University of Connecticut

Department of Economics Working Paper Series

Inflation Persistence and Structural Breaks:
The Experience of Inflation Targeting Countries and the US

by

Giorgio Canarella
University of Nevada, Las Vegas

Stephen M. Miller
University of Nevada, Las Vegas

Working Paper 2016-11
September 2016

365 Fairfield Way, Unit 1063
Storrs, CT 06269-1063
Phone: (860) 486-3022
Fax: (860) 486-4463
http://www.econ.uconn.edu/

Inflation Persistence and Structural Breaks: The Experience of Inflation Targeting Countries and the US

Giorgio Canarella[a]
giorgio.canarella@unlv.edu

and

Stephen M. Miller[a,*]
stephen.miller@unlv.edu

Abstract: This paper reports on a sequential three-stage analysis of inflation persistence using monthly data from 11 inflation targeting (IT) countries and, for comparison, the US, a non IT country with a history of credible monetary policy. First, we estimate inflation persistence in a rolling-window fractional integration setting using the semiparametric estimator suggested by Phillips (2007). Second, we use tests for unknown structural breaks as a means to identify effects of the regime switch and the global financial crisis on inflation persistence. We use the sequences of estimated persistence measures from the first stage as dependent variables in the Bai and Perron (2003) structural break tests. These results suggest that four countries (Canada, Iceland, Mexico, and South Korea) experience a structural break in inflation persistence that coincide with the implementation of the IT regime, and three IT countries (Sweden, Switzerland, and the UK), as well as the US experience a structural break in inflation persistence that coincides with the global financial crisis. Finally, we reapply the Phillips (2007) estimator to the subsamples defined by the breaks. We find that in most cases the estimates of inflation persistence switch from mean-reversion nonstationarity to mean-reversion stationarity.

Keywords: inflation persistence; inflation targeting; fractional integration; rolling window estimation; structural breaks

JEL classification: C14; E31; C22

a. Department of Economics; University of Nevada, Las Vegas; Las Vegas, Nevada 89154-6005

* Corresponding author

1. Introduction

One of the most important characteristics of the dynamics of inflation is its persistence. Policy makers view inflation persistence with great interest because of its central role in determining how monetary policy respomds to shocks over time (see, e.g., Gerlach and Tillmann, 2012; van der Cruijsen, et al., 2010). Persistence measures the speed at which shocks to inflation die out and the inflation rate returns to its target or its mean. With high inflation persistence, shocks to inflation exert long-lasting effects and may require a strong policy response to bring inflation under control. In the worst case, inflation may contain a unit root, which could lead to unpalatable choices for monetary policy, since inflation randomly drifts away in any direction, making it difficult for the monetary authorities to control inflation. In the best case, inflation may prove stationary, implying that it reverts to its initial level rapidly after a shock occurs. One area of the literature on inflation persistence considers whether shocks to the inflation process decay rapidly, as with non-integrated processes, or whether they decay much more slowly, as with fractionally integrated processes. In the latter case, inflation rates display evidence of persistence or long memory.[1]

New-Keynesian dynamic stochastic general equilibrium (DSGE) macroeconomic models that incorporate lags of inflation in the new Keynesian Phillips curve (NKPC)[2] identify inflationary expectations as the main determinant of inflation persistence, suggesting that the decline in inflation persistence may occur through enhanced anchoring of inflation expectations

---

[1] The empirical literature employs both autoregressive and fractionally integrated measures of inflation persistence. Examples of the autoregressive approach to estimate inflation persistence include Pivetta and Reis (2007), Gamber, Liebner and Smith (2012), Stock (2001), O'Reilly and Whelan (2005), Levin and Piger (2004), and Gerlach and Tillman (2012). Examples of fractional integration analysis include Hassler and Wolters (1995), Siklos (1999), Barkoulas, et al., (1999), Bos, et al., (1999, 2002), Kuttner and Posen (2001), Baillie, et al., (2002), Arize, et al., (2005), Levin, et al., (2004), Perusson (2004), Gadea and Mayoral (2005), Kumar and Okimoto (2007), Beechey and Österholm (2009), Yigit (2010), Meller and Nautz (2012), and Canarella and Miller (forthcoming).

[2] See Fuhrer and Moore (1995), Gali and Gertler (1999) and Christiano, et al., (2005) for some theoretical models that justify the inclusion of lags of inflation in the new Keynesian Phillips curves.

(Mishkin, 2007; Nautz and Strohsal, 2015). That is, a monetary policy that successfully anchors inflationary expectations reduces or eliminates inflation persistence, since well-anchored inflationary expectations are much less dependent on past inflation. That is, if the central bank announces that it will keep inflation at 2% over the medium to long run, and economic agents believe it, then inflation expectations should equal exactly 2%. [3] Unlike monetary targeting, the IT framework does not depend on the stability of the demand for money and, unlike foreign exchange targeting, does not require changes in interest rates, direct foreign exchange intervention, and the loss of independent monetary policy (Mishkin, 1998). The linkage between monetary policy and the stability of the demand for money, however, remains after the switch to IT. Thus, the debate about the stability of the money demand remains relevant even within the IT regime (Valadkhani, 2005).

Empirically verifying the success of IT in anchoring inflation expectations proves difficult, however, as data on inflation expectations appear with limited availability and relatively low frequency (see Levin and Piger, 2004). As an alternative, we can test the success of inflation targeting in anchoring inflation expectations by examining the time series properties of inflation (Yigit, 2010).

This paper reports the estimates of inflation persistence under structural breaks using monthly data from 11 IT countries and, for comparison, the US, a non IT country with a history of credible monetary policy. Our strategy unfolds through a sequential three-stage strategy. In

[3] Significant evidence exists that documents the fall of inflation persistence in the last twenty years in developed IT countries (Walsh, 2008; Kuttner and Posen, 2001; Benati, 2008). Benati (2008) estimates the parameters of a sticky-price DSGE model for the United Kingdom and Canada before and after the introduction of IT and finds that inflation persistence falls significantly during the respective IT periods. In other words, implementation of an explicit and credible IT is associated with lower persistence. Evidence from Canada, Sweden, and the UK supports the view that IT anchors expectations (Gurkaynak, et al., 2006, 2007). Evidence that inflation persistence declines in developing countries that formally adopt IT is scant and more elusive (Kuttner and Posen, 2001; Gurkaynak, et al., 2006, 2007).

the first stage, we estimate inflation persistence in a rolling-window fractional integration setting. The fractional integration approach possesses the advantage of not invoking the I(1) paradigm to represent nonstationary series. While the either $I(0)$ or $I(1)$ paradigm of inflation separates stationarity from nonstationarity, the fractional integration framework separates long memory and nonstationarity properties of inflation from its mean-reversion properties.[4] Conventional integer-order unit-root tests, whether their null is $I(1)$ or $I(0)$, may not detect the mean-reverting properties of the inflation rate series, and the common findings of unit-root behavior in inflation rate series may be an artefact of the testing methodology, which is known to have low power against fractional alternatives. We employ the semiparametric estimator suggested by Phillips (2007). The estimator is a modified log periodogram estimator (MLP) that extends the Geweke and Porter-Hudak (1983) estimator to the unit-root case.

As mentioned above, a successful IT policy can anchor expectations about the future of inflation. Within this context, we examine two hypotheses regarding the dynamics of inflation expectations. First, did the switching to the IT regime improve the formation of inflation expectations and therefore reduce or eliminate inflation persistence? Second, did inflation expectations become more loosely anchored in the recent period of financial turbulence and, therefore, increase inflation persistence? Financial crises create natural experiments that test the ability of the IT regime to maintain the anchoring of inflation expectations. A major financial crisis, such as that of 2008–09, provides a natural test of this anchoring, and an increase in inflation persistence during that crisis suggests a failure of the IT policy to anchor inflation expectations during the period of crisis.

[4] Ample evidence exists suggesting that inflation rates are mean-reverting, fractional-integration processes and do not exhibit unit-root behavior. Hassler and Wolters (1995), Baillie, et al., (2002), Baum, et al., (1999), Arize, et al., (2005) provide international evidence; Bos et al. (1999) provides G7 evidence; Kumar and Okimoto (2007) and Bos, et al., (2002) provide US evidence; Gadea and Mayoral (2005) provide OECD evidence; Meller and Nautz (2012) provide Euro-area evidence; and Canarella and Miller (forthcoming) provide IT countries evidence.

In the second stage, we use the sequences of estimated persistence measures from the first stage as dependent variables in the Bai and Perron (2003) structural break tests and test for unknown structural breaks. We adopt a statistical approach similar to the one employed in Piehl, et al., (2003), who suggest that we can use tests for unknown structural breaks to evaluate policy. We extend the approach of Piehl, et al., (2003), arguing that the evaluation of a change in regimes or a change in macroeconomic events of large proportions should depend on whether the change generates a structural change. This approach involves identifying any unknown structural break in the time series of inflation persistence. In this context, we can interpret a finding of a structural break that coincides with IT implementation or the global financial crisis as evidence supporting an effect of the regime switch or the global financial crisis on inflation persistence. On the other hand, a finding of no structural break or structural breaks at alternative dates would constitute evidence against any effect of these events.

Finally, in the third stage, we reapply the Phillips (2007) estimator to examine whether the regime switch resulting from the breaks and in particular the global financial crisis exert an effect on inflation persistence by reducing or increasing it.[5]

We summarize our findings as follows. First, the rolling-window estimates confirm that most inflation rate series exhibit a drop in persistence after the adoption of IT. This process,

---

[5] An alternative would be to estimate the structural breaks jointly with the persistence parameters. Several approaches have been proposed in the econometric literature. The problem of testing for a single change of the fractional parameter is discussed by Yamaguchi (2011) in a parametric (ARFIMA) setting. Typically, the sample is partitioned into two subsamples and the fractional integration parameter is estimated for each subsample. The test statistic is obtained by maximizing the difference of these estimates over all such partitions. A similar approach for detecting multiple breaks in the fractional integration parameter is used by Shimotsu (2006) in a more general semiparametric context. A related, though different, approach, based on the CUSUM statistic, is used by Sibbertsen and Kruse (2009) to test for a change in the fractional integration parameter from stationarity to nonstationarity, or viceversa. Still other approaches are discussed by Hassler and Meller (2014), Hassler, Rodrigues and Rubia (2014), Hassler and Scheithauer (2011), and Martins and Rodrigues (2010). The advantage of the joint estimation is its asymptotic efficiency. However, even though joint estimation is asymptotically efficient, this is not necessarily the case in small samples. Therefore, due to the advantage of conceptual and computational simplicity, we confine ourselves to the sequential estimation approach.

nevertheless, ends with the global financial crisis of 2008-2009, which reverses the decline in persistence in most IT countries. Recent evidence (Galati, et al., 2011; Autrup and Grothe, 2014) suggests that US and UK inflation expectations become less anchored since the outbreak of the financial crisis. This is consistent with our findings. We find that inflation persistence, a measure of anchoring, in several of our IT countries increased during the global financial crisis, implying that inflation expectations became de-anchored during the crisis. The length of the de-anchoring varies among our IT countries, but it generally appears as a temporary phenomenon, which is followed by a re-anchoring with further declines in inflation persistence.

