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# Inflation Persistence Before and After Inflation Targeting: A Fractional Integration Approach

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Abstract

We investigate the empirics of persistence in the inflation series for 13 OECD countries that explicitly adopted an inflation targeting (IT) regime before 1992. We estimate persistence in the pre- and post-IT periods using the modified log periodogram proposed by Kim and Phillips (2006, 2000) and Phillips (2007) and test for equality across the two periods. Our findings indicate that all inflation series show no evidence of unit-root behavior over the entire sample, and in the respective pre- and post-IT periods. Mean reversion and stationarity as well as mean reversion and nonstationarity exist in the pre-IT period, while mean reversion and stationarity characterize the post-IT period. Inflation exhibits fractional integration behavior over the entire sample period, the pre-IT period, and, in most cases, also in the post-IT period. The adoption of inflation targeting coincides with a structural break in all inflation series and marks a decrease in the point estimates of inflation persistence in most countries. For only about half of the countries, however, we can formally reject the null hypothesis of equality of inflation persistence against the alternative that inflation persistence declines in the post-IT period. Significant variations and asymmetries exist in inflation persistence across the countries in the sample, suggesting that the IT regime does not equalize persistence across the IT countries.

**Keywords:** persistence, modified log periodogram, inflation targeting, fractional integration

**JEL classification:** C14; E31; C22

## 1. Introduction

Since first adopted in New Zealand in 1990, inflation targeting (IT) has become the focal point of extensive research, especially since a growing number of central banks, in both developed and emerging countries, have shifted toward using IT to implement their monetary policy. A number of studies have emerged that attempt to measure the impact and effectiveness of IT on economic performance – lowering the average inflation rate, raising the average rate of output growth, and lowering inflation and/or output growth volatility (see, e.g., Ball and Sheridan, 2005; Batini and Laxton, 2007; Gonçalves and Salles, 2008; and more recently, Brito and Bystedt, 2010; Fang, Miller, and Lee, 2012).

Another strand of empirical literature examines the successfulness of IT in reducing inflation persistence. Inflation persistence is an unobservable variable that plays an important role in the conduct of monetary policy. Given an inflationary shock, inflation returns to its target more quickly (i.e., inflation exhibits less persistence), the more effectively the monetary authorities can reduce inflation fluctuations, all else equal (Fuhrer, 1995).[1] Since more persistent inflation takes longer to adjust to an inflation shock, the monetary authorities may need to implement a policy response to bring inflation back to its equilibrium level. On the other hand, with low inflation persistence, inflation reverts to its initial level more quickly after a shock occurs. In such a case, the response to the inflation shock may not require active monetary policy. In the worst case, inflation may follow a random walk (i.e., an I(1) process) making it impossible for central banks to bring it under control. In the best case, inflation may follow an

[1] Recent contributions on inflation persistence in the US include Kumar and Okimoto (2007), Pivetta and Reis (2007), and Mehra and Reilly (2009). Beechey and Österholm (2009), Batini (2006) and Meller and Nautz (2012) consider inflation persistence in the Euro area, while Gadea and Mayoral (2006) examine inflation persistence in 21 OECD countries.

I(0) process, implying that it reverts to its initial level soon after a shock occurs.[2] As a consequence, the optimal timing and size of monetary policy crucially depend on the knowledge of how shocks affect the dynamics of inflation. This is particularly important for countries with IT regime.

The IT literature maintains as a central proposition that establishing well-anchored inflation expectations supports a successful IT policy. Well-anchored expectations enable the central banks that adopt IT to maintain stability of output and employment in the short run, while ensuring the stability of prices in the long run. Erceg and Levin (2003) and Orphanides and Williams (2005) show that inflation proves less persistent if agents know more about central bank objectives. In the imperfect credibility model of Erceg and Levin (2003), the lack of credibility on the inflation target provides an important source of inflation persistence. Similarly, in the model of Orphanides and Williams (2005), well-anchored long-run inflation expectations lead to less persistent inflation than if the public remains uncertain about the long-run inflation objective. Recently, Yigit (2010) shows that inflation persistence declines if IT succeeds in reducing the heterogeneity of inflation expectations. Thus, the analysis of inflation persistence is particularly important for countries with an IT regime since the decline of inflation persistence reflects the improved credibility of monetary policy and suggests that inflation expectations have become better anchored than previously was the case. Siklos (1999) Kuttner and Posen (2001), Petursson (2005), Benati (2008), and Yigit (2010), however, provide mixed evidence as to whether IT reduces the persistence in inflation rates.

[2] Whether inflation follows a stationary or nonstationary process possesses theoretical implications, since a number of macroeconomic models (Dornbusch, 1976; Taylor, 1979, 1980; Calvo, 1983; and Ball, 1993) assume stationary inflation. Additionally, models such as Fuhrer and Moore (1995) and Blanchard and Galí, (2007) suggest that inflation persistence captures structural characteristics of the economy that do not likely respond to policy actions. Thus, a policy of IT should exert no effect on inflation persistence. Others such as Batini (2006), Beechey and Osterholm (2007), Benati (2008), and Mehra and Reilly (2009) present evidence that inflation persistence varies across monetary regimes.

A large number of studies in the existing literature model inflation as an autoregressive process and measure persistence using different metrics, such as the integer order of integration, the half-life of responses to shocks, the largest autoregressive root, and the sum of the autoregressive coefficients.[3] By contrast, this paper focuses on an alternative parameterization of inflation dynamics, examining inflation persistence within the context of a long-memory I(*d*) model, where *d* is the order of fractional integration. The I(*d*) approach provides a more powerful framework to detect persistence than the standard unit-root analysis. From an integer order of integration perspective, inflation exhibits either a stationary, I(0), or a random walk, I(1), process. In the first case, shocks do not persist. They are transitory and mean-reverting. In the second case, shocks are permanent and non mean-reverting. The developing interest in I(*d*) reflects the growing awareness of the limitations of the I(1)-(I(0) framework and the increasing interest in models used to describe long-memory phenomena, and patterns of mean-reversion and responses to shocks that do not exist in conventional unit-root analysis.[4]

This paper examines the relationship between the adoption of inflation targeting and inflation persistence. We employ fractional integration methods to characterize the behavior of inflation dynamics for 13 OECD countries that switched to IT in 1991 or earlier: Australia, Canada, Chile, Iceland, Israel, Mexico, New Zealand, Norway, South Africa, South Korea,

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[3] Examples include Nelson and Plosser (1982), Fuhrer and Moore (1995), Cogley and Sargent (2001), Stock (2001), Cecchetti and Debelle (2006), Pivetta and Reis (2007), and Zhang et al. (2008) for the U.S.; O'Reilly and Whelan (2005) and Beechey and Osterholm (2009) for Europe and Levin and Piger (2004) and Levin et al. (2004) for a group of OECD countries. Barsky (1987), Ball and Cecchetti (1990), and Brunner and Hess (1993) suggest that the U.S. inflation contains a unit root. The unit-root property appears to occur in a wide array of countries examined in O'Reilly and Whelan (2005) and Cecchetti et al. (2007).

[4] Baillie (1996) provides an extensive review of the concepts of fractional integration in economic series. Long-memory processes are defined in both time and frequency domains. In the time domain, a process $y_t$ exhibits long-memory, if its autocorrelation function $\rho(k)$, $k$ = 1,2,.., decreases at a hyperbolic rate rather than the exponential decay in a covariance-stationary ARMA process. In the frequency domain, the spectrum for a long-memory process diverges to infinity at the zero frequency. In practical applications, long-memory emerges when the series possesses a pole on a part of the spectrum close to the zero frequency (Granger and Joyeux, 1980).

Sweden, Switzerland, and the United Kingdom. For each country, we estimate the persistence of inflation in the respective pre- and post-IT periods, and then proceed to test for equality of the persistence parameters across these two periods. We find that (*i*) inflation exhibits a fractional integration behavior over the entire sample period and the pre-IT period, and, in most cases, in the post-IT period, (*ii*) the adoption of IT coincides with a structural break in the inflation series, (*iii*) significant variations and asymmetries exist in inflation persistence across the IT countries, (*iv*) only about half of the countries in the sample experience significantly lower persistence in the IT period, (*v*) in the IT period all inflation series exhibit mean reversion and stationarity, while mean reversion and stationarity as well as mean reversion and nonstationarity are present in the pre-IT period, and (*vi*) for one country, South Africa, persistence increases, switching from stationarity to nonstationarity, after the adoption of IT.

Yigit (2010) comes closest to our work. He shows that long memory in inflation series can reflect heterogeneous inflation expectations. Other explanations include money supply persistence, aggregation across heterogeneous production by firms, or aggregation of individual prices into an index of prices. Yigit (2010) then argues that the adoption of IT can reduce the heterogeneity of inflation expectations by reducing the extent of long memory, something that would not occur if the other explanations caused long memory in the inflation series. Yigit (2010) estimates the fractional integration parameter in the inflation process before and after the regime switch to IT using an autoregressive fractionally integrated moving average (ARFIMA) model with and without a generalized autoregressive conditional heteroskedasticity (GARCH) error structure. Our methodology differs from Yigit (2010). We estimate inflation persistence using the modified log periodogram (MLP) method proposed by Kim and Phillips (2006, 2000) and Phillips (2007). The method is robust to AR(1) and MA(1) specifications (Choi and Zivot, 2007), and, unlike the ARFIMA approach, is semiparametric and, therefore, does not require the specification of the ARMA model.[5]

The MLP method extends the basic log periodogram regression to include the important case of the unit root. That is, much applied econometric work use the log periodogram regression originally developed by Geweke and Porter-Hudak (1983). But, its validity is limited to the stationary region. Conversely, the modified log periodogram (MLP) regression proposed by Kim and Phillips (2006, 2000) and Phillips (2007) extends statistical estimation and inference to the nonstationary region and, therefore, can appropriately test a fractionally integrated process I($d$) against both I(0) and I(1) processes.

The remainder of the paper is organized as follows. Section 2 describes the data and presents results of unit-root and stationarity tests. The conflicting nature of the findings of these tests suggests that neither the I(0) nor the I(1) models prove entirely appropriate for the inflation data. Section 3 briefly recalls the fractional integration approach for measuring inflation persistence and describes the modified log periodogram (MLP) proposed Kim and Phillips (2006, 2000) and Phillips (2007), which we apply to the inflation series. Section 4 reports the estimates of the persistence parameters based on the entire sample and the two subsamples defined by the date of adoption of IT, and presents the findings of the Hassler and Olivares (2008) and Kumar and Okimoto (2007) tests of the null hypothesis of no change in the fractional parameter $d$. This, in turn, enables us to formally answer the main question of whether inflation

[5] The misspecification of the short-run dynamics of the ARFIMA model may invalidate the estimation of the fractional integration parameter (Gil-Alana, 2004). The task of identifying $p$ and $q$ for an ARMA process by a simple analysis of the autocorrelation and partial autocorrelation functions proves nearly impossible. Schmidt and Tscherning (1995), Crato and Ray (1996), and Smith et al. (1997) consider various information criteria and assess, by Monte Carlo simulations, the performance of these criteria for a fractionally integrated true model against ARMA and ARFIMA alternative models. Their results suggest that for an ARFIMA($p$, $d$, $q$) data generating process (DGP), the identification of the true model may not occur for small or moderate sample size.

persistence changes after the adoption of IT. Section 6 summarizes the main results and provides concluding remarks.

