

Inflation Targeting: New Evidence from Fractional Integration and Cointegration

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Abstract

We investigate inflation persistence in six inflation targeting (IT) countries from the global-economy perspective. This view maintains that inflation persistence in IT countries has declined mainly because of the decline of inflation persistence in the global economy. We provide empirical evidence on two yet unanswered questions. First, we investigate whether each IT country in the sample share a common persistence with Germany and the US, two non-IT countries with a relatively good inflation record, which we use as proxies for the global economy. This tests the weak-form global hypothesis. Second, for the countries that share common inflation persistence with Germany and the US, we examine whether the same long memory component drives their inflationary processes to converge in the long-run to a common stochastic equilibrium with Germany and the US. This tests the strong-form global hypothesis. Our findings cast doubt on the relevance of IT in the industrial, but not developing, countries in the sample, suggesting that the global economy probably played an important role in the decline of inflation persistence in industrial, but not developing, countries.

Keywords: inflation targeting, inflation persistence, fractional integration, cointegration

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1. Introduction

A great deal of recent empirical research has focused on inflation persistence.¹ A fair amount of empirical evidence, although not yet conclusive, suggests that inflation persistence has significantly declined in countries, especially industrial countries² that have adopted inflation targeting (IT).³ A partial list of such research includes Bratsiotis, *et al.* (2015), Baxa, *et al.* (2014), Canarella and Miller (2015, 2016), Gerlach and Tillman (2012), Yigit (2010), Benati (2008), Siklos (2008), Mishkin and Schmidt-Hebel (2007), and Kuttner and Posen (2001). Thus, researchers evaluate the success of IT, among other things, on its effect on inflation persistence. That is, a credible IT regime affects inflation expectations, resulting in fundamental changes in the stochastic process governing the dynamics of inflation. The literature frequently notes that changes in inflation expectations and changes in inflation controllability lead, in the final analysis, to a decline in inflation persistence (Chiquiar, *et al.*, 2010; Alogoskoufis, 1992; Agenor and Taylor, 1993). This theoretical paradigm, however, confines the empirical analysis exclusively to countries that have adopted IT and, in the empirical domain, simply relies on isolated tests of inflation persistence before and after IT implementation. This approach possesses certain ambiguities.

Two separate strands of literature converge to question the validity of adopting a closed-economy specification for policy analysis. On the one hand, the global inflation literature (e.g.,

¹ Inflation persistence means that the inflation rate converges slowly towards its long run equilibrium following a shock that pushes it off its equilibrium. Batini and Nelson (2001, 2002) offer three working definitions of inflation persistence: (a) positive serial correlation in inflation; (b) lags between systematic monetary policy actions and their peak effect on inflation; and (c) lagged response of inflation to non-systematic policy actions (i.e., policy shocks).

² The literature provides more ambiguous results about the effects of IT on inflation persistence in developing countries (Arize, *et al.*, 2005; Capistrán and Ramos Francia, 2006; Goncalves and Salles, 2008; Lin and Ye, 2009; Brito and Bystedt, 2010; and Gerlach and Tillman, 2012).

³ Bernanke, *et al.* (1999) describes the IT framework in detail. Amato and Gerlach (2002) note that the fundamental criteria for inflation targeting are (1) the public announcement of a numerical target for inflation, (2) the commitment to price stability as the overriding goal of monetary policy, (3) the use of an information-inclusive strategy, and (4) the adoption of transparency and accountability in monetary policy actions.

White, 2008; Hakkio, 2009; Ciccarelli and Mojon, 2010; Duncan and Martinez-Garcia, 2015) documents that the inflation process is a global phenomenon. During the last three decades, the world has experienced a remarkable process of disinflation, with average inflation rates in industrialized countries falling by ten percentage points with even sharper declines in developing countries. Inflation in industrial economies began to fall in the early 1980s while inflation in emerging economies began to fall in the 1990s. See, for example, Rogoff (2003) for details. Parallel to the decline in inflation rates, the forces of globalization of capital markets created common global linkages through commodities, trade, and finance, which increased the correlations between cross-countries inflation rates.

On the other hand, the bilateral country-pairing literature (Dueker and Fischer, 1996, 2006) argues that conclusions based solely on one-sided evidence from IT countries can produce misguided, or at least difficult to evaluate, evidence, since many non-IT countries also shared similar experiences. For instance, quite persuasive evidence exists that inflation persistence also declined in non-IT countries, mainly the US and the Euro Area (Meller and Nautz, 2012; Beechey and Österholm, 2009; Kumar and Okimoto, 2007; Benati, 2008; Gadea and Mayoral, 2005; Whelan and O'Reilly, 2005). Groeneveld, *et al.* (1998), Lee (1999), and Lin and Ye (2007) conclude that the effectiveness of the explicit IT approach does not clearly dominate alternative monetary strategies, since it may only be an artifact of the global disinflationary environment.

Does the success of IT simply reflect the global nature of the disinflationary process? This paper extends the available evidence on inflation persistence for a sample of IT countries by testing the validity of the global hypothesis. We analyze the behavior of inflation persistence in IT countries relative to the performance of inflation persistence in the global economy. Unlike

the bilateral country-pairing strategy employed by Dueker and Fischer (1996; 2006), however, which pairs Canada with the US, the UK with Germany, and Australia with New Zealand, we pair each IT country in our sample with the global economy, which we represent by Germany, the largest economy of the Euro Area, and the US, the world's largest economy. Germany and the US both enjoy a history of monetary policies deeply focused on controlling inflation, but have not made any explicit commitment to IT, and, thus, are not regarded as IT countries (Bernanke and Mishkin, 1997; Dueker and Fisher, 1996, 2006; Lee, 1999).

Empirically, we can distinguish two separate testable hypotheses regarding the global hypothesis -- the weak- and strong-form global hypotheses. Under the weak form, the global hypothesis requires that IT countries share a common persistence with Germany and the US. Their response to inflationary shocks should not differ from the response of Germany and the US. Under the strong form, the global hypothesis requires that IT countries that share a common persistence converge in the long-run to a common stochastic equilibrium with Germany and the US. We examine these issues from the perspective of both three developing IT countries (Chile, Israel, and Mexico) and three industrial IT countries (Canada, Sweden, and the UK). We choose only six IT countries because of the length of their IT experiences and their frequent use in previous empirical work.

We propose a multivariate empirical methodology based on the fractional integration approach. Following the recent empirical literature, we model inflation as a fractionally integrated process of order d , denoted by $I(d)$, and estimate inflation persistence by the fractional integration parameter d . An $I(d)$ process captures the long-range dependence, or long memory, of the data.⁴ Substantial empirical evidence suggests that neither $I(1)$ nor $I(0)$ processes

⁴ The degree of fractional integration d determines the persistence of $I(d)$. Fractionally integrated processes possess different characteristics depending on the value of d . In particular, when $d = 0$, then $I(d)$ is a short-memory

appropriately model inflation (Canarella and Miller, 2015, 2016; Meller and Nautz, 2012; Caporale and Gil-Alaña, 2011; Yigit, 2010; Kumar and Okimoto, 2007; Gadea and Mayoral, 2006). Importantly, $I(d)$ processes avoid the knife-edged dichotomy between $I(0)$ and $I(1)$ processes of the traditional time domain approach, which fails to separate the non-stationarity property of a time series from its mean-reversion property. Fractionally integrated processes allow the examination of more flexible patterns of responses to idiosyncratic shocks, in particular, long memory or non-stationarity with non-permanent, but slowly decaying shocks.

The degree of fractional integration d possesses substantial policy implications. For instance, the appropriate policy response to inflationary shocks depends on the degree to which their effect on inflation persists. As noted by Meller and Nautz (2012), if inflation persistence is too high, then shocks to inflation exert long-lived effects. As such, the lasting effect of shocks will make the central bank control of inflation much more difficult, thus, impeding the effectiveness of monetary policy, since it takes longer to achieve the policy goals. On the contrary, if inflation persistence is sufficiently low, steering inflation expectations becomes more effective and policy can quickly bring inflation expectations in line with its equilibrium. At one extreme, inflation may behave like a random-walk, $I(1)$, process, where the central banks cannot control it. At the other extreme, inflation may behave like a stationary, $I(0)$, process, where the inflation rate quickly reverts to its initial level. As such, central banks do not even need to respond to an inflation shock.

stationary process, with autocorrelations decaying geometrically. In contrast, when $0 < d < 1$, $I(d)$ is a fractionally integrated process, a long-memory process, characterized by long-range positive dependence and autocorrelations decaying hyperbolically. See Baillie (1996) and Granger and Joyeux (1980) for more details. Furthermore, when $0 < d < 0.5$, $I(d)$ is a mean-reverting stationary process, while when $0.5 \leq d < 1$, $I(d)$ is non-stationary, but still mean-reverting. In contrast, when $d \geq 1$, $I(d)$ is both non-stationary and non-mean-reverting. The unit-root case coincides with $d=1$. Finally, when $-0.5 < d < 0$, $I(d)$ is stationary, but anti-persistent, indicating that the process reverses itself more frequently than a random process.

