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Stock prices, inflation and inflation uncertainty in the U.S.: Testing the long-run relationship considering Dow Jones sector average indexes

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Abstract

We test for the long-run relationship between stock prices, inflation and its uncertainty for different U.S. sector stock indexes, over the recent time-span, namely 2006M1-2015M5. For this purpose we use a cointegration analysis with one structural break to capture the crisis effect, and two alternative measures of inflation uncertainty relying on a time-varying unobserved component model. In line with recent empirical studies we discover that in the long-run, the inflation and its uncertainty negatively impact the stock prices, opposed to the well-known Fisher effect. However, in the short-run the results are mixed, providing evidence for complex interdependences between stock prices, inflation and its uncertainty. Our results are robust regarding the use of a bounded or unbounded inflation trend for measuring the uncertainty, and a slight difference is noticed between different sector indexes.

Keywords: stock prices, inflation uncertainty, cointegration with structural breaks, U.S.

JEL codes: C22, E31, G15.

1. Introduction

Studies regarding the impact of inflation on asset prices have a long tradition in the financial field. The starting point is considered the “The Theory of Interest” by Fisher (1930), who advanced the idea that expected nominal return on an asset should equal the expected real return and the expected inflation (the so-called “Fisher effect”). However, the empirical literature hardly succeeded to demonstrate this effect, even if this was the common view before the 1970s. As such, noteworthy works demonstrated on contrary, that the inflation negatively impact the stock returns (Bodie, 1976; Jaffe and Mandelker, 1976; Nelson, 1976; Fama and Schwert, 1977), while subsequent papers find explanation for a potential negative relationship.

First, Modigliani and Cohn (1979) suggest that stock market investors fail to understand the inflation’s effects on the nominal cash flow, and during the periods of high inflation, the valuation errors induce an undervaluation in stocks, and *vice-versa*. Second, Feldstein (1980) shows that the inflation generates artificial capital gains, which are subject to taxation. Therefore, if the inflation increases the firms face higher tax liabilities. In this context, rational investors reduce common stock valuation to take into account the effect of inflation, which negatively affect the stock prices (tax-effects hypothesis). Third, Fama (1981) shows that the inverse relationship between real stock returns and expected inflation is generated by a positive relationship between equity market and the output growth, combined with a negative relationship between expected inflation and real economic activity (the proxy hypothesis). Fourth, Geske and Roll (1983) and Kaul (1987) explain the negative relationship using the argument of a counter-cyclical monetary policy. They suggest that positive shocks to real output generate monetary tightening which reduces the inflation, while triggering an increase in stock prices. Fifth, Hess and Lee (1999) explain the negative relationship as a combination of demand and supply disturbances. Sixth, Boudoukh and Richardson (1993) followed by others, provide a more complex explanation, arguing that stock returns are inversely related to realized and expected inflation in the short-run, but may be positively related to inflation in the long-run. However, in the same spirit, Sharpe (2002) draws a new perspective on the relationship between stock prices and inflation, underlining the existence of a potential negative relationship in the long-run also. He shows that a rise in expected inflation is accompanied by either a decline in expected long-run real earnings, or by a rise in the long-run real return, or both.

A different strand on literature put accent on the role of inflation uncertainty to explain the link between stock prices and inflation. Starting with [Levi and Makin \(1979\)](#) and [Kolluri \(1982\)](#), economist describe the Friedman effect¹ and its role in a Fisherian-type relationship. Starting from the Friedman's second reasoning which shows that the inflation uncertainty negatively impact the real output, [Hu and Willett \(2000\)](#) and [Park and Ratti \(2000\)](#) demonstrate that expectations of an output decline depress current stock prices, effect reinforced by greater economic uncertainty during high inflationary periods.

However, the existing empirical literature addressing the relationship between inflation and stock prices makes abstraction of inflation uncertainty, with few exceptions (for a discussion see [Azar, 2013](#)). Moreover, the mixed results and the lack of distinction between the short- and long-run horizons in estimating the nexus between stock prices, inflation and its uncertainty require supplementary investigations. Therefore, the present paper's contributions to the empirical literature are three fold.

First, we focus on the characteristics of the inflation and its uncertainty, which prove to be non-stationary processes, and on the endogeneity which exists between stock prices, inflation and its uncertainty. Given these evidence, we use a cointegration approach, focusing on the recent period and employing U.S. statistics over the period 2006M1 to 2015M5 (Dow Jones index). In order to underline the effect of the recent financial crisis we resort to the Gregory-Hansen cointegration test with one structural break ([Gregory and Hansen, 1996a](#)). The cointegration analysis allows to investigate the above-mentioned relationship at different time-horizons, and to explore the hypotheses advanced by [Sharpe \(2002\)](#). We focus on stock price level and not on returns as [Cochran and Defina \(1993\)](#) and [Alexakis et al. \(1996\)](#) did.

Second, different from previous studies which associate the inflation uncertainty with the inflation variability and use different generalized autoregressive conditional heteroskedasticity (GARCH)-type models, we consider that expected inflation is generally unobservable (see [Cukierman and Meltzer, 1986](#); [Kolluri and Wahab, 2008](#)). Further, the use of GARCH-type models to measure inflation uncertainty requires stationary inflation rates, which is not the case. Therefore, we use two alternatives measures of inflation uncertainty based on the unobserved component model. On the one hand, the [Stock and Watson's \(2007\)](#)

¹ The link between inflation and its uncertainty on the one hand, and between the inflation uncertainty and output on the other hand, became famous with the Friedman's Nobel lecture ([Friedman, 1977](#)). The first hypothesis of Friedman, showing the role of inflation uncertainty in explaining the level of inflation, was formalized by [Ball \(1992\)](#) (we call this Friedman – Ball hypothesis). Afterwards, several competing hypothesis were advanced and become famous, showing the positive impact of inflation on its uncertainty ([Cukierman and Meltzer, 1986](#)), or on contrary, a negative relationship where the inflation leads uncertainty ([Pourgerami and Maskus, 1987](#)), or where the inflation is leaded by its uncertainty ([Holland, 1995](#)).

approach, called the unobserved component model with stochastic volatility, where the inflation trend is not unbounded. On the other hand, we employ the recently proposed time-varying unobserved component model which allows for the trend inflation to be bounded in specific limits ([Chan et al., 2013](#)). The second approach is recommended for measuring uncertainty during inflation-targeting regimes. Although the Federal Reserve is not formally inflation targeting, it is believed to be an implicit inflation targeter ([Thornton, 2012](#)). [Bullard \(2012\)](#) argue that starting with January 2012, the Federal Open Market Committee (FOMC) explicitly named a numerical inflation target for the U.S. of 2%, as measured by the personal consumption expenditures price index. In this context it is believed that trend inflation and the long-term inflation expectations are not allowed to evolve in an unbounded fashion.