Second, we find that all inflation persistence processes undergo at least one structural break, but only four countries (Canada, Iceland, Mexico, and South Korea) experience a structural break that coincides with the implementation of IT. Of the remaining countries, we find no statistical discontinuity in the inflation persistence series, which signals that the IT policy did not exert an immediate effect on inflation persistence. Furthermore, only four countries, the US and three IT countries (Sweden, Switzerland, and the UK) experience a structural break in connection with the global financial crisis. The global financial crisis exposes the vulnerability of the IT regime in developed countries, and suggests that these three IT regimes as well as the US saw a de-anchoring of inflation expectations during the global financial crisis.

The remainder of the paper is organized as follows. Section 2 describes the data. Section 3 briefly discusses the concepts of a fractionally integrated process and long memory and outlines the methodology employed in this paper. Section 4 reports the main empirical findings. Finally, Section 5 concludes.

2. Data

We use monthly observations on the seasonally unadjusted Consumer Price Index (CPI all item;

2010 = 100) retrieved from the Main Economic Indicators (MEI) database (OECD.StatExtracts) for the following 11 OECD countries that adopted IT on, or before, 2001: Canada, Chile, Iceland, Israel, South Africa, South Korea, Mexico, Norway, Sweden, Switzerland, and United Kingdom. In addition, we use the US, an OECD non IT country, for comparison purposes. The data span the period from 1976:1 to 2013:6 for 450 observations. We seasonally adjust the CPI data using the X12 method of the US Census Bureau and measure inflation for each country as 1200 times the logarithmic first difference of the seasonally adjusted CPI series. We follow Bernanke, et al. (1999), Mishkin and Schmidt-Hebbel (2001), and Fracasso, et al., (2003) to recognize the start date of each IT regime. Of the 11 countries, six are industrial economies: Canada (1991:2), Iceland (2001:7), Norway (2001:3), Sweden (1993:1), Switzerland (2000:1), and the United Kingdom (1992:10); and five are emerging market economies: Chile (1991:1), Israel (1992:1), South Korea (1998:1), Mexico (1999:1), and South Africa (2000:2).[6] The date of the IT adoption dates appears in parentheses. Three of the emerging countries, Chile, Mexico, and Israel, experienced episodes of hyperinflation in last thirty years.[7]

Figure 1 plots the monthly CPI inflation series, from February 1976 to June 2013, with the date of IT implementation clearly marked by a vertical gridline. The series exhibit a great deal of variation. For the US, we clearly see the negative spike corresponding to the global financial crisis. For all 11 IT countries, the level of inflation clearly falls after the adoption of IT.

3. Methodology

[6] We do not consider Australia and New Zealand because these two countries do not report monthly CPI series. We also exclude Finland and Spain, which adopted IT in 1993 and 1994, respectively, but abandoned them on entering the EMU in 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.

[7] The monthly CPI inflation rate in Mexico reached an all-time high of 142.10% in February 1988; in Chile the highest monthly CPI inflation rate reached 143.44 in June 1976, while in Israel the highest inflation rate reached 301.67% in July 1985.

3.1 *Fractionally Integrated Processes and Measures of Persistence*

The fractionally integrated process y_t is defined as follows:

$$(1-L)^d(y_t-\mu)=\varepsilon_t\,, \tag{1}$$

where the fractional integration parameter d is any real number, L is the lag operator, y_t is the time series, such as inflation, μ is the unconditional mean of y_t, and ε_t is a stationary error with zero mean and constant variance. The operator $(1-L)^d$ is the fractional filter defined by means of the gamma function Γ

$$(1-L)^d = \sum_{k=0}^{\infty}\frac{\Gamma(k-d)L^k}{\Gamma(-d)\Gamma(k+1)}\,. \tag{2}$$

The fractional order of integration d plays a central role in the definition of fractionally integrated processes.[8] The long-run dynamics of an I(d) process are governed by the parameter d, which is our measure of persistence and, thus, determines the effect of shocks on the y_t process and governs the long-run dynamics of y_t. The macroeconomic literature stresses the cases of d = 0 and d = 1, but d can take on any real number.

Baillie et al (1996) detail the characteristics of an I(d) process. The process is stationary as long as $d<0.5$, but displays long memory for $0<d<1$. Long memory implies a form of serial dependence and persistence that cannot be captured by traditional ARMA processes. Long memory means that a significant dependence exists between observations widely separated in time and, therefore, the effects caused by shocks decay hyperbolically (shocks produce long-

[8] Researchers have used fractionally integrated models to study interest rates (Baum and Barkoulas, 2006; Dueker and Startz, 1998), aggregate output (Diebold and Rudebusch, 1989; Sowell, 1992), inflation (Baillie, et al., 1996), unemployment (Diebold and Rudebusch, 1989), monetary aggregates (Barkoulas, et al., 1999), and exchange rates (Cheung and Lai, 1993; Cheung, 1993; Baillie and Bollerslev, 1994; Diebold, et al., 1991). Long memory models have also been used to model the dynamics of the series in other fields, like hydrology and geophysics.

lasting effects and shocks arbitrarily far away in time still exhibit some influence on the dynamics of the process). For $0.5 \leq d < 1$ the process is nonstationary (even if the fractional parameter is significantly lower than 1) but possesses an infinite variance. As long as *d*<1, however, the process reverts to its mean, which means that any shock that affects the process at some point in time exerts a non-permanent effect on future path of the series. If *d*=0, the unique source of dynamics stems from the stationary dynamics of ε_t, and the process then exhibits no long memory. That is, the process is stationary with short memory or *I*(0). Any shock that affects the series only exerts a short-term effect, which completely disappears in the long run. This means that the correlation between consecutive observations fades out quickly and the series returns rapidly to its constant mean. In such case, we model the series as an ARMA(*p*, *q*). If *d*=1, on the other hand, the series exhibits a unit-root process, which we model as an ARIMA(*p*, 1, *q*) by differencing the unit-root process. Obviously, when $0 < d < 1$, conventional unit-root testing (*d=1* under the null) or stationarity testing (*d=0* under the null) may fail to detect mean reversion in the series and reach the wrong conclusion that the series experiences infinite memory (i.e., a unit root).

3.2 *The Modified Log Periodogram Regression Estimator*

Geweke and Porter-Hudak (1983) propose a semiparametric procedure to estimate the parameter *d* from *T* observation on a fractionally integrated process y_t in model (1). The estimator of *d*, denoted by $\hat{d}_{GPH}$, comes from regressing the logarithm of the periodogram of the time series on a constant and the logarithm of the Fourier frequencies in the neighborhood of the frequency zero. More specifically, define $I(\lambda_j) = \left|\xi(\lambda_j)\right|^2$ as the periodogram of the time series y_t at frequency λ_j, where $\xi(\lambda_j)$ is the discrete Fourier transform of y_t. Then the GPH estimator equals the negative of the OLS estimate of β_1 in the linear regression:

$$\ln\left(I(\lambda_j)\right) = \beta_0 + \beta_1 \ln\left|1 - e^{i\lambda_j}\right| + \nu_j \,, \tag{3}$$

where $I(\lambda_j)$ is the sample spectral density (periodogram) evaluated at the frequencies $\lambda_j = \frac{2\pi j}{T}$, $j = 1,...,m < T$, T is the sample size, and m is the bandwidth (i.e., the number of Fourier frequencies included in the log periodogram regression). Under Gaussian assumptions and in the stationary case, where $-0.5 < d < 0.5$, Robinson (1995) proves that $\hat{d}_{GPH}$ is consistent and asymptotically normally distributed as $\sqrt{m}\left(\hat{d}_{GPH} - d\right) \xrightarrow{d} N\left(0, \frac{\pi^2}{24}\right)$.

Phillips (2007) notes that the theory of statistical inference for the log periodogram regression only exists for the stationary 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. Phillips (2007) propose a method to estimate d using a modified log periodogram regression (MLP) estimator that accommodates the nonstationary range $d \geq 0.5$. 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 $I_v(\lambda_j) = \left|v\left(\lambda_j\right)\right|^2$ replace the GHP periodogram ordinates, 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) 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. Thus, the MLP estimator is especially useful in the nonstationary case when $d > 0.5$.[9]

4. Empirical Results

4.1. Rolling-Window Fractional Estimation Results

This section reports the results of the dynamics of inflation persistence using a rolling-window approach. The rolling-window approach examines whether inflation persistence in IT countries and the US changes in any notable manner. This, in turn, can highlight periods over which a pronounced decline or increase occurs in the persistence of the inflation rate series. The rolling-window fractional-integration approach estimates the inflation persistence parameter *d* using subsamples obtained by shifting the start and end points with a fixed window.

[9] See Phillips (2007) for a more detailed discussion, derivations and proofs.