Our empirical results suggest that long-memory and mean-reversion are prevalent features of the inflationary processes in the IT countries, Germany, and the US. We find that the inflationary processes in the industrial IT countries (Canada, Sweden, and the UK) exhibit stationarity and share a common degree of fractional integration with Germany and the US. Thus, they respond in a similar way to inflationary shocks as Germany and the US. Conversely, we find that the inflationary processes in the developing IT countries (Chile, Israel and Mexico) exhibit non-stationarity and do not share a common degree of fractional integration with Germany and the US. That is, their response to inflationary shocks takes much longer than in Germany and the US. For Canada, Sweden, and the UK, where we do not reject common inflation persistence with Germany and the US, we examine whether the same long-memory component drives these processes to converge in the long-run to a common stochastic equilibrium (i.e., they exhibit fractional cointegration).⁵ Integer cointegration techniques can only test the existence of a long-run relationship for I(1) series that are cointegrated, leading to an I(0) error correction term, which is not persistent. Recent studies, however, show that an I(d), fractionally integrated, rather than I(0), error-correction term may emerge in the cointegrating regression. This is particularly relevant for non-I(1) variables. The presence of fractional cointegration indicates that the inflationary processes converge in the long run to a common fractional stochastic equilibrium. This result implies that deviations from the long-run relationship shared by the series in question take a long time to dissipate before reaching their equilibrium level. In the post-IT sample, we find strong evidence of fractional cointegration for Canada, weak evidence for Sweden, and no evidence for the UK and the developing countries.

⁵ Fractional cointegration merges the methodology of cointegration (Granger, 1986) with the concepts of fractional integration. Importantly, fractional cointegration associates the existence of a long-run relationship with mean reversion in the error term, rather than requiring, as in integer cointegration, both mean reversion and stationarity. The presence of a long-run equilibrium relationship with d less than one implies that nonlinearities exist and that the equilibrium error may still exhibit mean-reversion.

We conjecture that, notwithstanding the IT regime, there exist common forces, such as two decades of favorable inflationary trends affecting the global economy, driving inflation persistence in Canada and Sweden, as well as in Germany and the US. These positive global developments (Ducker and Fisher, 1996; Ball and Sheridan, 2005) that benefitted Germany and the US may also have benefitted these industrial IT countries. This result diminishes the relevance, in our view, of the IT regime as an explanatory factor in the decline of inflation persistence in Canada and Sweden.

The rest of the paper is organized as follows. Section 2 describes the data. Section 3 analyzes the main empirical findings. Section 4 concludes.

2. Data and descriptive statistics

Our data consist of monthly observations on the seasonally unadjusted Consumer Price Index (CPI all item; 2010 = 100) in Canada, Chile, Germany, Israel, Mexico, Sweden, the UK, and the US. These data come from OECD *Main Economic Indicators: Historical Statistics*. The sample starts in 1976:01 and ends in 2013:06. We seasonally adjust the data using the Census X12 method, and compute the inflation rates as the monthly logarithmic difference of the seasonally adjusted CPI in each country, expressed as a percentage. The sample includes six OECD IT countries, Canada (1991:02), Chile (1991:01), Israel (1992:01), Mexico (1999:01), Sweden (1993:01), the United Kingdom (1992:10), and two non IT countries, Germany and the US. The adoption dates of IT appear in parentheses, and come from Bernanke, *et al.* (1999), Mishkin and Schmidt-Hebbel (2001), and Fracasso, *et al.* (2003). This sample of countries represents a heterogeneous profile of inflation. Canada and Sweden, for example, never experienced hyperinflation. In the 1970s, Chile saw an episode of hyperinflation (1973). In the 1980s and 1990s, inflation largely became a problem of developing countries. Israel in 1984-1985 and

Mexico in 1982-1993 experienced hyperinflation. We do not consider OECD countries where IT exhibits a short history (Finland, Spain), where IT was introduced relatively recently (South Korea, Norway, South Africa), or where countries do not publish monthly data (New Zealand, Australia).

Table 1 reports summary statistics of the IT countries for the entire sample and for the pre- and post-IT periods (Panel A) and for Germany and the US (Panel B) for the entire sample period and for pre- and post-1990 periods.⁶ Several observations emerge. First, for all IT countries, inflation remains more stable and less volatile under the IT regime than in previous periods. For example, Table 1 shows that the Canadian inflation rate declined since the adoption of IT. Before IT adoption, the average inflation was 6.61% with standard deviation equal to 4.07. Since IT adoption, however, the average inflation rate declined to 1.80% with a standard deviation of 3.28.

Second, the decline of the level and volatility of inflation also occurs for Germany and the US since the 90s, which coincides approximately with IT adoption. The evidence in Table 1, as Ball and Sheridan (2005) and Walsh (2009) note, however, does not constitute evidence of any causal link between IT and these improved inflation outcomes. That is, all OECD countries enjoyed lower and more stable inflation rates (see Walsh, 2009).

3. Empirical results

Our empirical analysis adopts a multi-step approach. First, we perform unconstrained estimation of the inflation series for an IT country, Germany, and the US. In this context, we utilize the panel log-periodogram estimator proposed by Robinson (1995), a multivariate

⁶ Because Germany and the US did not adopt IT, no natural break point for measuring the pre- and post-experience exists, and so the choice of 1990 is arbitrary. But it serves to illustrate how the era of IT adoption coincides with an era of lower and more stable inflation for both IT and non-IT countries. We clearly see this stylized fact in Table 1.

generalization of the Geweke and Porter-Hudak (1983) semiparametric procedure.⁷ This estimator incorporates the cross-sectional dependence of the series and explicitly models panel heterogeneity, since the fractional integration parameters can vary across countries. This exercise produces six distinct panels: (1) Canada, Germany, US; (2) Chile, Germany, US; (3) Israel, Germany, US; (4) Mexico, Germany, US; (5) Sweden, Germany, US; and (6) UK, Germany, US.

Second, we test each panel for a common order of fractional integration using a Wald-type test developed by Robinson (1995). This tests the weak-form global hypothesis. That is, we conduct a test of equality of inflation persistence between the inflation series of each IT country, Germany, and the US. The null hypothesis is $H_0 : d(IT) = d(US) = d(G)$, which we test against the two-sided alternative.

Third, we perform constrained estimation of the inflation series to generate estimates of common persistence for the panels where we do not reject the null hypothesis $H_0 : d(IT) = d(US) = d(G)$. That is, we estimate the common inflation persistence parameter by imposing the equality constraint $d(IT) = d(US) = d(G)$., which under the accepted hypothesis $H_0 : d(IT) = d(US) = d(G)$ implies a gain in efficiency.

Finally, using a generalized version of the ARFIMA model we examine whether fractional cointegration exists in the panels and whether this form of cointegration associates with long memory. This tests the strong-form global hypothesis.

⁷ In the empirical fractional integration literature, a few other papers employ the Robinson (1995) multivariate estimator. Bollerslev and Jubinski (1999) investigate the common long-run dependencies of trading volume and volatility. Andersen, *et al.* (2003) analyze the long-memory behavior of realized return volatilities of assets. Meller and Nautz (2012) analyze inflation persistence before and after the European Monetary Union. Michelacci and Zaffaroni (2000) estimate long-memory parameters for GDP per capita of a group of OECD countries.

3.1 Unrestricted Robinson estimates of inflation persistence

As previously mentioned, we first perform unrestricted estimation of inflation persistence in each of the six panels using the Robinson (1995) estimator. A brief summary of the estimator follows. For more details, see Robinson (1995).