## 2. Preliminary data analysis

We use monthly observations on the seasonally unadjusted Consumer Price Index (CPI) from January 1976 to June 2013 from the Main Economic Indicators (MEI) database (OECD.StatExtracts) for the following countries that adopted IT in 1991 or earlier: Canada, Chile, Iceland, Israel, Korea, Mexico, Norway, South Africa, Sweden, Switzerland, and United Kingdom. In addition, quarterly observations on the seasonally unadjusted CPI from the first quarter 1976 to the second quarter 2013 for Australia and New Zealand, for which the monthly series do not exist, also come from the OECD database. We deseasonalize the data using the Census X12 seasonally adjustment method. We measure inflation as the annualized percentage change of the CPI.[6] Allowing for differencing of series leaves 449 monthly and 149 quarterly usable observations.

We rely on Bernanke et al. (1999), Mishkin and Schmidt-Hebbel (2001), and Fracasso et al. (2003) to identify the start date of the IT regime. Of the 13 OECD countries, none are Euro area countries; eight are industrialized economies: Australia (September 1994), Canada (February 1991), Iceland (July 2001), New Zealand (March 1990), Norway (March 2001), Sweden (January 1993), Switzerland (January 2000), and the United Kingdom (October 1992); and five are emerging market economies: Chile (January 1991), Israel (January 1992), South Korea (January 1998), Mexico (January 1999), and South Africa (February 2000). The IT adoption dates appear in parentheses. We do not consider Finland and Spain, which adopted inflation targets in 1993 and 1994, respectively, but abandoned them upon entering EMU in

[6] We transform the seasonally adjusted price data $P_t$ into annualized monthly (quarterly) percentage rate of inflation using $\pi_t = 12 \times 100 \log\left(P_t / P_{t-1}\right)$ for monthly data and $\pi_t = 4 \times 100 \log\left(P_t / P_{t-1}\right)$ for quarterly data.

1998. Some discrepancies occur in the exact timing of adoption in some countries. The reason usually reflects the gradual adoption of the IT regime, which makes exact timing of adoption somewhat difficult. Ball and Sheridan (2005) provide discussion on the discrepancies between announcement and implementation of IT.

Figure 1 depicts the thirteen inflation series. The shaded area denotes the IT regime. Broadly speaking, the figures suggest that all thirteen inflation series move roughly in parallel, especially in the 1990s, and highlight the well-known trends of inflation in the last forty years. Specifically, we observe the sharp increase of inflation in the seventies, followed by the moderate reduction in the eighties and the generalized decline in average inflation in the nineties and thereafter.

Table 1 summarizes the first four moments of the empirical distribution of the inflation series for each of the 13 OECD countries for the entire sample and the two non-overlapping subsample periods, corresponding to the respective pre- and post-IT regimes.[7] The sample statistics highlight the substantial and pronounced differences that exist both between the countries and for any given country between the two periods and support two important stylized facts when comparing the pre- and post-IT periods. On average, the mean and the standard deviation of all 13 inflation series decline after the adoption of IT, as widely reported in the literature. In addition, in both the pre- and post-IT periods, the 13 inflation series exhibit higher peaks and fatter tails than the normal distribution, but with few exceptions, the skewness and kurtosis statistics are lower in the post-IT period. Thus, overall, it appears that IT substantially affects all four moments of the empirical distribution of the inflation series.

[7] Prior to switching to IT, six of the thirteen countries in the sample (Chile, Iceland, Israel, Norway, Sweden, and the UK) used exchange rates as nominal anchors in stabilization programs, while four (South Korea, Mexico, South Africa, and Switzerland) used the money supply. Australia, Canada and New Zealand did not specify a nominal anchor before switching to IT.

We also conduct unit-root tests on each of the inflation series, which helps to identify the long memory of the series. Baillie (1996) indicates that the ADF (Dickey and Fuller, 1979) test possesses low power in distinguishing long-memory from nonstationary series while together the PP (Phillips and Perron, 1988) and KPSS (Kwiatkowski et al., 1992) tests can provide evidence on the integration order of the series.[8] For this reason, Table 2 reports only the PP and KPSS tests.[9] They differ in their null hypothesis. The PP test takes stationarity as the alternative, where the KPSS take stationarity as the null. We consider two versions of the tests: level stationarity (reversion to a constant mean), and trend stationarity (reversion to a constant trend). These tests provide conflicting evidence, as both the PP and KPSS tests unambiguously reject their respective nulls for all series. That is, we find uniform evidence against the unit-root and the stationarity hypotheses, which casts doubt on the ARIMA framework as a characterization of the data, and suggests fractional integration as more appropriate univariate model. [10]

## 3. The Methodology of fractional integration

This section briefly discusses our fractional measure of persistence and its estimation methodology. This dichotomous view of the I(1)-I(0) framework does not include all possible

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[8] Baillie (1996) considers four possible outcomes when considering the PP and KPSS tests jointly. First, reject the PP, but not the KPSS statistic, which provides evidence of a stationary series. Second, reject the KPSS, but not the PP statistic, suggesting a unit-root process. Third, do not reject both statistics, suggesting insufficient information on the long-memory characteristics of the series. Fourth, reject both statistics, which provides evidence of a process between I(0) and I(1) or a fractionally integrated process.

[9] The PP test uses Newey–West (1987) standard errors to account for serial correlation and adopts lags chosen deterministically following the rule of thumb $\mathrm{int}[4(T/100)^{2/9}]$ (Newey and West, 1994). The KPSS test uses the quadratic spectral kernel. Andrews (1991) and Newey and West (1994) indicate that the latter kernel yields more accurate estimates of the long-run variance than the Bartlett kernel in finite samples. The bandwidth comes from an automatic bandwidth selection routine. Hobijn et al. (2004) find that the combination of the automatic bandwidth selection and the quadratic spectral kernel yields the best small sample test performance in Monte Carlo simulations.

[10] Sample autocorrelations also indicate long memory as all inflation series reveal a slow rate of decay. None exhibit exponential decay like stationary data. For Chile, Israel, and Mexico, the autocorrelations fall at a faster rate, whilst those corresponding to the remaining countries exhibit sinusoidal behavior that decays slowly. In either case, the autocorrelations exhibit a persistent pattern of moderately high values. Mexico displays the largest autocorrelation at lag 1 (0.92), followed by Israel (0.869) and Chile (0.864), while Switzerland displays the lowest autocorrelation at lag 30 (0.127), followed by Korea (0.154) and Australia (0.178).

data generating processes. In particular, it does not account for mean reverting series with a substantially lower rate of mean reversion than exponential decay. In fact, a slow mean reversion occurs when the effects of shocks disappear at a slow rate and, unlike I(0) processes, fractional integration models can accommodate this possibility. The fractional representation of inflation, in contrast to the unit-root representation, permits different characterizations of the dynamics of inflation and different patterns of responses to shocks. Low levels of *d* characterize weak persistence, while in the unit-root models, such persistence does not exist. On the other hand, high levels of *d* exhibit persistence with reversion to the mean, whereas in the unit-root models, such persistence exists, but without reversion to the mean. Consequently, I(*d*) processes cover a wider range of dynamic behaviors that do not exist *a priori* in the unit-root literature.

We write the fractionally integrated process $y_t$ integrated of order *d* as follows:

$$\Delta^d y_t = \varepsilon_t , \tag{1}$$

where $\Delta^d$ is the fractional differencing operator, $\varepsilon_t$ is a stationary process of order zero, I(0), with zero mean and spectral density $f_\varepsilon(\lambda)$, where $\lambda$ is the Fourier frequency. Hosking (1981) makes the concept of fractional differencing $\Delta^d$ operational by utilizing a binomial expansion of the difference operator:

$$\Delta^d = (1-L)^d = \sum_{k=0}^{\infty} \binom{d}{k} \left(-L\right)^k = 1 - dL - \frac{d(1-d)L^2}{2!} - d\left(1-d\right)\left(2-d\right)\frac{L^3}{3!} \ \ldots, \tag{2}$$

where *L* is the backward-shift operator. The parameter *d* in equation (1) measures the extent of the memory in $y_t$. If $\varepsilon_t$ is white noise, then equation (1) is a fractional white-noise process (Mandelbrot and Van Ness, 1968; Mandelbrot, 1971). If, on the other hand, $\varepsilon_t$ is a stationary and invertible ARMA(*p*, *q*) process, then equation (1) is an ARFIMA(*p*, *d*, *q*) process (Granger and Joyeux, 1980; Hosking, 1981).

The parameter $d$ governs the long-run dynamics of $y_t$. Time-series data with different fractional values behave in different ways. Kumar and Okimoto (2007) discuss several attractive features of using $d$ as a measure of persistence. First, the I($d$) process permits the comparison of highly persistent series. Second, the parameter $d$ does not affect the short-run dynamics of the data. The autoregressive approach to measure the long-run dynamics intimately relates to the first-order autocorrelation of the data, which responds to both short- and long-run dynamics. Third, unlike the local-to-unity parameter in the unit-root models, we can estimate $d$ consistently.

Baillie et al. (1996) and Sowell (1992) detail the characteristics of an I($d$) process. In general, $y_t$ is stationary, mean reverting, and exhibits long memory if $0 < d < 0.5$; whereas if $-0.5 < d < 0$, $y_t$ is stationary, mean reverting, but displays "intermediate" memory (i.e., anti-persistence). In this latter case, the autocorrelations are all negative except at lag 0. When $d$ = 0, $y_t$ is I(0). That is, the correlation between consecutive observations fades quickly and the series returns to its constant mean. This series is modeled as an ARMA($p$, $q$) and exhibits short-memory (weakly dependent), as opposed to long memory when $d > 0$. Contrary to short memory, long memory means a significant dependence between observations widely separated in time and, therefore, the effects caused by shocks decay slowly or, alternatively, shocks produce long lasting effects, that is, shocks arbitrarily far away in time still exhibit some influence on the dynamics of the process.