We use a 10-year rolling window estimation.[10] This window length corresponds to the empirical literature on inflation persistence. For example, Zhang and Clovis (2009), Tillman and Wolters (2015), Carlstrom and Fuerst (2008) employ a 10-year window. Pivetta and Reis (2007) use a 14-year window, and replicated most of their results with 12- and 10-year windows with no noticeable change. The choice of a 10-year window also reflected the need to have long enough samples to assess persistence. A smaller, say 5-year, rolling window increases significantly the variability of the estimates. We first estimate the inflation persistence parameter *d* using the first 10 years of data. We then update the data in 1-year increments (Kumar andOkimoto, 2007), and re-estimate *d* for the updated window. We repeat this procedure until the end of the sample period. That is, the first value of *d* uses the subsample 1976:2 to 1986:1, with the result reported in 1985, the second value of *d* uses the subsample 1977:2 to 1987:1, with the result reported in 1986, and so on, and the last value of *d* uses the subsample 2003:1 to 2013:6, with the result reported in 2012. We compute all estimates using the modified log periodogram (MLP) estimator of Phillips (2007) based on the bandwidth power of 0.80.[11] Because the bandwidth restricts the number of ordinates used in the spectral regression, this window includes approximately 40 percent of the total number of observations. Using the estimated parameter of inflation persistence in each rolling fractional integration estimation, we produce twelve time series of inflation persistence that illustrates its dynamic and time-varying nature.[12]

Figure 2 depicts the 10-year rolling-window fractional-integration estimates of inflation persistence and their associated 95-percent confidence bands against the end year of each

---

[10] A larger window length improves the precision and reliability of the estimates that makes the fulfillment of asymptotic properties more likely. Imposing the same parameter values over the longer sample period, however, worsens the bias.

[11] Supporting evidence, using $m = T^{0.70}$ and $m = T^{0.75}$ is available from the authors.

[12] Because of the arbitrariousness of the 10-year window, we also compute alternative window lengths of 8-year and 12-years. All findings, available from the authors, remain robust to the choice of the window.

window. For the US, we observe a general decline in inflation persistence, starting from the early 1980s to the start of the global financial crisis and the Great Recession. That decline may reflect the Great Moderation. After such declines, however, we often see minor increase.

Tables 1 provides a more detailed profile of the 11 IT persistence series and the US. Table 1 displays the MLP rolling-window estimates of inflation persistence, $\hat{d}_{MLP}$, and their associated standard errors, $se(\hat{d}_{MLP})$, the test statistics for the null of unit root (d =1), and the null of short memory stationarity (d = 0). For comparison purposes, we also report in the first row of Table 1 the estimates and test statistics using on the full samples. The full-sample results indicate long memory and stationarity characterize inflation in Canada, Iceland, Norway, South Africa, South Korea, Sweden, Switzerland, the UK, and the US. For Chile, Israel, and Mexico, on the other hand, the full sample analysis indicates nonstationarity but mean reversion.

The rolling window estimates, however, clearly indicate that the constant parameter model does not adequately capture the dynamics of inflation persistence. We find a variety of responses to the regime switching. Canada, Iceland, Norway, and Sweden experience a change in inflation persistence from long memory to short memory in relation to IT switching. This indicates that when announced, the targets instantly gain credibility. On the other hand, Chile, South Korea, South Africa, Israel, and the UK do not experience immediately a change in persistence. Inflation persistence in these countries maintains the long memory property before and immediately after the IT adoption. For Chile and South Korea, however, a decline in the persistence estimates occurs the year following the IT adoption. This indicates mild lag effects. For Israel and the UK, a decline only occurs four years after the IT announcement. This suggests strong lag effects. For these countries, obviously, the targets do not achieve instant credibility. South Africa is unusual in that inflation persistence after IT adoption exhibits sustained increases, with estimates in the non-stationary, mean-reverting interval. Mexico is the only country that experiences a change in inflation persistence from a unit-root process (i.e., a nonstationary and non-mean reverting process) to a nonstationary and mean-reverting process two years prior to the formal IT announcement. This suggests the presence of a strong anticipation effect: economic agents anticipated the IT switch and discounted it two years ahead of the formal announcement. Switzerland exhibits no change either. Inflation persistence in Switzerland exhibits short memory properties before and after the regime switch.

We reject the unit-root hypothesis for all windows except Mexico, where the unit root is accepted in 11 windows between 1992 and 2003. Except for a few intermittent rolling windows, namely Canada in the 1990 window, Sweden in the 1989 window, the UK in the 1990 window, and the US in the 1986 to 1989 windows, that display nonstationary but mean reverting behavior, all developed countries exhibit stationary and mean reverting behavior in all remaining windows. Conversely, the developing countries exhibit more frequently nonstationary mean reverting behavior, such as Chile in the 1985 window, Israel in the 1990 and 1993 windows, Mexico in the 1985 to 1991 windows, in the 1997 and 2004 windows, and in the 2006 and 2007 windows, South Africa in the 2003 to 2011 windows.

Kumar and Okimoto (2007) show that the rolling-window estimates of inflation persistence display a strong downward trend in the context of fractional integration.[13] Their study, however, uses a sample of monthly data that ends in 2003:04 and, therefore, cannot capture the possible effects of the global financial crisis on inflation persistence. Conversely, we find that the global financial crisis exerts an effect on inflation persistence. In fact, for a few

[13] The simple regression of the rolling-window estimates of inflation persistence against a constant and a linear trend indicates that Canada, Israel, South Korea, Mexico, and Norway possess a significant negative trend. The remaining series possess a negative, but insignificant, trend. These findings, however, ignore the potential presence of structural changes in the inflation persistence series.

countries, the global financial crisis interrupts the downward trend in inflation persistence. We interpret this finding as indirect evidence of a de-anchoring process of inflationary expectations during the global financial crisis. We identify three different groups of countries with different responses to the global financial crisis. In the first group, Chile, South Korea, Sweden, Switzerland, the United Kingdom, and the United States show inflation persistence changing from short memory in the 2007 window to long memory in the 2008 window. In Chile, persistence jumps from 0.243 in the 2007 window to 0.422 and 0.519 in the 2008 and 2012 windows, respectively. In South Korea, inflation persistence jumps to 0.393 in the 2008 window from 0.081 in the 2007 window. Sweden jumps from -0.040 in the 2007 window to 0.271 in the 2008 window and remains at that level throughout the remaining windows. Switzerland jumps from 0.083 in the 2007 window to 0.458 and 0.417 in the 2008 and 2009 windows, respectively. In the UK persistence exhibits a similar pattern, jumping from -0.001 in the 2007 window to 0.399 in the 2008 window, and remaining on average at this level in the following four windows. The remaining countries do not exhibit significant changes. In the second group, Canada, and Norway do not appear to respond to the global financial crisis, as inflation persistence exhibits short memory both in the 2007 and 2008 windows. In the third group, Iceland, Israel, Mexico, and South Africa also do not appear to respond to the global financial crisis, as inflation persistence exhibits long memory in the 2007 window and the 2008 window. Table 2 displays summary descriptive statistics of the rolling-window MLP estimates.

4.2. Tests for structural breaks

This section tests for structural breaks in the inflation persistence series generated by the prior analysis. We explore the frequency of persistent shocks by exploiting the structural break tests. This methodology makes no assumptions about the location of the break, such as the date of the regime switching, or even if one exists. Even if we know the exact time of IT adoption, we do not necessarily know when IT adoption exerted its effects. Piehl, et al., (2003) indicate that policy evaluation should rely on endogenous break tests.

First, following Piehl, et al., (2003), we employ structural break tests to evaluate IT adoption. The finding of a structural break that coincides with IT adoption supports the hypothesis that IT adoption exerted a significant effect on inflation persistence. On the other hand, a finding of no structural break or a structural break at an alternative date far from the IT adoption date provides evidence against IT adoption exerting an effect on inflation persistence, leading to the conclusion that the regime switch was ineffective, with regard to inflation persistence.

Second, we employ the structural break tests to assess whether no break point exists in inflation persistence during the IT regime. In particular, we use the structural break tests to assess whether the 11 IT countries in the sample and the US experienced a structural break that coincides with the global financial crisis. The finding of a structural break that coincides with the global financial crisis indicates that inflation expectations become de-anchored during the crisis, which weakens the IT commitment.

The literature on endogenous tests for structural breaks grew dramatically in recent years. See Perron (2005) for a comprehensive survey. We apply the Bai and Perron (2003) multiple structural change tests to the estimates of inflation persistence shown in Figure 2. That is, we treat the estimates from the fractional integration rolling windows as "data." We use the Bai-Perron method to test for structural breaks at an unknown point of time as well as to identify the number of breaks and corresponding dates of occurrence.

In practice, to apply the Bai and Perron (2003), we need to define the regression relationship of interest. We focus our analysis on a model that regresses inflation persistence on a constant and a linear trend and rely on three alternative versions of this basic model. First, we consider a model of a break in two parameters (breaks in the mean and trend). This model assumes that both the intercept and the trend are breaking variables. Second, we consider model of a break in one parameter (break in the mean), but allow the trend to enter the model as a non-breaking variable (i.e., the trend is constant across the regimes). Third, we consider a model of a break in one parameter (break in the mean) and exclude the trend (breaking or otherwise) from the model.

We determine the number of breaks and their location employing the SupF(ℓ+1/ℓ) test, which sequentially tests the null hypothesis of ℓ breaks against the alternative of ℓ+1 breaks. We set the maximum number of breaks at M=3 and use a trimming parameter of 0.15. Setting the maximum number of breaks to M=5 does not change the findings of a maximum of M=3. We implement the analysis using the three versions of the basic model, and report the findings corresponding to the version of the model with the highest goodness of fit, based on the adjusted R^2. Table 3 shows the sequential test results, together with the estimates of the number and the location of the breaks.

This approach supports the following inferences. The SupF(ℓ+1/ℓ) test suggests that we use the third model (only break in the mean and no trend) only for Norway. For Iceland and the UK, the SupF(ℓ+1/ℓ) indicates that the second model (break in the mean with non-breaking trend) is more appropriate. For Canada, Chile, Israel, Mexico, South Africa, South Korea, Switzerland, and the United States, the SupF(ℓ+1/ℓ) test recommends that we use the first model (breaks in the mean and trend). Sweden is a noteworthy case. The SupF(ℓ+1/ℓ) fails to find a structural break for Sweden. The no-break model of Sweden, however, delivers an adjusted R^2 of

-0.03. Re-estimating Sweden using the sequentially fixed number of breaks, we obtain two breaks, in 1989 and 2008, and an adjusted R^2 of 0.318 using the first version of the model. Based upon this additional finding, we conclude that all persistence series experience at least one structural break.