Let $X(t)$ denote the G dimensional vector with g^{th} element $X_g(t)$, $g = 1, \dots, G$. If each series is fractionally integrated, it follows that at frequencies λ close to 0, $f_{gg}(\lambda) \approx C_g \lambda^{-2d_g}$, where $f_{gg}(\lambda)$ is the spectral density at frequency λ , $g = 1, \dots, G$, satisfying the conditions $0 < C_g < \infty$ and $-0.5 < d_g < 0.5$. The parameters d_g , $g = 1, \dots, G$, are the degrees of fractional integration of $X_g(t)$, $g = 1, \dots, G$. The sample periodogram for $X_g(t)$, $t = 1, \dots, T$, where T is the sample size, is denoted by

$$I_g(\lambda) = \frac{1}{2\pi T} \left| \sum_{t=1}^T X_g(t) e^{it\lambda} \right|^2, \quad g = 1, \dots, G. \quad (4)$$

Robinson (1995) introduces the trimming parameter l such that the number of frequencies varies from $l+1$ to m rather than from 1 to m .⁸ Furthermore, Robinson (1995) shows that asymptotic efficiency improves by pooling J adjacent frequencies and estimating d from the modified periodogram ordinates.⁹ Following Robinson (1995), the log-periodogram at the k th Fourier frequency is given by

$$Y_{gk}^J = \log \left(\sum_{j=1}^J I_g(\lambda_{k+j-J}) \right), \quad g = 1, \dots, G, \quad k = l+J, l+2J, \dots, m, \quad l \geq 0, \quad J \geq 1, \quad (5)$$

⁸ The trimming parameter l specifies the number of initial ordinates excluded from the log-periodogram regression. Kunsch (1987) finds that such exclusion improves the properties of tests based on log-periodogram regressions. Hurvich, *et. al* (1998), however, show that the log-periodogram regression maintains desirable properties if all frequencies from 1 up to m are used.

⁹ If $J = 1$ is specified, no averaging is performed. With $J = 2$, the number of ordinates is halved; with $J = 3$, the number of ordinates is divided by three; and so on. For $J = 1$ and $l = 0$, the Robinson estimator coincides with the Geweke and Porter-Hudak estimator.

where J is fixed and $m - l$ is a multiple of J .

Robinson (1995) suggests the following log-periodogram regression

$$Y_{gk}^J = c_g^J - d_g 2 \ln \lambda_k + U_{gk}^J, \quad g = 1, \dots, G, \quad k = l+J, l+2J, \dots, m, \quad (6)$$

where U_{gk}^J is the vector of errors associated with $Y_{gk}^J = \log \left(\sum_{j=1}^J I_g(\lambda_{k+j-J}) \right)$. Then the OLS estimates

of $c^J = (c_1^J, \dots, c_G^J)'$ and $d^J = (d_1^J, \dots, d_G^J)'$ are given by \tilde{c}^J and \tilde{d}^J

$$\begin{bmatrix} \tilde{c}^J \\ \tilde{d}^J \end{bmatrix} = \text{vec}((Y^J)' Z^J ((Z^J)' Z^J)^{-1}), \quad (7)$$

where $Z^J = (Z_{l+J}, Z_{l+2J}, \dots, Z_m)$, $Y^J = (Y_1^J, \dots, Y_G^J)$, $Z_k = (1, -2 \ln \lambda_k)'$, and

$Y_g^J = (Y_{g,l+J}^J, Y_{g,l+2J}^J, \dots, Y_{g,m}^J)$ for m periodogram ordinates.

The Robinson (1995) estimator requires the specification of the trimming parameter l , the determination of the averaging of the periodogram over J frequencies, and the choice of the bandwidth parameter m . The empirical literature (e.g., Hassler and Olivares, 2013) suggests that researchers should estimate d with different values of m (or powers of T) to alleviate the concern that the results depend on the choice of the bandwidth. The bandwidth choice problem does not come with an easy solution, since it entails a bias-variance trade-off (Henry and Robinson, 1996). That is, the smaller the bandwidth m , the less likely that higher frequency dynamics (i.e., the short memory of the ARMA component) will contaminate the estimate of d . Smaller bandwidths, however, lead to smaller sample sizes and less reliable estimates.

We estimate each panel using $l = 0$ (Meller and Nautz, 2012) and a grid of values for J ($= 1, 2, 3, 4$) in combination with a grid of values for α ($= 0.70, 0.75, \text{ and } 0.80$). This range of values of α conforms to the empirical literature on inflation persistence (e.g., Canarella and

Miller, 2015; Kumar and Okimoto, 2007; Meller and Nautz, 2012; Kruse and Sibbertsen, 2010; Hassler and Wolters, 1995; Hassler and Meller, 2011; Hsu, 2005). Furthermore, since we employ monthly data, using much smaller bandwidths would produce standard errors too large to provide any meaningful information on the order of integration. Our empirical results, however, remain reasonably robust to changes in J and α . In particular, the differences across J for $\alpha = 0.70$, 0.75, and 0.80, as well as the differences across α for $J = 1, 2, 3, 4$ produce relatively minor differences and, in general, do not invalidate the qualitative characteristics of the series. Unreported results are available on request.

Table 2 summarizes the unrestricted Robinson estimates of the fractional integration exponents $d(\text{IT})$, $d(\text{G})$, and $d(\text{US})$ in each of the six panels for $\alpha = 0.70$ and $J = 2$. The Table reports results for the entire sample period and for the pre- and post-IT periods.¹⁰ The numbers in parentheses below the estimated parameters contain the t -statistics for the null hypothesis $H_0 : d = 0$ against the alternative hypothesis $H_1 : d > 0$. Several important findings emerge from Tables 2. The unrestricted estimates of the fractional integration parameter for the full sample overwhelmingly support the long-memory ($d > 0$) and mean-reversion ($d < 1$) properties of the inflation series, which implies that inflationary shocks dissipate at hyperbolic rates. Also, the overall evidence implies that the inflation series do not emerge from strict unit-root processes ($d = 1$), as the fractional integration parameters lie more than two standard errors below. Importantly, the evidence against $I(0)$ does not reflect a neglected deterministic trend. Second,

¹⁰ The log-periodogram estimator requires that the data do not contain deterministic components. If trends exist in the data, then the estimates can be biased into the non-stationary region (Sibbertsen, 2002). Preliminary analysis reveals that all inflation series contain significantly negative linear trends. Fractional integration does not eliminate the trend. Therefore, to satisfy the requirement of the log-periodogram approach, we detrend each of the inflation series by subtracting a country-specific local trend (see Shimotsu, 2006; Phillips, 2007; Bollerslev and Jubinski, 1999; Hassler, *et al.*, 2006). This filtering procedure also provides the additional advantage of preventing the trend from dominating the long-run dynamics of the inflation series. By controlling for the presence of low-frequency swings in the raw inflation data, it determines more appropriately the long-range dynamics of the series.

significant differences exist in the degree of mean-reversion among the IT countries. Inflation in the industrial countries -- Canada, Sweden, and the UK -- exhibits much less persistence than in the developing countries -- Chile, Israel, and Mexico. For Chile, Israel, and Mexico, the fractional integration parameter estimate exceeds 0.5, suggesting long-memory mean-reverting, but non-stationary process with infinite variance. Conversely, for industrial IT countries, the estimate of the fractional integration parameter falls in the open interval less than 0.5, suggesting long-memory, mean-reverting, but stationary process with finite variance. Solid evidence also exists that inflation in Germany and the US is a stationary, long-memory, mean-reverting process with finite variance.

Do the full-sample findings hold also for the pre- and post-IT period?¹¹ The unrestricted estimates of the fractional integration parameter for countries in the pre-IT period lie between zero and one, which is evidence of long memory ($d > 0$) and mean reversion ($d < 1$). As in the full-sample estimates, however, the degree of mean reversion varies substantially across the six IT countries. Inflation in Canada, Sweden, and the UK exhibits much less persistence than in Chile, Israel, and Mexico. For Chile, Israel, and Mexico, the pre-IT estimates of inflation persistence exceed 0.5, but fall below 1, by more than two standard errors, implying that inflation in these IT countries does not conform to a strict unit-root process, but to a long-memory, mean-reverting, non-stationary process. Conversely, for Canada, Sweden, and the UK, the pre-IT estimates are less than 0.5. For Canada and the UK, however, inflation exhibits a long-

¹¹ Sub-sample analysis provides a practical solution to several important empirical issues, such as exogenous changes in the economic system and structural instability of the coefficients of monetary policy rules (see Clarida, *et al.*, 1998, 2000). This approach, however, assumes that we know the timing of structural breaks, and that monetary policy setting does not evolve within each sample period. As such, this approach transfers the time-invariance assumption of the estimates from the full sample to the sub-samples. For IT countries, however, this approach seems plausible and adequate because monetary policy with respect to inflation and other macroeconomic variables change as a consequence of the implementation of IT. Endogenous break-point estimation is not straightforward in a fractional integration context (see Sibbertsen, 2004). The Bai-Perron methodology only applies to I(0) data.

memory, mean-reverting, stationary process, while for Sweden, it exhibits a short-memory stationary process, i.e. $I(0)$. Good evidence also exists that inflation in Germany and the US exhibits long-memory, mean-reverting, and stationary process.