The interesting case for many economic and financial series occurs when $0.5 \le d < 1$. This process reverts to its mean, but is a nonstationary process. Despite the influence of remote shocks, this series will revert to its mean in the long run. In other words, the series exhibits long, but transitory, memory. This differs substantially from the case $d \ge 1$, where the mean of the series exerts no influence on the long-run evolution of the series because recent and remote shocks completely dominate its movement. In this case, the series is nonstationary with infinite or permanent memory and does not revert to its mean. In particular, a unit-root process occurs when $d = 1$, corresponding to a unit root ( $\rho = 1$) in the autoregressive model. This series is modeled as an ARIMA(*p*, 1, *q*) by differencing the unit-root process In the different cases described above, only when $d < 1$ does a long-run equilibrium level of the series exist that is governed by the mean of the process. Obviously, when $0 < d < 1$, unit-root tests (testing $d$ = 1 under the null) or stationarity tests (testing $d$ = 0 under the null) may fail to detect mean reversion in the series and conclude that the series contain a unit root (Diebold and Rudebush, 1991 and Gonzalo and Lee, 2000).

Phillips (2007) notes that the theory of statistical inference for the log periodogram regression only exists for the stationary long-memory case with fractional integration parameter $-0.5 < d < 0.5$ and, therefore, does not address the case of $d \geq 0.5$. Generally speaking, seldom does any prior information exist about the range of $d$ before estimation. Thus, the determination of semiparametric estimators for $d \geq 0.5$ is important from both theoretical and practical reasons. In practical applications, however, researchers typically apply the log periodogram regression method to apparently nonstationary series by first differencing the data. The log periodogram estimator is not invariant to first differencing (Agiakloglou et al., 1993), so that a bias may exist due to over-differencing. Thus, absent prior information about the range of $d$ before estimation, the need exists for a more flexible estimation technique and inference for both the stationary and the nonstationary cases. Kim and Phillips (2006, 2000) and Phillips (2007) propose a method to estimate $d$ using a modified log periodogram regression (MLP) estimator that accommodates the nonstationary range $d \geq 0.5$ and is consistent when $0.5 < d = 1$. The MLP modification of the GPH estimator uses an exact representation of the discrete Fourier transform in the unit-root case. The regression involves a linear regression similar to equation (3), where the modified periodogram ordinates $I_v(\lambda_j) = \left|v(\lambda_j)\right|^2$ replace the periodogram ordinates $I_X(\lambda_s)$, and where $v(\lambda_j) = \xi(\lambda_j) + \frac{e^{i\lambda_j}}{1-e^{i\lambda_j}} \frac{y_T}{\sqrt{2\pi T}}$ and $y_T$ is the value of the final sample observation. Kim and Phillips (2006, 2000) show that the distribution of the MLP estimator, $\hat{d}_{MLP}$ , is $\sqrt{m}\left(\hat{d}_{MLP} - d\right) \to N\left(0, \frac{\pi^2}{24}\right)$. A semiparametric test statistic for the null of a unit root (i.e., $H_0 : d = 1$) uses the statistic (Phillips 2007):

$$z_d = \frac{\sqrt{m}\left(\hat{d}_{MLP} - 1\right)}{\pi / \sqrt{24}} \tag{4}$$

with critical values from the standard normal distribution.

## 4. Empirical results

### *4.1 Analysis of the entire sample*

We estimate the fractional integration parameter *d* for each of the 13 countries over the entire sample period using the MLP regression described in Section 3. Table 3 reports the estimates of inflation persistence, measured by its order of fractional integration *d*, together with its standard error, and the associated 95% confidence intervals. In addition, Table 3 provides the results of the dual tests of $H_0 : d = 0$ against $H_1 : d \neq 0$ and $H_0 : d = 1$ against $H_1 : d < 1$. The first tests short memory (corresponding to the I(0) hypothesis) versus long-memory alternatives; the second tests a unit-root versus long-memory alternatives. Throughout this subsection, we do not consider the possibility of changes in the persistence of inflation. We simply confirm that the order of fractional integration appropriately measures persistence before conducting formal tests of the equality of inflation persistence.

An important practical problem emerges with the log periodogram regression -- choosing the bandwidth *m*, defined as $m = T^{\alpha}$, where $0 < \alpha < 1$ and *T* is the sample size.[11] Hurvich et al. (1998) establish that the optimal *m* that minimizes the asymptotic mean squared error is of order $O(m^{0.80})$.[12] Since then, various researchers set $\alpha$ around 0.8. For instance, Maynard and Phillips (2001), Kellard and Sarantis (2008), and Phillips (2005) use 0.75, while Luu and Martens (2003), and Andersen et al. (2003) use 0.8. To circumvent the difficulties inherent to optimal selection rules, one can experiment with a few different values of $\alpha$. For instance, Kim and Phillips (2006) use $\alpha$ = 0.7 and 0.8, Kilic (2004) uses $\alpha$ = 0.5, 0.6, 0.7, and 0.8, while Choi and Zivot (2007) and Li (2002) use $\alpha$ = 0.7, 0.75, and 0.8. We adopt this common practice and document our findings for $\alpha$ = 0.7, 0.75, and 0.8. For our inflation series, similar values strike a satisfactory balance between the variance and bias of the estimates. Since we use monthly and quarterly data, the use of smaller bandwidths produces standard errors too large to provide meaningful information about the order of integration.

Another important issue concerns the detrending of the series before estimating the memory parameter. Since the presence of a trend can dominate the dynamics of the series, we follow Phillips' (2007) suggestion and apply the MLP estimator to the linearly detrended data. Table 3 presents estimates of *d* for each inflation series across the three different bandwidths using the MLP regression. In total, this table reports 39 relevant estimates of *d*. We do not difference the data, since the MLP regression does not require stationarity. The results from the

[11] From an empirical perspective, the estimates of *d* usually vary significantly with the choice of *m*. Too small an *m* leads to imprecise estimates because of a large standard deviation. Too large an *m* leads to a biased estimate of *d*. A large literature focuses on the choice of optimal bandwidth, (e.g., Delgado and Robinson, 1996 and Hurvich et al., 1998).

[12] Hurvich et al. (1998) find that the choice of $\alpha$ = 0.5, originally suggested by Geweke and Porter-Hudak (1983), and extensively used in empirical applications, leads to inferior performance to the optimal choice in reasonable samples.

analysis of the full sample provide solid evidence that fractional integration describes the inflation process. All estimates point in the direction of long memory. The estimates for South Korea, Mexico, New Zealand, Norway, South Africa, Switzerland and the UK exhibit a pattern of stability (less than 0.1 differences) across the bandwidths, while the estimates for Australia and Chile increase in the bandwidth from 0.7 to 0.75 and then drop back for 0.8, and the estimates for Canada, Iceland, Israel, and Sweden decrease (by 0.231 to 0.281) in the bandwidth. The positive point estimates of $d$ range between 0.213 (Norway with $\alpha = 0.70$) and 0.807 (Israel with $\alpha = 0.70$), which confirms the existence of long memory in the inflation series.

The tests of $H_0$: $d = 0$ and $H_0$: $d = 1$ reject both the I(0) and the I(1) hypotheses for $\alpha =$ 0.70, 0.75, and 0.80 and, thus, strongly support the fractional integration approach to detect changes in inflation persistence. The results also establish that significant heterogeneity in inflation persistence exists among the 13 countries. For Canada, South Korea, Norway, South Africa, Sweden, Switzerland, and the UK, the point estimates fall in the stationary and mean-reverting segment for $\alpha =$ 0.70, 0.75, and 0.80, while for Chile, Israel, Mexico, and New Zealand, the point estimates fall in the nonstationary, but still mean-reverting, segment for $\alpha =$ 0.70, 0.75, and 0.80. For Australia and Iceland, the estimates fall in the stationary and mean-reverting segment only for $\alpha$ = 0.70 and 0.80 and for $\alpha$ = 0.80, respectively, and in the nonstationary and non-mean-reverting segment for 0.75 and for 0.70 and 0.75, respectively. We also notice that in no case does the 95-percent confidence interval for $d$ include the nonstationary, non-mean-reverting range $d \geq 1$ . Except for Norway, South Africa, and Switzerland, however, the upper boundary of the 95-percent confidence intervals include the nonstationary mean-reverting range $0.5 \leq d < 1$.

Does the adoption of IT produce a structural break in the inflation series? Does inflation persist less when policymakers place greater emphasis on inflation and less emphasis on output? Does IT eliminate inflation persistence? The findings in Table 3 may prove misleading if the inflation series experience structural breaks at the date of IT adoption.

We conduct a structural break test using the Chow test (Chow, 1960). We treat the break date as known and fix the timing of the break exogenously to correspond to the date of adoption of IT, although we note that some uncertainty exists about the exact timing. A natural question, however, arises: Does the order of estimation matter (Choi and Zivot, 2007)? That is, should we test first for structural change and then for fractional integration or should we test first for fractional integration and then for structural change? In the first case, we employ the raw, unfiltered inflation series; in the second case, we employ the filtered, fractionally differenced inflation series. Since no guarantee exists that the results will match, we apply the test to both series.

We consider two models: the pure structural change in mean model and the partial structural change model with the AR coefficient of the lagged inflation not subject to structural change. We estimate both models by OLS with the break point exogenously selected to correspond to the date of IT adoption. Bai (1994) establishes the consistency of the OLS estimator for I(0) processes with a mean change. Kuan and Hsu (1998) extend this consistency result to I($d$) processes with mean change, provided that $-0.5 < d < 0.5$. Additionally, Kuan and Hsu (1998) demonstrate that for data that are I(1) or I($d$) with $0 < d < 0.5$, the OLS estimator suffers from the spurious change problem (i.e., it may spuriously suggest a change point when none exists). We note that the approach of Bai and Perron (1998, 2003), although frequently employed, does not explicitly allow for long memory.

Tables 4 and 5 present the results of using the log likelihood ratio version of the Chow test. The test, under the null hypothesis of no breaks at the IT adoption date, is distributed as a $\chi^2(1)$. The findings for the unfiltered data displayed in Table 4 consistently reject the null hypothesis of no break at the 1-percent level in both the pure and partial structural change models. The findings for the filtered (fractionally differenced) data prove mixed, however. In Table 5, the log-likelihood ratio test accepts the null hypothesis that the IT adoption date does not define a break in the inflation series for Australia, Chile, Iceland, Israel, New Zealand, and the UK. For the remaining inflation series, the test still rejects the null at the 5-percent level. Thus, Canada, South Korea, Mexico, Norway, South Africa, Sweden, and Switzerland exhibit evidence of structural change even after we adjust for long memory.

The results based on the *F*-test and the Wald test versions do not differ. At least in part, the results confirm the conjecture of Levin and Piger (2004) that breaks in inflation in the early 1990's coincide with the spreading of IT. Some reservations, however, apply to the test findings applied to the unfiltered data. In particular, the OLS estimator of breaks is not reliable for nonstationary I($d$) data with $0.5 \leq d < 1.5$ (Hsu and Kuan, 2008). The OLS estimator, however, does provide reliable estimates for stationary I($d$) data with $-0.5 < d < 0.5$. The estimates of $d$ for Canada, South Korea, Norway, South Africa, Sweden, Switzerland, and the UK fall in the stationary region and, therefore, we expect reliable Chow test results. The estimates of $d$ for the remaining countries (Chile, Iceland for $\alpha$ = 0.70, 0.75, Israel, Mexico, New Zealand and Australia for $\alpha$ = 0.75), however, fall in the nonstationary region, so we must interpret these results more cautiously.

