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Time Variation in Inflation Persistence:
New Evidence from Modelling US Inflation

Abstract

This article explores how inflation persistence relates to the conduct and goals of monetary policy by presenting a new approach to modelling US inflation persistence and the Fed's dual mandate. Our framework fills a gap in pre-existing models by more flexibly accounting for diverse dynamic properties and shocks. Estimating a Phillips Curve model augmented with inflation volatilities and expectations, we find that the degree of monthly inflation persistence is time variant since World War II. Variations in persistence continue to be observed regardless of the absolute level of inflation and the extent of the trade-off between inflation and unemployment. We demonstrate that inflation persistence varies in line with expectations formed by memories of past inflation. This supports the case for more flexible monetary policy at times, as in the 1980s or especially the present decade, when inflation is more persistent.

Keywords: Inflation persistence; Monetary Policy; Phillips curve; Integration.
JEL classification : E52; J64; C22; C52.

1. Introduction

Using micro/survey-based data on memories of past inflationary episodes, Ehrmann and Tzamourani (2012) and Malmendier and Nagel (2015) observe that the memory of the 1970s Great Inflation in the US only started to be dispelled in the 1990s. In a related phenomenon, a relatively small decline of inflation was observed in response to the Great Recession of 2007-09 compared to the 1979-85 period (Ball and Mazumder, 2011; Blanchard, 2016; Coibion and Gorodnichenko, 2015; Kiley, 2015; Stock, 2011; Watson, 2014). Such echoes of long memory in inflation data should be accounted for in monetary models. On the basis of this view, the literature on inflation long memory and volatility has made the case for modelling post-war US inflation dynamics with a fractional order of integration (FI) (Bos, Koopman and Ooms, 2014; Lovcha and Perez-Laborda, 2018).

This article contributes to that literature by presenting a statistical approach to examining the long-run level of US inflation in the context of the dual mandate of the Federal Reserve ("Fed") to "foster economic conditions that achieve both stable prices and maximum sustainable employment."[1] Modelling in previous studies is based on a single price stability goal. Our approach also takes account of two other factors: inflation memory affecting the slope of Phillips Curve (e.g. Blanchard, 2016); and, reflecting the missing disinflation in response to the Great Recession,

[1] https://www.chicagofed.org/research/dual-mandate/dual-mandate

the complex nexus of inflation and unemployment rates together with their time-varying volatilities. Our modelling provides the basis for an informal discussion of the policymaking implications of inflation persistence by estimating a Phillips Curve model augmented with inflation volatilities and expectations for the US economy.

To exploit fully the Phillips curve trade-off as recommended by Mavroeidis, Plagborg-Møller, and Stock (2014), our approach is predicated on past shocks driving the inflation dynamic rather than, as is the case with the New Keynesian Phillips Curve (NKPC) model, forward expectations. We examine monthly inflation data from 1957 to 2015 using the lags of inflation to overcome the problem of imperfectly rational expectations (Gali and Gertler, 1999). Unlike the traditional Phillips Curve model which restricts the sum of the coefficients of the lagged inflation rates to the equivalent of unity (Williams, 2006), we capture the degree of inflation persistence over time by using lagged inflation as a proxy for expectations in a form of univariate autoregressive fractionally integrated moving averages (ARFIMA) model with exponential generalized autoregressive conditional heteroskedastic (EGARCH) type innovations to reflect two factors: the postwar US inflation process not entirely or permanently following a random walk; and the time-varying volatility. Our modelling also includes supply shocks that may generate inflation, thereby reducing the bias of the coefficient towards NAIRU or the output gap.

Following this introduction, section two describes our method. Section three proceeds to estimate the parameters and discusses the results. Some concluding remarks are presented in section four.

2. Methodological approach

The empirical literature on post-war US inflation includes extensive discussions of the characteristics and determinants of inflation persistence – which we define following Willis (2003: 7) as "the speed with which inflation returns to baseline after a shock." Several studies highlight the time-varying nature of persistence (Clark and Davig, 2011; Cogley, Primiceri and Sargent, 2010; Cogley and Sargent, 2001; Levin and Piger, 2004; Stock and Watson, 2007, 2009; Watson, 2014). Other authors see persistence as high and unchanged (Pivetta and Reis, 2007), or dependent on the nature of the monetary policy regime (Benati, 2008). Erceg and Levin (2003), considering a single inflation target rather than a dual mandate as we do, develop a dynamic stochastic general equilibrium (DSGE) model in which private agents optimally extract information about the central bank's inflation target from the evolution of the interest rate. When the inflation target is frequently altered, inflation persistence is generated.

Table A.1 reports some conflictual findings on inflation persistence from various studies. Modelling inflation as a FI could contribute to lessen the divergence in these results (Hassler and Wolters, 1995; Lovcha and Perez-Laborda, 2018). We categorise the reaction of the US postwar inflation time series to shocks into three possible types: (i) persistence decays at an exponential rate (short memory), (ii) persistence decreases at a hyperbolic rate (long memory), or (iii) persistence is infinite (perfect memory). The three categories correspond to different degrees of integration of the US postwar inflation time series. A process with short memory is stationary (integrated with degree zero), a series with long memory is integrated to a fraction, and a series with perfect memory is integrated with degree 1 (Baillie, Chung and Tieslau, 1996; Baum, Barkoulas, and Caglayan, 1999; Gadea and Mayoral, 2006).