One can argue that the break dates can be misleading due to an initial transition period of credibility of inflation targets or to an initial anticipation prior to the implementation of the targets (Yigit, 2010). For this reason, we consider any structural break that falls in the band defined by 1-year movements of the regime change in both directions as coincident with the IT switching. From this perspective, Canada, Iceland, Mexico, and South Korea experience a structural break that broadly coincides with IT implementation. For the remaining eight countries, we do not identify any statistical discontinuity in the inflation persistence series in the band defined by the 1-year movements of the regime in both directions at the time of IT implementation. For Chile and Israel, the break takes place about four years after the regime switching. For Norway, South Africa, Sweden, and the UK, the break takes place approximately one to three years before the formal IT adoption date. Canada, Iceland, Norway, South Africa, and South Korea do not experience any further structural breaks after the implementation of IT. This, in turn, implies that the IT period in these countries corresponds to a time of relative stability of inflation persistence. Conversely, of the remaining six countries (Chile, Israel, Mexico, Sweden, Switzerland, and the UK), two countries (Chile and Israel) experience two structural breaks after IT implementation, and four countries (Mexico, Sweden, Switzerland, and the UK) experience one structural break after IT implementation.

The Bai-Perron tests identify only four countries, the US and three IT countries (Sweden, Switzerland, and the UK) that experience a structural break in connection with the global financial crisis. These shocks expose the vulnerability of the IT regime and US monetary policy. The IT framework, after all, may not prove more relevant on anchoring inflation expectations than other monetary regimes. Interestingly, none of the developing countries experience a structural break in connection with the global financial crisis.

Figure 3 highlights the structural changes in the mean and/or the trend of the 11 IT and US inflation persistence series. The vertical segment marks the official date of IT implementation. The graph at the top depicts the actual data and the fitted data, while the graph at the bottom depicts the residuals. Table 4 presents the OLS estimates of the regimes together with the *t*-statistics and the adjusted R^2.

Given the results in Table 3, we reapply the MLP test to the subsamples defined by the breaks. Table 5 displays the regime-specific MLP estimates of inflation persistence, $\hat{d}_{MLP}$, and their associated standard errors, $se(\hat{d}_{MLP})$, the test statistics for the null of unit root (d=1) and the null of short memory stationarity (d=0). There are 30 regimes, corresponding to 18 breaks in total. In all regimes, we reject the null hypothesis that $d = 1$. No switches from non–mean-reversion nonstationarity to mean-reversion stationarity exist. Switches from mean-reversion nonstationarity to mean-reversion stationarity, however, do exist (i.e., Chile, Israel, and Mexico from the first to the second regime, South Korea from the second to the third regime) and from mean-reversion stationarity to mean-reversion nonstationarity (Chile from the second to the third regime). For Chile and Israel, we reject the null hypothesis of d=0 for all three regimes. For Iceland and South Africa, we fail to reject the null hypothesis of d=0 in the first regime, but we can reject it in the second regime. Conversely, for Norway, South Korea, Sweden, Switzerland, and the UK, we reject the null hypothesis of d=0 in first regime, but cannot reject it in the second or third regimes (the second regime for Norway, the third regime for South Korea and the UK, and the second and third regimes for Sweden). The estimates of the fractional integration parameter show that when allowing for structural breaks, eight regimes no longer display long memory (Canada in the second regimes; Chile in the second regime; Iceland in the first regime; Norway in the second regime; South Africa in the first regime; South Korea in the third regime; Sweden in the first regime; Switzerland in the second regime; the UK in the third regime). The remaining regimes, including the US. exhibit long memory. Of these, 5 regimes exhibit nonstationary, but mean reversion (Chile in the first and third regimes; Israel in the first regime; Mexico in the first and second regimes), and the remaining 14 regimes display long memory, i.e., are stationary and mean reverting (Chile in the second regime; Iceland in the second regime; Israel in the second and third regimes; Mexico in the third regime; Norway in the first regime; South Korea in the first regime; Sweden in the second and third regime; Switzerland in the first regime; the UK in the first and second regimes; and the US in the first and second regimes). In sum, inflation persistence in these countries exhibits heterogeneous dynamics, suggesting that the response to inflationary shocks even in IT countries is more sophisticated than the conventional IT framework can explain. We need to examine three countries, in particular, that experienced hyperinflation in the sample period: Chile in1973, Israel in 1984-1985, and Mexico in 1982-1993. These are the only countries that exhibit nonstationarity in inflation persistence. That is, the "memory" of hyperinflation may not easily fade away, even after the adoption of the IT regime, leading to nonstationary persistence.

5. Conclusions

We employ a three-stage approach to the analysis the dynamics of inflation persistence in 11 IT countries and the US. First, we implement rolling-window-fractional-integration tests. The rolling-window estimates indicate that most inflation rate series exhibit a drop in persistence after the adoption of IT. The fall in persistence, however, does not always synchronize with the IT adoption date. Canada, Iceland, Norway, and Sweden experience a change in inflation persistence from long memory to short memory in relation to IT switching. This indicates that when announced, the targets instantly gain credibility. On the other hand, Chile, South Korea, South Africa, Israel, and the UK do not experience immediately a change in persistence. Inflation persistence in these countries maintains the long memory property before and immediately after the IT adoption. For Chile and South Korea, however, a decline in the persistence occurs the year following IT adoption. This indicates mild lag effects. For Israel and the UK, a decline only occurs four years after the IT announcement. This, on the other hand, suggests strong lag effects. For these countries, obviously, the targets do not achieve instant credibility. South Africa is unusual in that inflation persistence after IT adoption exhibits sustained increases, with estimates in the non-stationary, mean-reverting interval. Mexico is the only country that experiences a change in inflation persistence from a unit-root process (i.e., a nonstationary and non-mean reverting process) to a nonstationary and mean-reverting process two years prior to the formal IT announcement. This suggests the presence of strong anticipation effects: economic agents anticipated the IT switch and discounted it two years ahead of the formal announcement. Switzerland exhibits no change either. Inflation persistence in Switzerland exhibits short memory properties before and after the regime switch.

The response of inflation persistence to the global financial crisis causes heterogeneous effects. The global financial crisis, however, terminates this decline in most inflation rate series. We find evidence of a de-anchoring process of inflationary expectations during the global financial crisis, where about half of the inflation series exhibit an increase from 2007 to 2008. The increase, however, generally appears transitory.

Second, we test for structural breaks in the inflation persistence series generated by the first-stage analysis. We extend Piehl, et al., (2003) and argue that the evaluation of a change in regimes or the evaluation of a change in macroeconomic events of large proportions should depend on whether the change generates a structural change. We find that all inflation persistence processes undergo at least one structural break, but only four countries (Canada, Iceland, Mexico, and South Korea) experience a structural break that coincides with the implementation of IT. Of the remaining countries, we find no statistical discontinuity in the inflation persistence series at IT implementation, which signals that the IT policy did not exert an immediate effect on inflation persistence. Furthermore, only four countries, the US and three IT countries (Sweden, Switzerland, and the UK) experience a structural break in connection with the global financial crisis.

Finally, we reapply the Phillips (2007) estimator to the subsamples defined by the breaks. We find that in most cases the estimates of inflation persistence switch from mean-reversion nonstationarity to mean-reversion stationarity.

In conclusion, we show that most IT industrial countries in our sample experience decreasing persistence over time, while for developing countries, some show decreasing, others show increasing, and still others present highly persistent inflationary processes. In terms of macroeconomic policies, we think that these results are important for developed and developing countries. For the developing countries, more than for the developed countries, the effectiveness of the IT regime relates to the capacity of their central banks to enact macroeconomic policies that focus on diminishing inflation persistence. Conversely, for the developed countries, more than for the developing countries, financial crises represent an obstacle to well-anchoring of inflation expectations, making inflation persistence more vulnerable to financial crisis.

References:

Agiakloglou, C., Newbold, P., and Wohar, M., 1993. Bias in an Estimator of the Fractional Difference Parameter. *Journal of Time Series Analysis*, 14, 235–246.

Arize, A. C., Malindretos J., and Nam K., 2005. Inflation and structural change in 50 developing countries. *Atlantic Economic Journal* 33, 461–471.

Autrup, S. L. and Grothe, M., 2014. Economic Surprises and Inflation Expectations: Has Anchoring of Expectations Survived the Crisis? European Central Bank Working Paper Series No 1671.

Bai, J., and Perron, P., 2003. Computation and analysis of multiple structural change models. *Journal of Applied Economet*rics 18, 1–22.

Baillie R. T., and Bollerslev, T. , 1994. Cointegration, fractional cointegration, and exchange rate dynamics. *Journal of Finance* 49, 737–745.

Baillie, R. T., Chung, C. F., and Tieslau, M. A., 1996. Analyzing inflation by the fractionally integrated ARFIMA-GARCH model. *Journal of Applied Econometrics* 11, 23-40.

Baillie, R.T., Han, Y.W., and Kwon, T.-G., 2002. Further long memory properties of inflationary shocks. *Southern Economic Journal* 68, 496-510.

Ball, L., and Sheridan, N., 2005. Does inflation targeting matter? In *The Inflation Targeting Debate* B. Bernanke and M. Woodford (eds). Chicago: University of Chicago Press, 249-276.

Barkoulas, J., Baum, C., and Caglayan, M., 1999. Fractional monetary dynamics. *Applied Economics* 13, 1393-1400.

Baum C. F., and Barkoulas, J. T., 2006. Dynamics of intra-EMS interest rate linkages Journal of Money, Credit, and Banking 38 (2), 469-482.

Baum, C. F., Barkoulas, J. T., and Caglayan, M., 1999. Persistence in international inflation rates *Southern Economic Journal* 65, 900-913.

Beechey, M., and Österholm, P., 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.

Bernanke, B., Laubach, T., Mishkin, F., and Posen A. S., 1999. *Inflation Targeting: Lessons from the International Experience*. Princeton, NJ: Princeton University Press.

Bos, C. S., Franses, P. H. B. F., and Ooms, M., 1999. Long memory and level shifts: Reanalyzing inflation rates. *Empirical Economics* 24, 427–449.

Bos, C., Franses, P. H., and Ooms, M., 2002. Inflation, forecast intervals and long memory regression models. *International Journal of Forecasting* 18, 243–264.