### *4.2 Analysis of the pre- and post-IT periods*

Tables 6 and 7 report the MLP estimates and their standard errors obtained for the sub-sample periods, corresponding to the periods before and after the IT adoption dates, respectively.

Clearly, the results confirm the presence of fractional integration for the majority of the series in both the pre- and post-IT periods. The fractional property of inflation is not spurious, since it does not vanish when permitting mean breaks.

For the pre-IT period reported in Table 6, the point estimates of $d$ do not change much across bandwidths, except for Australia, Canada, Iceland, and Israel where $d$ varies by more than 0.20 (Switzerland's range is 0.197). In all cases, however, the estimates fall in the open interval between zero and one, and exhibit reasonable standard errors. Except for South Africa, we can reject the null hypothesis of $d$ = 0 for all countries. For South Africa, the null hypothesis of $d$ = 0 is not rejected for $\alpha$ = 0.70, 0.75, and 0.80. Similarly, we can reject the null hypothesis of $d$ = 1 for all countries, except Australia, Canada, Israel, and Mexico for some, generally the lower, bandwidths. The estimates fall in the nonstationary, but mean-reverting, segment for Australia, Canada, Chile, Israel, and Mexico for all bandwidths, while the estimates fall in the stationary, mean-reverting segment ($0 < d < 0.5$) for Norway, South Africa, Sweden, Switzerland, and the UK for all bandwidths. For Iceland, South Korea and New Zealand, the estimates fall in the nonstationary mean-reverting segment for lower bandwidths and in the stationary mean-reverting segment for higher bandwidths. From these results, we find that South Africa experiences the most stationary series, whereas Australia and Mexico experience the most nonstationary series. Moreover, the only countries that experience the nonstationary and non-mean reverting range $d \geq 1$ at the 95-percent confidence interval include Australia with $\alpha$ = 0.70, 0.75, and 0.80, Canada with $\alpha$ = 0.70, and 0.75, and Israel and Mexico with $\alpha$ = 0.70. Other countries, however, exhibit confidence intervals for $d$ that span the stationary/nonstationary boundary, such as Chile, Iceland, South Korea, New Zealand, and the UK for all bandwidths; Norway for $\alpha$ = 0.70 and 0.75; Sweden for $\alpha$ = 0.75; and Switzerland for $\alpha$ = 0.70. These results are consistent with the empirical literature (e.g., see Kouretas and Wohar, 2012; Gadea and Mayoral, 2006; and Baum at al., 1999). A comparison of our pre-IT findings with the findings of Yigit (2010) for the common set of countries (i.e., Australia, Canada, Israel, New Zealand, Sweden and the UK) indicates that his long-memory parameter is lower, except for Canada where inflation persistence is an I(0) process.

The findings from the post-IT period reported in Table 7 differ distinctly. The point estimates of $d$ for the post-IT period are smaller than those for the pre-IT period, suggesting a higher tendency for mean-reversion. Also, the variability of $d$ estimates falls with Norway exhibiting the highest range of 0.192. The estimates prove much less sensitive to the choice of the bandwidth and almost no estimates of $d$ fall in the nonstationary, but mean-reverting, segment ($0.5 < d < 1$) except for Chile and South Africa for all bandwidths; and Iceland for $\alpha$ = 0.75. We reject the null hypothesis of $d$ = 1 for all series across $\alpha$ = 0.70, 0.75 and 0.80 and no instance occurs where the 95-percent confidence interval for $d$ includes the nonstationary and non-mean reverting range $d \geq 1$. On the other hand, the findings show substantial asymmetries in inflation persistence across the IT states, where the point estimates of $d$ differ in persistence even among the IT countries with mean reverting inflation dynamics. Thus, a shock to inflation likely generates different dynamics for each country due to the differences in persistence. We reject the null hypothesis of $d$ = 0 for Australia, Chile, Iceland, Israel, Mexico, New Zealand (except for $\alpha$ = 0.70), South Africa, Sweden, and Switzerland for $\alpha$ = 0.70, 0.75, and 0.80. Shocks to the series are more persistent in Chile, Iceland, Israel, Mexico, New Zealand, South Africa, and Switzerland (except for $\alpha$ = 0.80), where the upper boundary of their 95-percent confidence intervals includes the nonstationary, but mean-reverting, region. For Canada, South Korea, Norway, and the UK, instead, we cannot reject the null hypothesis of $d$ = 0 for $\alpha$ = 0.70, 0.75, and 0.80. Norway and South Korea, however, provide an interesting case.

The lower bands of 95-percent confidence intervals for Norway and South Korea are in the stationary region, but display substantial evidence of antipersistence. Time-series analysis generally does not experience antipersistence, and the analysis of inflation with respect to long-memory properties typically focuses on persistence. The data in the post-IT period, however, expose the possibility of antipersistence in the dynamics of inflation. Both persistent and antipersistent processes are mean-reverting and show hyperbolic decay. A persistent process, however, implies that a positive (negative) movement generally follows a positive (negative) movement. In contrast, an antipersistent process implies that a positive (negative) movement generally follows a negative (positive) movement. In other words, the persistent process trends, whereas the antipersistent process reverses itself more often than a white-noise process. These results suggest that the inflation dynamics in Norway and South Korea are complex, with many mechanisms interacting to generate a stronger pattern of mean reversion. Yigit (2010) finds estimates of antipersistence for Canada and Israel, but the results are significant only for Israel in the ARFIMA-GARCH(1,1) specification. A comparison of our post-IT findings with the findings of Yigit (2010) for the common set of countries is warranted. For Canada and the UK, our results do not differ much from Yigit (2010): inflation is an I(0) process. Similarly, for New Zealand and Sweden, our results do not differ from Yigit (2010): inflation is a long-memory process. For Australia and Israel, however, we find that inflation is a long-memory process, while Yigit (2010) finds that an I(0) process, using the ARFIMA specification. Conversely, using the ARFIMA-GARCH specification, Yigit (2010) finds that Australia is I(0) while Israel follows a strongly antipersistent behavior process.

In sum, we can generally characterize the inflation series in the pre- and post-IT periods by fractional integration behavior, with an order of integration distinguishable for the most part from both zero and one. Also, the point estimates of $d$ in the post IT period generally decrease for all countries for $\alpha$ = 0.70, 0.75 and 0.80, except South Africa. For Norway and South Korea the decrease spans the antipersistent region. The point estimates of the order of integration for South Africa for $\alpha$ = 0.70, 0.75 and 0.80 move from a stationary range in the pre-IT period to a nonstationary, but still mean-reverting, range in the post-IT period, with the 95-percent confidence interval encompassing the stationarity/nonstationarity boundary. [13]

South Africa produces results counter to the rest of our sample. These findings, however, are consistent with Burger and Marinkov (2008) who, using a recursive unrestricted VAR, conclude that IT in South Africa did not contribute to lower inflation persistence. Similarly, Kaseeram and Contogiannis (2011) find that South Africa did not experience a significant decline in inflation persistence since the adoption of the IT regime. These findings are also consistent with Gupta and Ulwilingiye (2012) and Gupta et al. (2010). In contrast, Rangasamy (2009), using the standard autoregressive approach, finds a significant drop in inflation persistence in the years following the adoption of IT. The resolution of this empirical puzzle lies beyond the scope of this paper.[14]

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[13] Although Phillips (2007) suggests that we remove deterministic linear trends from the series before the application of the MLP estimator, we also estimate the persistence of the inflation series before and after the IT adoption without detrending the data at the suggestion of a referee. For all countries, the estimates of $d$ without detrending rise relative to their values with detrended data in the pre-IT period. In the post-IT period, the estimates rise only for Canada, Israel, and Mexico. For the remaining countries, the estimates instead fall slightly using the detrended model. The Kumar-Okimoto and Hassler-Olivares test results are also generally invariant to detrending. The main exceptions are Australia (the Kumar-Okimoto test results reverse from insignificant to significant), Chile (both test results reverse from insignificant to significant), and South Africa (both test results reverse from significant to insignificant). Without detrending, the difference in the estimates of persistence for the pre- and post-IT periods are much smaller than we previously found, resulting in the failure to reject the null hypothesis of equality. We note, however, that the regression of the rate of inflation on time in these three countries as well as in the remaining ten countries produces a significantly negative linear trend estimate. We interpret this to imply that ignoring the trend in the MLP regressions leads to misspecification problems.

[14] The frequency of observations does not affect our findings, since similar results also occur using quarterly data.

### *4.3 Tests against a break in d*

The results of the previous analysis suggest that the point estimates of inflation persistence are generally lower in the post-IT period. To examine this conclusion formally and more rigorously, however, we test the null hypothesis that no significant change in inflation persistence occurs in each of the 13 countries following the adoption of IT. We test this hypothesis against the alternative hypothesis that a significant decline in inflation persistence occurs. Relying exclusively on the IT adoption date may not identify the correct timing of the possible declines in inflation persistence. In fact, we could assume that the actual transition process from NIT to IT follows a gradual, rather than an instantaneous, process, but trying to identify that process of adjustment proves a demanding task. Consequently, we simply use the two samples of data, the pre- and post-IT samples, to test the difference in *d* between the two periods. This may not provide the most powerful test to detect a decline in persistence (i.e., the tests are rather conservative in the sense that they may fail to detect a decline in inflation persistence). If we reject the null of no decline, however, this constitutes strong evidence in favor of a decline in inflation persistence.