Third, existing studies, however, do not focus on stock markets sector indexes. Nevertheless, the impact of inflation and its uncertainty on stock prices might differ for various categories of prices (i.e. free-market established prices versus regulated prices, raw material versus final consumption goods, etc.). Therefore, we use in our analysis the Dow Jones Composite Average index and three sector indexes namely, industrial, transportations and utilities.

The rest of the paper is structure as follows. Section 2 presents a brief literature review of papers addressing the link between stock prices and inflation, and between stock prices, inflation and its uncertainty. Section 3 describes the research methodology. Section 4 presents the data and the empirical results, while the last section concludes.

2. Empirical literature review

The empirical evidence suggests the presence of complexities, regarding the stock prices – inflation nexus, and it is heavily oriented toward the U.S. case. While oldest empirical studies examined this relationship at relatively short horizons, subsequent works focused on longer time-horizons ([Boudoukh and Richardson, 1993](#); [Solnik and Solnik, 1997](#); [Schotman and Schweitzer, 2000](#)), and find evidence for the Fisher effect. On contrary, other studies like [Engsted and Tanggaard \(2002\)](#) discover that the Fisher effect diminishes with the horizon's increases.

Recent empirical works focus on nonlinearities which may exist between stock prices, inflation and its uncertainty ([Boyd, 2001](#); [Kim, 2003](#); [Liu et al., 2005](#); [Maghyereh, 2006](#); [Karagianni and Kyrtsoy, 2011](#)). While for example [Liu et al. \(2005\)](#) employ a regime-switching model, [Maghyereh \(2006\)](#) resorts to a nonparametric cointegration test. Further,

Karagianni and Kyrtsov (2011) use a recurrence quantification analysis and a series of tests for structural breaks and nonlinear causality, documenting negative nonlinear linkages between the inherent dynamics of inflation and stock returns. Further, several studies combine the time-frequency domain resorting to wavelets. Kim and In (2005) use a six level decomposition of the U.S. inflation and stock returns and show that there is a positive relationship at the shortest and longest time-scales, while at intermediate scales stock return and inflation are negatively correlated. More recently, Tiwari et al. (2015) use a continuous wavelet transform and document a positive relations for higher time scales, which however lacks in robustness when a different measure of inflation is used.

Few studies investigate the link between stock prices, inflation and its uncertainty (see Alexakis et al., 1996; Azar, 2013). Along these, Alexakis et al. (1996) examines the link between inflation uncertainty and stock prices for a group of developed and emerging economies over the period 1980M1 to 1993M12, and report a negative relationship. Recently, Azar (2013) shows that inflation uncertainty dominates the inflation in explaining stock prices in the case of the U.S. over the time-span 1950M1-2011M3. However, both variables become redundant when other fundamental variables are included in the regression.

In line with these papers we investigate the nexus between stock prices, inflation and its uncertainty for the U.S. However, different from previous works, we focus on the long-run relationship and perform a sector-level analysis. Moreover, we use two newly proposed measures of inflation uncertainty, based on time-varying unobserved component models.

3. Methodology

3.1. Measuring the inflation uncertainty

While many researches associate the inflation uncertainty with the inflation volatility and use different GRACH-type approaches for computing the volatility, other studies relate the inflation uncertainty to the inflation gap. In the second case the unobserved component (UC) model is employed.

3.2.1. Stock and Watson's (2007) approach

Stock and Watson (2007) propose a generalized form of the UC model in which the variances of the permanent and transitory disturbances evolve randomly over time. The new model, called the unobserved component model with stochastic volatility (UC-SV), is the following:

$$\pi_t = \tau_t + \eta_t, \text{ where } \eta_t = \sigma_{\eta,t} \zeta_{\eta,t} \quad (1)$$

$$\tau_t = \tau_t + \varepsilon_t, \text{ where } \varepsilon_t = \sigma_{\varepsilon,t} \zeta_{\varepsilon,t} \quad (2)$$

$$\ln \sigma_{\eta,t}^2 = \ln \sigma_{\eta,t-1}^2 + v_{\eta,t} \quad (3)$$

$$\ln \sigma_{\varepsilon,t}^2 = \ln \sigma_{\varepsilon,t-1}^2 + v_{\varepsilon,t} \quad (4)$$

where: π_t is the level of inflation, τ_t is the inflation stochastic trend, and the η_t represents the inflation gap, with the property that $\lim_{j \rightarrow \infty} E_t[\eta_{t+j}] = 0$.

In this framework, a constant parameter model supposes $\tau_t = \tau$ and $\eta_t = \rho_1 \eta_{t-1} - \dots - \rho_p \eta_{t-p} + \varepsilon_t$, with $\varepsilon_t \sim N(0, \sigma_\eta^2)$. A particular, simply case of this model, is a local trend model with stochastic volatility in the inflation equation, which imposes that any unit root behavior manifests in the inflation trend component:

$$\begin{aligned} \tau_t &= \tau_{t-1} + \varepsilon_t^\tau \\ \eta_t &= \varepsilon_t \exp\left(\frac{h_t}{2}\right) \\ h_t &= h_{t-1} + \varepsilon_t^h \end{aligned} \quad (5)$$

where: $\varepsilon_t^\tau \sim N(0, \sigma_\tau^2)$, $\varepsilon_t \sim N(0, 1)$ and $\varepsilon_t^h \sim N(0, \sigma_h^2)$, errors which are independent of one another at all leads and lags.

Defined in this manner, the trend inflation is a driftless random walk, which evolves in an unbounded fashion. [Stock and Watson \(2007\)](#) generalize the model allowing for the inflation trend to vary over time. Therefore, the stochastic volatility in the innovation to the inflation trend is:

$$\varepsilon_t^\tau \sim N(0, \exp(g_t)), \quad (6)$$

where: $g_t = g_{t-1} + \varepsilon_t^g$ and $\varepsilon_t^g \sim N(0, \sigma_g^2)$.