Econometric approaches include the largest autoregressive root of inflation (Taylor, 2000; Cogley and Sargent, 2001), a Markov regime-switching model of inflation (Evans and Wachtel, 1993), a structural FI vector autoregressive model (FIVAR) (Lovcha and Perez-Laborda, 2018), a stochastic volatility (SV) model (Bos, Koopman and Ooms, 2014), quantile regression techniques (Gaglianone, Guillén and Figueiredo, 2018) and changes in the vector autoregression impulse response functions interpreted through the use of a structural macroeconomic model (Boivin and Giannoni, 2006) allowing for structural breaks (Levin and Piger, 2004).

Marques (2004) has evidenced that the impulse response function is not a useful measure of persistence as it is an infinite-length vector. To overcome this difficulty, Andrews and Chen (1994) show that one can rely on the sum of the autoregressive coefficients as a measure of persistence and this finding underpins the univariate approach. Therefore, we employ an empirical model of a univariate AR(n) process to capture inflation persistence.

Batini and Nelson (2001: 383) identify three different types of inflation persistence: "positive serial correlation in inflation", "lags between systematic monetary policy actions and their (peak) effect on inflation", and "lagged responses of inflation to non-systematic policy action (i.e. policy shocks)". Inflation is (highly) persistent when inflation responds slowly to a shock, and not (very) persistent when the response speed is high (Equation 1):

$$\pi_t = \mu_0 + \sum_{i=1}^{n} \phi_i \pi_{t-i} + \varepsilon_t \qquad (1)$$

where π_t stands for inflation, μ_0 is the intercept, n and ϕ_i are the order and the coefficients of AR terms, and ε_t is the disturbance term, which is serially uncorrelated and follows a Gaussian distribution with mean zero and σ_i^2.

To characterise the significant autocorrelation between observations of a time series dynamic, Granger (1980) and Granger and Joyeux (1980) develop an ARFIMA model with the flexibility of allowing fractional orders of integration. Equation (2) has been modelled combining autoregressive and moving averages (ARMA). The ARMA (n, m) is the following:

$$\phi(L)\pi_t = \mu + \theta(L)\varepsilon_t \qquad (2)$$

where L is the lag operator $L^k \pi_t = \pi_{t-k}$, μ is the regressor, $\phi(L) = 1 - \sum_{i=1}^{n} \phi_i L^i, \theta(L) = 1 + \sum_{i=1}^{m} \theta_i L^i$ and both $\phi(L)$ and $\theta(L)$'s roots lie outside the unit circle. A time series π_t follows an ARFIMA (n, d, m) process, which can be expressed as:

$$\phi(L)(1-L)^d(\pi_t-\mu)=\theta(L)\varepsilon_t \qquad (3)$$

where $(1-L)^d$ accounts for the long memory and is defined as:

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

with Γ denoting the Gamma function. The fractionally differencing parameter d, lying between zero and unity, measures the speed of that inflation's convergence to equilibrium after a shock to an I(d) process (as defined below). When $d = 0$, the series is an I(0) process with short-run behaviour, in which the effects of shocks fade at an exponential rate of decay so that the series quickly regains its equilibrium. When $d = 1$, the series is an I(1) process: following a shock, the series does not revert to its mean and the persistence in response to shocks is infinite. When $0 < d < 1$, the series lies between the distinctive I(0) and I(1), and may thus be identified as an I(d) process with long-run dependence, in which persistence dies out hyperbolically – that is the series takes a considerable time to reach mean reversion after a shock (Granger and Joyeux, 1980; Baillie, 1996). Specifically, when $0 < d < 0.5$, the series is stationary with mean reversion; when $0.5 \leqslant d < 1$, the series is non- stationary with mean reversion; when $d \geqslant 1$, the series is non-stationary and non-reverting; in the case of $-0.5 < d < 0$, the series is stationary and "reverses itself more frequently than random process" (Canarella and Miller, 2017).

2.1. Modelling the Nairu

We characterize the features of inflation behaviour in the framework of the so-called 'Triangle model.' Underlying this model developed by Gordon (1997) is the concept of rational expectations as a determinant of short-run inflationary behaviour where inflation is formulated as a function of three components: inertia, that is the influence of past inflation on expectations of future inflation encompassing the formation of all expectations including explicit or implicit wage and price contracts (for instance see Bhattarai, 2016 for evidence from OECD countries); demand, as inflation is affected by aggregate demand driven by unemployment or the output gap; and supply shocks such as oil and commodity price changes affecting costs (this final factor in 'Gordon's Triangle' had previously tended to be overlooked as a driver of expectations). The model is expressed as follows:

$$\pi_t = \phi'(L)\pi_{t-1} + \delta(L)D_t + \psi(L)X_t + \theta(L)\varepsilon_t \qquad (4)$$

with the inflation rate π_t, being the dependent variable, lags on the inflation rate represent inertia π_{t-1} , $\phi'(L) = 1 - \sum_{i=1}^{n} \phi_i' L^i$. These lagged inflation rates represent inflation inertia by restricting the sum of the coefficients of the lagged inflation rates to the equivalent of unity. D_t is an index of excess supply/ demand (normalized so that $D_t=0$ indicates the absence of excess supply/demand), $\delta(L) = 1 + \sum_{i=1}^{m} \delta_i L^i$, X_t is a vector of supply shock variables (normalized so that X_t=0 indicates an absence of supply shocks), $\psi(L) = \sum_{i=1}^{m} \psi_i L^i$, and ε_t is a serially uncorrelated error term. $\phi'(L)$, $\delta(L)$,and $\psi(L)$ roots lie outside the unit circle.