The post-IT period estimates display significant systematic changes relative to the pre-IT estimates. The long memory ($d > 0$) and mean-reversion ($d < 1$) properties still exist for the developing countries. Israel and Mexico display estimates still significantly greater than 0.5, although less than 1, while Chile display a persistence not significantly different from 0.5. This implies that inflation, even after the adoption of IT, remains a non-stationary process in Chile, Israel, and Mexico. In contrast, Canada and Sweden, but not the UK, display estimates significantly less than 0.5. Thus, in the post-IT sample, we find strong evidence of stationarity for Canada and Sweden, weak evidence of stationarity for the UK, and strong evidence of non-stationarity for Chile, Israel, and Mexico. Noticeably, inflation in the developing countries exhibits more persistence than in the industrial countries both before and after IT implementation. In addition, we observe a significant decline in inflation persistence for Chile and Mexico in the post-IT sample compared to the pre-IT sample, but not in Israel. Among the industrial countries, only Canada experiences a decline in inflation persistence. Sweden experiences a switch from short-memory in the pre-IT sample to long-memory in the post-IT sample. Conversely, Germany and the US experience a substantial decline in inflation persistence in the post-IT period in comparison to the pre-IT sample.

Overall, these findings suggest that substantial heterogeneity exists in the stochastic processes of inflation between developed and developing countries, both before and after the IT adoption, and, most importantly, across the IT countries even after the IT adoption, which indicates that IT does not equalize inflation persistence. Although inflation exhibits a stationary

process in all six countries in both the pre- and post-IT periods, a shift from non-stationarity to stationarity does not occur in the developing countries as a result of the IT adoption, although some decline in inflation persistence did occur for Chile and Israel. For the industrial IT countries, only Canada experiences a significant decline in inflation persistence. In contrast, Sweden and the UK experience an increase in inflation persistence. Inflation in Sweden shifts from $I(0)$ in the pre-IT period to long memory in the post-IT period, while in the UK, inflation persistence increases across the two regimes, but remains within the long-memory stationary range. Germany and the US successfully reduced inflation persistence despite no explicit IT regime.¹²

Our findings provide some noteworthy policy implications. Previous research argues that the implementation of a stable monetary policy regime with a well-defined nominal anchor such as inflation targeting contributes to a decrease of inflation persistence. Although Chile, Israel, and the UK adopted IT more than two decades ago and Mexico almost two decades ago, inflation persistence did not change dramatically in these countries and remained at high levels when compared to Canada and Sweden, a sign that, at least partially, the IT policy did not successfully anchor inflation expectations in those countries.

3.2 Tests of the weak-form global hypothesis

Now, we test whether the inflation series in each trivariate panel share a common degree of fractional integration (i.e., show similar long-memory characteristics and share the same stochastic properties). As previously noted, Robinson (1995) developed a formal Wald-type test of the hypothesis of equality of fractional integration orders. In our case, we test the equality of

¹² Because the analysis implicitly assumes that Germany and the US share a common inflation persistence, we also re-estimated each panel under the constraint $d(\text{US})=d(\text{G})$. The results indicate that only small parameter variations occur. Results are available on request.

the fractional orders of the series in each of the six panels (i.e., $H_0 : d(IT) = d(G) = d(US)$) by applying to equation (10) the linear homogeneous restriction $H_0 : PD = 0$

$$\text{where } P = \begin{bmatrix} 1 & -1 & 0 \\ 1 & 0 & -1 \end{bmatrix} \text{ and } D = \begin{bmatrix} d(IT) \\ d(G) \\ d(US) \end{bmatrix}.$$

Under the null hypothesis $H_0 : PD = 0$, the Wald-type statistic

$$W = (\hat{d}^J)' P' \left[(0, P) \left\{ (Z^J)' Z^J (\tilde{\Omega}^J)^{-1} \right\} \begin{pmatrix} 0 \\ P' \end{pmatrix} \right]^{-1} P \hat{d}^J \quad (10)$$

is asymptotically distributed as a chi-square with 2 degrees of freedom. Failure to reject the null provides evidence of equality of the orders of fractional integration of the inflation series.

Table 3 report the Wald-type statistics for the full sample and the pre- and post-IT sample periods. For Canada, Sweden and the UK, we fail to reject the null hypothesis of common persistence at the 5-percent level in all three sample periods. On the other hand, for Mexico, we reject the null hypothesis of common persistence at the 5-percent level in all three sample periods. For Chile and Israel, we reject the hypothesis of common persistence at the 5-percent level for the full sample and the post-IT sample, but fail to reject it in pre-IT sample.

In sum, the Wald-type test rejects at the 5-percent level the null hypothesis of common inflation persistence for Chile, Israel, and Mexico for the entire sample period and, most importantly, for the post IT-period. These findings indicate that over the entire sample and in the post-IT sample, the inflationary processes of Chile, Israel, and Mexico possess time-series properties that differ substantially from the inflation processes of Germany and the US. Their response to shocks, although still hyperbolic in nature, take longer than the response of Germany and the US. In contrast, we fail to reject the hypothesis of common persistence for Canada and

Sweden for the entire sample period and the post-IT period at the 5-percent level. This result is robust to both the averaging parameter and the bandwidth.¹³

3.3 *Restricted Robinson estimates of inflation persistence*

The Robinson (1995) estimator allows for cross-equations restrictions, such that all or some of the G series share a common fractional integration parameter. That is, $d_g = \delta$, $g = 1, \dots, G$, or, in matrix form, $d = Q\delta$, where $Q = (1, 1, \dots, 1)'$ is a $G \times 1$ vector and δ is a scalar representing the unknown common fractional integration parameter. The restricted OLS estimator that imposes a common degree of fractional integration \hat{c} and \hat{d} is expressed as

$$\begin{bmatrix} \hat{c}^J \\ \hat{d}^J \end{bmatrix} = \left\{ \begin{bmatrix} I_G & 0 \\ 0 & Q' \end{bmatrix} (Z^J)' Z^J (\tilde{\Omega}^J)^{-1} \begin{bmatrix} I_G & 0 \\ 0 & Q \end{bmatrix} \right\}^{-1} \begin{bmatrix} I_G & 0 \\ 0 & Q' \end{bmatrix} \text{vec} \left((\tilde{\Omega}^J)^{-1} (Y^J)' Z^J \right). \quad (9)$$

When no restrictions are imposed, we set $Q = I_G$ and obtain the unrestricted OLS estimator. See equation (9).

Table 4 reports the restricted estimates of inflation persistence obtained by imposing the constraint $d(\text{IT})=d(\text{G})=d(\text{US})$. We conduct the estimation only for the cases where we fail to reject the null of equality at the 5-percent level. When we reject the null, we simply add N/A as a reminder that no evidence exists of a common fractional integrated parameter.¹⁴ Restricting fractional integration parameter to a common value across each IT country, Germany, and the

¹³ Our analysis focused on the dynamic behavior of detrended data, since detrending controls for the presence of low-frequency swings in the raw data. As a robustness test, however, we also analyzed the raw data for evidence of fractional integration and common persistence. We conducted this robustness check to guard against the possibility that the results depend on the detrending procedure. We find qualitatively the same results: inflation rates are fractionally integrated and only the three industrialized countries share a common degree of long-range hyperbolic decay in the IT sample. The detailed results are available from the authors. In general, using the raw data instead of the detrended data increases the magnitude of the estimates. This, however, does not change the main conclusions of the tests. Thus, the estimates of inflation persistence and the tests of common persistence are robust to the use of the original data.

¹⁴ We exclude Mexico, however, because the tests of equality consistently reject the null hypothesis for each of the three sample periods.

US results in estimates of common inflation persistence that maintain the fundamental characteristics of the corresponding unrestricted estimates. That is, for the industrial IT countries, we find a common parameter that suggests long-memory, mean-reversion, and stationarity in the full-sample and post-IT periods. For the developing countries, the estimates of common inflation persistence maintain long-memory, mean-reversion, and non-stationarity, or borderline non-stationarity, characteristics of the corresponding unrestricted estimates.

Because the evidence repeatedly shows that Canada, Sweden, and the UK individually share a common order of fractional integration with Germany and the US (i.e., satisfy the weak-form global hypothesis), we consider whether this property also holds collectively. Applied to the entire sample, the “augmented” Robinson test of common persistence between the three industrial countries, as a group, Germany, and the US fails to reject the null. The corresponding estimates of common persistence are, respectively, 0.385 (t-statistic = 9.682), 0.327 (t-statistic = 8.159), 0.343 (t-statistic = 7.493), and 0.369 (t-statistic = 7.076). Conversely, when we apply the “augmented” Robinson test of common persistence between the three developing countries, as a group, Germany, and the US, we reject the hypothesis at any reasonable level of significance.

