.55***
I(0)-I(1) estimated date 1972Q2		1.13	1.36	5.45	33.87***	70.29***	156.28***	30.93***	62.54***	140.10***
I(1)-I(0) estimated date 1985Q1		0.84	0.91	3.81	30.91***	62.53***	140.00***	27.78***	54.63***	123.12***
Brazil Inflation	111	18.37***	86.84***	182.12***	925.17***	2263.73***	4535.90***	925.17***	2263.73***	4535.90***
I(0)-I(1) estimated date 1963Q2		5.88**	15.11***	43.44***	516.65***	905.15***	2181.90***	381.26***	612.70***	1514.55***
I(1)-I(0) estimated date 1973Q4		3.44	7.31**	22.78**	380.62***	612.18***	1510.75***	266.54***	389.79***	983.53***
Canada Inflation	200	3.74*	8.89**	26.60**	29.47***	76.47***	162.52***	29.47***	76.47***	162.52***

I(0)-I(1) estimated date 1997Q4		2.62	5.16**	17.02*	26.97***	66.52***	145.41***	25.76***	62.69***	137.56***
I(1)-I(0) estimated date 1982Q3		2.22	4.11*	13.91*	25.75***	62.68***	137.51***	24.39***	58.52***	128.82***
Chile Inflation	200	54.93**	258.32***	526.23***	176.49***	1220.85***	2451.28***	176.49***	1220.85***	2451.28***
I(0)-I(1) estimated date 1972Q2		17.50***	44.59***	124.70***	96.24***	470.18***	1144.34***	70.14***	313.22***	782.56***
I(1)-I(0) estimated date 1993Q4		10.22***	21.49***	65.20***	70.02***	312.95***	780.51***	48.32***	195.61***	499.29***
China Inflation	108	0.64	0.35	1.86	4.49*	8.03**	22.66**	4.49	8.03**	QQ
I(0)-I(1) estimated date 1974Q4		0.29	0.10	0.68	4.04*	6.80**	19.84**	3.82	6.33*	18.57*
I(1)-I(0) estimated date 1963Q2		0.20	0.06	0.43	3.82*	6.33**	18.56**	3.58	5.83*	17.17*
Denmark Inflation	200	2.43	3.56*	11.96	47.08***	111.41***	232.42***	47.08***	111.41***	232.42***
I(0)-I(1) estimated date 1997Q3		1.69	2.03	7.55	44.25***	101.07***	215.03***	42.84***	96.97***	206.85***
I(1)-I(0) estimated date 1986Q2		1.42	1.61	6.13	42.84***	96.96***	206.79***	41.25***	92.42***	197.57***
Finland Inflation	200	7.93***	24.94***	57.65***	23.85***	58.14***	125.85***	23.85***	58.14***	125.85***
I(0)-I(1) estimated date 1973Q2		5.50**	14.21***	36.35***	17.45***	35.56***	85.01***	14.83***	28.85***	69.89***
I(1)-I(0) estimated date 1983Q3		4.63**	11.25***	29.54***	14.82***	28.84***	69.80***	12.24***	22.64***	55.45***
France Inflation	200	7.25**	18.19***	44.16***	24.68***	52.69***	114.08***	24.68***	52.69***	114.08***
I(0)-I(1) estimated date 1979Q1		1.23	1.20	4.75	20.20***	38.44***	88.69***	18.19***	33.61***	78.22***
I(1)-I(0) estimated date 1985Q2		0.54	0.39	1.74	18.18***	33.60***	78.15***	16.08***	28.77***	67.42***
Germany Inflation	67	0.08	0.04	0.25	19.67***	29.45***	66.37***	19.67***	29.45***	66.37***
I(0)-I(1) estimated date 1971Q2		0.05	0.02	0.12	17.16***	23.76***	55.91***	15.98***	21.68***	51.33***
I(1)-I(0) estimated date 1960Q4		0.04	0.01	0.09	15.98***	21.68***	51.30***	14.70***	19.50***	46.39***
Greece Inflation	200	2.80	3.22	13.55*	60.34***	689.75***	1389.10***	60.34***	689.75***	1389.10***
I(0)-I(1) estimated date 1997Q4		2.27	2.33	10.41	58.48***	656.56***	1335.46***	57.53***	642.92***	1309.49***

