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Asymmetric inflation dynamics: Evidence from quantile regression analysis

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ABSTRACT

This paper applies the regression quantile approach developed by Koenker and Xiao (2004) to investigate the dynamic behavior of inflation in 12 OECD countries. By analyzing the behavior in a wide range of quantiles, this method allows us to quantify the influence of various sizes of shocks that hit the inflation, and is able to capture possible asymmetric adjustment of the inflation towards to its long-run equilibrium. It therefore sheds new lights on the inflation dynamics compared with the conventional unit root methodologies. Our results suggest that generally, the inflation rates are not only mean-reverting but also exhibit asymmetries in their dynamic adjustments, in which large negative shocks tend to induce strong mean reversion, and on the contrary, large positive shocks do not. Policy implications related to the empirical findings are also provided.

1. Introduction

The behavior of macroeconomic variables has received considerable attention in the literature. In particular, inflation rate is one of the most important variables due to its key roles in many macroeconomic models (e.g., Levin and Piger, 2003; Angeloni et al., 2006), and to the association with monetary policies implemented by the authorities whose objective is to deliver price stability (e.g., Zhang and Clovis, 2010). In addition, as argued by, inter alia, Lee and Wu (2001), Ho (2009), and Henry and Shields (2004), the dynamic behavior of inflation rate has a number of economic implications.

In a large body of empirical literature, the time-series properties of inflation are investigated by using unit root testing procedure assuming constant dynamics, which focus on the conditional central tendency (e.g., Rose, 1988; Ng and Perron, 2001; Tsong and Lee, 2010). Under such assumption, the speed of inflation adjustment towards its equilibrium would be constant no matter how far the inflation is above or below its long-run level or how big the negative or positive shock hitting the inflation. As a result, the aforementioned studies at most can only distinguish inflation between a unit-root and a stationary process, lacking the ability to further elaborate on the inflation dynamics. Another important observation is that the distribution of inflation is often leptokurtic, differing significantly in shape from the Normal (Charemza and Hristova, 2005). In this case, as pointed out by Koenker and Xiao (2004), the commonly-used unit root tests can exhibit rather poor power performance, tending to bias test results in favor of a unit root.

In this paper, we seek to re-investigate the inflation dynamics by employing the quantile regression inference developed by Koenker and Xiao (2004), and to provide complementary information on the inflation adjustments. Such approach is used in recent empirical studies on real exchange rates (Nikolaou, 2008) and on nominal interest rates (Koenker and Xiao, 2004). The methodology has several advantages over the conventional counterparts. First, instead of only concentrating on the

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constant speed of adjustment with conditional mean function, this method allows for different and possibly asymmetric speed of adjustment at various quantiles. Consequently, the possibility of sign asymmetry, different adjustment mechanism towards the long-run equilibrium for positive or negative shocks, can be detected. Also, the tendency of mean reversion can be quantified based on the size of the shock hitting the inflation. As noted by Boldin (1999), taking into account possible asymmetric dynamics for a series can improve the performance of monetary policy.

Second, the quantile unit root tests can provide an alternative way to study local persistence in time series. To be specific, within certain quantiles, the inflation rate can display unit-root behavior, but within the others, its mean-reverting property may be sufficient to insure global mean reversion. By contrast, the inflation rate can merely be distinguished globally between nonstationarity and mean reversion with the conventional counterparts.

Finally, it offers a more flexible and refined analysis of the inflation dynamics by relaxing the assumption that the inflation rate follows a particular distribution. Moreover, the shocks analyzed are actual, whose magnitudes are determined endogenously by the data. Due to better data descriptions accommodating potential heavy-tailed behavior in inflation, this method can lead to substantial power gains over the conventional least squared-based counterparts as argued by Koenker and Xiao (2004). As a result, strong evidence in favor of mean-reverting inflation rate, as expected, is more likely to be uncovered.