3.4 Tests of the strong-form global hypothesis

The results of the Wald-type tests for equality of integration orders provide information to detect some similarities in the time dependence structure of inflation in IT countries, Germany, and the US, but provides no evidence regarding the existence of a common factor between the three series. We consider next the possibility that these processes that share the same long-memory parameter with Germany and the US also exhibit the same long-memory component (i.e., they converge in the long run to a common fractional stochastic equilibrium). For that purpose, we perform tests of fractional cointegration. That is, we estimate the degree of fractional

cointegration between each IT country, Germany, and the US and examine whether this generalized version of cointegration associates with short memory or long memory.

Fractional cointegration has received much recent attention. Integer cointegration results in an $I(0)$ relation between $I(1)$ variables. Fractional cointegration analysis is more complex than integer cointegration, as it relaxes both the $I(1)$ requirement of the variables and the $I(0)$ requirement of the error correction term. According to Engle and Granger (1987), cointegration arises when the integration order b of the residuals of the cointegration regression is less than the common integration order d of the constituent series. Depending on the values of d , b , and the reduction parameter $d - b$, three different forms of fractional cointegration can exist, each reflecting how the equilibrium errors respond to shocks (Aloy and De Truchis, 2012; De Truchis and Keddad, 2013). All three cases share the necessary condition $b < d$.

First, “strong” fractional cointegration occurs when the constituent series are non-stationary but the residual series is stationary (i.e., $0.5 < d \leq 1.0$, $b < 0.5$, and $d - b > 0.5$). This results in weak persistence of the deviations from the long-run equilibrium. This case includes the conventional integer cointegration case (i.e. $I(1)/I(0)$ or $d=1$ and $b=0$), that involves a strong adjustment mechanism. Second, “weak” fractional cointegration occurs when both the constituent series and the residual series are non-stationary (i.e., $0.5 < d \leq 1.0$, $b > 0.5$, but $d - b < 0.5$), which results in high persistence of deviations from equilibrium due to the non-stationarity of the error correction term. Finally, “stationary” fractional cointegration occurs when both the constituent series and the residual series are stationary (i.e., $-0.5 < d < 0.5$,

$b < 0.5$, and $d - b < 0.5$), which results in a weak persistence and stationary cointegration errors.¹⁵

In our case, as Table 4 clearly indicates, we deal with the possibility of stationary fractional cointegration, since the common fractional integration parameter is less than 0.5. We estimate jointly the long-run equilibrium relationship and the fractional integration parameter of the residuals using a variant of the fractional cointegration model suggested by De Truchis and Keddad (2013). That is,

$$(1-L)^b (IT(t) - \alpha_0 - \alpha_1 G(t) - \alpha_2 US(t)) = \varepsilon(t), \quad (12)$$

where the residual series is $I(b)$ and the constituent series of $IT(t)$, $G(t)$, and $US(t)$ are $I(d)$. We test the hypothesis that the memory of the residual series is of lower order than the common memory of the constituent series.¹⁶ This involves testing $H_0 : b = 0$ against the alternative $H_1 : b > 0$, and $H_0 : d - b = 0$ against the alternative $H_1 : d - b < 0$.

Tables 5 to 7 report the results of the joint estimates of the cointegrating regressions (12) and the memory parameter of the residuals using the ARFIMA(0,b,0) exact maximum likelihood approach, as proposed by Sowell (1990) for the three sample periods. We normalize the regressions on the IT country, which preserves the weak exogeneity of $G(t)$ and $US(t)$. The sign and the value of the long-run coefficients provide information for how Canada, Sweden, and

¹⁵ Spurious regression occurs when $0.5 \leq d \leq 1.0$, $0.5 \leq b \leq d$, and $d - b = 0$, while standard cointegration coincides with $d = 1$, $b = 0$, and $d - b = 1$.

¹⁶ The distribution theory for fractional cointegration is still poorly understood. Robinson (1994) shows that that conventional estimators of the cointegrating regression, such as OLS, are inconsistent when $b > 0$ (i.e., when the cointegrating relation does not completely purge the regression errors of memory). As an alternative, Robinson (1994) introduces the narrow-band least square (NBLS) estimator for the bivariate case, using the periodogram and the real part of the cross periodogram. See also Robinson and Marinucci (2003) and Christensen and Nelson (2006). The distribution theory of the NBLS, however, is not complete for all relevant parameter regions. Robinson and Marinucci (2003) provide the limiting distribution of the NBLS estimator for the case $d > 0.5$ and $b > 0$. Christensen and Nielsen (2006) provide the limiting distribution for the case $d > 0$, $b \geq 0$, and $d + b < 0.5$. Furthermore, it has been developed only the case of bivariate cointegration applications.

the UK adjust to shocks emanating from Germany and the US. The long-run coefficient for the US is always significant in all three samples, while the long-run coefficient for Germany is not significant for Canada, but significant for Sweden and the UK. The US exerts the greatest influence on Canada and the UK, while Germany exerts the greatest influence on Sweden and the UK.

We do not report the test results for Chile and Israel for the pre-IT sample because we reject the null hypotheses that the estimated coefficients on Germany and the US significantly differ from zero at the 5-percent level.¹⁷ The lack of statistical significance of these parameter estimates imply the absence of long-run co-movement between the two inflations rates. We also reject the hypothesis that a long-run co-movement exists between Canada and Germany in all three samples. Excluding Germany, however, exerts little effect on the overall results. The estimates of the memory parameter of the residuals for the full sample, the pre-IT sample, and the post-IT sample are 0.105 (3.604), 0.163 (2.477), and 0.097 (2.286), respectively. The estimates of the coefficient on the US are 0.535 (13.957), 0.231 (2.440), and 0.713 (22.353), respectively. In all three samples, we reject the null hypothesis $H_0 : d - b = 0$ in favor of the one-sided alternative $H_1 : d - b > 0$. Thus, Canada maintains a stationary fractional cointegration only with the US. A similar problem emerges with Sweden and the UK, but only in the pre-IT sample. Estimating the model for Sweden without German inflation does not change the estimate on the US nor the significance of the memory parameter, but slightly increases the estimate of the coefficient on the US inflation. The estimate of the memory parameter of the residuals is 0.073 (1.423) and the estimate of the coefficient on US inflation is 0.397 (3.126). Conversely,

¹⁷ The null hypothesis that these coefficients do not jointly significantly differ from zero is also not rejected. For Chile, the F test is $F(2,170) = 0.898$ (p-value = 0.408), while for Israel the F test is $F(2,186) = 0.338$ (p-value = 0.713).

estimating the model for the UK without the US inflation increases slightly both the coefficient estimate on German inflation and the memory estimate. The estimate of the memory parameter of the residuals is 0.278 (7.337) and that on German inflation is 0.483 (3.955).¹⁸

We conduct the hypothesis testing on b as if the residuals are observed and obtain the estimate of d from the value of the common fractional integration parameters from Table 4. The estimates of b , as well as the common estimates of d of the three constituent series, are in the stationary region. From the residual analysis, we cannot reject the null hypothesis that in the full sample period Canada, Sweden, and the UK exhibit a stationary fractional cointegration with Germany and the US, as the residuals show a lower order of fractional integration than the constituent series themselves, $b < d$. We reject the null hypothesis $H_0 : d - b = 0$ in favor of $H_1 : d - b > 0$ at the 1-percent level, suggesting that the stationary fractional cointegration is characterized by long memory, while we reject the null hypothesis $H_0 : d - b = 0.5$ in favor of $H_1 : d - b < 0.5$ also at the 1-percent level. Similar results apply to the pre-IT sample period. The only difference is that the stationary fractional cointegration between Sweden, Germany, and the US is characterized by short memory. In the post-IT period, however, although the estimated residuals from the cointegrated regressions exhibit a lower order than the common fractional integration parameter of the three constituent series, we fail to reject the null hypothesis $H_0 : d - b = 0$ at the 5-percent level for the UK.

¹⁸ We also considered, following Aloy and De Truchis (2012), four other estimators: the ordinary least squares (OLS) and several modifications to OLS, namely, the fully modified least squares (FMOLS) by Phillips and Hansen (1990), the dynamic least squares (DOLS) by Phillips and Loretan (1991), Saikkonen (1991) and Stock and Watson (1993), and the canonical cointegrating regression estimator (CCR) by Park (1992). The estimates of the long-run coefficients remain robust across the four estimation techniques and do not differ in any substantial way from the ones detailed in Tables 5 to 7. For this reason, we do not report them, but they are available on request. We obtained the residuals from the four cointegrating regressions and applied the Robinson estimator. The results, as expected, slightly vary, but not in any substantial manner that questions the finding in Tables 5 to 7.