Carlstrom, C.T., and Fuerst, T.S., 2008. Explaining Apparent Changes in the Phillips Curve: TheGreat Moderation and Monetary Policy. Federal Reserve Bank of Cleveland, *Economic Commentary* (February).

Canarella, G., and Miller, S. M., forthcoming. Inflation Persistence Before and After Inflation Targeting: A Fractional Integration Approach. *Eastern Economic Journal*.

Cheung Y.-W., and Lai K., 1993. A fractional cointegration analysis of purchasing power parity. *Journal of Business and Economic Statistics* 11, 103–112.

Cheung, Y.-W., 1993. Tests for fractional integration: A Monte Carlo investigation. *Journal of Time Series Analysis* 14, 331–345.

Christiano, L., Einchenbaum, M., and Evans, C., 2005. Nominal Rigidities and the Dynamics Effects of Shocks to Monetary Policy. *Journal of Political Economy* 113, 1-45.

Diebold, F.X., Husted, S. and Rush, M., 1991. Real exchange rates under the gold standard. *Journal of Political Economy* 99, 1252-1271.

Diebold, F.X., and Rudebusch, G.D., 1989. Scoring the Leading Indicators. *Journal of Business* 64, 369-392.

Dueker, M., and Startz, R., 1998. Maximum-likelihood estimation of fractional cointegration with an application to U.S. and Canadian bond rates. *Review of Economics and Statistics* 80, 420-426.

Fracasso, A., Genberg, H., and Wyplosz, C., 2003. How do central banks write? An evaluation of inflation targeting. Central Banks Geneva Reports on the World Economy Special Report 2, Center for Economic Policy Research, London.

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

Gadea, M. D., and Mayoral, L., 2005. The Persistence of Inflation in OECD Countries: A Fractionally Integrated Approach. *The International Journal of Central Banking* 2, 51-104.

Galati, G., Poelhekke, S., and Zhou, C., 2011. Did the Crisis Affect Inflation Expectations? *International Journal of Central Banking* 7, 167-207.

Gali, J., and Gertler, M., 1999. A Structural Econometric Analysis. *Journal of Monetary Economics* 44, 195-222.

Gamber, E. N., Liebner, J. P., and Smith, J. K., 2012. Inflation persistence: Revisited. Working Paper, Center of Economic Research, George Washington University.

Gerlach, S. and Tillmann, P. 2012. Inflation targeting and inflation persistence in Asia, *Journal of Asian Economics* 23, 360-373.

Geweke, J., and Porter-Hudak, S., 1983. The estimation and application of long memory time series models. *Journal of Time Series Analysis* 4, 221–237.

Gürkaynak, R. S., Levin, A.T., and Swanson, E.T., 2006. Does inflation targeting anchor long run inflation expectations? Evidence from long-term bond yields in the US, UK and Sweden. *Federal Reserve Bank of San Francisco Working Paper Series*.

Gürkaynak, R. S., Levin, A.T., Marder, A.N., and Swanson, E.T., 2007. Inflation targeting and the anchoring of inflation expectations in the Western Hemisphere. In F. S. Mishkin and K. Schmidt-Hebbel (eds), *Monetary Policy under Inflation Targeting*, Central Bank of Chile.

Hassler, U., and Meller, B., 2014. Detecting multiple breaks in long memory the case of US inflation. *Empirical Economics* 46, 653-680.

Hassler, U., P.M.M. Rodrigues, P.M.M., and Rubia, A., 2014. Persistence in the banking industry: Fractional integration and breaks in memory. *Journal of Empirical Finance* 29, 95-112.

Hassler, U., and Wolters, J., 1995. Long memory in inflation rates: International evidence. *Journal of Business and Economic Statistics* 13:1, 37–45.

Hosking, J., 1981. Fractional differencing. *Biometrika* 68, 165–176.

Hurvich, C. M., Deo, R., and Brodsky, J., 1998. The mean squared error of Geweke and Porter Hudak's estimator of the memory parameter of a long memory time series. *Journal of Time Series Analysis* 19, 19–46.

Kim, C. S., and Phillips, P. C. B., 2006. Log periodogram regression; The nonstationary case. Cowles Foundation Discussion Paper No. 1587.

Kumar, M. S., and Okimoto, T., 2007. Dynamics of persistence in international inflation rates. *Journal of Money, Credit and Banking* 39, 1457-1479.

Kuttner, K. N., and Poser, A. S., 2001. Beyond bipolar: A three-dimensional assessment of monetary frameworks. *International Journal of Finance and Economics* 6, 369-387.

Levin, A. T., Natalucci, F. M., and Piger, J. M., 2004. The macroeconomic effects of inflation targeting. Federal Reserve Bank of St. Louis, *Review* 86, 51–80.

Levin, A. T., and Piger, J. M., 2004. Is inflation persistence intrinsic in industrial economies. European Central Bank Working Paper Series No. 334.

Martins, L. F., and Rodrigues, P. M. M., 2014. Testing for persistence in fractionally integrated models: An application to world inflation rates. *Computational Statistics and Data Analysis* 76, 502-522.

Meller, B., and Nautz, D., 2012. Inflation persistence in the Euro area before and after the European Monetary Union. *Economic Modelling* 29, 1170–1176.

Mishkin, S. F., 2007. Inflation Dynamics. Working Paper Series National Bureau of Economic Research, NBER, No 13147.

Mishkin, F. S., and Schmidt-Hebbel, K., 2001. One decade of inflation targeting in the world: What do we know and what do we need to know? National Bureau of Economic Research, Working Paper No. 8397.

Mishkin, F.S., 1998. The Dangers of Exchange Rate Pegging in Emerging-Market Countries. *International Finance* 1, 81-101.

Nautz, D., and Strohsal, T., 2015. Are US Inflation Expectations Re-Anchored? *Economics Letters* 127, 6-9.

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

Phillips, P. C. B., 2007. Unit root log periodogram regression. *Journal of Econometrics* 138, 104-124.

Perron, P., 2005. Dealing with structural breaks. In Mills T.C., and Patterson, K., (eds) Palgrave Handbook of econometrics, v. 1, *Econometric Theory*. Palgrave Macmillan.

Piehl, A. M., Cooper, S. J., Braga, A. A., and Kennedy, D. M., 2003. Testing for structural breaks in evaluation of programs. *Review of Economics and Statistics* 85, 550-558.

Pivetta, F., and Reis, R., 2007. The persistence of inflation in the United States. *Journal of Economic Dynamics and Control* 31, 1326-1358.

Robinson, P. M., 1995. Log periodogram regression of time series with long memory dependence. *The Annals of Statistics* 23, 1048–1072.

Shimotsu, K., 2006. Simple (but effective) tests of long memory versus structural breaks. Working paper, Queen's University.

Sibbertsen, P., and Kruse, R., 2009. Testing for a break in persistence under long-range dependencies. Journal of Time Series Analysis 30, 263-285.

Siklos, P. L., 1999. Inflation-Target Design: Changing Inflation Performance and Persistence in Industrial Countries. Federal Reserve Bank of St. Louis *Review* 81, 47–58.

Sowell, F., 1992. Modeling long-run behavior with the fractional ARIMA models. *Journal of Monetary Economics* 29, 277-302.

Stock, J. H., 2001. Comment on 'Evolving post-World War II U.S. inflation dynamics'. *NBER Macroeconomics Annual 16*, 379–387.

Tillmann, P., and Wolters, M.H., 2015. The changing dynamics of US inflation persistence: a quantile regression approach. *Studies in Nonlinear Dynamics & Econometrics*. 19, 161-182.

Valadkhani, A.,2005. Modelling Demand for Broad Money in Australia. *Australian Economic Papers* 44, 47-64.

van der Cruijsen, C. A. B, Eijffinger, S. C. W., and Hoogduin, L. H., 2010. Optimal central bank transparency. *Journal of International Money and Finance* 29, 1482‑1507.

Walsh, C., 2008. Inflation targeting: what have we learned? The John Kuszczak Memorial Lecture, prepared for the conference *International experience with the conduct of monetary policy under inflation targeting*. Bank of Canada, 22–23 July.

Yamaguchi, K., 2011. Estimating a change point in the long memory parameter. *Journal of Time Series Analysis* 32, 304-314.

Yigit, T., 2010. Inflation targeting: An indirect approach to assess the direct impact. *Journal of International Money and Finance* 29, 1357-68.

Zhang, C., and Clovis, J., 2009. Modeling China Inflation Persistence. *Annals of Economics and Finance* 10, 89-110.