Our empirical results suggest that the inflation rates in 12 OECD countries are globally meaning-reverting based on the quantile inference, in sharp contrast to the counterparts from the univariate unit root tests focusing on the average behavior. Moreover, both the sign and size of the shock, in general, have marked impact on the speed of inflation adjustment towards its long-run equilibrium. Specifically, the negative inflation shocks with large magnitude can induce strong mean-reverting tendencies, while the positive counterparts found at medium or extreme quantiles could have infinite lives for inflation rates. As a result, asymmetric mean reversion towards the equilibrium for inflation rates is prevalent. This implies that for effectively curbing inflation, policymakers should keep alert to monitoring potential inflation increase and take precautionary measures to anchor inflation expectations. When the inflation is hit by negative shocks, however, the authorities may prefer not to intervene in the inflation dynamics such that the inflation could revert back to its long-run equilibrium level eventually.

The remainder of the paper is organized as follows. In Section 2, we review the studies on the inflation dynamics. Section 3 presents the methodology used in this study. The empirical results are collected in Section 4. Section 5 concludes the paper.

2. Literature review

Since the pioneer work of Nelson and Plosser (1982), the issue that whether most macroeconomic time series can be described as a unit root process has motivated a large body of literature. The empirical studies, however, still cannot reach a consensus. With conventional univariate unit root or stationarity tests, Evans and Lewis (1995), Nelson and Schwert (1977), Crowder and Wohar (1999), Camarero et al. (2000), MacDonald and Murphy (1989), Ball and Cecchetti (1990), and Crowder and Hoffman (1996) fail to find strong evidence in favor of mean reversion in inflation rates. By employing the state-of-the-art univariate unit root tests with good asymptotic size and power properties, Ng and Perron (2001) still cannot uncover sufficient evidence supporting that shocks to inflation are short-lived. However, Rose (1988) rejects the unit-root null in US inflation, and Edwards (1988) reports strong persistence in Latin American inflation rates. In addition, Baillie et al. (1996) find that inflation rates in developed countries are fractionally integrated and mean-reverting.

It is well-documented in the literature that traditionally univariate unit root tests have low power against persistent alternatives, especially for small sample sizes encountered in empirical studies. To alleviate this problem, the panel-type tests and covariate tests are used in recent studies. The former tests exploit cross-sectional information to increase power, while the latter tests incorporate related covariates which contain valuable information to boost power. For example, Culver and Papell (1997) implement the panel unit root test of Levin et al. (2002) to the data of 13 OECD countries, and reject the unit-root null in inflation rate. By taking into account the cross-sectional dependency among individual countries, Lee and Wu (2001) provide strong empirical evidence to support the mean reversion of inflation rate with the test proposed by Im et al. (2003). In addition, Basher and Westerlund (2008) apply a battery of panel tests to the same data set in Culver and Papell (1997), and their results suggest that inflation rates are stationary after controlling for cross-sectional dependency and structural break. However, with the nonlinear IV panel unit root test suggested by Chang (2002) and Ho (2009) finds that the inflation rates may accelerate. In other words, shocks to inflation appear to be infinitely persistent. By carrying out a battery of covariate tests proposed by Hansen (1995), Elliott and Jansson (2003), Jansson (2004), and Tsong and Lee (2010) provide strong evidence supporting that the inflation process in 15 OECD countries displays mean reversion.

Another strand of literature further concentrates on the changes in the degree of inflation persistence, instead of only using unit root tests to distinguish the inflation process as either an $I(0)$ or $I(1)$ process. Taylor (2000) estimates the largest autoregressive root (LAR) and suggests that the US inflation persistence has been declined after the Volcker–Greenspan era. Considering a possible structural break, Levin and Piger (2003) showed that high inflation persistence is not intrinsic in industrialized countries. Similarly, with the estimation of a Bayesian VAR model, Cogley and Sargent (2001, 2005) claimed that the US inflation persistence has experienced a significant decline. Kumar and Okimoto (2007) investigate the dynamics of inflation persistence using long memory approach and find that there has been a marked decrease in US inflation persistence over the past two decades. By employing an ARMA model with time-varying autoregressive parameter, Beechey and

Österholm (2009) show that inflation persistence has fallen remarkably in the Euro area after January 1999 and inflation no longer displays non-stationary behavior. In addition, Chinese inflation over the post-1997 period tends to revert more quickly to its long-run level than over the pre-1997 period as is exhibited in Zhang and Clovis (2010).