If in the estimation of equation (4) the sum of the coefficients on the lagged inflation values equals unity, then there is a "natural rate" of the demand variable consistent with a constant rate of inflation. Current and lagged values of the unemployment gap are used as a proxy for the excess supply/demand parameter D_t, where the unemployment gap is defined as the difference between the actual rate of unemployment and a time varying natural rate (or NAIRU) as the literature provides evidence that the NAIRU "is not carved in stone, the NAIRU can move" (Gordon, 1997:11; Watson, 2014). Supply shocks are included to reduce the bias of the coefficient towards NAIRU or the output gap.

2.2. Linking inflation expectations and persistence

The inflation rates in equation (5) are modelled as an ARFIMA process, without restricting the sum of the coefficients of the lagged inflation rates to the equivalent of unity. If the estimated value of d is very close to unity, the unit root restriction will then be imposed to estimate again.

$$\phi(L)(1-L)^d \pi_t = \delta(L)(U - U^*)_t + \psi(L)X_t + \zeta(L)h_t^{\frac{1}{2}} + \theta(L)\varepsilon_t \qquad (5)$$

where d represents the inflation persistence driving factor, which is between zero and unity, U denotes the observed unemployment rate and U^* the unobserved

NAIRU. To obtain the time series of U^*in equation (5), we apply the Hodrick-Prescott filter with $\lambda = 14400$ (Hodrick and Prescott, 1997) to smooth the actual unemployment process. We are aware of the criticism that has been directed to the use of the Hodrick-Prescott filter as for instance by James Hamilton (2017). However, and as stated by Stephen Williamson:

> "there are typically medium-run changes in growth trends (e.g. real GDP grew at a relatively high rate in the 1960s, and at a relatively low rate from 2000-2012). If we are interested in variation in the time series only at business cycle frequencies, we should want to take out some of that medium-run variation. This requires that we somehow allow the growth trend to change over time. That's essentially what the HP filter does".[2]

The slope coefficient δ shows the trade-off between inflation and unemployment. ψ stands for the impact of supply shocks on inflation. Low estimated values of $|\delta|$ and ψ indicate that inflation has a weaker link with unemployment and with supply shocks.

Meanwhile autoregressive conditional heteroscedasticity (ARCH) effects may be present in the inflation dynamics which could affect the inflation level. Also, the inflation level may affect the volatility of inflation, which could in turn affect the inflation level. ζ captures the in-mean effects implying how the inflation level is

[2] http://newmonetarism.blogspot.co.uk/2012/07/hp-filters-and-potential-output.html

affected by the volatility of inflation, and γ reflects the impact of the inflation level on the volatility of inflation. The innovations ε_t are assumed Gaussian with mean zero and standard deviation $h_t^{1/2}$ conditional on information set up to time *t*-1, following a Nelson's (1991) EGARCH process which does not impose a positivity restriction on equation (6)

$$ln(h_t) = \omega_0 + \alpha(L)g(z_t) + \beta(L)lnh_t + \gamma(L)\pi_{t-1} \qquad (6)$$

where $g(z_t) = \theta_1 z_t + \theta_2[|z_t| - E|z_t|]$, z_t denotes a sequence of iid random variables with zero mean and unit variance, and $\varepsilon_t = z_t h_t^{1/2}$ as defined by Engle (1982); the formulation for $g(z_t)$ allows for the sign and size of z_t to have distinct impacts on the inflation volatility – that is, a negative innovation may differ from a positive innovation; h_t is the conditional variance, the log specification making sure that the conditional variance is positive. $\phi(L) = 1 - \sum_{i=1}^{n} \phi_i L^i$, $\alpha(L) = \sum_{i=1}^{q} \alpha_i L_i$, $\beta(L) = \sum_{i=1}^{p} \beta_i L^i$, and all the roots of $\phi(L), \alpha(L), and\ \beta(L)$, lie outside the unit circle. The parameter θ_1 captures the leverage effects when $\theta_1 < 0$ and lnh_t responds symmetrically to z_t when $\theta_1 = 0$. Note, $E|z| = \sqrt{\frac{2}{\pi}}$ under the assumption that ε_t is normally distributed and the MLE is computed by the following logarithm likelihood function

$$L(\mu, d, \phi, \delta, \theta, \omega, \alpha, \beta, \gamma) = -\frac{T}{2}\log 2\pi - \frac{1}{2}\sum_{t=1}^{T}(\log h_t + \frac{\varepsilon_t}{h_t})$$

By capturing both exogenous and intrinsic effects, our model (equations 5 and 6), is flexible and capable of linking inflation expectations and persistence.