3.2.2. *Chan et al. 's (2013) approach*

The unbounded model of [Stock and Watson \(2007\)](#) might be inconsistent if the monetary authority decides to intervene if inflation moves outside of a desirable range, that is, if an inflation targeting regime is in place. To overcome this limit, [Chan et al. \(2013\)](#) build upon [Cogley et al. \(2010\)](#), who propose a different specification for the time-varying inflation process, which allows for stochastic volatility in the state equation and not in the inflation trend.

The [Cogley et al. 's \(2010\)](#) inflation specification is:

$$\pi_t = \phi_{0t-1} + \phi_{1t-1} \pi_{t-1} + \varepsilon_t \exp\left(\frac{h_t}{2}\right), \quad (7)$$

where: $h_t = h_{t-1} + \varepsilon_t^h$ and $\phi_t = \phi_{t-1} + \varepsilon_t^\phi$ represents the vector of coefficients.

The authors restrict the autoregressive parameter to lies inside the unit circle, and the inflation gap becomes:

$$\eta_t = \tau_{t+1} - \frac{\phi_{0t}}{1-\phi_{1t}} \quad (8)$$

where: $\eta_{t+1} = \phi_{1t} \left(\pi_t - \frac{\phi_{0t}}{1-\phi_{1t}} \right) + \varepsilon_{t+1} \exp \left(\frac{h_{t+1}}{2} \right)$.

Chan et al. (2013) advance in their turn an unobserved components model with an autoregression in the transitory component, considering the first order autoregression case:

$$(\pi_t - \tau_t) = \rho_t(\pi_{t-1} - \tau_{t-1}) + \varepsilon_t \exp \left(\frac{h_t}{2} \right) \quad (9)$$

where: $\tau_t = \tau_{t-1} + \varepsilon_t^\tau$, $h_t = h_{t-1} + \varepsilon_t^h$ and $\rho_t = \rho_{t-1} + \varepsilon_t^\rho$, while $\varepsilon_t \sim N(0,1)$ and $\varepsilon_t^h \sim N(0, \sigma_h^2)$.

For bounding the inflation trend, the authors propose an innovation in the state equation:

$$\varepsilon_t^\tau \sim TN(a - \tau_{t-1}, b - \tau_{t-1}; 0, \sigma_\tau^2) \quad (10)$$

where: $TN(a, b; \mu, \sigma^2)$ is the Gaussian distribution with the mean μ and the variance σ^2

truncated to the interval (a, b) and $E_t[\tau_{t+1}] = \tau_t + \sigma_\tau \left[\frac{\phi\left(\frac{a-\tau_t}{\sigma_\tau}\right) - \phi\left(\frac{b-\tau_t}{\sigma_\tau}\right)}{\Phi\left(\frac{b-\tau_t}{\sigma_\tau}\right) - \Phi\left(\frac{a-\tau_t}{\sigma_\tau}\right)} \right]$ is the conditional expectation with $a \leq \tau_t \leq b$.

Chan et al. (2013) propose the following innovation in the inflation gap (ρ_t):

$$\varepsilon_t^\rho \sim TN(a_\rho - \rho_{t-1}, b_\rho - \rho_{t-1}; 0, \sigma_\rho^2) \quad (11)$$

where: the conditional expectation became $E_t[\rho_{t+1}] = \rho_t + \sigma_\rho \left[\frac{\phi\left(\frac{a_\rho - \rho_t}{\sigma_\rho}\right) - \phi\left(\frac{b_\rho - \rho_t}{\sigma_\rho}\right)}{\Phi\left(\frac{b_\rho - \rho_t}{\sigma_\rho}\right) - \Phi\left(\frac{a_\rho - \rho_t}{\sigma_\rho}\right)} \right]$, with

$a_\rho \leq \rho_t \leq b_\rho$, and the bounds are generated in order to satisfy the inflation gap long-run zero convergence condition (ρ_t is limited to lies inside the unit circle or within constant limits).

3.2. Gregory-Hansen cointegration test

Our general equation is:

$$LNDJ = c + \alpha_1 I + \alpha_2 U + \varepsilon_t \quad (12)$$

where: $LNDJ$ is the natural log of the DJ composite and sector average indexes (DJCA, DJIA, DJTA, DJUA); c is the intercept; I is the CPI inflation; U represents the inflation uncertainty (U1, U2); ε_t are the error terms.

In order to capture the long-run relationship between variables, usually c and α are considered time-invariant. However, Gregory and Hansen (2006a) consider that, if the

cointegration holds over some periods of time, it may shift toward a new long-run relationship. They treat the timing and shifts as unknown and allows changes in the intercept and slope, defining the following dummy variable:

$$\varphi_{t\tau} = \begin{cases} 0, & \text{if } t \leq [n\tau] \\ 1, & \text{if } t > [n\tau] \end{cases} \quad (13)$$

The cointegration test proposed by [Gregory and Hansen \(2006a\)](#) accommodates a single endogenous break. The authors propose three models with assumptions about structural breaks in the intercept and slope. Afterwards, in [Gregory and Hansen \(2006b\)](#) a forth model is advanced, allowing for breaks in the trend also. Based on these four models, our equations become:

Model 1 – [Gregory and Hansen's \(1996a\)](#) test: Level shift (GH-LS)

$$LNDJ = c_1 + c_2\varphi_{t\tau} + \alpha_1 I + \beta\alpha_2 U + \varepsilon_t \quad (14)$$

Model 2 – [Gregory and Hansen's \(1996a\)](#) test: Level shift with trend (GH-LST)

$$LNDJ = c_1 + c_2\varphi_{t\tau} + \beta t + \alpha_1 I + \alpha_2 U + \varepsilon_t \quad (15)$$

Model 3 – [Gregory and Hansen's \(1996a\)](#) test: Regime shift (GH-RS)

$$LNDJ = c_1 + c_2\varphi_{t\tau} + \alpha_1 I + \alpha_2 I\varphi_{t\tau} + \alpha_3 U + \alpha_4 U\varphi_{t\tau} + \varepsilon_t \quad (16)$$