Define $\hat{d}_1$ and $\hat{d}_2$ as the estimates of *d* in the pre- and post-IT subsamples. Kumar and Okimoto (2007) propose an approximate upper bound of the test statistic for the null hypothesis $H_o : d_1 = d_2$ against the alternative $H_o : d_1 > d_2$

$$\frac{\hat{d}_1 - \hat{d}_2}{\sqrt{\mathrm{Var}\left(\hat{d}_1\right) + \mathrm{Var}\left(\hat{d}_2\right)}} \to N\left(0,1\right), \qquad (6)$$

assuming $\mathrm{Cov}\left(\hat{d}_1, \hat{d}_2\right) > 0$ so that

$$\mathrm{Var}\left(\hat{d}_1 - \hat{d}_2\right) = \mathrm{Var}\left(\hat{d}_1\right) + \mathrm{Var}\left(\hat{d}_2\right) - 2\mathrm{Cov}\left(\hat{d}_1, \hat{d}_2\right) \leq \mathrm{Var}\left(\hat{d}_1\right) + \mathrm{Var}\left(\hat{d}_2\right).$$

Hassler and Olivares (2008) also suggest another test, based on Shimotsu (2006), where they derive the asymptotic variance, as a result of the asymptotic independence of the estimators, as follows

$$\operatorname{Var}\left(\hat{d}_1 - \hat{d}_2\right) = \frac{1}{4m_1} + \frac{1}{4m_2} = \frac{m_1 + m_2}{4m_1 m_2} \tag{7}$$

where $m_1$ and $m_2$ denote the bandwidths from the respective subsamples (Hassler and Meller, 2014). The Hassler-Olivares asymptotic test is then defined as

$$2\frac{\sqrt{m_1 m_2}}{\sqrt{m_1 + m_2}}\left(\hat{d}_1 - \hat{d}_2\right) \to N\left(0,1\right). \tag{8}$$

Table 8 reports the results of the Kumar-Okimoto and Hassler-Olivares tests for differences in the pre- and post-IT estimates of $d$ corresponding to $\alpha$ = 0.70, 0.75, and 0.80. The Kumar-Okimoto test indicates a significant decline in persistence for only four countries: Canada, South Korea, Mexico, and Norway for all bandwidths. For Australia, Chile, Iceland, New Zealand, Sweden, and Switzerland, the test fails to reject the null hypothesis for all bandwidths, while for Israel and the UK the test is sensitive to the choice of $\alpha$. We reject the null hypothesis for Israel only for $\alpha$ = 0.70, while for the UK only for $\alpha$ = 0.75 and 0.80. For South Africa, the test paradoxically indicates a significant increase in persistence.

The Hassler-Olivares test findings, conversely, indicate a significant decline in persistence for Australia, Canada, Israel, South Korea, Mexico, Norway, and the UK for all bandwidths. For Chile, New Zealand, Sweden, and Switzerland, the test suggests that no significant decline in persistence occurs for all bandwidths, while for Iceland the test suggests a significant decline only for $\alpha$ = 0.70. For South Africa, the Hassler-Olivares test reaches the same conclusion of the Kumar-Okimoto test, but with the stronger *t*-statistics.[15] Since Sweden and Switzerland exhibit relatively low estimates for *d* in the first subsample, the tests cannot reject the null hypothesis of no decline in inflation persistence in the second subsample. Thus, taken as a whole, the findings provide only partial evidence for the decline in the persistence of inflation in the 13 IT countries.[16] Compared to the Kumar-Okimoto test findings, the Hassler-Olivares test findings provide more evidence of a significant decrease in persistence in the post-IT period.

Finally, we document the robustness of the results to alternative specifications of the adoption dates.[17] Yigit (2010) argues that the original break dates could be misleading due to an initial credibility transition. In addition, there are various issues concerning the exact timing of IT adoption in some countries. Chile, Australia, and New Zealand are three cases in point (Petursson, 2005). For example, for Chile, we define the starting date of IT as January 1991

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[15] Since 1980, monetary policy implemented three distinct regimes in South Africa. From the 1980 to 1989, monetary policy did not successfully contain inflation. From 1990 to 2000, a significant improvement in the performance of the South African Reserve Bank (SARB) occurred as average inflation fell from 13.06 percent to 8.87 percent. Since the SARB did not pursue an explicit inflation target, we can characterize this period as implicit IT. From 2000 till the present, the SARB pursues low inflation with an average inflation equal to 5.21 percent. This period differs from the second in that the SARB pursues an official and explicitly stated inflation target. Consequently, we conduct additional tests to assess the importance of the second period. First, we include the second period in the IT period. The results do not differ qualitatively from those reported in the tables. We can reject the nulls of $d = 0$ and $d = 1$ at the 1-percent level, and the MLP regression estimates of $d$ vary from 0.539 for $\alpha = 0.70$ to 0.420 for $\alpha = 0.80$. Second, we test whether the explicit IT regime proved marginally more successful than the implicit targeting regime. We do not find that that occurs. We reject the null of $d = 1$ at the 1-percent level, but we cannot reject the null of $d = 0$ for $\alpha = 0.70$, 0.75, and 0.80. The MLP estimates of $d$ vary from 0.170 for $\alpha = 0.70$ to 0.289 for $\alpha = 0.80$. Thus, persistence in the explicit IT regime, where the SARB wanted to keep inflation within the target band of 3 to 6 percent, exceeds its actual level under the assumption that the SARB followed the more eclectic and heterogeneous policy of the previous period. This perverse outcome could occur, *ceteris paribus*, because of the wide targeting band. A lower and narrower target band could improve the credibility of the SARB (Gupta and Ulwilingiye, 2012), causing inflationary expectations to converge to a focal point (Demertzis and Viegi, 2008), resulting in an increases in monetary policy credibility and a decline in inflation persistence.

[16] The Kumar-Okimoto and Hassler-Olivares test results are also generally invariant to detrending. The main exceptions are Australia (the Kumar-Okimoto test results reverse from insignificant to significant), Chile (both test results reverse from insignificant to significant), and South Africa (both test results reverse from significant to insignificant). Without detrending, the difference in the estimates of persistence for the pre- and post-IT periods are much smaller than we previously found, resulting in the failure to reject the null hypothesis of equality.

[17] We thank an anonymous referee for this suggestion.

(Fracasso et al., 2003), which is the start of the first calendar year of the new regime, rather than September 1990 (Truman, 2003), when the Central Bank of Chile first announced an inflation target. For Australia, we define the starting date of IT as September 1994 (Bernanke et al. , 1999), which is when an exact numerical target was first announced, rather than April 1993 (Schaechter et al., 2000) when the Reserve Bank of Australia announce the adoption of the new framework. For New Zealand, we define the starting date of IT as March 1990 (Mishkin and Schmidt-Hebbel, 2001), which marks the publication of the first Policy Targets Agreement between the Minister of Finance and the Governor of the Reserve Bank of New Zealand, rather than April 1988 (Fracasso et al., 2003) when the IT policy was first announced in the New Zealand's Government budget statement. Thus, to accommodate alternative starting dates and allow for the possibility of an initial credibility transition (Yigit, 2010), we check the robustness of our results to 1- and 2-year changes of the starting dates in both directions.

The findings, not displayed for economy of space, generally correspond to those presented in Tables 6 and 7. The decline in persistence still occurs in all the countries, with the exception of South Africa. The estimates of persistence are generally stable, with the exception of Australia and Canada, where estimates prove sensitive to the pre- and post- IT re-sampling. For example, extending the pre-IT sample by adding 1 or 2 years reduces the estimates of $d$ for Australia from 0.905 (for $\alpha$ = 0.70) to 0.645 and 0.523, respectively, and for Canada from 0.960 (for $\alpha$ = 0.70) to 0.453 and 0.304, respectively. Similarly, reducing the pre-IT sample by subtracting 1 or 2 years reduces the estimate for $d$ for Canada to 0.422 and 0.326 (for $\alpha$ = 0.70), respectively. The Kumar-Okimoto and Hassler-Olivares tests for equality reflect these changes. For example, for Australia the Hassler-Olivares test rejects the equality of persistence, as indicated in Table 8. For this country, reducing the pre-IT sample by 1 or 2 years does not affect the test results, but increasing the sample by 1 or 2 years reverses the test outcome (i.e., the test no longer rejects the null of equality). Similarly, for Canada the Kumar-Okimoto test rejects the equality of persistence, as indicated by Table 8. This outcome, however, completely reverses by 1- and 2-year re-sampling of the regime change in both directions.

## 5. Conclusions

This paper contributes to the debate on the effectiveness of IT by analyzing inflation persistence before and after IT adoption in 13 OECD countries using a fractional integration approach.. This relatively new approach in econometrics permits different characterization of the dynamic of inflation. Low levels of *d* characterize weak persistence, while in the unit-root model such persistence does not exist. On the other hand, high levels of *d* exhibit persistence with mean-reversion, whereas in the unit-root model such persistence exists, but without mean reversion. The empirical evidence suggests that the MLP approach to fractional integration successfully models inflation. We find mean reversion and stationarity as well as mean reversion and nonstationarity in the pre-IT period, while mean reversion and stationarity uniformly characterize the post-IT period. Inflation exhibits fractional integration over the entire sample period and the pre-IT period, and in most cases, also in the post-IT period. These variations, in turn, suggest that IT affects inflation persistence. But, do the variations prove significant? If IT successfully decreases the variability and heterogeneity of inflation expectations and increases the credibility of monetary policy, then evidence of persistence after the regime switch should disappear or significantly fall. Instead, we find that the success of IT in reducing inflation persistence is mixed, as we can formally reject the null hypothesis of equality of inflation persistence against the alternative that inflation persistence declines in the post-targeting period for only about half the countries.

That said, the parameter estimates display important differences before and after IT. We observe that most countries exhibit significant inflation persistence in the pre-IT regime. We reject the hypothesis that inflation in the pre-IT period conforms to an I(1) process for all 13 countries and, similarly, reject the hypothesis that it conforms to an I(0) process, with the exception of South Africa, which displays I(0) dynamics. Most countries show confidence interval estimates for *d* that span the stationary mean-reverting and nonstationary but mean-reverting boundaries, with a few countries displaying the upper band of the confidence interval estimates in the nonstationary, non-mean-reverting range. Norway and South Korea are special cases as the lower bands of their 95-percent confidence interval suggest the possibility of antipersistence, which implies a faster rate of mean-reversion than a white-noise process. The evidence of decline of inflation persistence in the post IT period, however, proves mixed.

Our results show that in all countries, the orders of fractional integration significantly differ from one (i.e., the evidence of non-mean-reverting inflation no longer exists in the 13 countries that adopted an IT regime). We find, however, that even after the adoption of IT, inflation persistence still remains to some degree in Australia, Chile, Iceland, Israel, Mexico, New Zealand, South Africa, and Switzerland, while in South Korea and Norway we cannot exclude the possibility of antipersistence.

Among the countries where inflation persistence still remains, we find different rates of mean reversion. Australia, Iceland, Israel, Mexico, New Zealand, and Switzerland exhibit point estimates of *d* in the stationary range, while for Chile and South Africa, the estimates fall in the nonstationary range.

Finally, shocks to inflation in Canada, South Korea, Norway, Sweden, and the UK prove temporary and revert quickly to the mean, while in Australia, Iceland, Israel, Mexico, New

Zealand, and Switzerland, shocks revert to the mean, but over a period of time that depends on the degree of persistence. In Chile and South Africa, on the other hand, the time required to revert to the mean is even longer.