On the contrary, Stock (2001) estimated the LAR with rolling window estimation method, and concluded that there is no indication of a clear reduction in US inflation persistence. Similar conclusions are drawn by Pivetta and Reis (2007), but with both Bayesian and rolling window approach to the estimation of the LAR. With the employment of rolling regressions on split samples, Batini (2006) finds that European inflation is rather inertial; more importantly, the inflation persistence seems to have varied only marginally over the past 30 years. Similar results are presented in O'Reilly and Whelan (2005) as well, in which the estimates of persistence parameter in Euro area are generally close to one and rather stable over time.

3. Empirical methodology

In this section we briefly describe Koenker and Xiao's (2004) quantile regression framework employed in the present paper, which enjoys power gains over the augmented Dickey–Fuller (ADF) test when the shock exhibits heavy-tailed behavior. More importantly, this testing procedure enables us to explore the speed of mean reversion for a series under different magnitudes and signs of the shock. To be precise, this methodology is capable of unveiling possibly different mean-reverting patterns by explicitly testing for a unit root at different quantiles, in which the underlying series is hit by a shock with various sizes and signs.

Let us consider the ADF regression model given by:

$$y_t = \alpha_1 y_{t-1} + \sum_{j=1}^{q} \alpha_{j+1} \Delta y_{t-j} + u_t, \quad t = 1, 2, \ldots, n, \tag{1}$$

where $y_t = \pi_t - \mu$, with π_t and μ denoting the inflation rate and its long-run equilibrium value, respectively; u_t is iid random variable with zero mean and constant variance. In this model, the AR coefficient α_1 measures the persistence of y_t. If $\alpha_1 = 1$, then y_t follows a unit root process, and if $|\alpha_1| < 1$, then the behavior of y_t displays mean reversion. Following Koenker and Xiao (2004) and based on Eq. (1), the τ th conditional quantile of y_t, conditional on the past information set $\Im_{t-1}$, can be expressed as a linear function of y_{t-1} and lagged values of Δy_t as follows:

$$Q_{y_t}(\tau|\Im_{t-1}) = x_t' \alpha(\tau), \tag{2}$$

where $x_t = (1, y_{t-1}, \Delta y_{t-1}, \ldots, \Delta y_{t-q})'$ and $\alpha(\tau) = (\alpha_0(\tau), \alpha_1(\tau), \ldots, \alpha_{q+1}(\tau))'$, with $\alpha_0(\tau)$ denoting the τ th quantile of u_t. It is important to note that $\alpha_1(\tau)$ measures the speed of mean reversion of y_t within each quantile, and is dependent on the τ th quantile under investigation.

For a given τ, the parameter vector $\alpha(\tau)$ in Eq. (2) is estimated by minimizing sum of asymmetrically weighted absolute deviations:

$$\min \sum_{t=1}^{n} (\tau - I(y_t < x_t' \alpha(\tau))) |y_t - x_t' \alpha(\tau)|, \tag{3}$$

where I denotes an indicator function, i.e., $I = 1$ if $y_t < x_t'\alpha(\tau)$, and $I = 0$, otherwise. Given the solution of Eq. (3), denoted by $\hat{\alpha}(\tau)$, Koenker and Xiao (2004) suggest testing the time-series properties of y_t within the τ th quantile by using the following t ratio statistic:

$$t_n(\tau) = \frac{\hat{f}(F^{-1}(\tau))}{\sqrt{\tau(1-\tau)}} (Y_{-1}' P_X Y_{-1})^{1/2} (\hat{\alpha}_1(\tau) - 1), \tag{4}$$

where $\hat{f}(F^{-1}(\tau))$ is a consistent estimator of $f(F^{-1}(\tau))$, with f and F representing the density and distribution function of u_t in Eq. (1), Y_{-1} is the vector of lagged dependent variables (y_{t-1}), and P_X is the projection matrix onto the space orthogonal to $X = (1, \Delta y_{t-1}, \ldots, \Delta y_{t-q})$. According to Koenker and Xiao (2004), $\hat{f}(F^{-1}(\tau))$ can be written as $\hat{f}(F^{-1}(\tau)) = (\tau_i - \tau_{i-1}) / x_t'(\hat{\alpha}(\tau_i) - \hat{\alpha}(\tau_{i-1}))$ with $\tau_i \in \Gamma$. We choose $\Gamma = \{0.1, 0.2, \ldots, 0.9\}$ in our empirical study. By employing the test statistic $t_n(\tau)$, we can examine the unit root properties of the series by looking at its behavior in each quantile. In other words, this not only enables us to take a closer look at the dynamics of the series, but also to investigate possibly different mean reverting behavior when the series is hit by different magnitudes and signs of shock at different quantiles. By contrast, the widely-used unit root tests only focus on the conditional central tendency and lack the ability to elaborate on such behavior. In addition, with $\hat{\alpha}_1(\tau)$ we calculate the half-lives (HLs) of a shock hitting the inflation rate within the quantile with the formula $\ln(0.5)/\ln(\hat{\alpha}_1(\tau))$.[2]

[2] As noted in Murray and Papell (2005), OLS and the median-unbiased method can be used to estimate the AR coefficient α_1 in Eq. (1). Then HLs can be calculated with the formula $\ln(0.5)/\ln(\hat{\alpha}_1)$ or through impulse response function. Though the median-unbiased method is more robust in conditional mean specification, to the best of our knowledge, its property is still unknown for quantile autoregression models. Therefore, this method is not employed in this paper. We thank the referee for this insightful comment.

Another approach to generally analyze the unit root behavior based on the quantile framework involves examining the nonstationary properties over a range of quantiles, instead of only focusing on the selected quantile. For this purpose, Koenker and Xiao (2004) recommend the quantile Kolmogorov–Smirnov (KS) test, which is given by

$$QKS = \sup|t_n(\tau)|, \tag{5}$$

where $t_n(\tau)$ is the t ratio statistic defined in Eq. (4). In practice, we calculate $t_n(\tau)$ at $\tau \in \Gamma$, and thus the QKS test can be constructed by taking maximum over Γ.

The limiting distributions of the $t_n(\tau)$ and QKS tests are nonstandard, and depend on nuisance parameters. Koenker and Xiao (2004) suggest using a re-sampling procedure to approximate their small-sample distributions as described below.

(1) Fit the following q-order autoregression with Δy_t by ordinary least squares (OLS):

$$\Delta y_t = \sum_{j=1}^{q} \hat{\beta}_j \Delta y_{t-j} + \hat{u}_t, \tag{6}$$

and obtain estimates $\hat{\beta}_j$ for $j = 1, 2, \ldots, q$, as well as the residuals $\hat{u}_t$.

(2) Draw a bootstrap sample of u_t^* with replacement from the empirical distribution of the centered residuals $\tilde{u}_t = \hat{u}_t - (n-q)^{-1}\sum_{t=q+1}^{n} \hat{u}_t$.

(3) Generate the bootstrap sample of Δy_t^* recursively using the fitted autoregression given by

$$\Delta y_t^* = \sum_{j=1}^{q} \hat{\beta}_j \Delta y_{t-j}^* + u_t^*, \tag{7}$$

with $\hat{\beta}_j$ being OLS estimates in Eq. (6), and initial values $\Delta y_j^* = \Delta y_j$ for $j = 1, 2, \ldots, q$.