Model 4 – [Gregory and Hansen's \(1996b\)](#) test: Regime shift with trend change (GH-RST)

$$LNDJ = c_1 + c_2\varphi_{t\tau} + \beta_1 t + \beta_2 t\varphi_{t\tau} + \alpha_1 I + \alpha_2 I\varphi_{t\tau} + \alpha_3 U + \alpha_4 U\varphi_{t\tau} + \varepsilon_t \quad (17)$$

The test statistics computation in [Gregory and Hansen \(1996a\)](#) is based on the Ordinary Least Square (OLS) approach and each model yields the residuals $\hat{\varepsilon}_{t\tau}$, where the subscript τ shows that the residual sequence depends on the choice of the change point τ . The first-order serial correlation coefficient is:

$$\hat{\rho}_\tau = \sum_{t=1}^{n-1} \hat{\varepsilon}_{t\tau} \hat{\varepsilon}_{t+1\tau} / \sum_{t=1}^{n-1} \hat{\varepsilon}_{t\tau}^2 \quad (18)$$

The second-stage residuals are defined as $\hat{v}_{t\tau} = \hat{\varepsilon}_{t\tau} - \hat{\rho}_\tau \hat{\varepsilon}_{t-1\tau}$, while the estimate of long-run variance of $\hat{v}_{t\tau}$ is $\hat{\sigma}_\tau^2 = \hat{\gamma}_\tau(0) + 2\hat{\lambda}_\tau$. The estimate of the bias-corrected first-order serial correlation coefficient is:

$$\hat{\rho}_\tau^* = \sum_{t=1}^{n-1} (\hat{\varepsilon}_{t\tau} \hat{\varepsilon}_{t+1\tau} - \hat{\lambda}_\tau) / \sum_{t=1}^{n-1} \hat{\varepsilon}_{t\tau}^2 \quad (19)$$

[Gregory and Hansen \(1996a\)](#) propose three test statistics, namely two Phillips statistics and one ADF statistic. The Phillips statistics are the following:

$$Z_\alpha(\tau) = n(\hat{\rho}_\tau^* - 1) \quad (20)$$

$$Z_t(\tau) = (\hat{\rho}_\tau^* - 1) / \hat{s}_\tau \quad (21)$$

where: $\hat{s}_\tau^2 = \hat{\sigma}_\tau^2 / \sum_{t=1}^{n-1} \hat{\varepsilon}_{t\tau}^2$.

The ADF statistics is calculated by regressing $\Delta\hat{\varepsilon}_{t\tau}$ upon $\hat{\varepsilon}_{t-1\tau}$ and $\Delta\hat{\varepsilon}_{t-1\tau}, \dots, \Delta\hat{\varepsilon}_{t-K\tau}$ for suitably chosen lag truncation K . As such, the ADF statistics is the t -statistics for the regressor $\hat{\varepsilon}_{t-1\tau}$, denoted by:

$$ADF(\tau) = tstat(\hat{\varepsilon}_{t-1\tau}) \quad (22)$$

4. Data and empirical findings

4.1. Data

Data are extracted from FED St. Louis (FRED database), having a monthly frequency and covering the period 2006M1 to 2015M5. The DJ average indexes data are available in the FRED database starting with 2005M12 only. This constraint forced us to limit the analysis to the indicated time-span, which however covers the pre- and post-crisis period. The CPI inflation is considered, and four DJ average indexes expressed in natural log (Industrial, Utility, and Transportation). The DJCA index is a combination of all other three major DJ sector average indexes. Each DJ sector average index is made up of a select group of prominent stocks. DJIA includes 30 companies acting in the manufacture of industrial and consumer goods but also financial services, entertainment and information technology. DJIA is one of the most representative indexes of the U.S. stock market and by extension, of the U.S. economy. DJTA is the oldest American index which includes different transportation companies. Given the fact that transportations represent the link between the manufacturers of goods and the customers, it needs to confirm the DJIA trend. Finally, the DJUA index refers to electricity and natural gas companies. In the U.S. the utilities' prices have been significantly "deregulated" starting with the 1990s. Compared to the other sector indexes, the utility index composition has remained largely unaltered over time.

4.2. Unit root tests results

For checking the presence of unit roots in our series, we resort to the classic Augmented Dickey–Fuller test (ADF) and Phillips–Perron test (PP). The ADF test has a low power against the null and usually if the test indicates that the variables are stationary in first difference, there is no point to use alternative tests that have more power against the null. However, because the Gregory-Hansen's approach is an extension of unit root tests with structural breaks, we also use the Zivot and Andrews's test with one structural break (ZA) for assessing the presence of unit roots.

Table 1 shows that the ADF and PP tests cannot reject the null of unit root presence in level, for any of the selected variables. On contrary the stationarity is documented in the first difference for all variables, although to a smaller extent for the [Chan et al.'s \(2013\)](#) inflation uncertainty variable (U2). The ZA test presents similar findings except for the stock prices for which the stationarity is documented in level also. All in all we conclude that our variables are non-stationary in level but stationary in first difference and we can use the cointegration analysis.

Table 1. Unit root tests

Variables	ADF test		PP test	
	<i>Level</i>	<i>First difference</i>	<i>Level</i>	<i>First difference</i>
DJCA	-0.20	-9.92***	-0.21	-10.7***
DJIA	0.05	-8.49***	0.12	-8.38***
DJTA	-0.28	-8.82***	-0.64	-8.90***
DJUA	-0.78	-8.73***	-1.03	-8.70***
I	-1.79	-6.22***	-2.50	-6.10***
U1	3.32	-4.33***	1.41	-4.50***
U2	-0.91	-1.84	-1.74	-2.84*
Variables	ZA test			
	<i>Level</i>	<i>Break</i>	<i>First difference</i>	<i>Break</i>
DJCA	-9.64***	2007M9	-10.3***	2007M10
DJIA	-8.96***	2007M9	-10.3***	2007M10
DJTA	-4.79*	2008M9	-9.88***	2009M4
DJUA	-6.29***	2008M8	-9.20***	2009M4
I	-4.30	2008M10	-6.95***	2009M8
U1	-1.01	2007M10	-6.91***	2013M11
U2	-3.48	2010M7	-6.01***	2009M4

Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively and rejecting H_0 of unit root presence; (ii) the natural log of stock indexes is considered.