Table 1: Rolling-window estimates and tests of inflation persistence

Window	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$
	Canada				Chile				Iceland			
Full Sample	0.248	0.072	3.425	-13.472	0.703	0.060	11.669	-5.328	0.386	0.070	5.494	-10.999
1985	0.210	0.111	1.900	-8.353	0.755	0.109	6.910	-2.591	0.242	0.133	1.822	-8.012
1986	0.164	0.105	1.563	-8.844	0.447	0.090	4.942	-5.853	0.372	0.113	3.282	-6.642
1987	0.157	0.112	1.406	-8.912	0.438	0.114	3.832	-5.939	0.315	0.157	1.998	-7.248
1988	0.242	0.123	1.966	-8.015	0.406	0.107	3.779	-6.282	0.284	0.129	2.198	-7.576
1989	0.325	0.118	2.746	-7.138	0.401	0.092	4.367	-6.338	0.302	0.133	2.264	-7.387
1990	0.736	0.136	5.418	-2.793	0.298	0.107	2.774	-7.422	0.407	0.173	2.353	-6.271
1991	0.087	0.096	0.900	-9.660	0.240	0.083	2.889	-8.035	0.487	0.136	3.574	-5.424
1992	0.172	0.122	1.414	-8.753	0.112	0.117	0.954	-9.391	0.490	0.125	3.915	-5.396
1993	0.163	0.119	1.366	-8.852	0.048	0.087	0.547	-10.073	0.462	0.095	4.859	-5.695
1994	0.187	0.146	1.284	-8.594	0.398	0.109	3.650	-6.369	0.468	0.123	3.823	-5.622
1995	0.129	0.127	1.019	-9.206	0.387	0.109	3.549	-6.483	0.460	0.129	3.578	-5.707
1996	0.125	0.130	0.961	-9.250	0.329	0.100	3.271	-7.100	0.400	0.168	2.375	-6.343
1997	0.269	0.114	2.363	-7.733	0.339	0.113	3.001	-6.988	0.545	0.129	4.239	-4.808
1998	0.222	0.088	2.525	-8.227	0.150	0.122	1.220	-8.992	0.463	0.136	3.392	-5.682
1999	0.161	0.102	1.579	-8.878	0.190	0.113	1.683	-8.569	0.443	0.112	3.965	-5.889
2000	0.322	0.109	2.943	-7.168	0.308	0.150	2.053	-7.319	0.304	0.104	2.916	-7.366
2001	0.048	0.129	0.370	-10.071	0.085	0.130	0.655	-9.674	0.263	0.137	1.925	-7.792
2002	0.364	0.113	3.228	-6.730	0.158	0.111	1.422	-8.908	0.204	0.112	1.819	-8.414
2003	-0.082	0.120	-0.682	-11.442	0.069	0.117	0.588	-9.845	0.324	0.099	3.268	-7.151
2004	0.102	0.116	0.876	-9.502	0.163	0.121	1.346	-8.855	0.341	0.104	3.266	-6.974
2005	0.253	0.123	2.051	-7.902	0.249	0.109	2.282	-7.948	0.265	0.098	2.717	-7.769
2006	-0.028	0.141	-0.194	-10.867	0.263	0.130	2.016	-7.793	0.315	0.124	2.54	-7.248
2007	0.024	0.130	0.181	-10.327	0.243	0.145	1.681	-8.003	0.277	0.134	2.071	-7.641
2008	0.185	0.102	1.817	-8.621	0.422	0.120	3.518	-6.114	0.386	0.096	4.008	-6.498
2009	0.048	0.122	0.393	-10.068	0.252	0.133	1.893	-7.910	0.335	0.124	2.703	-7.037
2010	-0.010	0.103	-0.102	-10.687	0.406	0.101	4.036	-6.283	0.499	0.114	4.372	-5.303
2011	0.033	0.111	0.296	-10.228	0.439	0.118	3.725	-5.933	0.439	0.140	3.147	-5.930
2012	0.040	0.093	0.429	-10.262	0.519	0.116	4.484	-5.144	0.468	0.113	4.139	-5.692

Table 1: Rolling-window estimates and tests of inflation persistence (continued)

Window	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$
	Israel				Mexico				Norway			
Full Sample	0.526	0.068	7.726	-8.500	0.694	0.063	10.996	-5.476	0.258	0.055	4.695	-13.292
1985	0.474	0.150	3.149	-5.563	0.585	0.113	5.187	-4.394	0.371	0.139	2.671	-6.652
1986	0.454	0.134	3.376	-5.778	0.547	0.130	4.199	-4.786	0.431	0.126	3.418	-6.012
1987	0.437	0.126	3.458	-5.952	0.602	0.118	5.080	-4.214	0.401	0.124	3.232	-6.333
1988	0.479	0.118	4.053	-5.51	0.742	0.115	6.436	-2.725	0.295	0.135	2.182	-7.456
1989	0.466	0.113	4.120	-5.643	0.775	0.116	6.673	-2.381	0.160	0.124	1.294	-8.880
1990	0.514	0.106	4.865	-5.137	0.806	0.114	7.055	-2.047	0.165	0.137	1.207	-8.830
1991	0.449	0.107	4.200	-5.832	0.736	0.125	5.910	-2.788	0.378	0.121	3.133	-6.569
1992	0.452	0.113	4.003	-5.797	0.861	0.124	6.958	-1.466	0.287	0.096	2.964	-7.539
1993	0.540	0.088	6.148	-4.866	0.857	0.092	9.266	-1.512	0.442	0.135	3.266	-5.899
1994	0.493	0.080	6.142	-5.361	0.874	0.086	10.127	-1.332	0.328	0.125	2.617	-7.107
1995	0.236	0.090	2.609	-8.078	0.877	0.115	7.634	-1.298	0.498	0.147	3.392	-5.299
1996	0.221	0.106	2.081	-8.236	1.002	0.113	8.881	0.024	0.37	0.122	3.017	-6.661
1997	0.146	0.127	1.143	-9.036	0.696	0.084	8.295	-3.216	0.031	0.123	0.254	-10.244
1998	0.378	0.132	2.864	-6.581	0.852	0.110	7.732	-1.568	0.018	0.118	0.151	-10.386
1999	0.362	0.094	3.831	-6.752	0.966	0.124	7.772	-0.355	0.123	0.115	1.070	-9.271
2000	0.357	0.103	3.474	-6.804	1.01	0.123	8.237	0.108	0.269	0.119	2.259	-7.730
2001	0.497	0.109	4.538	-5.323	0.948	0.100	9.446	-0.548	0.137	0.116	1.180	-9.126
2002	0.383	0.121	3.180	-6.523	0.952	0.118	8.081	-0.508	0.701	0.137	5.110	-3.156
2003	0.372	0.138	2.693	-6.640	0.829	0.095	8.688	-1.804	0.135	0.103	1.312	-9.145
2004	0.344	0.093	3.698	-6.938	0.721	0.089	8.137	-2.948	-0.071	0.135	-0.532	-11.337
2005	0.381	0.093	4.105	-6.548	0.455	0.113	4.040	-5.766	-0.003	0.102	-0.029	-10.608
2006	0.328	0.103	3.191	-7.103	0.534	0.093	5.742	-4.927	0.419	0.104	4.005	-6.141
2007	0.395	0.121	3.273	-6.402	0.571	0.086	6.676	-4.532	-0.013	0.114	-0.114	-10.715
2008	0.386	0.137	2.824	-6.489	0.332	0.142	2.332	-7.069	0.071	0.108	0.661	-9.819
2009	0.393	0.099	3.972	-6.419	0.496	0.091	5.427	-5.333	0.159	0.161	0.983	-8.894
2010	0.354	0.080	4.444	-6.83	0.246	0.113	2.174	-7.974	0.110	0.108	1.019	-9.406
2011	0.280	0.104	2.695	-7.615	0.311	0.127	2.446	-7.284	0.032	0.115	0.280	-10.233
2012	0.361	0.094	3.842	-6.832	0.357	0.164	2.174	-6.870	0.227	0.113	2.008	-8.256

Table 1: Rolling-window estimates and tests of inflation persistence (continued)

Window	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0: d=0$	$H_0: d=1$
	South Africa				South Korea				Sweden			
Full Sample	0.291	0.065	4.466	-12.694	0.345	0.063	5.477	-11.734	0.219	0.062	3.497	-13.999
1985	0.178	0.105	1.693	-8.686	0.239	0.158	1.512	-8.050	0.091	0.108	0.842	-9.612
1986	-0.128	0.144	-0.888	-11.934	0.348	0.102	3.427	-6.896	0.106	0.111	0.954	-9.454
1987	-0.058	0.111	-0.524	-11.193	0.435	0.118	3.696	-5.977	0.203	0.086	2.371	-8.426
1988	-0.160	0.107	-1.487	-12.269	0.383	0.112	3.405	-6.527	0.105	0.120	0.877	-9.464
1989	-0.108	0.108	-0.998	-11.724	0.600	0.123	4.865	-4.234	0.550	0.124	4.426	-4.762
1990	-0.063	0.110	-0.577	-11.251	0.529	0.115	4.585	-4.983	0.379	0.099	3.838	-6.562
1991	-0.018	0.106	-0.173	-10.771	0.153	0.126	1.212	-8.956	0.185	0.173	1.067	-8.618
1992	-0.007	0.105	-0.071	-10.656	0.160	0.135	1.183	-8.885	0.446	0.107	4.152	-5.858
1993	0.029	0.136	0.215	-10.266	0.131	0.148	0.884	-9.188	0.170	0.093	1.839	-8.773
1994	-0.037	0.134	-0.278	-10.970	0.092	0.143	0.645	-9.602	0.251	0.105	2.386	-7.918
1995	0.101	0.099	1.012	-9.508	0.118	0.131	0.896	-9.330	0.292	0.122	2.390	-7.483
1996	0.196	0.114	1.717	-8.494	0.013	0.117	0.114	-10.435	0.244	0.101	2.420	-7.999
1997	0.138	0.113	1.216	-9.114	0.446	0.143	3.112	-5.854	0.314	0.089	3.521	-7.255
1998	0.297	0.109	2.723	-7.426	0.470	0.102	4.615	-5.603	0.223	0.071	3.122	-8.217
1999	0.307	0.123	2.492	-7.327	0.275	0.114	2.406	-7.664	0.108	0.096	1.128	-9.432
2000	0.350	0.098	3.566	-6.870	0.38	0.123	3.102	-6.553	0.074	0.164	0.454	-9.789
2001	0.414	0.153	2.706	-6.194	0.221	0.104	2.132	-8.241	-0.144	0.113	-1.271	-12.101
2002	0.307	0.112	2.744	-7.322	0.244	0.136	1.787	-7.995	0.431	0.114	3.786	-6.015
2003	0.500	0.127	3.920	-5.281	0.239	0.120	1.986	-8.048	0.245	0.100	2.460	-7.982
2004	0.571	0.102	5.587	-4.534	0.361	0.126	2.871	-6.762	0.151	0.115	1.312	-8.975
2005	0.546	0.114	4.778	-4.794	0.279	0.113	2.467	-7.624	0.080	0.101	0.797	-9.728
2006	0.605	0.110	5.487	-4.175	0.408	0.116	3.502	-6.262	0.080	0.110	0.724	-9.733
2007	0.637	0.107	5.927	-3.830	0.081	0.103	0.786	-9.720	-0.040	0.114	-0.348	-10.995
2008	0.507	0.128	3.964	-5.207	0.393	0.117	3.370	-6.420	0.271	0.086	3.142	-7.707
2009	0.659	0.133	4.957	-3.596	0.054	0.117	0.460	-10.006	0.173	0.097	1.775	-8.748
2010	0.590	0.127	4.635	-4.330	0.219	0.096	2.274	-8.256	0.244	0.098	2.477	-7.999
2011	0.641	0.128	4.980	-3.790	0.141	0.113	1.247	-9.083	0.361	0.112	3.206	-6.761
2012	0.432	0.105	4.104	-6.071	0.017	0.112	0.153	-10.507	0.321	0.114	2.811	-7.254