(4) A bootstrap sample of y_t^* can be obtained based on

$$y_t^* = y_{t-1}^* + \Delta y_t^*, \tag{8}$$

with $y_1^* = y_1$.

(5) With the re-sample y_t^*, compute the bootstrap counterparts of $\hat{\alpha}_0(\tau)$, $\hat{\alpha}_1(\tau)$, the $t_n(\tau)$ and QKS tests, denoted by $\hat{\alpha}_0^*(\tau)$, $\hat{\alpha}_1^*(\tau)$, $t_n^*(\tau)$ and QKS*, respectively.

(6) Repeat Steps 2 to 5 NB times. NB is 500 in this paper.

(7) Compute the empirical distribution function of the NB values of $\hat{\alpha}_0^*(\tau)$, $\hat{\alpha}_1^*(\tau)$, the $t_n^*(\tau)$ and QKS* tests, and use these empirical distribution functions as an approximation to the cumulative distribution functions of the respective tests under the null.

(8) Using the bootstrap p-value to make inference. Also, the bootstrap confidence intervals for $\hat{\alpha}_0(\tau)$ and $\hat{\alpha}_1(\tau)$ can be accordingly obtained from the empirical distribution functions of $\hat{\alpha}_0^*(\tau)$ and $\hat{\alpha}_1^*(\tau)$, respectively.

4. Empirical investigation

4.1. Data description and preliminary results

Quarterly observations of consumer price indices (CPI) are retrieved from the International Monetary Fund's IFS data base for a period ranging from 1957Q1 to 2010Q1. The 12 OECD countries considered include Australia, Denmark, France, Germany, Ireland, Italy, Japan, New Zealand, Spain, Sweden, the United Kingdom (UK), and the United States (US). The CPI is transformed into annual inflation rates π_t with

$$\pi_t = 400(\ln(\mathrm{CPI_t}) - \ln(\mathrm{CPI}_{t-1})). \tag{9}$$

Table 1 reports the first four sample moments of the inflation rates and the results of the Jarque–Bera (JB) normality tests. Spain has the largest long-run mean for inflation (7.157), while Germany has the smallest counterpart (2.740). The largest and smallest sample standard deviations are 6.407 and 2.513 for Ireland and Germany, respectively. Note also that the sample correlation between these inflation means and standard deviations is as high as 0.815, conforming to the results that high inflation may accompany high inflation uncertainty (e.g., Okun, 1971; Davis and Kanago, 1988; Daal et al., 2005, to name a few). More importantly, all the inflation rates exhibit fat tail and non-normality since the JB test overwhelmingly rejects the null hypothesis of normality with extremely small p-values, consistent with the results in Charemza and Hristova (2005). The significant evidence of non-normality in the inflation rates lends strong support to the employment of the quantile regression approach in this study as emphasized by Koenker and Xiao (2004).

Though our focus is on the results of quantile regression, the counterparts of the conventional unit root tests such as DF-GLS and MZ$_\alpha$-GLS are also included for the sake of comparison. Since it is well documented in the literature that the inflation rate might be described as an MA process with a negative root, to reduce serious size distortions with more correct model specification, the lag length for the two tests are selected by the modified Akaike information criterion (MAIC) proposed by Ng and Perron (2001) with maximum lag set at 16.

Table 2 collects these testing results. In general, the two univariate unit-root tests fail to uncover broad evidence in favor of stationary inflation rates, in line with previous studies such as Rapach and Weber (2004) and MacDonald and Murphy (1989), among others. The DF-GLS test rejects the unit-root null at the 5% level for Australia and Germany, with one additional rejection for the US at the 10% level. For the MZ$_\alpha$-GLS test, only Australia and Germany are rejected at the 5% and 10%

Table 5
Results for quantile unit root test on inflation rates for eurozone countries.