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**Table 1: Descriptive Statistics for Inflation of 13 IT Countries: The Pre- and Post-IT Sub-Samples and the Full Sample**

| Country | Period | Mean | St. Dev. | Skewness | Kurtosis |
|---|---|---|---|---|---|
| **Australia** | 1976QII-1994QII | 7.001 | 3.933 | 0.201 | 4.337 |
| | 1994QIII-2013QII | 2.669 | 2.228 | 2.395 | 15.338 |
| | 1976QII-2013QII | 4.791 | 3.841 | 0.946 | 4.325 |
| **Canada** | 1976M02-1991M01 | 6.611 | 4.074 | 1.072 | 6.427 |
| | 1991M02-2013M06 | 1.805 | 3.281 | -0.280 | 4.578 |
| | 1976M02-2013M06 | 3.731 | 4.316 | 0.602 | 5.507 |
| **Chile** | 1976M02-1990M12 | 27.424 | 24.036 | 2.674 | 11.828 |
| | 1991M01-2013M06 | 5.242 | 5.748 | 0.739 | 4.710 |
| | 1976M02-2013M06 | 14.085 | 19.173 | 3.407 | 19.399 |
| **Iceland** | 1976M02-2001M06 | 19.199 | 22.224 | 1.553 | 5.907 |
| | 2001M07-2013M06 | 5.517 | 5.941 | 1.153 | 6.966 |
| | 1976M02-2013M06 | 14.811 | 19.679 | 2.039 | 8.181 |
| **Israel** | 1976M02-1991M12 | 56.632 | 52.312 | 1.623 | 6.004 |
| | 1992M01-2013M06 | 4.409 | 6.055 | 0.547 | 3.485 |
| | 1976M02-2013M06 | 26.624 | 43.009 | 2.600 | 11.008 |
| **Mexico** | 1976M02-1998M12 | 32.497 | 26.232 | 1.650 | 5.854 |
| | 1999M01-2013M06 | 5.044 | 3.216 | 0.809 | 5.041 |
| | 1976M02-2013M06 | 21.858 | 24.580 | 2.044 | 7.672 |
| **New Zealand** | 1976QII-1989QIV | 11.369 | 5.518 | 0.923 | 6.150 |
| | 1990QI-2013QII | 2.228 | 1.926 | 1.196 | 7.315 |
| | 1976QII-2013QII | 5.601 | 5.747 | 1.463 | 5.717 |
| **Norway** | 1976M02-2001M02 | 5.372 | 4.508 | 0.543 | 0.543 |
| | 2001M03-2013M06 | 1.733 | 4.562 | 1.138 | 9.661 |
| | 1976M02-2013M06 | 4.172 | 4.834 | 0.567 | 6.131 |
| **South Africa** | 1976M02-2000M01 | 11.305 | 7.487 | 0.178 | 4.927 |
| | 2000M02-2013M06 | 5.217 | 4.723 | -0.130 | 4.611 |
| | 1976M02-2013M06 | 9.122 | 7.240 | 0.448 | 4.758 |
| **South Korea** | 1976M02-1997M12 | 7.943 | 8.614 | 1.627 | 6.131 |
| | 1998M01-2013M06 | 2.890 | 3.817 | 0.944 | 6.760 |
| | 1976M02-2013M06 | 5.850 | 7.458 | 2.032 | 8.611 |
| **Sweden** | 1976M02-1992M12 | 7.700 | 6.100 | 1.208 | 6.361 |
| | 1993M01-2013M06 | 1.391 | 3.351 | 1.565 | 24.879 |
| | 1976M02-2013M06 | 4.243 | 5.727 | 1.507 | 8.015 |
| **Switzerland** | 1976M02-1999M12 | 2.577 | 3.250 | 1.066 | 4.944 |
| | 2000M01-2013M06 | 0.635 | 2.854 | -0.219 | 3.563 |
| | 1976M02-2013M06 | 1.876 | 3.247 | 0.711 | 5.029 |
| **UK** | 1976M02-1992M09 | 7.631 | 6.239 | 2.869 | 17.887 |
| | 1992M10-2013M06 | 2.160 | 2.228 | 0.077 | 4.324 |
| | 1976M02-1992M09 | 4.597 | 5.239 | 5.239 | 22.466 |

**Table 2: KPSS and PP Unit-Root Tests of Inflation: Level and Trend Stationarity**

| Country | Level Stationarity | | Trend Stationarity | |
|---|---|---|---|---|
| | **KPSS** | **PP** | **KPSS** | **PP** |
| **Australia** | 2.350* | -5.391* | 0.373* | -8.128* |
| **Canada** | 4.580* | -14.817* | 0.695* | -18.151* |
| **Chile** | 4.160* | -6.123* | 0.487* | -7.241* |
| **Iceland** | 4.900* | -10.979* | 0.937* | -15.342* |
| **Israel** | 4.030* | -4.671* | 0.460* | -6.533* |
| **Mexico** | 3.470* | -3.611* | 0.412* | -4.671* |
| **New Zealand** | 2.380* | -3.884* | 0.413* | -5.982* |
| **Norway** | 4.830* | -13.425* | 0.506* | -16.569* |
| **South Africa** | 4.280* | -16.812* | 0.299* | -19.408* |
| **South Korea** | 2.780* | -11.932* | 0.332* | -13.794* |
| **Switzerland** | 2.420* | -15.700* | 0.170* | -16.862* |
| **Sweden** | 5.210* | -14.823* | 0.432* | -18.764* |
| **UK** | 4.140* | -10.653* | 0.814* | -13.582* |

**Note:** The critical values for the KPSS test with a null hypothesis of stationary series the following: 0.347, 0.463, and 0.739 for the 10-, 5-, and 1-percent significance levels without trend and 0.119, 0.146, and 0.216 for the 10-, 5-, and 1-percent significance levels with trend. The KPSS critical values remain the same for monthly and quarterly data. The critical values for the PP test with a null hypothesis of nonstationary series equal the following for quarterly data: -2.577, -2.887, and -3.494 for the 10-, 5-, and 1-percent significance levels without trend and -3.144, -3.444, and -4.024 for the 10-, 5-, and 1-percent significance levels with trend. The critical values for the PP test with a null hypothesis of nonstationary series equal the following for monthly data: -2.570, -2.872, and -3.444 for the 10-, 5-, and 1-percent significance levels without trend and -3.130, -3.422, and -3.982 for the 10-, 5-, and 1-percent significance levels with trend. * mean significant at the 1-percent level.

**Table 3: MLP Regression Estimates and Tests for Stationarity and Unit Root: The Full Sample**

| Country | $\alpha$ | $\hat{d}$ | se($\hat{d}$) | $H_0$: $d$=0 | $H_0$: $d$=1 | 95% CI |
|---|---|---|---|---|---|---|
| **Australia** | 0.70 | 0.387 | 0.093 | 4.159* | -5.495* | [0.205, 0.569] |
| | 0.75 | 0.582 | 0.106 | 5.509* | -4.221* | [0.374, 0.790] |
| | 0.80 | 0.434 | 0.098 | 4.418* | -6.488* | [0.242, 0.626] |
| **Canada** | 0.70 | 0.493 | 0.104 | 4.725* | -6.664* | [0.289, 0.697] |
| | 0.75 | 0.370 | 0.090 | 4.132* | -9.668* | [0.194, 0.546] |
| | 0.80 | 0.248 | 0.072 | 3.425* | -13.473* | [0.107, 0.389] |
| **Chile** | 0.70 | 0.588 | 0.062 | 9.487* | -5.412* | [0.466, 0.710] |
| | 0.75 | 0.756 | 0.069 | 11.001* | -3.754* | [0.621, 0.891] |
| | 0.80 | 0.703 | 0.060 | 11.670* | -5.328* | [0.585, 0.821] |
| **Iceland** | 0.70 | 0.623 | 0.097 | 6.415* | -4.959* | [0.433, 0.813] |
| | 0.75 | 0.535 | 0.081 | 6.631* | -7.138* | [0.376, 0.694] |
| | 0.80 | 0.386 | 0.070 | 5.495* | -11.000* | [0.249, 0.523] |
| **Israel** | 0.70 | 0.807 | 0.097 | 8.312* | -2.536* | [0.617, 0.997] |
| | 0.75 | 0.674 | 0.083 | 8.086* | -5.008* | [0.511, 0.837] |
| | 0.80 | 0.526 | 0.068 | 7.726* | -8.501* | [0.393, 0.359] |
| **Mexico** | 0.70 | 0.761 | 0.084 | 9.020* | -3.136* | [0.596, 0.926] |
| | 0.75 | 0.754 | 0.074 | 10.137* | -3.771* | [0.609, 0.899] |
| | 0.80 | 0.694 | 0.063 | 10.996* | -5.476* | [0.571, 0.817] |
| **New Zealand** | 0.70 | 0.522 | 0.140 | 3.717* | -4.284* | [0.248, 0.796] |
| | 0.75 | 0.629 | 0.121 | 5.202* | -3.754* | [0.392, 0.866] |
| | 0.80 | 0.573 | 0.107 | 5.367* | -4.888* | [0.363, 0.783] |
| **Norway** | 0.70 | 0.213 | 0.078 | 2.726* | -10.334* | [0.060, 0.366] |
| | 0.75 | 0.295 | 0.068 | 4.322* | -10.820* | [0.162, 0.428] |
| | 0.80 | 0.258 | 0.055 | 4.695* | -13.293* | [0.150, 0.366] |
| **South Africa** | 0.70 | 0.231 | 0.088 | 2.625* | -10.104* | [0.059, 0.403] |
| | 0.75 | 0.316 | 0.079 | 3.986* | -10.511* | [0.161, 0.471] |
| | 0.80 | 0.291 | 0.065 | 4.467* | -12.694* | [0.164, 0.418] |
| **South Korea** | 0.70 | 0.342 | 0.092 | 3.716* | -8.648* | [0.162, 0.522] |
| | 0.75 | 0.329 | 0.075 | 4.370* | -10.304* | [0.182, 0.476] |
| | 0.80 | 0.345 | 0.063 | 5.477* | -11.734* | [0.222, 0.468] |
| **Sweden** | 0.70 | 0.350 | 0.090 | 3.881* | -8.545* | [0.174, 0.526] |
| | 0.75 | 0.248 | 0.072 | 3.431* | -11.552* | [0.107, 0.389] |
| | 0.80 | 0.219 | 0.062 | 3.497* | -14.000* | [0.097, 0.341] |
| **Switzerland** | 0.70 | 0.293 | 0.081 | 3.625* | -9.283* | [0.134, 0.452] |
| | 0.75 | 0.302 | 0.081 | 3.746* | -10.714* | [0.143, 0.461] |
| | 0.80 | 0.239 | 0.064 | 3.713* | -13.642* | [0.114, 0.364] |
| **UK** | 0.70 | 0.451 | 0.101 | 4.461* | -7.209* | [0.253, 0.649] |
| | 0.75 | 0.451 | 0.083 | 5.460* | -8.435* | [0.288, 0.614] |
| | 0.80 | 0.469 | 0.075 | 6.281* | -9.516* | [0.322, 0.616] |

**Note:** $T^{\alpha}$ equals the bandwidth and $\alpha$ equals the power of the bandwidth. $\hat{d}$ equals the estimate of the persistence parameter. $se(\hat{d})$equals the standard error of the estimate of the persistence parameter. The confidence interval (CI) defines the range for the 95-percent confidence level. * means significant at the 1-percent level.