I(1)-I(0) estimated date 1970Q1		2.06	2.04	9.24	57.52***	642.89***	1309.31***	56.43***	627.46***	1279.42***
Ireland Inflation	200	4.52*	19.61***	47.90***	34.74***	87.37***	184.31***	34.74***	87.37***	184.31***
I(0)-I(1) estimated date 1972Q4		3.22	11.64***	31.24***	24.70***	51.09***	120.10***	20.85***	41.17***	98.02***
I(1)-I(0) estimated date 1984Q2		2.74	9.37**	25.76**	20.66***	40.64***	96.85***	16.97***	31.73***	76.45***
Italy Inflation	200	16.70***	52.76***	113.68***	99.57***	694.01***	1397.62***	99.57***	694.01***	1397.62***
I(0)-I(1) estimated date 1972Q3		5.77**	10.32***	29.84***	75.30***	447.15***	983.95***	65.09***	307.84***	825.93***
I(1)-I(0) estimated date 1996Q4		3.50	5.24**	16.34*	65.04***	370.69***	824.93***	54.82***	298.54***	671.47***
Japan Inflation	200	4.49*	10.63***	28.61***	23.06***	81.42***	172.28***	23.06***	81.42***	172.28***
I(0)-I(1) estimated date 1972Q4		3.06	5.89**	17.65**	20.53***	67.81***	148.87***	19.32***	62.73***	138.41***
I(1)-I(0) estimated date 1977Q2		2.56	4.61*	14.20*	19.31***	62.72***	138.34***	17.99***	57.31***	126.98***
Korea Inflation	151	0.67	0.88	6.28	63.80***	329.56***	668.16***	63.80***	329.56***	668.16***
I(0)-I(1) estimated date 1985Q3		0.35	0.32	2.74	56.20***	269.90***	569.69***	52.60***	247.90***	526.14***
I(1)-I(0) estimated date 1969Q2		0.25	0.21	1.88	52.58***	247.86***	525.85***	48.65***	224.64***	478.91***
Luxembourg Inflation	200	1.51	2.28	10.41	27.24***	76.68***	162.52***	27.24***	76.68***	162.52***
I(0)-I(1) estimated date 1997Q4		0.84	0.92	4.96	22.31***	56.00***	126.44***	20.10***	48.98***	111.56***
I(1)-I(0) estimated date 1983Q4		0.64	0.63	3.55	20.09***	48.97***	111.46***	17.78***	41.94***	96.20***
Malaysia Inflation	200	1.02	1.43	8.69	16.43***	27.46***	63.41***	16.43***	27.46***	63.41***
I(0)-I(1) estimated date 1997Q4		0.88	1.13	7.20	11.74***	16.19***	41.58***	13.13***	19.74***	48.07***
I(1)-I(0) estimated date 1975Q2		0.82	1.03	6.62	9.84***	12.92***	33.64***	12.00***	17.61***	43.10***
Mexico Inflation	200	112.85***	380.60***	770.79***	7.30**	23.48***	55.80***	112.85***	380.60***	770.79***
I(0)-I(1) estimated date 1972Q4		79.99***	224.37***	499.84***	1.92	2.88	10.44	67.30***	177.62***	406.57***
I(1)-I(0) estimated date 1997Q1		68.04***	180.15***	411.27***	0.95	1.17	4.50	54.62***	136.44***	316.08***