4.3. Cointegration test

4.3.1. Stock and Watson's (2007) uncertainty results

The results of the Gregory-Hansen cointegration tests for the [Stock and Watson's \(2007\)](#) inflation uncertainty (U1) are presented in Table 2. Several findings can be advanced. First, for all the indexes except for transportation (DJTA), two out of four models (level shift with trend and regime shift with trend changes) confirm the cointegration relationship. Second, there is a large agreement between the ADF and Phillips statistics. Third, it appears that the cointegration between stock prices, inflation and inflation uncertainty is stronger for the DJIA and DJCA as compared to DJUA, which can be explained by the fact that energy prices are not completely "deregulated". Finally, in the case of DJTA, the cointegration appears for the level shift model, both in the case of Phillips and ADF tests, but only at 10% and respectively 5% level of significance. However, there is no consensus between the two Phillips tests,

which put in question the long-run relationship between our variables in the case of the DJTA index. The last observation shows that the transportation sector is not very affected by the level of inflation and its uncertainty probably because the fuel prices are not highly correlated with overall consumption prices in the U.S.

Table 2. Gregory-Hansen cointegration tests – Stock and Watson’s (2007) inflation uncertainty

<i>Models</i>	ADF	Break	Z_t	Break	Z_a	Break
DJCA						
<i>GH-LS</i>	-3.20	2008M8	-3.68	2008M8	-16.7	2008M8
<i>GH-LST</i>	-9.09***	2007M5	-7.29***	2010M5	-70.0***	2010M5
<i>GH-RS</i>	-3.11	2010M1	-3.57	2010M1	-17.2	2010M1
<i>GH-RST</i>	-9.06***	2012M3	-7.32***	2007M5	-71.5***	2007M5
DJIA						
<i>GH-LS</i>	-3.19	2008M10	-3.68	2008M8	-16.6	2008M8
<i>GH-LST</i>	-8.72***	2007M5	-5.81***	2010M12	-50.8*	2010M12
<i>GH-RS</i>	-3.07	2010M1	-3.52	2010M1	-17.7	2010M1
<i>GH-RST</i>	-9.18***	2008M9	-7.43***	2007M5	-73.5**	2007M5
DJTA						
<i>GH-LS</i>	-4.85*	2013M3	-5.25**	2013M3	-34.0	2013M3
<i>GH-LST</i>	-4.84	2008M4	-5.16*	2013M3	-34.5	2013M3
<i>GH-RS</i>	-4.86	2013M3	-5.27*	2013M3	-34.1	2013M3
<i>GH-RST</i>	-4.70	2011M5	-5.00	2008M10	-34.7	2008M10
DJUA						
<i>GH-LS</i>	-3.64	2012M9	-3.91	2012M8	-19.4	2012M8
<i>GH-LST</i>	-6.05***	2008M11	-5.86***	2008M10	-53.5*	2008M10
<i>GH-RS</i>	-3.84	2008M4	-4.03	2008M5	-22.6	2008M5
<i>GH-RST</i>	-6.91***	2008M9	-6.94***	2008M10	-67.3*	2008M10

Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively and rejecting H_0 of no cointegration; (ii) the natural log of stock indexes is considered.

4.3.2. Chan et al. ’s (2013) uncertainty results

In order to check for the robustness of these findings and to compare the results in the case of alternative uncertainty definition, we proceed to test the cointegration relationship using the U2 uncertainty measure proposed by Chan et al. (2013). Table 3 shows similar results for the DJCA, DJIA and DJUA, where the cointegration is noticed for the level shift with trend model, and for the regime shift with trend model. However, different from the U1 results, for the U2 in the case of DJTA, the cointegration appears for the regime shift model, but only at 10% level of significance. All in all we conclude that our results are robust.

Table 3. Gregory-Hansen cointegration tests – Chan et al.'s (2013) inflation uncertainty

<i>Models</i>	<i>ADF</i>	<i>Break</i>	<i>Z_t</i>	<i>Break</i>	<i>Z_u</i>	<i>Break</i>
DJCA						
<i>GH-LS</i>	-4.10	2007M12	-4.03	2008M1	-23.7	2008M1
<i>GH-LST</i>	-8.98***	2007M6	-7.23***	2007M11	-70.0***	2007M11
<i>GH-RS</i>	-3.75	2007M12	-3.77	2007M7	-21.6	2007M7
<i>GH-RST</i>	-11.2***	2007M5	-8.13***	2007M7	-82.2***	2007M7
DJIA						
<i>GH-LS</i>	-4.05	2007M12	-3.97	2008M1	-23.4	2008M1
<i>GH-LST</i>	-8.65***	2007M6	-5.93***	2007M9	-52.0*	2007M9
<i>GH-RS</i>	-3.75	2007M12	-3.82	2007M7	-21.4	2007M7
<i>GH-RST</i>	-11.8***	2007M7	-10.7***	2007M5	-114.***	2007M5
DJTA						
<i>GH-LS</i>	-4.36	2013M12	-4.49	2013M8	-30.6	2013M8
<i>GH-LST</i>	-4.29	2008M10	-4.53	2008M11	-33.0	2008M11
<i>GH-RS</i>	-5.42*	2013M4	-5.48*	2013M5	-40.8	2013M5
<i>GH-RST</i>	-5.58	2011M5	-5.64	2011M6	-44.0	2011M6
DJUA						
<i>GH-LS</i>	-2.80	2013M6	-3.07	2013M11	-14.0	2013M11
<i>GH-LST</i>	-6.05***	2008M11	-5.85***	2008M8	-53.5*	2008M8
<i>GH-RS</i>	-3.29	2013M5	-3.47	2013M5	-17.2	2013M5
<i>GH-RST</i>	-6.61***	2009M1	-6.34**	2009M4	-55.5	2009M4
<i>Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively and rejecting H0 of no cointegration; (ii) the natural log of stock indexes is considered.</i>						

4.4. Cointegration equations

Tables 2 and 3 show that two out of four models reject the null hypothesis of no cointegration for DJCA, DJIA and DJUA. Therefore, in order to select the best model we proceed to the estimation of cointegration equations for these two models, using the Engle–Granger method and the OLS technique. For DJTA index the cointegration equation is computed for that model for which both the ADF and Phillips statistics confirm the existence of the long-run relationship. The results for U1 are presented in Table 4 below.