**Table 4: Chow Tests of Structural Change Using Unfiltered (Raw) Data**

| Country | Breakpoint | Pure structural change | Partial structural change |
|---|---|---|---|
| **Australia** | 1994QIII | 57.147* | 14.171* |
| **Canada** | 1991M02 | 159.106* | 90.629* |
| **Chile** | 1991M01 | 174.224* | 13.582* |
| **Iceland** | 2001M07 | 50.078* | 13.747* |
| **Israel** | 1992M01 | 201.221* | 17.673* |
| **Mexico** | 1999M01 | 158.073* | 8.596* |
| **New Zealand** | 1990QI | 133.983* | 26.111* |
| **Norway** | 2001M03 | 60.182* | 24.114* |
| **South Africa** | 2000M02 | 79.870* | 50.046* |
| **South Korea** | 1998M01 | 53.149* | 18.734* |
| **Sweden** | 1993M01 | 160.909* | 90.588* |
| **Switzerland** | 2001M01 | 38.768* | 20.530* |
| **UK** | 1992M10 | 141.281* | 40.893* |

**Note:** The 10-, 5-, and 1-percent significance levels for the $\chi^2(1)$ distribution occur for 2.706, 3.841, and 6.635, respectively. * mean significant at the 1-percent level.

**Table 5: Chow Tests of Structural Change Using Filtered (Fractionally Differenced) Data**

| Country | Breakpoint | $\alpha$ | Pure Structural Change | Partial Structural Change |
|---|---|---|---|---|
| **Australia** | 1994QIII | 0.70 | 2.218 | 2.070 |
| | | 0.75 | 0.078 | 0.060 |
| | | 0.80 | 1.201 | 0.213 |
| **Canada** | 1991M02 | 0.70 | 3.041ǂ | 5.384** |
| | | 0.75 | 9.180* | 14.254* |
| | | 0.80 | 26.631* | 33.04* |
| **Chile** | 1991M01 | 0.70 | 0.019 | 0.015 |
| | | 0.75 | 0.222 | 0.301 |
| | | 0.80 | 0.133 | 0.177 |
| **Iceland** | 2001M07 | 0.70 | 0.079 | 0.091 |
| | | 0.75 | 0.246 | 0.318 |
| | | 0.80 | 1.652 | 1.832 |
| **Israel** | 1992M01 | 0.70 | 0.047 | 0.036 |
| | | 0.75 | 0.462 | 0.455 |
| | | 0.80 | 3.463ǂ | 2.868ǂ |
| **Mexico** | 1999M01 | 0.70 | 0.388 | 0.368 |
| | | 0.75 | 0.422 | 0.398 |
| | | 0.80 | 0.861 | 1.601 |
| **New Zealand** | 1990QI | 0.70 | 2.727ǂ | 2.312 |
| | | 0.75 | 0.694 | 0.676 |
| | | 0.80 | 1.485 | 1.357 |
| **Norway** | 2001M03 | 0.70 | 10.212* | 9.074* |
| | | 0.75 | 4.459** | 4.842** |
| | | 0.80 | 6.548* | 6.537** |
| **South Africa** | 2000M02 | 0.70 | 14.053* | 17.294* |
| | | 0.75 | 5.872** | 8.961* |
| | | 0.80 | 7.673* | 11.035* |
| **South Korea** | 1998M01 | 0.70 | 5.039** | 4.676** |
| | | 0.75 | 5.603** | 5.084** |
| | | 0.80 | 4.917** | 4.585** |
| **Sweden** | 1993M01 | 0.70 | 7.352* | 10.706* |
| | | 0.75 | 20.838* | 25.126* |
| | | 0.80 | 27.443* | 30.919* |
| **Switzerland** | 2001M01 | 0.70 | 3.834ǂ | 4.813** |
| | | 0.75 | 3.514ǂ | 4.490** |
| | | 0.80 | 6.380** | 7.138* |
| **UK** | 1992M10 | 0.70 | 0.781 | 1.391 |
| | | 0.75 | 0.781 | 1.391 |
| | | 0.80 | 0.534 | 1.029 |

**Note:** The 10-, 5-, and 1-percent significance levels for the $\chi^2(1)$ distribution occur for 2.706, 3.841, and 6.635, respectively. ǂ, **, and * mean significant at the 10-, 5-, and 1-percent levels.

**Table 6: MLP Regression Estimates and Tests for Stationarity and Unit Root: The Pre-IT Sub-Sample**

| Country | $\alpha$ | $\hat{d}$ | se($\hat{d}$) | $H_0$: $d$=0 | $H_0$: $d$=1 | 95% CI |
|---|---|---|---|---|---|---|
| **Australia** | 0.70 | 0.905 | 0.300 | 3.014* | -0.665 | [0.317, 1.493] |
| | 0.75 | 0.708 | 0.271 | 2.612* | -2.233** | [0.177, 1.239] |
| | 0.80 | 0.617 | 0.242 | 2.547** | -3.267* | [0.143, 1.091] |
| **Canada** | 0.70 | 0.960 | 0.160 | 6.003* | -0.381 | [0.646, 1.274] |
| | 0.75 | 0.746 | 0.140 | 5.324* | -2.778* | [0.472, 1.020] |
| | 0.80 | 0.573 | 0.125 | 4.580* | -5.283* | [0.328, 0.818] |
| **Chile** | 0.70 | 0.722 | 0.133 | 5.414* | -2.639* | [0.461, 0.983] |
| | 0.75 | 0.610 | 0.112 | 5.442* | -4.217* | [0.390, 0.830] |
| | 0.80 | 0.596 | 0.093 | 6.424* | -5.002* | [0.414, 0.778] |
| **Iceland** | 0.70 | 0.661 | 0.127 | 5.214* | -3.885* | [0.412, 0.910] |
| | 0.75 | 0.471 | 0.105 | 4.475* | -7.000* | [0.265, 0.677] |
| | 0.80 | 0.383 | 0.096 | 3.977* | -9.475* | [0.195, 0.571] |
| **Israel** | 0.70 | 0.768 | 0.143 | 5.375* | -2.259** | [0.488, 1.048] |
| | 0.75 | 0.540 | 0.124 | 4.350* | -5.125* | [0.297, 0.783] |
| | 0.80 | 0.538 | 0.100 | 5.381* | -5.850* | [0.342, 0.734] |
| **Mexico** | 0.70 | 0.886 | 0.108 | 8.186* | -1.258 | [0.674, 1.098] |
| | 0.75 | 0.799 | 0.093 | 8.626* | -2.567** | [0.617, 0.981] |
| | 0.80 | 0.776 | 0.081 | 9.568* | -3.292* | [0.617, 0.935] |
| **New Zealand** | 0.70 | 0.569 | 0.187 | 3.041* | -2.687* | [0.202, 0.936] |
| | 0.75 | 0.492 | 0.157 | 3.136** | -3.544* | [0.184, 0.800] |
| | 0.80 | 0.498 | 0.132 | 3.778* | -3.834* | [0.239, 0.757] |
| **Norway** | 0.70 | 0.354 | 0.105 | 3.373* | -7.400* | [0.148, 0.560] |
| | 0.75 | 0.400 | 0.090 | 4.424* | -7.938* | [0.224, 0.576] |
| | 0.80 | 0.308 | 0.077 | 4.023* | -10.568* | [0.157, 0.459] |
| **South Africa** | 0.70 | 0.158 | 0.124 | 1.279 | -9.467* | [-0.085, 0.401] |
| | 0.75 | 0.127 | 0.098 | 1.301 | -11.308* | [-0.065, 0.319] |
| | 0.80 | 0.149 | 0.088 | 1.703ǂ | -12.728* | [-0.023, 0.321] |
| **South Korea** | 0.70 | 0.602 | 0.096 | 6.286* | -4.339* | [0.414, 0.790] |
| | 0.75 | 0.538 | 0.091 | 5.926* | -5.815* | [0.360, 0.716] |
| | 0.80 | 0.433 | 0.080 | 5.443* | -8.194* | [0.276, 0.590] |
| **Sweden** | 0.70 | 0.249 | 0.121 | 2.064** | -7.498* | [0.012, 0.486] |
| | 0.75 | 0.279 | 0.114 | 2.461** | -8.181* | [0.056, 0.502] |
| | 0.80 | 0.245 | 0.095 | 2.582* | -9.856* | [0.059, 0.431] |
| **Switzerland** | 0.70 | 0.436 | 0.113 | 3.846* | -6.346* | [0.215, 0.657] |
| | 0.75 | 0.303 | 0.095 | 3.184* | -9.022* | [0.117, 0.489] |
| | 0.80 | 0.239 | 0.077 | 3.109* | -11.389* | [0.088, 0.390] |
| **UK** | 0.70 | 0.399 | 0.106 | 3.769* | -5.926* | [0.191, 0.607] |
| | 0.75 | 0.480 | 0.101 | 4.747* | -5.900* | [0.282, 0.678] |
| | 0.80 | 0.374 | 0.088 | 4.257* | -8.105* | [0.202, 0.546] |

Note: See Table 3. ǂ, **, and * means significant at the 10-, 5-, and 1-percent levels.