Netherlands Inflation	200	54.93***	258.32***	526.23***	176.49***	1220.85***	2451.28***	176.49***	1220.85***	2451.28***
I(0)-I(1) estimated date 1980Q4		17.50***	44.59***	124.70***	96.24***	470.18***	1144.34***	70.14***	313.22***	782.56***
I(1)-I(0) estimated date 1984Q2		10.22***	21.49***	65.20***	70.02***	312.95***	780.51***	48.32***	195.61***	499.29***
New Zealand Inflation	200	10.38***	51.22***	111.67***	31.21***	106.50***	222.58***	31.21***	106.50***	222.58***
I(0)-I(1) estimated date 1974Q3		6.44**	24.61***	61.25***	18.97***	48.66***	119.09***	15.24***	37.01***	91.67***
I(1)-I(0) estimated date 1990Q2		5.14**	18.15***	46.73***	14.61***	34.84***	86.99***	11.41***	25.68***	64.66***
Philippine Inflation	200	2.87	12.09***	32.50***	8.67***	25.34***	60.14***	8.67**	25.34***	60.14***
I(0)-I(1) estimated date 1983Q3		2.58	10.29***	28.47***	6.86**	17.55***	44.86***	7.40**	20.07***	49.47***
I(1)-I(0) estimated date 1985Q1		2.46	9.62**	26.82**	6.07**	15.01***	38.72***	6.94**	18.52***	45.80***
Portugal Inflation	200	15.07***	47.56***	104.08***	22.96***	72.54***	154.56***	22.96***	72.54***	154.56***
I(0)-I(1) estimated date 1997Q4		7.87***	17.55***	45.97***	20.77***	61.93***	136.24***	19.71***	57.90***	127.93***
I(1)-I(0) estimated date 1986Q1		5.80**	11.60***	31.82***	19.70***	57.90***	127.87***	18.53***	53.56***	118.75***
Singapore Inflation	187	9.38***	60.82***	130.98***	10.25***	19.12***	46.46***	10.25***	60.82***	130.98***
I(0)-I(1) estimated date 1969Q2		8.47***	51.94***	115.08***	6.40**	9.11**	25.71**	8.78**	48.44***	108.19***
I(1)-I(0) estimated date 1970Q4		8.07***	48.64***	108.57***	5.00**	6.64**	19.10**	8.25**	44.77***	100.35***
Spain Inflation	200	15.07***	47.56***	104.08***	22.96***	72.54***	154.56***	22.96***	72.54***	154.56***
I(0)-I(1) estimated date 1997Q4		7.87***	17.55***	45.97***	20.77***	61.93***	136.24***	19.71***	57.90***	127.93***
I(1)-I(0) estimated date 1986Q1		5.80**	11.60***	31.82***	19.70***	57.90***	127.87***	18.53***	53.56***	118.75***
Sweden Inflation	200	12.62***	58.62***	126.79***	8.22***	27.85***	65.18***	12.62***	58.62***	126.79***
I(0)-I(1) estimated date 1989Q4		7.82***	28.11***	69.42***	5.87**	16.41***	42.73***	7.57**	27.58***	67.34***
I(1)-I(0) estimated date 1967Q4		6.24**	20.72***	52.93***	4.92**	13.10***	34.56***	6.15**	21.25***	52.49***
Thailand Inflation	171	0.35	0.20	3.31	6.13**	10.92**	30.11***	6.13**	10.92**	30.11**

I(0)-I(1) estimated date 1990Q1		0.24	0.11	2.06	4.17*	5.97**	18.59**	3.47	4.72	14.90
I(1)-I(0) estimated date 1966Q4		0.20	0.09	1.66	3.41	4.61*	14.60*	2.76	3.53	11.29
UK inflation	200	5.86**	30.58***	70.54***	29.62***	86.69***	182.97***	29.62***	86.69***	182.97***
I(0)-I(1) estimated date 1973Q3		4.30*	18.98***	47.72***	27.39***	76.62***	165.79***	26.29***	72.70***	157.84***
I(1)-I(0) estimated date 1981Q2		3.71*	15.57***	40.02***	26.28***	72.69***	157.78***	25.05***	68.40***	148.92***
US Inflation	200	2.40	16.60***	42.77***	35.51***	100.52***	210.44***	35.51***	100.52***	210.44***
I(0)-I(1) estimated date 1967Q4		0.64	2.18	8.09	32.86***	89.00***	190.95***	31.56***	84.50***	181.91***
I(1)-I(0) estimated date 1981Q3		0.34	0.94	3.82	31.56***	84.49***	181.85***	30.10***	79.57***	171.77***