Country	τ	0.1	0.2	0.3	0.4	0.5	0.6	0.7	0.8	0.9
France	$\alpha_0(\tau)$	−2.680	−1.471	−1.023	−0.559	−0.155	0.217	0.723	1.243	2.332
	p-value	0.000**	0.000**	0.000**	0.000**	0.154	0.095*	0.001**	0.000**	0.000**
	$\alpha_1(\tau)$	0.603	0.775	0.806	0.868	0.916	0.927	0.974	0.956	1.114
	p-value	0.002**	0.002**	0.000**	0.010**	0.200	0.188	0.742	0.604	0.992
	Half-lives	1.371	2.713	3.219	4.905	7.924	9.122	26.342	15.427	∞
	QKS/*p*-value	4.048/0.000**	Start date for euro launch = 1999Q1							
Germany	$\alpha_0(\tau)$	−2.934	−1.624	−1.208	−0.721	−0.164	0.640	1.039	1.543	2.942
	p-value	0.000**	0.000**	0.000**	0.000**	0.261	0.001**	0.000**	0.000**	0.000**
	$\alpha_1(\tau)$	0.301	0.350	0.344	0.352	0.385	0.387	0.360	0.309	0.191
	p-value	0.000**	0.000**	0.000**	0.000**	0.000**	0.000**	0.000**	0.000**	0.000**
	Half-lives	0.577	0.660	0.649	0.663	0.726	0.730	0.678	0.590	0.419
	QKS/*p*-value	6.570/0.000**	Start date for euro launch = 1999Q1							
Ireland	$\alpha_0(\tau)$	−4.212	−2.896	−1.960	−1.355	−0.268	0.517	1.342	3.084	4.966
	p-value	0.000**	0.000**	0.000**	0.000**	0.242	0.069*	0.001**	0.000**	0.000**
	$\alpha_1(\tau)$	0.554	0.623	0.680	0.756	0.818	0.939	1.021	1.187	1.176
	p-value	0.006**	0.000**	0.000**	0.002**	0.014**	0.572	0.902	1.000	0.986
	Half-lives	1.172	1.467	1.794	2.480	3.443	11.082	∞	∞	∞
	QKS/*p*-value	5.402/0.000**	Start date for euro launch = 1999Q1							
Italy	$\alpha_0(\tau)$	−2.784	−1.866	−1.176	−0.552	0.016	0.347	0.813	1.711	2.856
	p-value	0.000**	0.000**	0.000**	0.009**	0.469	0.071*	0.006**	0.000**	0.000**
	$\alpha_1(\tau)$	0.795	0.792	0.839	0.862	0.959	0.944	1.014	1.066	1.042
	p-value	0.002**	0.002**	0.006**	0.006**	0.442	0.314	0.916	0.984	0.926
	Half-lives	3.022	2.976	3.949	4.675	16.556	12.002	∞	∞	∞
	QKS/*p*-value	3.859/0.012**	Start date for euro launch = 1999Q1							
Spain	$\alpha_0(\tau)$	−3.934	−2.775	−1.826	−1.247	−0.507	0.376	1.170	1.919	5.121
	p-value	0.000**	0.000**	0.000**	0.000**	0.078*	0.170	0.006**	0.001**	0.000**
	$\alpha_1(\tau)$	0.726	0.787	0.733	0.782	0.829	0.885	0.886	0.921	1.012
	p-value	0.042**	0.008**	0.000**	0.006**	0.034**	0.170	0.252	0.642	0.882
	Half-lives	2.162	2.900	2.227	2.815	3.686	5.660	5.745	8.435	∞
	QKS/*p*-value	4.147/0.004**	Start date for euro launch = 1999Q1							

Notes: See notes below Table 3.

may prefer not to intervene in the inflation through monetary policies since it can return to its long-run equilibrium level eventually.

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