3. Results

3.1. The Data

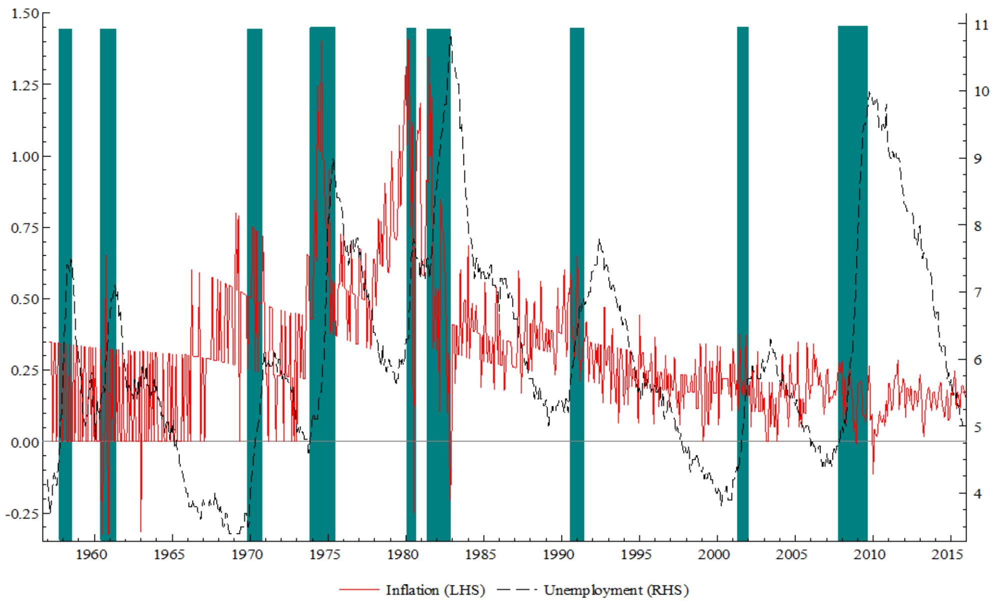
We use monthly data, starting in January 1957, taking into account the regime change identified by Barsky (1987), and running to December 2015. We measure inflation by the natural log difference of the core CPI (consumer price index for all urban consumers, all items less food and energy) – that is, 100ΔlogCPI. The core CPI and the civilian unemployment rate are seasonally adjusted and taken from the US Bureau of Labor Statistics. The data were downloaded from the Federal Reserve Economic Database (FRED)[3]. Our estimates are generated by using the package 'Time Series Modelling v4.49 by James Davidson'[4]. Supply shocks are proxied by the oil price measured by the relative spot price of the West Texas Intermediate (WTI) blend of crude oil obtained from the Dow Jones & Company data service. The path of the oil price relative to core inflation influences the stance of the Fed's monetary policy in the short-run, responding to incoming temporary supply shocks (Mishkin, 2007). For examining the Fed's key short-run objective of stabilizing the

3 Both inflation and unemployment data can be found at: https://fred.stlouisfed.org/series/CPIAUCSL; for unemployment https://fred.stlouisfed.org/series/LNU03008636 and https://fred.stlouisfed.org/series/CNP16OV

4 http://www.timeseriesmodelling.com/

economy, we divide the data into six subsamples according to the business cycles defined by the National Bureau of Economic Research (NBER) (table A.2). In this way, the chosen procedure has the effect of "smoothing out the peaks and valleys in output and employment around their long-run growth path" (FRBSF, 2004: 5) (Figure 1).

FIGURE 1. US INFLATION, UNEMPLOYMENT AND BUSINESS CYCLE



Notes:
Green bars are the recessions in the US as defined by NBER.
LHS is left-hand scale and RHS is right-hand scale.
Sources: Bureau of Labor Statistics, National Bureau of Economic Research.

As shown in Table 1, average US monthly inflation rates were notably lower in the 1960s, 1990s, 2000s and 2010s than in the 1970s and 1980s. ARCH effects are present in the residuals in all sample periods except for the 2000s sample.

TABLE 1 - US MONTHLY INFLATION DESCRIPTIVE STATISTICS

Sample	Obs.	Mean	Std. Dev	ARCH_LM(2)
1960s	156	0.209	0.215	12.949 [0.000]
1970s	120	0.544	0.272	60.096

				[0.000]
1980s	120	0.461	0.293	58.740 [0.000]
1990s	120	0.255	0.118	15.263 [0.000]
2000s	120	0.177	0.085	1.045 [0.355]
2010s	72	0.141	0.064	5.903 [0.004]
1957:01-2015:12	708	0.304	0.251	565.69 [0.000]

Notes:
Obs. and Std. Dev denote the number of observations and standard deviations respectively.
The numbers in brackets are *p*-values.

Table 2 highlights that in the 2010s, average unemployment which stood at 7.574 percent was higher than in 1980s and the highest level in three decades – and with the highest standard deviation.