The value of b , which corresponds to the integration order of the residual series, yields information about the speed of adjustment toward equilibrium. The greater is the difference between d and b , the greater is the reduction of the dimensionality, and the smaller are the shocks between each IT country, Germany, and the US that lead to a persistent deviation from the long-run equilibrium.

In summary, we observe stationary fractional cointegration for Canada, Sweden, and the UK at the 1-percent level for the full sample; for Canada, Sweden, and the UK at the 1-percent level for the pre-IT sample; and for Canada at the 1-percent level and Sweden at the 10-percent level for the post-IT sample.¹⁹ Thus, in the post-IT sample, we find strong evidence of stationary fractional cointegration for Canada, weak evidence of stationary fractional cointegration for Sweden, and no evidence of stationary fractional cointegration of the UK and the three developing countries. Clearly, inferences on inflation persistence may prove misleading, if we ignore the possibility of stationary fractional cointegration.

4. Conclusion

The empirical results offer new insights on the dynamics of inflation persistence in IT countries. Most literature on inflation persistence in IT countries takes a unilateral approach, relying exclusively on tests before and after the implementation of IT. The period of IT implementation, however, also corresponds to a period that has witnessed profound changes in the world economy, including the rise of the phenomenon of globalization and the international decline of inflation. Untangling the positive effects of IT from the positive effects of the global economic environment may prove difficult. But, neglecting these global developments may lead to overstating the domestic effects of IT implementation, since the success of IT countries may

¹⁹ Tables 5 to 7 also report the test of the null hypothesis $H_0 = d - b = 0.5$ against the one-sided alternative $H_0 = d - b < 0.5$. The tests confirm that the stationary version of the fractional cointegration results.

largely reflect the global nature of the disinflationary process. This, in turn, could lead to erroneous inferences about how a unilateral change in monetary policy affects the dynamics of the economy.

Thus, this paper investigates inflation persistence in six inflation targeting (IT) countries from the perspective of the global economy. This view maintains that inflation persistence in IT countries has declined mainly because of the decline of inflation persistence in the global economy. The argument is mainly rooted in the global inflation literature, which documents that inflation in the last three decades has largely been a global phenomenon, and in the bilateral country-pairing literature, which argues that conclusions based solely on one-sided evidence from IT countries are difficult to evaluate when many non-IT countries also shared similar experiences.

Within this perspective, we provide empirical evidence on two yet unanswered questions. First, we investigate whether each IT country in the sample share a common persistence with Germany and the US, two non-IT countries with a relatively good inflation record, which we use as proxies for the global economy. This tests the weak-form global hypothesis. Using a multivariate frequency-domain approach, we find that inflation in the industrial countries in the sample (Canada, Sweden, and the UK) exhibits a long-memory stationary process that shares a common degree of persistence with Germany and the US. Thus, they respond in similar way to shocks as Germany and the US. In contrast, we find that inflation in the developing countries (Chile, Israel, Mexico) exhibits a long-memory non-stationary process that does not share a common persistence with Germany and the US.

Second, given the common inflation persistence across an IT country, Germany, and the US, we examine whether the same long-memory component drives the inflationary processes of

Canada, Sweden, and the UK to converge in the long-run to a common stochastic equilibrium with Germany and the US. This tests the strong-form global hypothesis. We find strong evidence of stationary fractional cointegration between each industrial IT country (Canada, Sweden, and the UK) and Germany and the US over the full sample period. This cointegration is characterized by long memory. We also find strong evidence of stationary fractional cointegration between each industrial IT country (Canada, Sweden, and the UK) and Germany, and the US over the pre-IT sample period. Except for Sweden, this stationary fractional integration is also characterized by long memory. For the post-IT sample period, however, we find strong evidence for Canada, weak evidence for Sweden, but no evidence for the UK. We conjecture that notwithstanding the IT regime, there exists common forces driving inflation persistence in Canada and Sweden, as well as Germany and the US. These findings cast doubt on the relevance of IT in our sample of industrial countries and suggests that the global economy may importantly affect the decline of inflation persistence in these countries.

The overall findings suggest several conclusions. First, we provide additional evidence that inflation in all six IT countries is a mean-reverting, fractionally integrated process that does not depend on the monetary regime. Second, we find substantial evidence that inflation persistence in developing IT countries inherently differs from that of industrial IT countries, Germany, and the US. Notably, inflation conforms to stationary processes in the industrial countries, Germany, and the US, while it conforms to non-stationary processes in the developing countries, even after IT implementation. Third, we show that in the post-IT period, the three industrial IT countries share a common persistence with Germany and the US and, consequently, respond in the same way to inflationary shocks. That is, the IT industrial countries exhibit dynamics of the inflation rate that harmonizes with the dynamics of the global economy. Finally,

we provide evidence that in two of the three industrial IT countries inflation is cointegrated with Germany and the US, and that this form of cointegration associates with long memory. Thus, our results supports the conjecture that the global disinflation process of the last two decades may drive inflation persistence in two of the three industrial IT countries. No evidence of cointegration exists for the UK in the post-IT period, and no evidence of common persistence and cointegration exists for the three developing IT countries.

Our empirical analysis also answer three additional questions. Does inflation persistence become lower for both industrial and developing countries after the adoption of the IT regime? Did inflation persistence remain stable throughout our sample period for all countries? Does inflation persistence in countries experiencing higher inflation in the recent past (such as Israel and Mexico) still remain higher than in the other countries?

Our results provide additional evidence that inflation in all six IT countries as well as Germany and the US is a mean-reverting, fractionally integrated processes that do not depend on the monetary regime. Contrary to the findings of Levin and Piger (2002), O'Reilly and Whelan (2005), and Pivetta and Reis (2006) inflation persistence did not remain stable throughout the sample period, particularly for the industrial countries. Our results also contrast with the unit-root view of inflation, as argued by Mishkin (1992), Culver and Papell (1997), and Crowder and Hoffman (1996). Inflation exhibits, however, more persistent dynamics in the developing countries than in the industrial countries, including Germany and the US. Evidence also exists that IT implementation changed the process of forming inflationary expectations and, in particular, how well-anchored these expectations are, as reflected in the decline of the estimates of inflation persistence for some, but not all, IT countries. That is, inflation persistence in some IT countries did not noticeably decline. Substantial declines, on the other hand, did occur in

inflation persistence in Germany and the US. Importantly, we find that Israel and Mexico who experienced high inflation and high inflation persistence before IT do not display lower inflation persistence after the adoption of IT. Perhaps “memory” of high inflation does not easily fade away, even after the adoption of IT. In sum, the statistical evidence is mixed. These mixed results suggest that the response to inflationary shocks even in countries that adopt IT follows a more sophisticated process than the conventional IT framework can explain.

References:

- Agenor, P. R., and Taylor, M. P., 1993. Analysing credibility in high-inflation countries: A new approach. *Economic Journal* 103, 329–36.
- Alogoskoufis, G., 1992. Monetary accommodation, exchange rate regimes and inflation persistence. *Economic Journal* 102, 461–80.
- Aloy, M., and De Truchis, G., 2012. Estimation and testing for fractional cointegration. AMSE Working Papers 1215, Aix-Marseille School of Economics, Marseille, France.
- Amato, J. D., and Gerlach, S., 2002. Inflation targeting in emerging market and transition economies: lessons after a decade. *European Economic Review* 46, 781–90.
- Arize, A. C., Malindretos J., and Nam K., 2005. Inflation and structural change in 50 developing countries. *Atlantic Economic Journal* 33, 461–471.
- Bai, J., and Perron, P., 2003. Computation and analysis of multiple structural change models. *Journal of Applied Econometrics* 18, 1–22.
- Baillie, R. T., 1996. Long-memory processes and fractional integration in econometrics. *Journal of Econometrics* 73, 5–59.
- Ball, L., and Sheridan, N., 2005. Does inflation targeting matter? In: B. Bernanke and M. Woodford (eds), *The Inflation Targeting Debate*. Chicago: University of Chicago Press, 249-276.
- Bandi, F., and Perron, B., 2006. Long memory and the relation between implied and realized volatility. *Journal of Financial Econometrics* 4, 636-670.
- Baum, C. F., Barkoulas, J. T., and Caglayan, M., 1999. Persistence in international inflation rates. *Southern Economic Journal* 65, 900-913.
- Batini, N. , 2002. Euro Area inflation persistence. ECB Working paper No 201.