Table 1: Rolling-window estimates and tests of inflation persistence (continued)

Window	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0:$ $d=0$	$H_0:$ $d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0:$ $d=0$	$H_0:$ $d=1$	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0:$ $d=0$	$H_0:$ $d=1$
	Switzerland				United Kingdom				United States			
Full Sample	0.239	0.064	3.712	-13.642	0.469	0.075	6.280	-9.515	0.391	0.062	6.294	-10.915
1985	0.338	0.085	3.969	-7.004	0.274	0.120	2.290	-7.674	0.479	0.115	4.157	-5.509
1986	0.251	0.082	3.071	-7.924	0.261	0.086	3.050	-7.815	0.627	0.137	4.561	-3.942
1987	0.282	0.083	3.391	-7.592	0.220	0.098	2.233	-8.252	0.673	0.131	5.118	-3.459
1988	0.125	0.086	1.457	-9.252	0.216	0.104	2.075	-8.292	0.588	0.108	5.446	-4.355
1989	0.257	0.124	2.066	-7.860	0.401	0.126	3.188	-6.340	0.617	0.131	4.698	-4.053
1990	0.255	0.119	2.149	-7.876	0.547	0.155	3.521	-4.795	0.482	0.102	4.742	-5.481
1991	0.323	0.106	3.060	-7.159	0.273	0.120	2.266	-7.693	0.199	0.180	1.109	-8.469
1992	0.241	0.111	2.184	-8.022	0.331	0.110	3.012	-7.076	0.361	0.141	2.555	-6.754
1993	0.174	0.124	1.395	-8.741	0.251	0.113	2.212	-7.922	0.500	0.115	4.359	-5.294
1994	0.312	0.138	2.261	-7.275	0.306	0.106	2.893	-7.338	0.401	0.092	4.361	-6.333
1995	0.111	0.133	0.828	-9.407	0.264	0.127	2.071	-7.788	0.376	0.089	4.205	-6.599
1996	0.193	0.125	1.538	-8.538	0.118	0.119	0.994	-9.323	0.277	0.105	2.649	-7.644
1997	0.314	0.106	2.974	-7.252	0.127	0.114	1.110	-9.236	0.323	0.104	3.105	-7.165
1998	0.048	0.124	0.389	-10.064	0.195	0.118	1.649	-8.518	0.306	0.103	2.957	-7.344
1999	0.161	0.117	1.373	-8.869	0.142	0.124	1.150	-9.071	0.190	0.159	1.200	-8.563
2000	0.058	0.127	0.460	-9.959	0.202	0.132	1.533	-8.435	0.137	0.137	0.997	-9.128
2001	0.224	0.104	2.150	-8.212	0.304	0.120	2.539	-7.362	0.175	0.160	1.090	-8.725
2002	0.311	0.109	2.860	-7.285	0.043	0.132	0.323	-10.124	0.083	0.114	0.726	-9.699
2003	0.154	0.115	1.332	-8.952	-0.009	0.124	-0.075	-10.675	0.166	0.122	1.362	-8.817
2004	0.104	0.134	0.776	-9.476	0.109	0.134	0.813	-9.424	0.216	0.120	1.807	-8.288
2005	0.127	0.145	0.874	-9.230	0.030	0.131	0.230	-10.257	0.167	0.114	1.469	-8.811
2006	0.139	0.125	1.112	-9.105	0.003	0.125	0.020	-10.549	-0.322	0.131	-2.452	-13.981
2007	0.083	0.148	0.559	-9.700	-0.001	0.098	-0.011	-10.588	-0.065	0.162	-0.401	-11.262
2008	0.458	0.135	3.386	-5.727	0.399	0.096	4.170	-6.351	0.312	0.154	2.026	-7.277
2009	0.417	0.146	2.861	-6.171	0.298	0.108	2.773	-7.421	0.270	0.140	1.936	-7.715
2010	0.165	0.128	1.283	-8.833	0.424	0.124	3.429	-6.093	0.316	0.114	2.776	-7.233
2011	0.311	0.169	1.840	-7.286	0.379	0.126	3.012	-6.564	0.282	0.113	2.500	-7.594
2012	0.278	0.110	2.530	-7.722	0.312	0.097	3.217	-7.349	0.318	0.088	3.626	-7.292

Note: The table displays the 10-year rolling-window estimates of inflation persistence, ($\hat{d}_{MLP}$), their associated standard errors, ($se(\hat{d}_{MLP})$), and the test statistics for the null of unit root ($H_0: d=1$) and the null of short memory ($H_0: d=0$). We obtain the rolling window estimates using Phillips (2007) modified log periodogram based on the bandwidth power of 0.8. For comparison, the first row reports the full sample estimates (1976:2-2013:6) and related statistics.

Table 2: Rolling-window estimates summary

Country	Mean	Maximum	Minimum	St. Dev.
Canada	0.166	0.736	-0.082	0.157
Chile	0.304	0.755	0.048	0.157
Iceland	0.377	0.545	0.204	0.092
Israel	0.390	0.540	0.146	0.092
South Korea	0.265	0.600	0.013	0.159
Mexico	0.698	1.010	0.246	0.224
Norway	0.231	0.702	-0.072	0.186
South Africa	0.265	0.659	-0.160	0.275
Sweden	0.211	0.550	0.144	0.151
Switzerland	0.222	0.458	0.048	0.106
United Kingdom	0.229	0.547	-0.009	0.142
United States	0.302	0.673	-0.322	0.213

Note: The table reports four summary statistics of the rolling-window fractional-integration estimates.

Table 3: Bai-Perron structural break test : SupF(l+1/l) statistics

Country	SupF(1/0)	SupF(2/1)	SupF(3/2)	No. of Breaks	Break Date
Canada (1)	14.030*	5.728		1	1991
Chile (1)	30.467*	24.001*	4.913	2	1994 2004
Iceland (2)	44.499*	2.988		1	2000
Israel (1)	28.293*	14.181*	3.784	2	1995 2002
Mexico (1)	109.536*	14.603*	4.087	2	1997 2003
Norway (3)	12.173*	3.595		1	1997
South Africa (1)	14.457*	9.433		1	1989
South Korea (1)	13.540*	50.461*	5.461	2	1991 1997
Sweden	5.792			0	none
Switzerland (1)	20.061*	6.762		1	2008
UK (2)	49.811*	22.954*	2.879	2	1989 2008
United States (1)	50.578*	2.088		1	2008

Note: (1) = break in mean and trend; (2) = break in mean and non breaking trend; (3) = break in mean only. The 5-percent critical values for the SupF(ℓ+1/ℓ) are 8.58, 10.13, and 11.14 for ℓ = 1,2,3, respectively, for the model with break in the mean or the model with break in the mean with no-breaking trend, and 11.47, 12.95, and 14.03 for ℓ = 1,2,3, respectively, for the model with a break in the mean and a break in the trend (Bai and Perron, 2003). The maximum number of breaks is set at M=3 and a trimming parameter of 0.15 is used to determine the minimal number of observations in each segment h = [0.15T] where T is the sample size. We allow error distributions to differ across breaks. We test for three different specification (see Table 14): break in mean and trend, break in mean with a non-breaking trend, and break in mean without a trend.

Table 4: Estimates of the Bai-Perron structural breaks

Country	Constant			Trend			Adj R^2
	Regime 1	Regime 2	Regime 3	Regime 1	Regime 2	Regime 3	
Canada	0.077	0.246		0.091	-0.007		0.447
	(0.912)	(3.541)		(3.268)	(-1.822)		
Chile	0.638	0.718	-0.549	-0.072	-0.035	0.038	0.817
	(14.546)	(6.965)	(-2.968)	(-7.839)	(-4.731)	(4.774)	
Iceland	0.298	0.007		0.015	0.015	0.015	0.613
	(12.191)	(0.126)		(5.859)	(5.859)	(5.859)	
Israel	0.453	-0.276	0.434	0.004	0.045	-0.003	0.759
	(25.361)	(-1.586)	(5.864)	(1.457)	(3.428)	(-0.997)	
South Africa	0.101	-0.214		-0.094	0.034		0.883
	(1.382)	(-4.509)		(-2.451)	(12.216)		
So. Korea	0.268	0.321	0.649	0.061	-0.024	-0.019	0.671
	(6.087)	(5.395)	(5.271)	(4.224)	(-3.599)	(-3.214)	
Mexico	0.575	0.236	1.646	0.035	0.046	-0.051	0.868
	(20.751)	(0.875)	(6.083)	(8.371)	(2.492)	(-4.325)	
Norway	0.344	0.147					0.258
	(11.549)	(3.049)					
Sweden	0.105	0.462	-0.446	0.013	-0.019	0.028	0.318
	(2.663)	(5.385)	(-0.994)	(0.657)	(-3.159)	(1.608)	
Switzerland	0.281	1.491		-0.007	-0.046		0.404
	(8.621)	(2.164)		(-2.912)	(-1.693)		
United Kingdom	0.277	0.486	0.929	-0.022	-0.022	-0.022	0.741
	(42.038)	(10.104)	(10.502)	(-6.668)	(-6.668)	(-6.668)	
United States	0.636	0.239		-0.031	0.002		0.728
	(12.872)	(1.441)		(-7.891)	(0.361)		