Table 2 - US unemployment and unemployment gap descriptive statistics

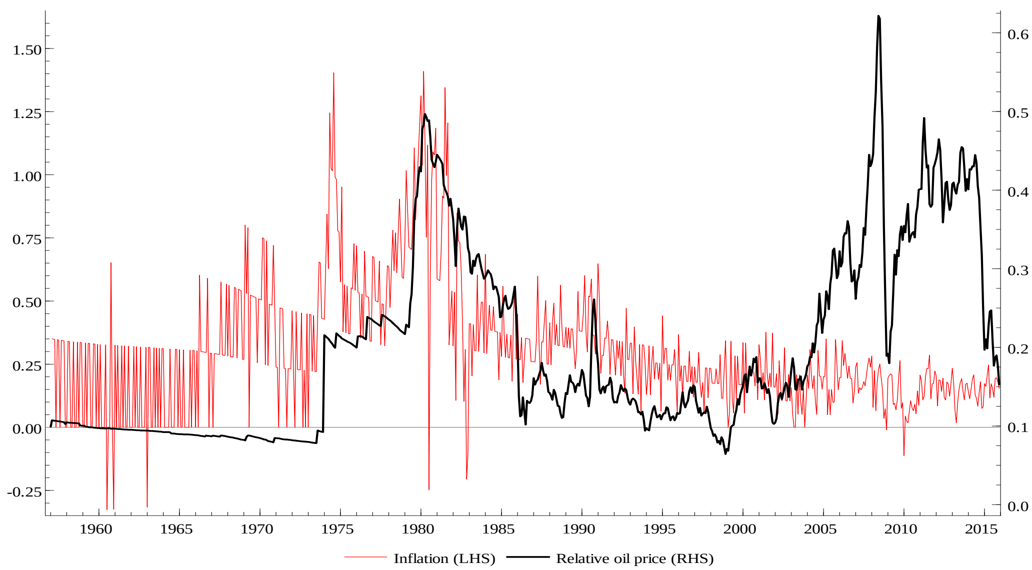
Sample	Obs.	Mean		Std. Dev	
		Unemployment	Gap	Unemployment	Gap
1960s	156	4.953	-0.325	1.129	5.668
1970s	120	6.218	0.116	1.159	6.504
1980s	120	7.273	0.217	1.475	5.396
1990s	120	5.763	0.261	1.045	2.598
2000s	120	5.541	-0.182	1.441	5.293
2010s	72	7.574	0.106	1.523	2.719
1957:01-2015:12	708	6.064	0.009	1.571	5.088

Notes:
Obs. and Std. Dev denote the number of observations and standard deviations respectively.

Figure 2 plots the inflation rate and relative oil prices. Before 1974, the oil price remained low and stable, and had no apparent effects on inflation. Twice during the 1970s, the oil price increased sharply (the so-called oil shocks) and correlation was observed between inflation rates and relative oil prices. Subsequently no correlation was observed – either when the oil price fell in the mid-1980s (Trehan, 2005; Hooker, 2002) or when the oil price rose in 2004-08, and again from 2009 through 2011: in none of these episodes was there a corresponding inflation response (respectively, downwards or upwards). This observation is in line with the findings of Mallick and Mohsin (2016), which evidenced that the US inflationary shocks are driven by monetary variables.

Kiley (2015) observes that the average pace of inflation since 2008 is almost similar for the overall and core CPIs and deduces that the oil price is not a factor in explaining the average pace of inflation; in contrast with Coibion and Gorodnichenko (2015) who explain the 'missing' disinflation during the Great Recession as being the result of the rise in oil prices during 2009-2011. On balance, therefore, the influence of the oil price, although reduced, may persist to some degree (Gordon, 2011).

FIGURE 2. US INFLATION AND OIL PRICES



Notes:
LHS is left hand scale and RHS is right hand scale.
Sources: Bureau of Labor Statistics, Dow Jones & Company.

To identify such a stylised fact as the persistence of inflation dynamics, several unit root tests are employed. The Phillips-Perron (PP) test is used for the null hypothesis of a unit root against the alternative of stationarity. In contrast, the null hypothesis in the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) test is that the series is stationary, that is I(0), which is based on the statistic $\eta = \sum_{t=1}^{T} S_t^2/(T^2 s_0)$ with $S_t = \sum_{i=1}^{t} u_i$ and s_0 being an estimator of the residual spectrum at frequency zero. Unlike the two threshold tests, the HML (Harris, McCabe and Leybourne, 2008) test is for the null hypothesis of short memory against long memory alternatives, that is the test of I(0) against I(d).

Table 3 reports three-unit root tests for inflation – PP, KPSS and the HML tests. The PP test statistics for three subsamples (1960s, 1990s and 2000s) are significant at the 1 percent level, while the KPSS statistics imply that the tests reject the null of stationarity at the 1 percent level for the 1960s,1980s and 1990s, at 5 percent for the 1970s and at 10 percent for the 2000s and 2010s subsamples. The unit root tests in Table 3 show that the postwar US inflation process does not entirely follow a random walk all the time. HML tests reject the null of inflation following an I(0) process for all the sub periods. Our results suggest that the inflation process is best described as I(d), rather than I(1) or I(0), and that an ARFIMA is the proper methodology to assess the integrability of this series.

TABLE 3. UNIT ROOT TESTS FOR INFLATION

Sample	PP H_0: I(1) $Z(t_{\hat{\alpha}})$	KPSS H_0: I(0) η_μ	HML H_0: I(0) $\hat{S}_k$
1960s	-11.686	0.970	3.218
1970s	-5.311	0.514**	3.458
1980s	-5.690	0.784	3.408

1990s	-9.095	0.966	3.671
2000s	-10.120	0.212***	3.761
2010s	-4.835	0.323***	2.808

Notes:
$Z(t_{\hat{\alpha}})$ and η_μ are Phillips-Perron adjusted statistic and LM statistic respectively, using the Parzen Kernel estimation method with Newey-West bandwidth and drift. $\hat{S}_k$ is HML statistic with c=1 and L=0.66. The statistics are all significant at 1 percent level except for those with asterisks.
**Significant at 5 percent level.
***Significant at 10 percent level.

3.2. Estimates

The results of estimating equations 5 and 6 by maximizing the log-likelihood function are presented in Table 4. The robustness of these estimates is demonstrated by using alternative time periods and models with different lagged inflation levels and volatility. The preferred specification is selected using the Akaike information criterion (AIC).