In the case of the three indexes where the cointegration is documented by the GH-LST and GH-RST models, we notice that the regime shift model with trend (GH-RST) better explains the influence of inflation and its uncertainty on stock prices. In all the cases the trend and the dummy variable are significant, meaning that this relationship is nonlinear over time. Nevertheless, several differences appear when we compare the DJCA, DJIA and DJUA indexes. On the one hand, for the DJIA both the inflation and its uncertainty negatively impact the stock prices. On the other hand, in the case of DJCA only the inflation has a negative impact, while in the case of DJUA the uncertainty is the one influencing the stock prices. In all the situations the impact is negative, showing that investors are sensitive to macroeconomic environment and to the uncertainty regarding the macroeconomic context (i.e. the general level of prices), pointing against the Fisher effect. This evidence also point in

the favor of the Friedman – Ball and Cukierman – Meltzer hypotheses, which sustain that the level of inflation and its uncertainty move in the same direction. Finally, in the case of DJTA a different situation appears. For the GH-LS model the inflation has a positive impact on the stock prices, while the uncertainty has a negative influence.

Table 4. Cointegration equations between stock prices, inflation and Stock and Watson's (2007) inflation uncertainty

DJCA	GH-LS (Dummy 2008M8)	GH-LST (Dummy 2010M5)	GH-RS (Dummy 2010M1)	GH-RST (Dummy 2007M5)
Intercept		8.210*** (0.004)		8.250*** (0.045)
Dum × Intercept		0.000 (0.001)		-0.041 (0.046)
Trend		0.008*** (0.000)		0.007*** (0.000)
Dum × Trend				0.000** (0.000)
I		-0.000 (0.000)		-0.005** (0.002)
Dum × I				0.005*** (0.002)
U1		-0.000 (0.001)		-0.009 (0.023)
Dum × U1				0.009 (0.023)
DJIA	GH-LS (Dummy 2008M8)	GH-LST (Dummy 2010M12)	GH-RS (Dummy 2010M1)	GH-RST (Dummy 2007M5)
Intercept		9.283*** (0.005)		9.483*** (0.045)
Dum × Intercept		0.001 (0.002)		-0.204*** (0.045)
Trend		0.009*** (0.000)		0.007*** (0.000)
Dum × Trend				0.001*** (0.000)
I		-0.000 (0.000)		-0.010*** (0.002)
Dum × I				0.010*** (0.002)
U1		-0.000 (0.002)		-0.075*** (0.023)
Dum × U1				0.075*** (0.023)
DJTA	GH-LS (Dummy 2013M3)	GH-LST (Dummy 2013M3)	GH-RS (Dummy 2013M3)	GH-RST (Dummy 2008M10)
Intercept	8.951*** (0.081)			
Dum × Intercept	0.391*** (0.034)			
Trend				
Dum × Trend				
I	0.065*** (0.087)			
Dum × I				
U1	-0.312*** (0.038)			
Dum × U1				
DJUA	GH-LS (Dummy 2012M8)	GH-LST (Dummy 2008M10)	GH-RS (Dummy 2008M5)	GH-RST (Dummy 2008M10)
Intercept		6.188*** (0.048)		6.450*** (0.143)
Dum × Intercept		-0.405*** (0.023)		-0.677*** (0.164)
Trend		0.006*** (0.000)		0.009*** (0.001)
Dum × Trend				-0.003** (0.001)
I		0.001 (0.004)		0.004 (0.014)
Dum × I				-0.000 (0.015)
U1		-0.068*** (0.020)		-0.212** (0.085)
Dum × U1				0.148 (0.090)

Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively; (ii) standard errors in brackets; (iii) Dum means that the dummy variable is unity after this date and zero otherwise; (iv) the break is considered based on the Z_t test.

We continue our analysis with the cointegration equations for U2. Table 5 shows similar results in the case of DJCA, DJIA and DJUA indexes. Moreover, in this case the GH-

RST model documents the negative influence of both the inflation and the inflation uncertainty, for all three indexes. Consequently, when the inflation uncertainty is computed based on a bounded inflation trend, the role of macroeconomic context in explaining the stock prices is stronger. The results confirm than the [Fama \(1981\)](#) argument who sustain that stock prices are correlated with business cycles, being negatively impacted by the level of inflation.

Table 5. Cointegration equations between stock prices, inflation and Chan et al.'s (2013) inflation uncertainty

DJCA	GH-LS (Dummy 2008M1)	GH-LST (Dummy 2007M11)	GH-RS (Dummy 2007M7)	GH-RST (Dummy 2007M7)
Intercept		8.211*** (0.002)		8.286*** (0.011)
Dum × Intercept		0.000 (0.001)		-0.077*** (0.012)
Trend		0.008*** (0.000)		0.007*** (0.000)
Dum × Trend				0.000** (0.000)
I		-0.000 (0.000)		-0.013*** (0.002)
Dum × I				0.013*** (0.002)
U2		-0.000 (0.000)		-0.080*** (0.014)
Dum × U2				0.080*** (0.014)
DJIA	GH-LS (Dummy 2008M1)	GH-LST (Dummy 2007M9)	GH-RS (Dummy 2007M7)	GH-RST (Dummy 2007M5)
Intercept		9.286*** (0.003)		9.483*** (0.045)
Dum × Intercept		0.003 (0.002)		-0.204*** (0.045)
Trend		0.008*** (0.000)		0.007*** (0.000)
Dum × Trend				0.001*** (0.000)
I		-0.001** (0.000)		-0.010*** (0.002)
Dum × I				0.010*** (0.002)
U2		-0.002* (0.001)		-0.075*** (0.023)
Dum × U2				0.075*** (0.023)
DJTA	GH-LS (Dummy 2013M8)	GH-LST (Dummy 2008M11)	GH-RS (Dummy 2013M5)	GH-RST (Dummy 2011M6)
Intercept			8.621*** (0.028)	
Dum × Intercept			-0.104 (0.157)	
Trend				
Dum × Trend				
I			-0.017** (0.009)	
Dum × I			0.126** (0.057)	
U2			-0.179*** (0.013)	
Dum × U2			0.906*** (0.206)	
DJUA	GH-LS (Dummy 2013M11)	GH-LST (Dummy 2008M8)	GH-RS (Dummy 2013M5)	GH-RST (Dummy 2009M4)
Intercept		6.014*** (0.026)		6.165*** (0.025)
Dum × Intercept		-0.429*** (0.021)		-0.629*** (0.052)
Trend		0.006*** (0.000)		0.011*** (0.000)
Dum × Trend				-0.003*** (0.000)
I		0.006 (0.005)		-0.024*** (0.006)
Dum × I				0.031*** (0.010)
U2		0.005 (0.010)		-0.253*** (0.016)
Dum × U2				0.268*** (0.021)

Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively; (ii) standard errors in brackets; (iii) Dum means that the dummy variable is unity after this date and zero otherwise; (iv) the break is considered based on the Z_t test.