Figure 1.1

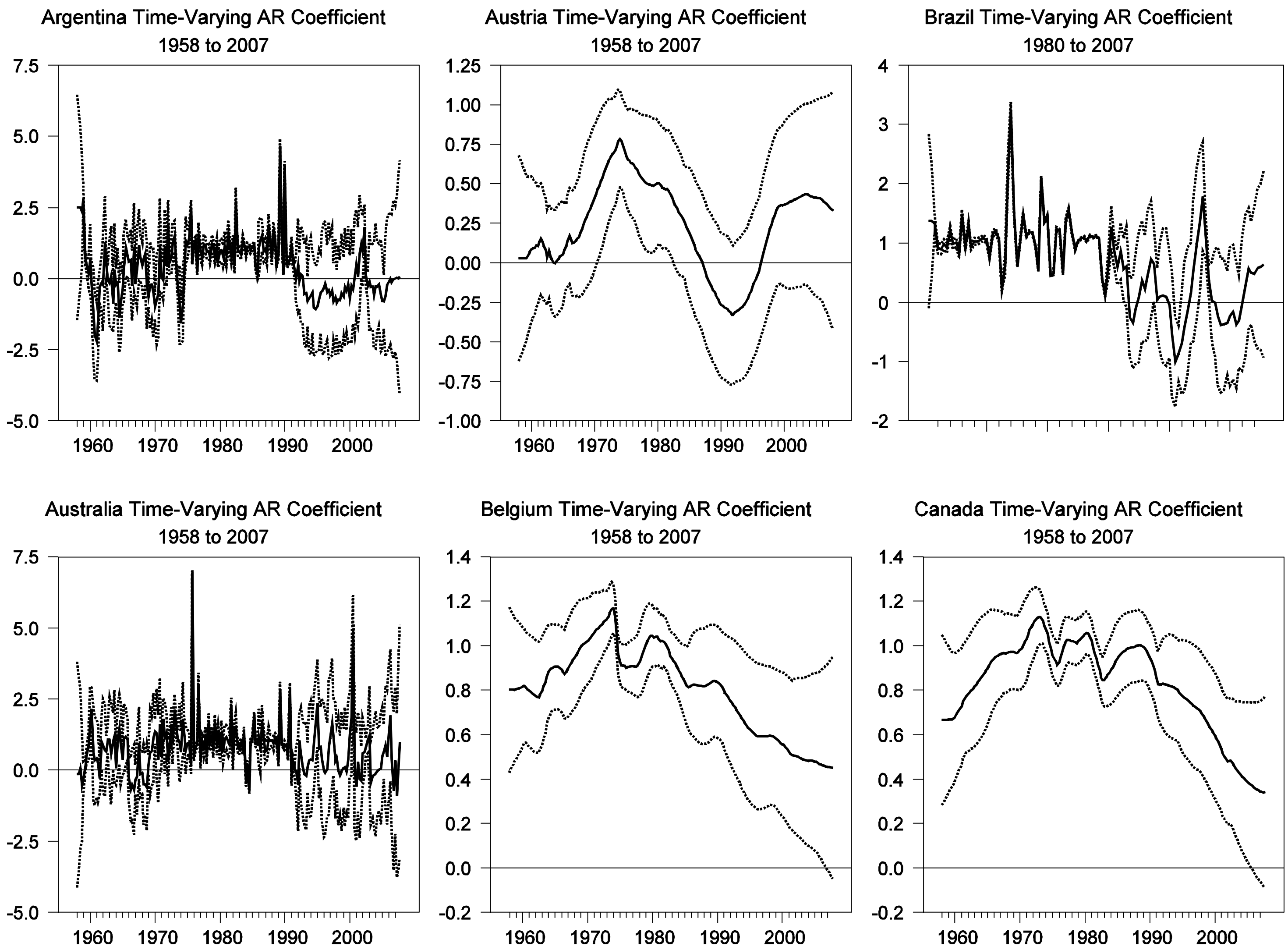
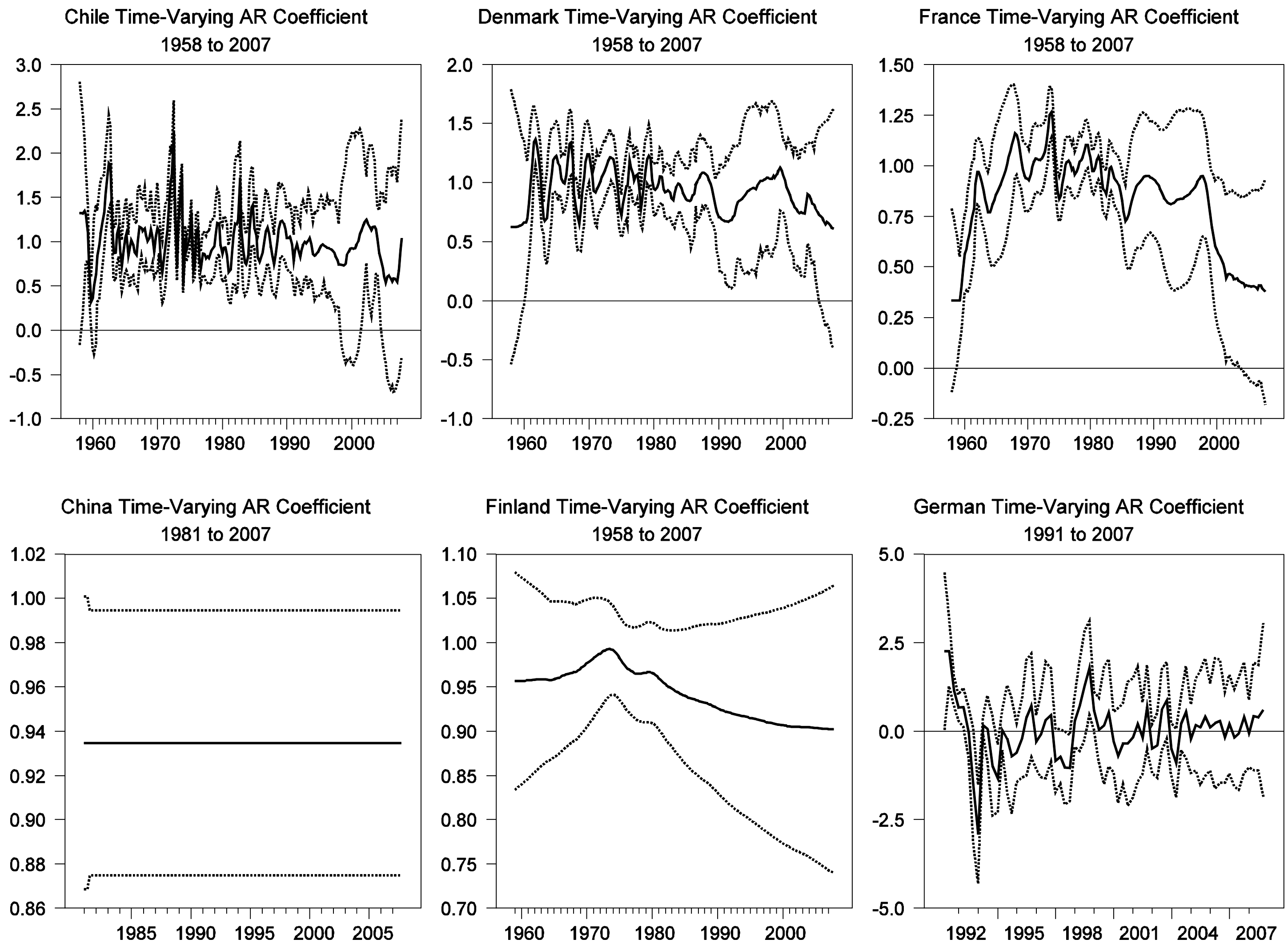