Our model considered both long memory and conditional heteroscedasticity. The estimated values of *d* are all between zero and 0.5 and significantly distant from 0 or 1 with small standard error, implying that each subsample of the inflation process exhibits a long memory feature. The estimations of ζ and γ do not support significant interactions between the inflation level and inflation volatility, and therefore are not reported. All the roots of ϕ(L), α(L), and β(L), lie outside the unit circle, satisfying our model's specifications.

TABLE 4 - PHILLIPS CURVE-EGARCH-IN-MEAN-LEVEL, ESTIMATION RESULTS

	1960s	1970s	1980s	1990s	2000s	2010s
d	0.129**	0.23**	0.26*	0.209*	0.203**	0.321**
	(0.061)	(0.118)	(0.072)	(0.068)	(0.1)	(0.138)
ϕ	$0.216^{(5)}$***	$0.302^{(6)}$**	$0.149^{(12)}$**	$-0.318^{(1)}$*	$0.255^{(8)}$**	$-0.426^{(12)}$*
	(0.114)	(0.127)	(0.075)	(0.104)	(0.109)	(0.102)

δ	-0.044	-0.141*	-0.161*	$0.016^{(2)}$	0.035	-0.068
	(0.037)	(0.029)	(0.047)	(0.045)	(0.025)	(0.048)
ψ	0.149	$0.242^{(1)}$*	0.032	0.083**	0.024***	-
	(0.14)	(0.03)	(0.075)	(0.037)	(0.014)	
ζ	-	$0.216^{(1)}$	$-0.293^{(1)}$	-	-	-
		(0.513)	(0.587)			
α	0.809*	$0.222^{(2)}$*	0.877*	$0.205^{(3)}$	-	-
	(0.115)	(0.09)	(0.081)	(0.253)		
β	0.888*	0.554**	-	-	-	-
	(0.166)	(0.295)				
Γ	-	-	-	-	-	-
Q(12)	33.746	5.976	17.424	14.034	4.497	6.248
	[0.010]	[0.014]	[0.010]	[0.010]	[0.034]	[0.012]
$Q^2(12)$	4.888	7.176	14.4	13.123	9.849	2.706
	[0.978]	[0.893]	[0.346]	[0.438]	[0.706]	[0.999]
AIC	-307.799	-210.351	-191.506	-139.125	-137.914	56.319

Notes*:*
Standard errors and *t* probabilities are given respectively in parentheses and brackets.
*Significant at 1 percent level.
**Significant at 5 percent level.
***Significant at 10 percent level.
Q (12) and Q^2 (12) are the Box Pierce tests based on residuals and squared residuals.
ϕ only reports the last lag of the AR term.
The superscript denotes the number of lagged terms.
In this formulation, θ_2 is set to be 1.

It is worth noting that the monthly inflation dynamic in this paper exhibits ARCH effects, which is captured through our EGARCH-in-mean-level approach. As in table 4, the coefficients of α and β are well presented up to 1980s. In addition, no significant interactions between the inflation rate and its volatilities could be found. This may indirectly affect the inflation-unemployment trade-off factor. This finding corroborates the conclusion of various researchers that changes in inflation do not influence the relation between inflation and unemployment (Mourougane and Ibaragi, 2004; Mihailov, Rumler and Scharler, 2011; Blanchard, 2016).

The estimation results of the slope coefficient δ are negative, demonstrating the Phillips curve effect in each individual decade except for the 1990s and 2000s samples. Notably, the slope coefficient δ has changed as persistence *d* varies over decades. The estimations of ψ show the impact of the oil price on inflation over the various decades. Our findings suggest that the order of inflation integration, from the highest to the lowest were as follows: 2010s, 1980s, 1970s, 1990s, 2000s and 1960s.

During both 2010s and 1980s samples, inflation was moving relatively slowly in response to a shock. A marked contrast emerges, therefore, between those two periods, in which inflation displayed a higher degree of persistence, and all the other periods. Comparing these two episodes of relatively more persistent inflation, the 2010s sample stands out, as inflation did not decline as much as during the 1980s sample despite the sharp recessions common to both periods.

In the 2010s, *d* reached 0.321, δ is negative at -0.068, monthly inflation and unemployment rates averaged respectively 0.14 percent and 7.57 percent and inflation did not respond to the oil price. Persistent inflation in the 1980s, with *d* amounting to 0.26, led to a steep Phillips curve with δ reaching -0.161, that is a unit decrease of unemployment deviating from NAIRU was followed respectively by a 16 percent increase in inflation, the oil price had a minimal effect with ψ= 0.032.

Most authors agree that the differences between those two periods may be attributed to inflation expectations having become strongly anchored at low levels by the 2010s. While during the 1980s the memory of the high inflation of the previous decade had resulted in inflation expectations being what is conventionally described as “unanchored”, though it may be more precise to say that expectations were set (anchored) at high levels. Watson (2014) for instance explains the missing disinflation and the relatively higher persistence by inflation expectations being more anchored during 2007-2013 than during 1980-1985. Blanchard (2016: 33) remarked that “at very low rates of inflation, people may not focus on inflation, and thus may not adjust expectations in response to movements in inflation”. This remark is supported by studies of countries with very low inflation such as Switzerland or Japan (Mourougane and Ibaragi, 2004; Mihailov, Rumler and Scharler, 2011).