Different from Table 4, in Table 5 the GH-RS model employed in the case of DJTA index shows that both inflation and its uncertainty negatively impact the stock prices. Once again the findings of previous papers are validated, and the fact that the inflation and its uncertainty move together, provide indirect evidence in the favor of Friedman – Ball and Cukierman – Meltzer hypotheses.

After estimating and interpreting the cointegration relationships, we proceed to the estimation of the short-run relationship between stock prices, inflation and inflation uncertainty in the U.S., using an Error Correction Model (ECM).

4.5. Error correction model

To estimate the short-run dynamic equation we obtain the residuals from the representatives models (the GD-RST model for DJCA, DJIA, DJUA and LS/RS model for DJTA). After obtaining residual series (called ECM), we can estimate the error correction models for the relationship between stock prices, inflation and inflation uncertainty. Table 6 presents the results for U1.

Table 6. ECM – Stock and Watson’s (2007) inflation uncertainty

	DJCA: GH-RST (Dummy 2007M5)		DJIA: GH-RST (Dummy 2007M5)		DJTA: GH-RS (Dummy 2013M3)		DJUA: GH-RST (Dummy 2008M10)	
Intercept	0.008***	(0.000)	0.009***	(0.000)	0.008	(0.005)	0.004	(0.003)
ECM _{t-1}	-0.859***	(0.079)	-1.008***	(0.078)	-0.041	(0.049)	-0.200***	(0.068)
ΔI	0.001	(0.000)	-0.000	(0.000)	0.018*	(0.010)	0.018***	(0.006)
$\Delta U1$	0.003	(0.005)	0.002	(0.005)	0.085	(0.075)	0.083*	(0.048)
R ²	0.524		0.622		0.058		0.125	

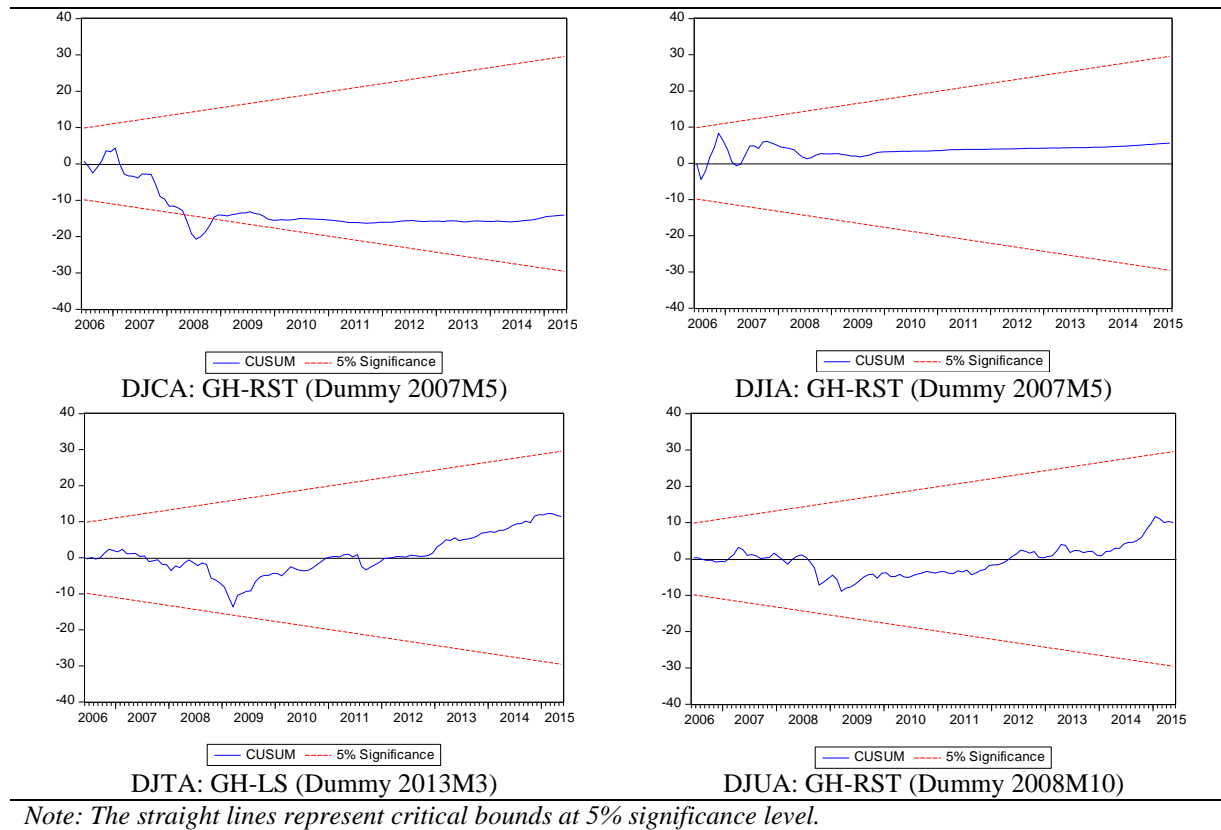
*Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively; (ii) standard errors in brackets; (iii) ECM_{t-1} is the lagged error correction term.*

We notice that for DJCA and DJIA indexes the inflation and its uncertainty have no influence on the stock prices in the short-run. However, a positive short-run influence is documented for the DJTA and DJUA indexes. In the case of the DJUA index the short-run result contrasts the long-run evidence. While in the long-run the inflation and its uncertainty negatively impact the stock prices, in the short-run the opposite phenomenon is recorded. The result can be explained by the fact that energy prices are connected to the general level of prices in the short-run. An increased demand of consumption goods on the market leads to higher prices and an increased production, which trigger the energy prices, and *vice-versa*. Further, the speculative traders contribute to this mechanism, being guided by the present economic context. On contrary, the institutional long-run investors, which are sensitive to the

macroeconomic fundamentals, perceive the level of inflation and its increased uncertainty as a threat for the economy. This negative sentiment determines a decreased interest for the stock market if the inflation increases. Finally, the role of business cycles cannot be neglected in explaining these results.