Argentina Time-Varying AR Coefficient
1958 to 2007
Austria Time-Varying AR Coefficient
1958 to 2007
Brazil Time-Varying AR Coefficient
1980 to 2007
Australia Time-Varying AR Coefficient
1958 to 2007
Belgium Time-Varying AR Coefficient
1958 to 2007
Canada Time-Varying AR Coefficient
1958 to 2007
7.5
5.0
2.5
0.0
-2.5
-5.0
1.25
1.00
0.75
0.50
0.25
0.00
-0.25
-0.50
-0.75
-1.00
4
3
2
1
0
-1
-2
1.4
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
1960
1970
1980
1990
2000

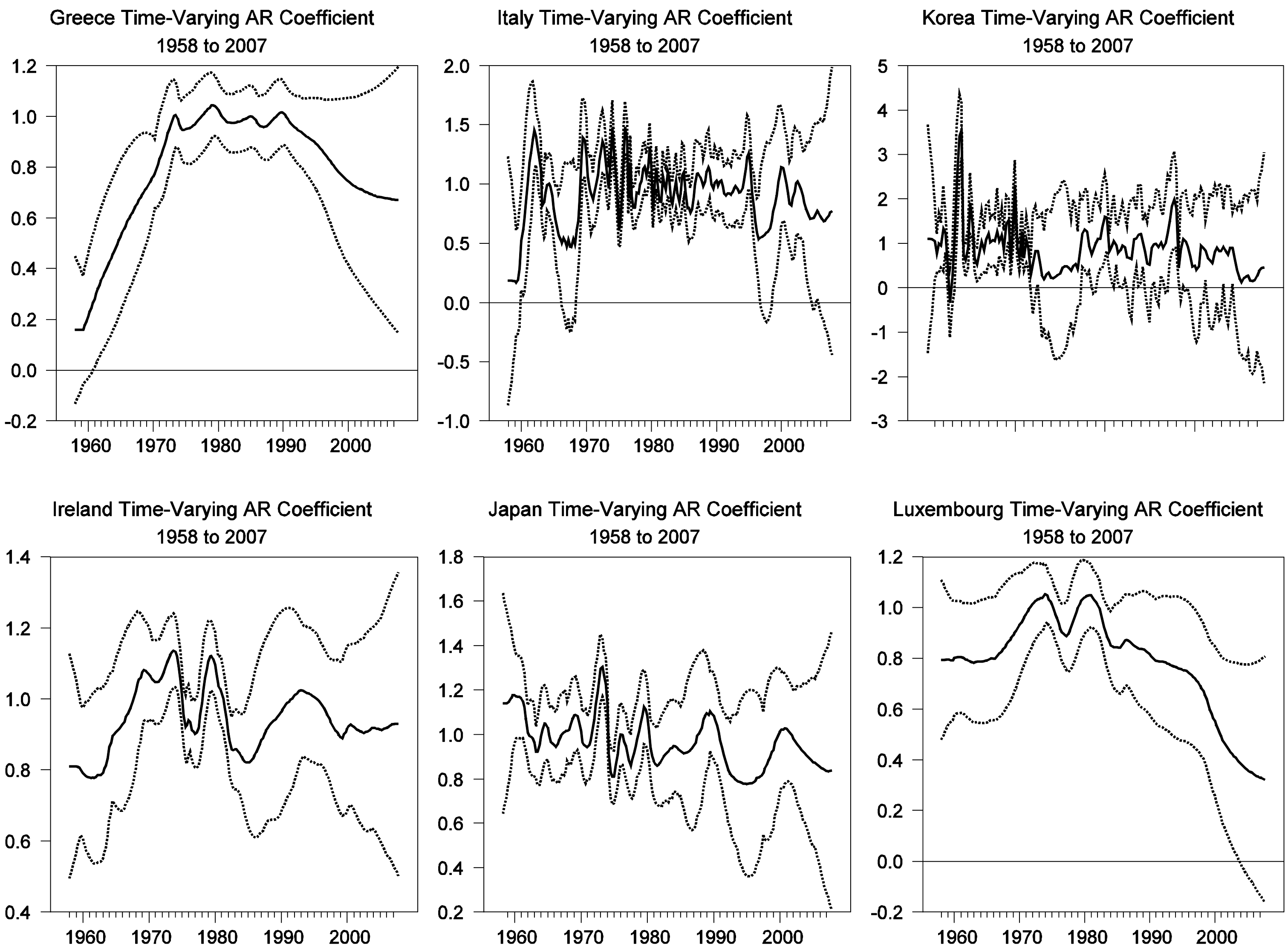
Figure 1.2



Chile Time-Varying AR Coefficient
1958 to 2007
Denmark Time-Varying AR Coefficient
1958 to 2007
France Time-Varying AR Coefficient
1958 to 2007
China Time-Varying AR Coefficient
1981 to 2007
Finland Time-Varying AR Coefficient
1958 to 2007
German Time-Varying AR Coefficient
1991 to 2007

Figure 1.3

1



Greece Time-Varying AR Coefficient
1958 to 2007
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
1960
1970
1980
1990
2000
Italy Time-Varying AR Coefficient
1958 to 2007
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
1960
1970
1980
1990
2000
Korea Time-Varying AR Coefficient
1958 to 2007
5
4
3
2
1
0
-1
-2
-3
Ireland Time-Varying AR Coefficient
1958 to 2007
1.4
1.2
1.0
0.8
0.6
0.4
1960
1970
1980
1990
2000
Japan Time-Varying AR Coefficient
1958 to 2007
1.8
1.6
1.4
1.2
1.0
0.8
0.6
0.4
0.2
1960
1970
1980
1990
2000
Luxembourg Time-Varying AR Coefficient
1958 to 2007
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
1960
1970
1980
1990
2000