In the 1980s, after enduring stagflation in the 1970s, the Fed announced a major policy shift by tightening the money supply to lower inflation in 1979 (see the historical accounts of the events by Bordo and Siklos, 2015). This policy move was followed by a short recession lasting six months. Monthly inflation and unemployment rates averaged respectively 0.46 percent and 7.3 percent respectively (tables 2 and 3). There were long lags between the introduction of the new monetary policy regime comprising more aggressive counter-inflationary policies and the actual fall in the inflation rate (Hardouvelis and Barnhart, 1989). Several authors suggest a link between the relatively high inflation persistence and inadequate Fed credibility (Erceg and Levin, 2003).

In the 1970s sample, the estimated d amounts to 0.23, the PC was steep with δ reaching -0.141 that is a unit decrease of unemployment deviating from NAIRU was followed by a 14 percent increase in inflation. The inflation process was strongly affected by the increase in the oil price (in 1973-1974 and 1979), a unit rise of the oil price resulted in about a 0.24 increase in inflation.

Our findings concur with the literature using several different econometric approaches for the period 1990s, 2000s and 1960s where a relatively lower degree of persistence is observed as for instance Erceg and Levin (2003), Taylor (2000), Evans and Wachtel (1993), Cogley and Sargent (2001, 2005). Most authors attribute the change in inflation persistence to the anchoring of inflation expectations by the central bank's commitment to an inflation target (Benati, 2008; 2015; Cogley and Sargent, 2005 and Williams, 2006).

In the 1990s, lower persistence is observed with $d = 0.209$, inflation exhibits a positive correlation with unemployment with $\delta = 0.016$ and is not responsive to the oil price with $\psi = 0.083$. Our estimates suggest that the coefficient of the Phillips curve is not significant. Inflation began to decline at the end of the second recession in November 1982. During the period, January 1984-December 1989, monthly inflation and unemployment rates averaged respectively 0.36 percent and 6.44 percent, signifying that the US had re-entered a period of low inflation (Belton Jr. and Cebula, 1998).

In October 1989, the Fed announced Zero-Inflation Resolution (Greenspan, 1989) as a way of confirming its determination to maintain price stability. The longest economic expansion of any ten-year period since the end of WWII ensued, with monthly inflation and unemployment rates averaging 0.255 percent and 5.76 percent respectively. The estimated inflation persistence in the 2000s declined very little with d=0.203, δ is positive and amounts to 0.035, and inflation is hardly responsive to the oil price with ψ= 0.024. Although inflation exhibited a positive correlation with unemployment in the 1990s and 2000s, our estimates suggest that the coefficient of the Phillips curve is not significant over the last three decades. This does not necessarily mean the failure of the Phillips curve as empirically speaking, this insignificance of the coefficient can be explained by the difficulty of rejecting the null of a zero coefficient. One explanation might be that the inflation-unemployment trade-off weakens in a low-inflation environment (e.g., Ball, Mankiw and Romer, 1988); but Stock and Watson (2010: 32) remark that a smaller slope parameter at low inflation levels is not robustly confirmed by statistical tests.

We observe the lowest degree of persistence (d =0.129) for the sample of the 1960s as well as the Phillips curve having a flattening tendency with δ = -0.044. That was a period when improving the medium-to long-term inflation stabilization tradeoff with depressed real activity was not much of a policy concern. Policymakers thought that a particular choice of inflation implied a particular choice of unemployment. The Phillips Curve was a level-level relation in the sense that a certain level of unemployment will correspond to a certain level of inflation. But the absence of expectations in the models of academics and policy-makers does not mean that public expectations were not an actual factor in real life. The impact of the oil price on inflation was estimated as being insignificant. From 1957 through to the end of 1960s, two recessions occurred (1958 and 1961), lasting eight months and ten months respectively (table A.2). After the second recession ended in February 1961, there was an expansion lasting 106 months. During this period, the Fed successfully maintained inflation at a low and stable level with an average monthly inflation rate of about 0.21 percent (table 2) and an average monthly unemployment rate of about 5 percent (table 3).

4. Concluding remarks

While we have limited our analysis to a statistical approach, our findings are consistent with the literature supporting a decline in inflation persistence during the periods following the Great Inflation and a relatively higher degree of persistence in the 2010s sample. Our results suggest, in line with the theoretical interpretation of d, that the relatively higher value of d in the 2010s relative to the 1980s was due to inflation expectations being more anchored in the 2010s.

But if we look at the entire set of our results, we find that inflation expectations were more anchored in the 1970s, when d=0.23, than in the 1990s (d=0.209) and 2000s (d=0.203). Here we see the effect on expectations of the memory of inflation experienced in the preceding decade(s). This would mean, paradoxically, that the Arthur Burns Fed enjoyed relatively higher credibility in a high inflation environment than the Alan Greenspan Fed in a low inflation environment. Taken as a whole, our results – including the firmer anchoring of inflation expectations by the 2010s (d=0.321) – are at odds with other authors' findings that inflation persistence declined after the 1990s (Beechey and Osterholm, 2009; Benati, 2008; Gadea and Mayoral, 2006).

In the introduction, we highlighted the important feature of our statistical modelling as being to encompass – over and above the varying shape of the Phillips curve – the complex cointegration of inflation and unemployment rates together with their time varying volatilities. Reviewing the estimates coming out of this modelling exercise, our findings suggest that the cointegration-based measure could indicate the scope for increased/reduced policy flexibility in line with higher/lower inflation persistence when pursuing the dual employment and price stability goal in a long-term perspective. This possibility looks especially applicable to the post-2010 sample, when the zero-lower bound for nominal interest rates became a relevant factor. Such circumstances may offer monetary policy makers the scope to take fuller advantage of firmly anchored inflation expectations.