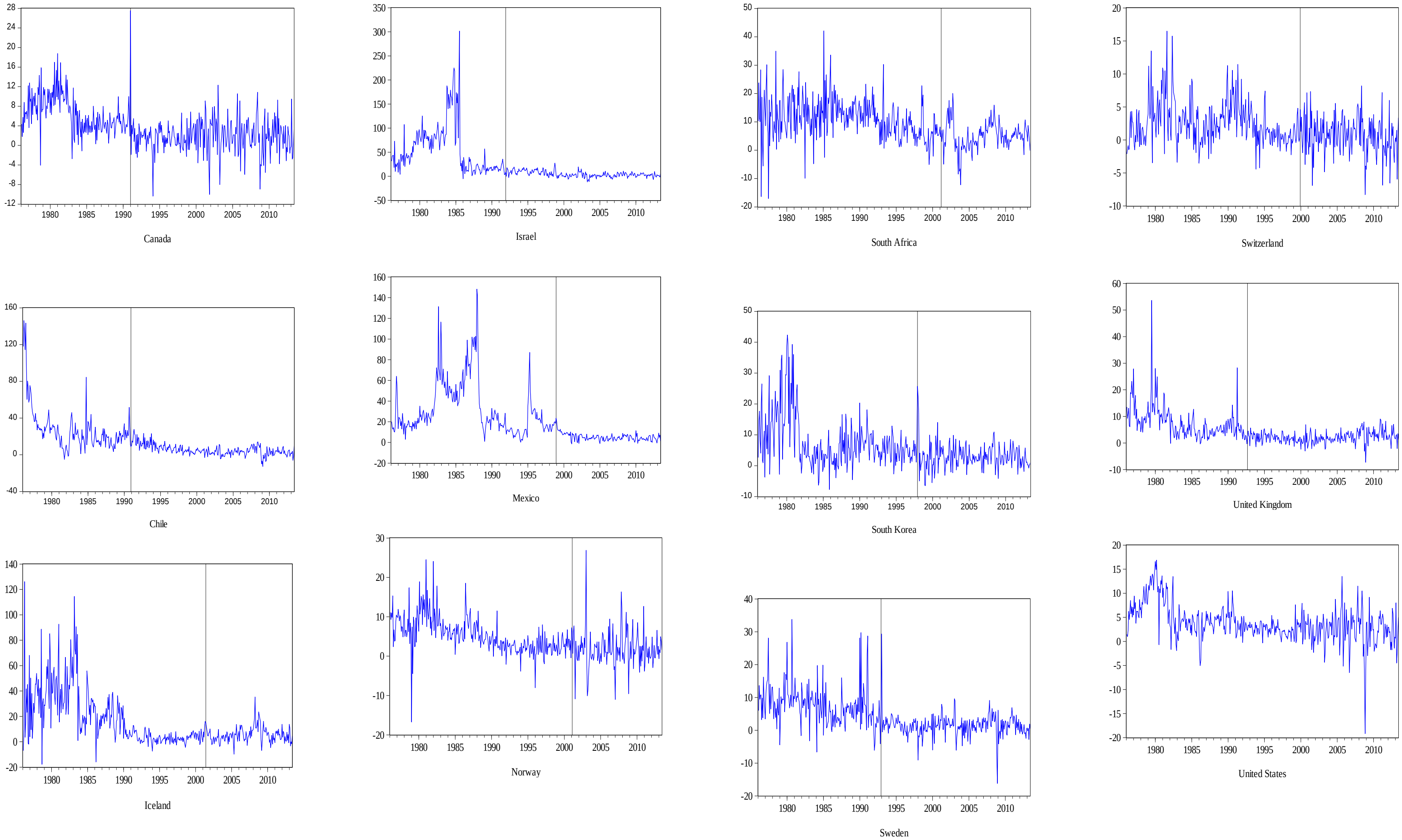
Note: We consider three alternative models -- break in mean and trend, break in mean with a non-breaking trend, and break in mean without a trend. The Table reports the findings corresponding to the model with the highest goodness of fit (Adj. R^2). The break dates are reported in Table 3, except for Sweden. Sweden is estimated using the sequentially fixed number of breaks approach.. This method delivers two breaks: 1989 and 2008.

Table 5: Phillips log periodogram estimates of the Bai-Perron subsamples

Country	Regime	$\hat{d}_{MLP}$	$se(\hat{d}_{MLP})$	$H_0 : d = 0$	$H_0 : d = 1$
Canada	1976:2-1992:1	0.209	0.104	2.015	-10.091
	1992:2-2013:6	0.162	0.094	1.716	-11.973
Chile	1976:2-1995:1	0.736	0.091	8.183	-3.578
	1995:2-2005:1	0.162	0.121	1.346	-8.855
	2005:2-2013:6	0.529	0.124	4.245	-4.638
Iceland	1976:2-2001:1	0.184	0.117	1.564	-6.966
	2001:2-2013:6	0.468	0.092	5.071	-6.094
Israel	1976:2-1996:1	0.534	0.085	6.263	-6.493
	1996:2 -2004:1	0.471	0.129	3.631	-5.088
	2004:2-2013:6	0.366	0.101	3.623	-6.481
Mexico	1976:2-1998:1	0.772	0.081	9.433	-3.296
	1998:2-2004:1	0.586	0.132	4.418	-3.529
	2004:2-2013:6	0.311	0.118	2.629	-7.047
Norway	1976:2-1998:1	0.339	0.086	3.936	-9.547
	1998:2-2013:6	-0.122	0.119	-1.022	-14.109
South Africa	1976:2-1990:1	-0.126	0.124	-1.012	-13.601
	1990:2-2013:6	0.451	0.099	4.524	-8.116
South Korea	1976:2-1992:1	0.378	0.106	3.569	-7.935
	1992:2-1998:1	0.434	0.123	3.512	-4.830
	1998:2-2013:6	0.067	0.934	0.714	-11.732
Sweden	1976:2-1989:1	0.175	0.113	1.538	-9.626
	1989:2-2008:1	0.278	0.088	3.139	-9.804
	2008:2-2013:6	0.302	0.113	2.665	-5.753
Switzerland	1976:2-2009:1	0.371	0.066	5.599	-10.687
	2009:2-2013:6	0.081	0.127	0.632	-6.877
United Kingdom	1976:2-1990:1	0.390	0.094	4.147	-7.365
	1990:2-2009:1	0.295	0.068	4.314	-9.581
	2009:2-2013:6	0.415	0.237	1.752	-4.369
United States	1976:2-2009:1	0.499	0.072	6.923	-8.511
	2009:2-2013:6	0.414	0.178	2.326	-4.378

Note: The table reports the regime-specific estimates of inflation persistence ($\hat{d}_{MLP}$), their associated standard errors ($se(\hat{d}_{MLP})$), and the tests statistics of for the null of unit root ($H_0 : d = 1$) and the null of short memory ($H_0 : d = 0$).

Figure 1: Inflation rates

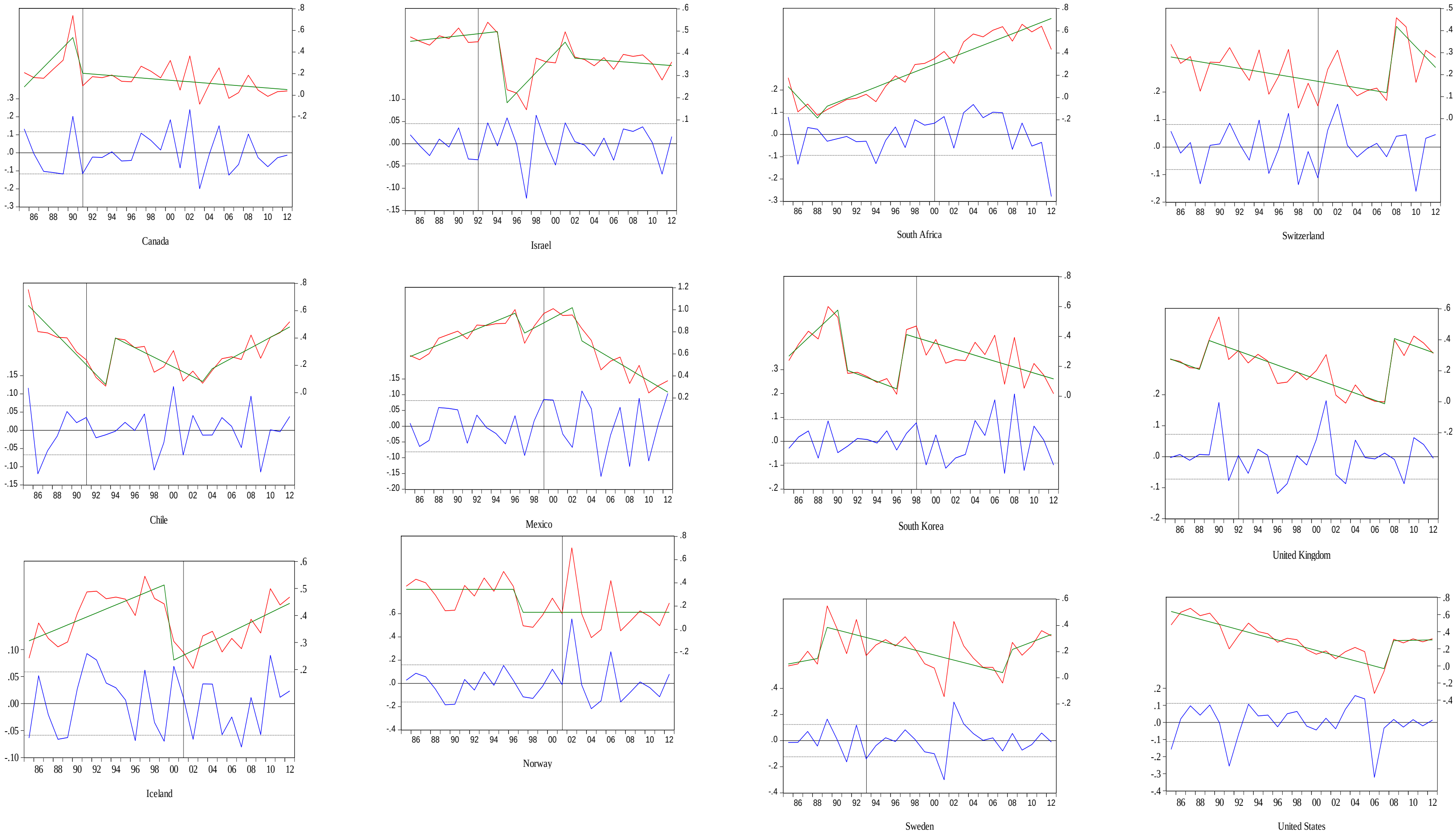


Note: Vertical line denotes IT adoption date.

Figure 2: Phillips modified log periodogram regression estimates (95% confidence band based on 10-year rolling window)

Note: Vertical line denotes IT adoption date.

Figure 3: Bai-Perron structural changes in the inflation persistence series.



Note: Vertical line denotes IT adoption date