Figure 1.4

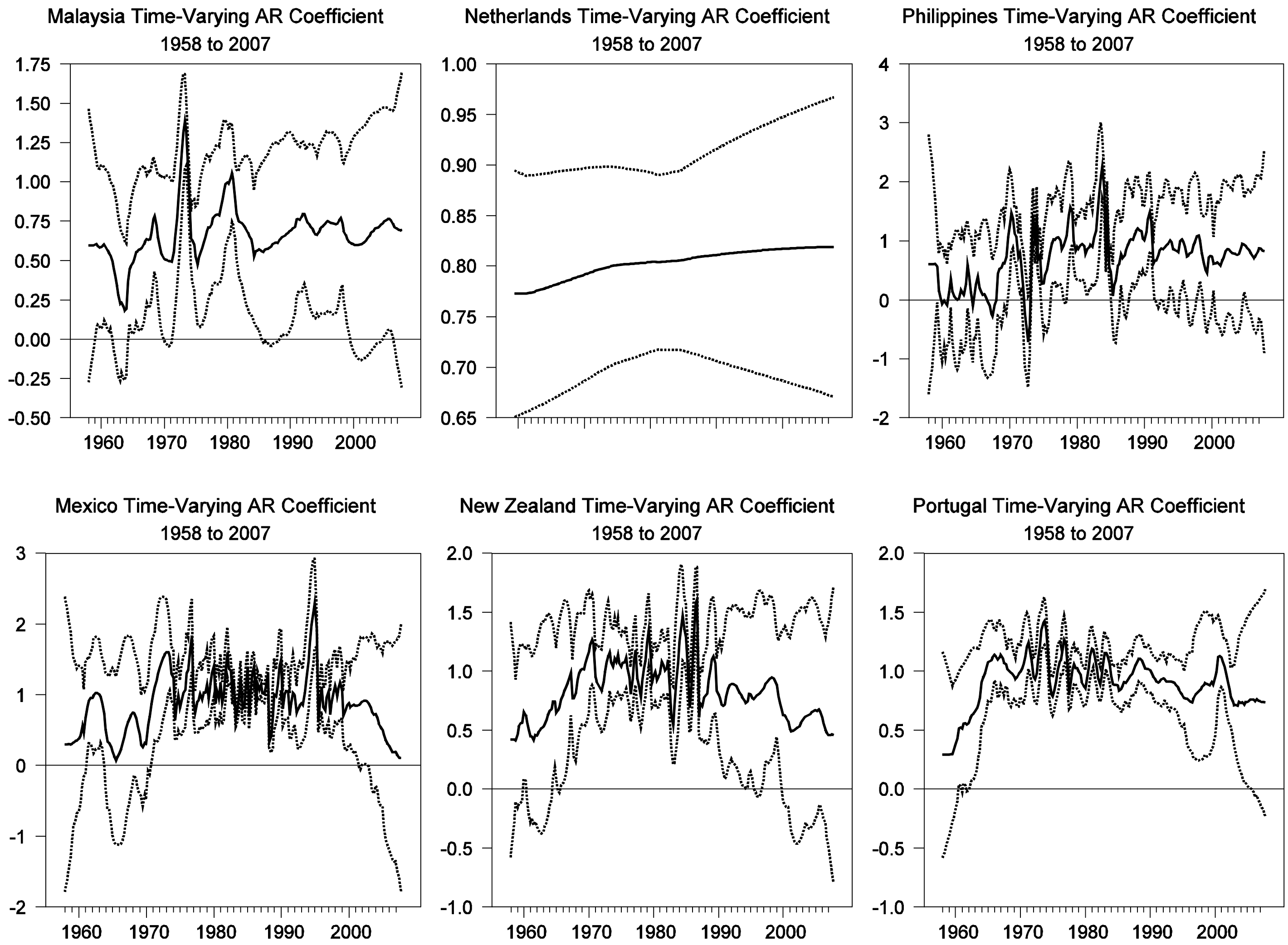