- 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.
- Bollerslev, T., and Jubinski, D., 1999. Equity trading volume and volatility: Latent information arrivals and common long-run dependencies. *Journal of Business and Economic Statistics* 17, 9–21.
- Bratsiotis, G. J., Madsen, J. and Martin, C., 2015. Inflation targeting and inflation persistence. *Economic and Political Studies* 3, 3-17.
- Brito, R. D., and Bystedt, B., 2010. Inflation targeting in emerging economies: Panel evidence. *Journal of Development Economics* 91, 198-210.
- Baxa, J., Horváth, R., Vašíček, B., 2014. How does monetary policy change? Evidence on inflation-targeting countries. *Macroeconomic Dynamics* 18, 593-630.
- Canarella, G., and Miller, S. M., 2015. Inflation persistence before and after inflation targeting: A fractional integration approach. *Eastern Economic Journal*, in press.
- Canarella, G., and Miller, S. M. 2016. Inflation persistence and structural breaks: The experience of inflation targeting countries and the US. Available at SSRN: <http://ssrn.com/abstract=2486872> or <http://dx.doi.org/10.2139/ssrn.2486872> .
- Caporale, G. M., and Gil-Alaña, L. A., 2011. Multi-factor Gegenbauer processes and European inflation rates. *Journal of Economic Integration* 26, 386-409.
- Calvo, G. A., 1983. Staggered prices in a utility maximization framework. *Journal of Monetary Economics* 12, 383–393.
- Cecchetti, S., and Ehrmann, M., 2000. Does inflation targeting increase output volatility? An international comparison of policymakers' preferences and outcomes. Working Paper 69. Central Bank of Chile.
- Chiquiar, D., Noriega, A. E., and Ramos-Francia, M., 2010. A time series approach to test a change in inflation persistence: The Mexican experience. *Applied Economics* 42, 3067–3075
- Ciccarelli, M., and Mojon, B., 2010. Global inflation. *The Review of Economics and Statistics* 92, 524-535.

- Clarida, R., Galí, J., and Gertler, M., 1998. Monetary policy rules in practice: Some international evidence, *European Economic Review* 42, 1033-1067.
- Clarida, R., Galí, J., and Gertler, M., 2000. Monetary policy rules and macroeconomic stability: Evidence and some theory. *Quarterly Journal of Economics* 115, 147- 180.
- Corvoisier, S., and Mojon, B., 2005. Breaks in the mean of inflation: How they happen and what to do with them? ECB Working Paper 451.
- Cunado, J., Gil-Alaña, L. A., and Perez de Gracia, F., 2005. A test for rational bubbles in the Nasdaq stock index. A fractionally integrated approach. *Journal of Banking and Finance* 29, 2633-2654.
- De Truchis, G., and Keddad, B., 2013. Southeast Asian monetary integration: new evidences from fractional cointegration of RER. *Journal of International Financial Markets, Institutions and Money* 26, 394-412
- Duncan, R., and Martínez-García, E., 2015. Forecasting Local Inflation with Global Inflation: When Economic Theory Meets the Facts. Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute Working Paper No. 235.
- Engle, R. F. and Granger, C. W. J., 1987. Co-integration and error correction: Representation, estimation and testing. *Econometrica* 55, 251–276.
- Fiess N., and Byrne, J., 2007. Euro Area inflation: Aggregation bias and convergence. Department of Economics, University of Glasgow Working Paper.
- 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., 2010. Inflation persistence. In: B. M. Friedman and M. Woodford (ed.), *Handbook of Monetary Economics*, edition 1, volume 3, chapter 9, pp. 423-486 Elsevier.
- Fuhrer, J. C., and Moore, G., 1995. Inflation persistence. *Quarterly Journal of Economics* 110, 127-160.
- 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.
- Goncalves, C. E. S., and Salles, J. M., 2008. Inflation targeting in emerging economies: What do the data say? *Journal of Development Economics* 85, 312-318.

- Granger, C. W. J., and Joyeux, R., 1980. An introduction to long-memory models and fractional differencing. *Journal of Time Series Analysis* 1, 15–39.
- Harvey, C., 1995. Predictable risk and returns in developing markets. *Review of Financial Studies* 8, 773-816.
- Hassler, U., 2011. Estimation of fractional integration under temporal aggregation. *Journal of Econometrics* 162, 240-247.
- Hassler, U., and Meller, B., 2011. Detecting multiple breaks in long memory: the case of US inflation. Discussion Paper Series 1: Economic Studies No. 26/2011, Deutsche Bundesbank.
- Hassler, U., and Wolters, J., 1995. Long memory in inflation rates: International evidence. *Journal of Business and Economic Statistics* 13:1, 37–45.
- Hassler, U., Marmol, F., and Velasco, C., 2006. Residual log-periodogram inference for long run relationships. *Journal of Econometrics* 130, 165- 207.
- Hauser, M. A., Pötscher, B. M., and Reschenhofer, E., 1999. Measuring persistence in aggregate output: ARMA models, fractionally integrated ARMA models and nonparametric procedures. *Empirical Economics* 24, 243-269.
- Hosking, J., 1981. Fractional differencing. *Biometrika* 68, 165–176.
- Hsu, C.-C., 2005. Long memory or structural changes: An empirical examination on inflation rates. *Economics Letters* 88, 289–294.
- 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.
- Kellard, N., and Sarantis, N., 2008. Can exchange rate volatility explain persistence in the forward premium? *Journal of Empirical Finance* 15, 714-728.
- Kellard, N., Dunis, C. and Sarantis, N., 2010. Foreign exchange, fractional cointegration and the implied realized volatility relation. *Journal of Banking and Finance* 34, 882–891.
- Kim, C. S., and Phillips, P. C. B., 1999. Modified log periodogram regression. Yale University, mimeo.
- 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., and Posen, A., 1999. Does talk matter after all? Inflation targeting and central bank behavior. Federal Reserve Bank of New York, Staff Reports 88.
- Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., and Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series have a unit root. *Journal of Econometrics* 54, 159-178.
- Leiderman, L., and Svensson, L. E. O., (Eds.) 1995. *Inflation Targets*. CEPR, London.
- 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.
- Lin, S., and He, H., 2009. Does inflation targeting make a difference in developing countries? *Journal of Development Economics* 89, 118-123.
- Lobato, I., and Velasco, C., 2000. Long memory in stock-market trading volume. *Journal of Business and Economic Statistics* 18, 410-427.
- Marinucci, D., and Robinson, P. M., 2001. Semiparametric fractional cointegration analysis. *Journal of Econometrics* 105, 225-247.
- Marques, C. R., 2004. Inflation persistence: Facts or artefacts? Working Paper, European Central Bank.
- Meller, B., and Nautz, D., 2012. Inflation persistence in the Euro area before and after the European Monetary Union. *Economic Modelling* 29, 1170–1176.
- Michelacci, C. and Zaffaroni, P., 2000. (Fractional) Beta convergence. *Journal of Monetary Economics* 45, 129-153.
- Mihov, I., and Rose, A. K., 2008. Is old money better than new? Duration and monetary regimes. *Economics: The Open-Access, Open-Assessment E-Journal* 2, 1-24.
- 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.
- Nelson, C. R., and Plosser, C. I., 1982. Trends and random walks in macroeconomics time series: Some evidence and implications. *Journal of Monetary Economics* 10, 139-162.
- Nielsen, M. O., 2007. Local Whittle analysis of stationary fractional cointegration and the implied–realized volatility relation. *Journal of Business and Economic Statistics* 25, 427-446.

- Noriega, A. E., Capistrán, C., Ramos-Francia, M., 2013. On the dynamics of inflation persistence around the world, *Empirical Economics* 44, 1243-1265.
- O'Reilly, G., and Whelan, K., 2005. Has Euro-area inflation persistence changed over time? *The Review of Economics and Statistics* 87, 709-720.
- Park, J. Y. , 1992. Canonical cointegrating regression. *Econometrica* 60, 119-143.
- Phillips, P. C. B., 1999. Unit root log periodogram regression. Unpublished working paper No. 1244, Cowles Foundation for Research in Economics, Yale University.
- Phillips, P., and Hansen, B., 1990. Statistical Inference in Instrumental Variables Regression with I(1) Processes. *Review of Economic Studies* 57, 99-125.
- Phillips, P. C. B., and Loretan, M., 1991. Estimating long run economic equilibria. *Review of Economic Studies* 58, 407-436.
- Phillips, P. C. B., and Perron, P., 1988. Testing for a unit root in a time series regression, *Biometrika* 75, 335-346.
- Phillips, P. C. B., 2007. Unit root log periodogram regression. *Journal of Econometrics* 138, 104–124.
- 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.
- Rogoff, K. S., 2003. Globalization and global disinflation. In *Monetary Policy and Uncertainty: Adapting to a Changing Economy*, Federal Reserve Bank of Kansas City Jackson Hole Symposium
- Saikkonen, P., 1991. Asymptotically efficient estimation of cointegrating regressions. *Econometric Theory* 7, 1-21
- Siklos, P. L., 2008. Inflation targeting around the world. *Emerging Markets Finance and Trade* 44, 17–37.
- Sibbertsen, P., 2004. Long memory versus structural breaks: An overview. *Statistical Papers* 45, 465–515.
- Souza, L. R., 2007. Temporal aggregation and bandwidth selection in estimating long memory. *Journal of Time Series Analysis* 28, 701-722.