Our model fills a gap since it has the flexibility to describe diverse dynamic properties, and to accommodate possible intrinsic as well as exogenous shocks. Thanks to these features, our modelling exercise has come up with a particularly interesting empirical finding that measures of credibility and inflation persistence are not always fully aligned. In other words, a holistic view of our results raises questions about the actual operation over time of the logical link between inflation expectations becoming more firmly anchored and inflation becoming more or less responsive to shocks – i.e. changes in persistence. Further research might fruitfully explore this timing misalignment between the credibility of monetary policy and inflation persistence.

Continuing research on the basis of our new approach to modelling inflation persistence might also benefit from two more detailed refinements. First, our analysis of US post-war inflation dynamics could be deepened by extending our single-equation reduced form into a stylised model, as this would allow for a more theoretical discussion of the time variation in inflation persistence. Second, our approach to persistence should be extended to the "new Keynesian Triangle Phillips Curve" featuring supply shock variables as in Malikane (2014), in order to check the robustness of the results to model specification. Finally, this article used core CPI for the measure of inflation, while a natural extension would be to consider different specifications. As argued by, for instance, Mallick and Mohsin (2016), a distinction between durable and non-durable goods inflation may provide better insight into inflation dynamics.

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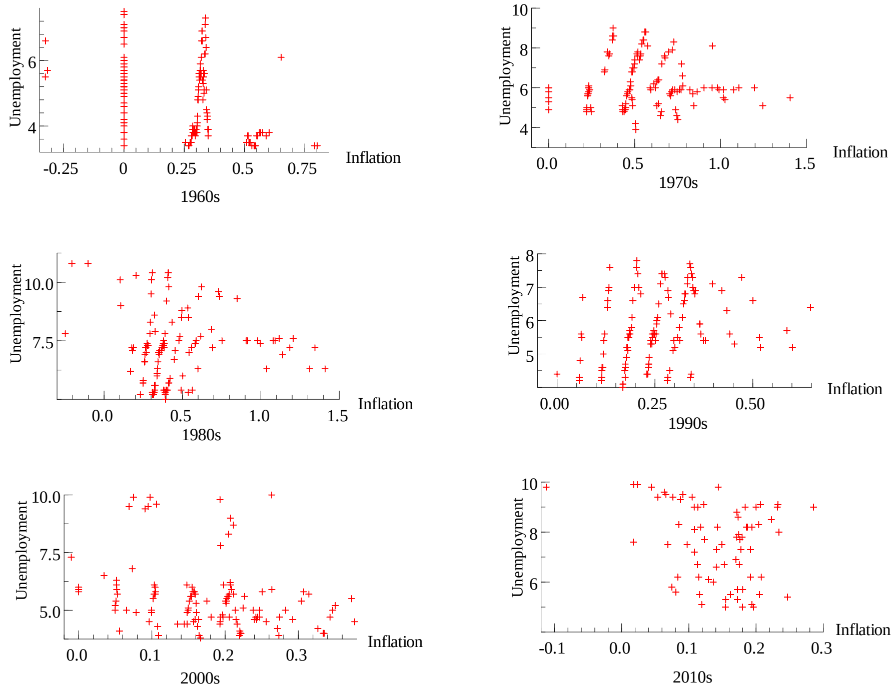
Appendices

FIGURE A.1 – INFLATION AND UNEMPLOYMENT FOR EACH SUBSAMPLE



SOURCE: BUREAU OF LABOR STATISTICS

Table A.1: Post-World War II US inflation persistence

Samples	Persistence	Authors
1960-1978	High	Barsky (1987); Brainard and Perry (2000); Taylor (2000); Kim, Nelson and Piger (2001); Zeng (2014)
1960s-mid 1980s	High	Fuhrer (2010)
1965-1993	Very high	Fuhrer and Moore (1995)
1965-2001	High and unchanged	Pivetta and Reis (2007)
Late 1960s-1970s	High	Cogley and Sargent (2001)
1970s	High	Beechey and Österholm (2012)
Late 1970s-1980s	High	Cogley and Sargent (2001)
Volcker-Greenspan era	Low	Brainard and Perry (2000); Taylor (2000); Kim, Nelson and Piger (2001); Beechey and Österholm (2012)
1980-2006	Low and changed	Williams (2006)
1980-2009	Low and changed	Zeng (2014)
Since early 1980s	A decline in inflation persistence	Cogley, Primiceri, and Sargent (2010); Kurozumi and Zandweghe (2018)

1984-2003	Low and changed	Levin and Piger (2004)
Since mid-1980s	High or low and changed	Fuhrer (2010)
1990s	Low	Cogley and Sargent (2001)
1965-early 1980s	High	Widely agreed
Since early 1980s	High or low, changed or unchanged	Disputed

Source: Authors' own elaboration

TABLE A.2 – US BUSINESS CYCLE (1957-2009)

Dates		Duration (Months)	
Peak	Trough	Contraction	Expansion
August 1957	April 1958	8	39
April 1960	February 1961	10	24
December 1969	November 1970	11	106
November 1973	March 1975	16	36
January 1980	July 1980	6	58
July 1981	November 1982	16	12
July 1990	March 1991	8	92
March 2001	November 2001	8	120
December 2007	June 2009	18	73

Source: Business Cycle Dating Committee, NBER. http://www.nber.org/cycles.html