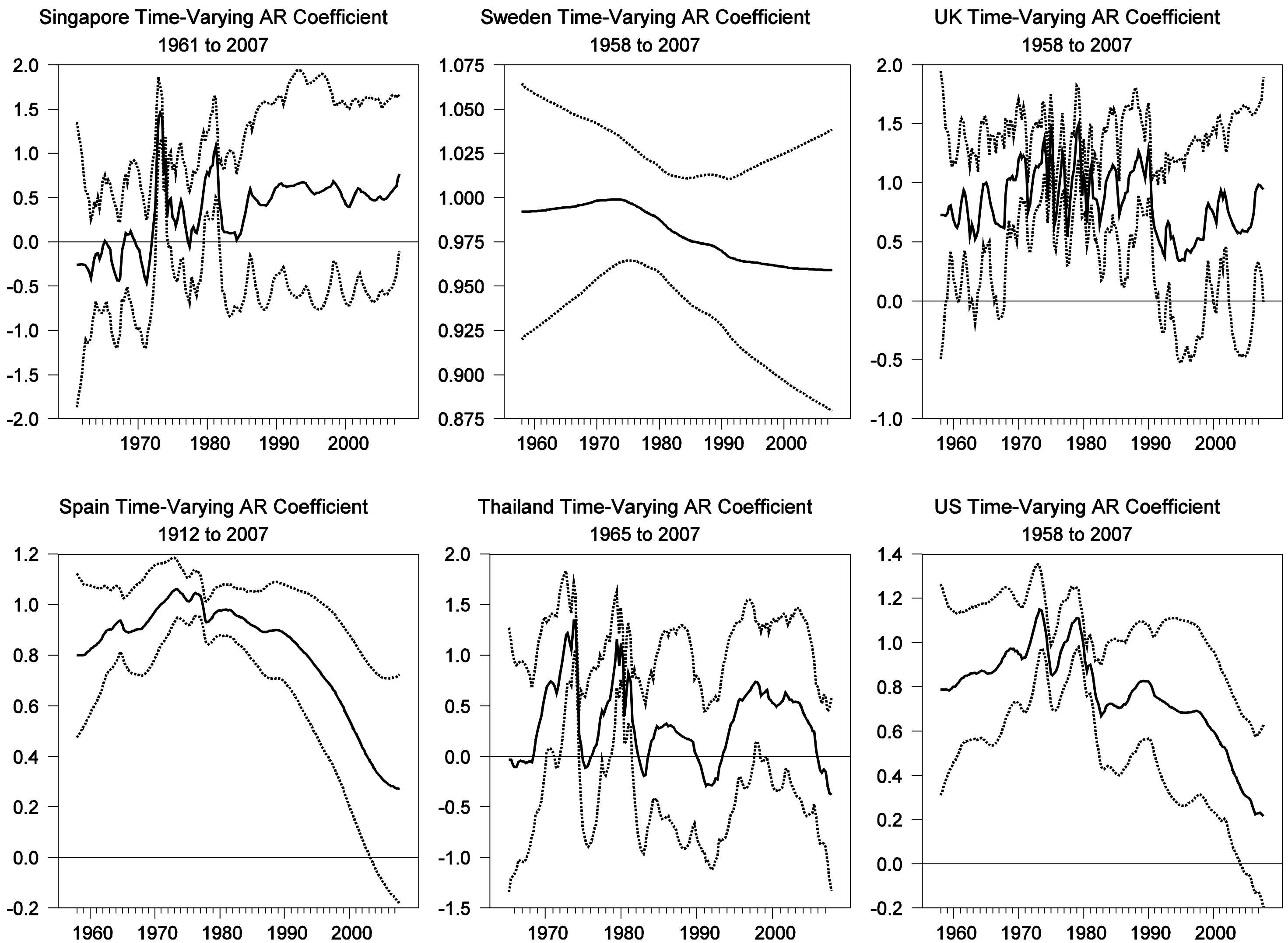