- Sowell, F., 1992, Modeling long-run behavior with the fractional ARIMA models. *Journal of Monetary Economics* 29, 277-302.
- Stock, J. H. and Watson, M. W., 1993. A simple estimator of cointegrating vectors in higher order integrated systems. *Econometrica* 61, 783–820.
- Stock, J. H., 2001. Comment on ‘Evolving post-World War II US inflation dynamics’. *NBER Macroeconomics Annual*. 16, 379–387.
- Taylor, J., 1999. Staggered price and wage setting in macroeconomics. J. Taylor and M. Woodford (Eds), *Handbook of Macroeconomics*, vol 1.b. North-Holland.
- Velasco, C., 1999. Non-stationary log periodogram regression. *Journal of Econometrics* 91, 325-371.
- Walsh, C. E., 2009. Inflation targeting: What have we learned? *International Finance* 12, 195–233.
- Whelan, K., and O’Reilly, G., 2005. Has Euro-Area inflation persistence changed over time. *The Review of Economics and Statistics* 87, 709-720.
- White, W. R., 2008. Globalization and the determinants of domestic inflation. BIS Working Paper, Bank for International Settlements.
- Yigit, T., 2010. Inflation targeting: An indirect approach to assess the direct impact. *Journal of International Money and Finance* 29, 1357–1368.

Table 1: Inflation rates: descriptive statistics

Country	Sample Period	Mean	St. Dev.	Skewness	Kurtosis
<u>IT countries</u>					
Chile	1976:02-1990:08	27.462	24.166	2.685	11.816
	1990:09-2013:06	5.541	6.560	1.888	12.519
	1976:02-2013:06	14.085	19.173	3.407	19.399
Israel	1976:02-1991:12	56.631	52.312	1.622	6.004
	1992:01-2013:06	4.408	6.055	0.546	3.485
	1976:02-2013:06	26.624	43.009	2.600	11.008
Mexico	1976:02-1998:12	30.688	25.903	1.730	6.164
	1999:01-2013:06	4.257	2.490	0.060	2.610
	1976:06-2013:06	21.858	24.58	2.044	7.672
Canada	1976:02-1991:01	6.611	4.074	1.072	6.427
	1991:02-2013:06	1.805	3.281	-0.280	4.578
	1976:02-2013:06	3.731	4.316	0.602	5.507
Sweden	1976:02-1992:12	7.700	6.101	1.209	6.361
	1993:01-2013:06	1.391	3.352	1.566	24.879
	1976:02-2013:06	4.244	5.728	1.508	8.015
UK	1976:02-1992:09	7.631	6.239	2.869	17.887
	1992:10-2013:06	2.160	2.228	0.077	4.324
	1976:02-1992:09	4.597	5.239	5.239	22.466
<u>Non-IT countries</u>					
Germany	1976:02-1990:12	2.984	2.760	0.782	4.509
	1991:01-2013:06	1.929	2.695	1.622	12.043
	1976:02-2013:06	2.348	2.764	1.225	8.239
US	1976:02-1990:12	5.887	3.942	0.367	3.232
	1991:01-2013:06	2.448	3.039	-1.414	13.151
	1976:02-2013:06	3.817	3.811	0.043	6.709

Note: Inflation is measured as the monthly logarithmic difference expressed as a percentage. Data are from OECD *Main Economic Indicators: Historical Statistics* and are seasonally adjusted using the Census X12 method.

Table 2: Unrestricted Robinson estimates of inflation persistence

Country	Full Sample			Pre-IT Sample			Post-IT Sample		
	$d(IT)$	$d(G)$	$d(US)$	$d(IT)$	$d(G)$	$d(US)$	$d(IT)$	$d(G)$	$d(US)$
Chile	0.537 (6.649)	0.266 (3.292)	0.297 (3.676)	0.714 (6.558)	0.433 (3.985)	0.421 (3.864)	0.490 (5.893)	0.327 (3.937)	0.193 (2.331)
Israel	0.738 (7.406)	0.266 (2.668)	0.297 (2.979)	0.679 (5.166)	0.465 (3.543)	0.442 (3.364)	0.737 (9.444)	0.172 (2.205)	0.185 (2.372)
Mexico	0.713 (7.303)	0.266 (2.724)	0.297 (3.042)	0.766 (7.921)	0.253 (2.621)	0.461 (4.759)	0.652 (5.585)	0.151 (1.288)	0.073 (0.631)
Canada	0.374 (4.116)	0.266 (2.921)	0.297 (3.261)	0.458 (3.687)	0.384 (3.092)	0.435 (3.502)	0.283 (3.144)	0.342 (3.807)	0.199 (2.211)
Sweden	0.299 (3.411)	0.266 (3.030)	0.297 (3.383)	0.138 (1.347)	0.442 (4.302)	0.435 (4.235)	0.346 (4.097)	0.157 (1.864)	0.182 (2.154)
UK	0.398 (4.239)	0.266 (2.831)	0.297 (3.161)	0.379 (3.403)	0.405 (3.639)	0.431 (3.864)	0.434 (4.296)	0.237 (2.351)	0.201 (1.997)

Note: The numbers in parenthesis are t -statistics.

Table 3: Tests of equality of inflation persistence

Country	Full Sample	Pre-IT Sample	Post-IT Sample
Chile	6.748	4.632	6.368
Israel	14.044	1.970	34.144
Mexico	13.074	14.215	14.478
Canada	0.758	0.184	1.284
Sweden	0.091	5.491	2.948
UK	1.084	0.106	3.064

Note: The table reports Robinson's (1995) chi-square test statistic for the null hypothesis $H_0 : d(IT) = d(G) = d(US)$. The test statistic has 2 degrees of freedom. The critical values for chi-square with two degrees of freedom are: 4.605 (10 percent), 5.991 (5 percent) and 9.219 (1 percent).

Table 4: Restricted Robinson estimates of common inflation persistence

Country	Full Sample	Pre-IT sample	Post-IT Sample
Chile	N/A	0.523 (8.119)	N/A
Israel	N/A	0.529 (6.973)	N/A
Canada	0.312 (5.981)	0.426 (6.041)	0.275 (5.315)
Sweden	0.287 (5.726)	0.338 (5.536)	0.228 (4.653)
UK	0.321 (5.934)	0.405 (6.401)	0.291 (4.952)

Note: The number in parenthesis are t -statistics. The estimates are obtained by imposing the restriction $d(IT)=d(G)=d(US)$. N/A stands for "not applicable."

Table 5: Tests of stationary fractional cointegration: the full sample

Country	Germany	US	b	$H_0 : d - b = 0$	$H_0 : d - b = 0.5$
Canada	0.031 (0.506)	0.526 (13.07)	0.104 (3.551)	0.207 (7.066)	0.306 (10.391)
Sweden	0.194 (3.585)	0.352 (5.256)	0.115 (4.132)	0.167 (5.962)	0.333 (11.870)
UK	0.256 (4.239)	0.186 (2.816)	0.271 (12.75)	0.067 (3.171)	0.433 (20.390)

Note: The number in parenthesis are t -statistics. The value of d equals the estimate from Table 4.

Table 6: Tests of stationary fractional cointegration: the pre-IT sample

Country	Germany	US	b	$H_0 : d - b = 0$	$H_0 : d - b = 0.5$
Canada	0.185 (1.165)	0.187 (1.872)	0.139 (2.101)	0.286 (4.295)	0.214 (3.303)
Sweden	0.074 (0.307)	0.420 (3.254)	0.078 (1.520)	0.259 (5.058)	0.241 (4.051)
UK	0.445 (3.452)	0.194 (1.448)	0.248 (5.755)	0.156 (3.613)	0.344 (7.861)

Note: The number in parenthesis are t -statistics. The value of d equals the estimate from Table 4.

Table 7: Tests of stationary fractional cointegration: the post-IT sample

Country	Germany	US	b	$H_0 : d - b = 0$	$H_0 : d - b = 0.5$
Canada	0.016 (0.241)	0.717 (20.311)	0.096 (2.191)	0.178 (4.041)	0.322 (8.741)
Sweden	0.469 (8.019)	0.268 (4.329)	0.159 (4.557)	0.068 (1.962)	0.432 (12.532)
UK	0.083 (1.967)	0.209 (4.472)	0.288 (7.971)	0.002 (0.068)	0.498 (14.243)

Note: The number in parenthesis are t -statistics. The value of d equals the estimate from Table 4.