**Table 7: MLP Regression Estimates and Tests for Stationarity and Unit Root: The Post-IT Sub-Sample**

| Country | $\alpha$ | $\hat{d}$ | se($\hat{d}$) | $H_0$: $d$=0 | $H_0$: $d$=1 | 95% CI |
|---|---|---|---|---|---|---|
| **Australia** | 0.70 | 0.429 | 0.129 | 3.324* | -3.981* | [0.176, 0.682] |
| | 0.75 | 0.421 | 0.138 | 3.053* | -4.516* | [0.151, 0.691] |
| | 0.80 | 0.343 | 0.118 | 2.897* | -5.702* | [0.112, 0.574] |
| **Canada** | 0.70 | 0.126 | 0.115 | 1.100 | -9.636* | [-0.099, 0.351] |
| | 0.75 | 0.091 | 0.099 | 0.921 | -11.515* | [-0.103, 0.285] |
| | 0.80 | 0.081 | 0.088 | 0.917 | -13.367* | [-0.091, 0.253] |
| **Chile** | 0.70 | 0.600 | 0.116 | 5.173* | -4.407* | [0.373, 0.827] |
| | 0.75 | 0.591 | 0.108 | 5.483* | -5.184* | [0.379, 0.803] |
| | 0.80 | 0.558 | 0.090 | 6.169* | -6.471* | [0.382, 0.734] |
| **Iceland** | 0.70 | 0.382 | 0.108 | 3.520* | -5.454* | [0.170, 0.594] |
| | 0.75 | 0.521 | 0.106 | 4.916* | -4.782* | [0.313, 0.729] |
| | 0.80 | 0.461 | 0.093 | 4.978* | -6.114* | [0.279, 0.643] |
| **Israel** | 0.70 | 0.401 | 0.113 | 3.552* | -6.472* | [0.180, 0.622] |
| | 0.75 | 0.332 | 0.093 | 3.583* | -8.335* | [0.150, 0.514] |
| | 0.80 | 0.374 | 0.082 | 4.568* | -8.951* | [0.213, 0.535] |
| **Mexico** | 0.70 | 0.268 | 0.119 | 2.245** | -6.946* | [0.035, 0.501] |
| | 0.75 | 0.325 | 0.110 | 2.948* | -7.214* | [0.109, 0.541] |
| | 0.80 | 0.335 | 0.089 | 3.756* | -8.168* | [0.161, 0.509] |
| **New Zealand** | 0.70 | 0.296 | 0.159 | 1.861ǂ | -5.377* | [-0.016, 0.60] |
| | 0.75 | 0.311 | 0.136 | 2.290** | -5.889* | [0.044, 0.578] |
| | 0.80 | 0.356 | 0.130 | 2.732* | -6.110* | [0.101, 0.611] |
| **Norway** | 0.70 | -0.184 | 0.125 | -1.477 | -10.610* | [-0.429, 0.061] |
| | 0.75 | -0.102 | 0.109 | -0.944 | -11.141* | [-0.316, 0.112] |
| | 0.80 | 0.008 | 0.107 | 0.071 | -11.372* | [-0.202, 0.218] |
| **South Africa** | 0.70 | 0.507 | 0.133 | 3.807* | -4.550* | [0.246, 0.768] |
| | 0.75 | 0.551 | 0.115 | 4.803* | -4.696* | [0.326, 0.776] |
| | 0.80 | 0.541 | 0.110 | 4.923* | -5.455* | [0.325, 0.757] |
| **South Korea** | 0.70 | 0.084 | 0.155 | 0.545 | -8.802* | [-0.220, 0.388] |
| | 0.75 | 0.074 | 0.121 | 0.613 | -10.210* | [-0.163, 0.311] |
| | 0.80 | 0.050 | 0.103 | 0.483 | -11.945* | [-0.152, 0.252] |
| **Sweden** | 0.70 | 0.217 | 0.081 | 2.695* | -8.368* | [0.058, 0.376] |
| | 0.75 | 0.248 | 0.072 | 3.455* | -9.239* | [0.107, 0.389] |
| | 0.80 | 0.223 | 0.065 | 3.452* | -10.905* | [0.096, 0.350] |
| **Switzerland** | 0.70 | 0.398 | 0.148 | 2.698* | -5.552* | [0.108, 0.688] |
| | 0.75 | 0.383 | 0.129 | 2.970* | -6.457* | [0.130, 0.636] |
| | 0.80 | 0.257 | 0.107 | 2.405** | -8.825* | [0.047, 0.467] |
| **UK** | 0.70 | 0.155 | 0.116 | 1.334 | -9.036* | [-0.072, 0.382] |
| | 0.75 | 0.176 | 0.101 | 1.734ǂ | -10.121* | [-0.022, 0.374] |
| | 0.80 | 0.113 | 0.082 | 1.373 | -12.527* | [-0.048, 0.274] |

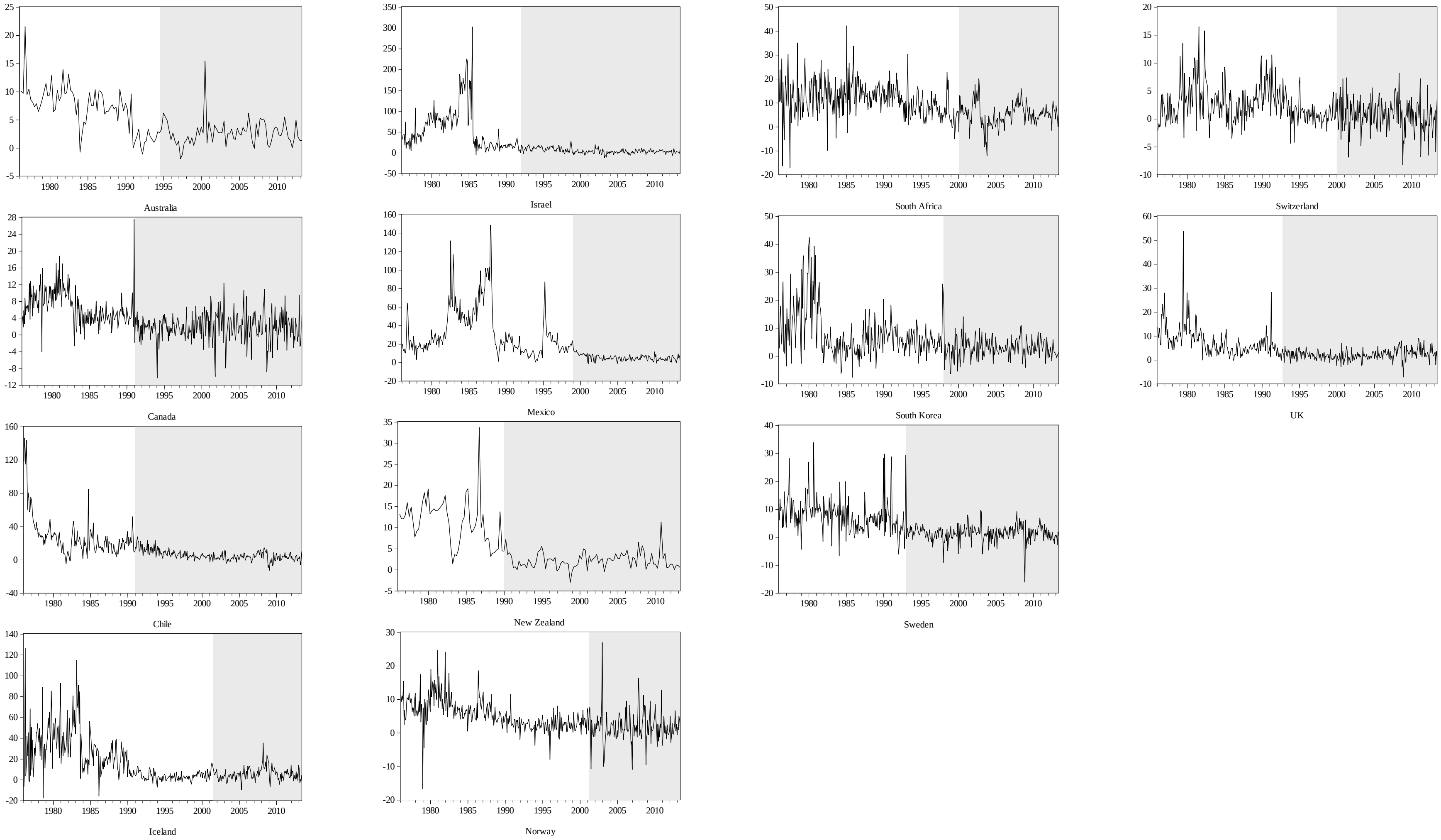
Note: See Table 3. ǂ, **, and * means significant at the 10-, 5-, and 1-percent levels.

**Table 8: Kumar-Okimoto and Hassler-Olivares Tests of Equality of *d* Pre- and Post-IT: $H_0 : d_1 - d_2 = 0$ vs. $H_1 : d_1 - d_2 > 0$.**

| **Country** | $\alpha$ | $\hat{d}_1 - \hat{d}_2$ | **Kumar-Okimoto Test** | **Hassler-Olivares Test** |
|---|---|---|---|---|
| **Australia** | 0.70 | 0.476 | 1.458 | 3.010* |
| | 0.75 | 0.287 | 0.944 | 2.009** |
| | 0.80 | 0.274 | 1.018 | 2.140** |
| **Canada** | 0.70 | 0.834 | 4.233* | 7.692* |
| | 0.75 | 0.655 | 3.820* | 6.947* |
| | 0.80 | 0.492 | 3.218* | 5.948* |
| **Chile** | 0.70 | 0.122 | 0.691 | 1.125 |
| | 0.75 | 0.019 | 0.122 | 0.200 |
| | 0.80 | 0.038 | 0.294 | 0.461 |
| **Iceland** | 0.70 | 0.279 | 1.674ǂ | 2.501** |
| | 0.75 | -0.050 | -0.335 | -0.511 |
| | 0.80 | -0.078 | -0.584 | -0.913 |
| **Israel** | 0.70 | 0.367 | 2.014** | 3.405* |
| | 0.75 | 0.208 | 1.342 | 2.216** |
| | 0.80 | 0.164 | 1.268 | 1.994** |
| **Mexico** | 0.70 | 0.618 | 3.846* | 5.700* |
| | 0.75 | 0.474 | 3.291* | 4.982* |
| | 0.80 | 0.441 | 3.665* | 5.332* |
| **New Zealand** | 0.70 | 0.273 | 1.112 | 1.692ǂ |
| | 0.75 | 0.181 | 0.871 | 1.254 |
| | 0.80 | 0.142 | 0.766 | 1.084 |
| **Norway** | 0.70 | 0.538 | 3.296* | 4.870* |
| | 0.75 | 0.502 | 3.551* | 5.171* |
| | 0.80 | 0.300 | 2.276** | 3.527* |
| **South Africa** | 0.70 | -0.349 | -1.919ǂ | -3.193* |
| | 0.75 | -0.424 | -2.806* | -4.426* |
| | 0.80 | -0.392 | -2.783* | -4.676* |
| **South Korea** | 0.70 | 0.518 | 2.841* | 4.793* |
| | 0.75 | 0.464 | 3.065* | 4.933* |
| | 0.80 | 0.383 | 2.937* | 4.661* |
| **Sweden** | 0.70 | 0.032 | 0.220 | 0.299 |
| | 0.75 | 0.031 | 0.230 | 0.331 |
| | 0.80 | 0.022 | 0.191 | 0.270 |
| **Switzerland** | 0.70 | 0.038 | 0.204 | 0.348 |
| | 0.75 | -0.080 | -0.499 | -0.835 |
| | 0.80 | -0.018 | -0.137 | -0.215 |
| **UK** | 0.70 | 0.244 | 1.553 | 2.269** |
| | 0.75 | 0.304 | 2.128** | 3.250* |
| | 0.80 | 0.261 | 2.170** | 3.195* |

Note: See Table 3. ǂ, **, and * means significant at the 10-, 5-, and 1-percent levels.

## Figure 1: Monthly Inflation Rates in 13 OECD IT Countries




**Note:** The shaded area denotes the IT regime.