Malaysia Time-Varying AR Coefficient
1958 to 2007
1.75
1.50
1.25
1.00
0.75
0.50
0.25
0.00
-0.25
-0.50
1960 1970 1980 1990 2000
Netherlands Time-Varying AR Coefficient
1958 to 2007
1.00
0.95
0.90
0.85
0.80
0.75
0.70
0.65
Philippines Time-Varying AR Coefficient
1958 to 2007
4
3
2
1
0
-1
-2
1960 1970 1980 1990 2000
Mexico Time-Varying AR Coefficient
1958 to 2007
3
2
1
0
-1
-2
1960 1970 1980 1990 2000
New Zealand Time-Varying AR Coefficient
1958 to 2007
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
1960 1970 1980 1990 2000
Portugal Time-Varying AR Coefficient
1958 to 2007
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
1960 1970 1980 1990 2000

Figure 1.5


Singapore Time-Varying AR Coefficient
1961 to 2007
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
-1.5
-2.0
1970
1980
1990
2000
Sweden Time-Varying AR Coefficient
1958 to 2007
1.075
1.050
1.025
1.000
0.975
0.950
0.925
0.900
0.875
1960
1970
1980
1990
2000
UK Time-Varying AR Coefficient
1958 to 2007
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
1960
1970
1980
1990
2000
Spain Time-Varying AR Coefficient
1912 to 2007
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
1960
1970
1980
1990
2000
Thailand Time-Varying AR Coefficient
1965 to 2007
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
-1.5
1970
1980
1990
2000
US Time-Varying AR Coefficient
1958 to 2007
1.4
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
1960
1970
1980
1990
2000