We proceed further to test for the stability of the analyzed relationship. In this respect we use a CUSUM test for the coefficient stability (Figure 1) and a CUSUM SQUARES test for the residuals stability (see Appendix). The results show that all coefficients are stable in the case of ECM models, but instability in the residuals is recorded around the start-up of the global crisis.

Figure 1. CUSUM test – Stock and Watson’s (2007) inflation uncertainty



The last part of our analysis is dedicated to the robustness check for the results of ECM models, using this time U2. The coefficient of ECM_{t-1} is negative and significant which prove the existence of the long-run relationship between our variables. The results are similar for the DJCA and DJIA indexes, and expected for the DJTA index, where a different model applies compared to U1 (GH-RS instead of GH-LS). However, in the case of DJUA the findings are different from those reported in Table 6. The new results show a negative short-run influence of the inflation uncertainty, in agreement with the long-run situation.

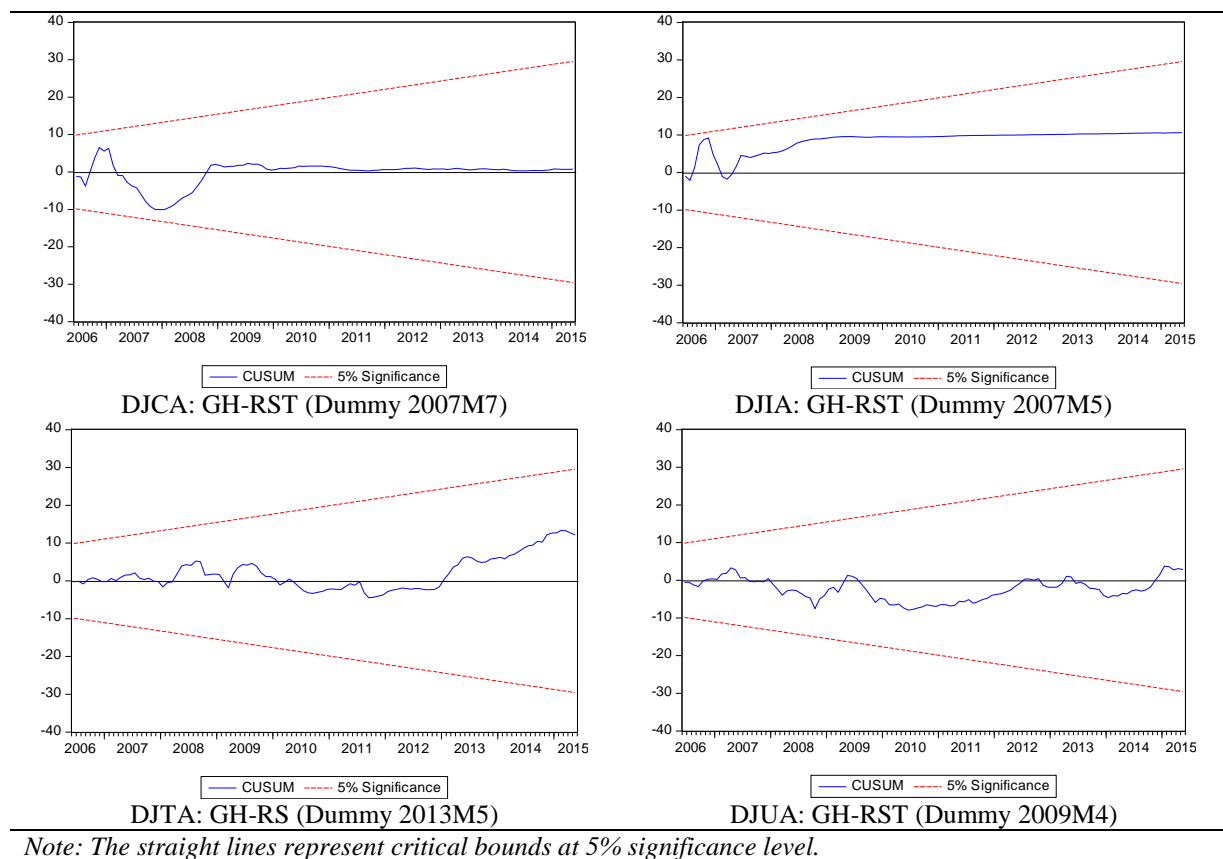
Table 7. ECM – Chan et al.’s (2013) inflation uncertainty

	DJCA: GH-RST (Dummy 2007M7)	DJIA: GH-RST (Dummy 2007M5)	DJTA: GH-RS (Dummy 2013M5)	DJUA: GH-RST (Dummy 2009M4)
Intercept	0.008*** (0.000)	0.009*** (0.000)	0.007 (0.004)	0.003 (0.003)
ECM _{t-1}	-1.091*** (0.073)	-1.003*** (0.078)	-0.168*** (0.055)	-0.345*** (0.078)
ΔI	0.001* (0.000)	-0.001 (0.000)	0.000 (0.010)	0.005 (0.006)
$\Delta U2$	0.001 (0.002)	-0.001 (0.003)	-0.152*** (0.041)	-0.081*** (0.028)
R ²	0.678	0.621	0.206	0.678

Notes: (i) *, ** and *** implies significance at 10%, 5% and 1% levels respectively; (ii) standard errors in brackets; (iii) ECM_{t-1} is the lagged error correction term.

In this case also the CUSUM test documents the stability of coefficients (Figure 2) but a relative instability of the residuals in crisis times (Appendix).

Figure 2. CUSUM test – Chan et al.’s (2013) inflation uncertainty



On the whole, we document that the short-run relationship between stock prices, inflation and inflation uncertainty is less consistent compared to the long-run relationship. Further, in the case of the short-run equation the empirical evidence is mixed and the results are influenced by the way inflation uncertainty is defined. However, in the long-run there is a clear and robust relationship between our variables and in general the inflation and its uncertainty negatively impact the stock prices, results in agreement with the findings reported

by [Cochran and Defina \(1993\)](#), [Alexakis et al. \(1996\)](#) and [Karagianni and Kyrtsov \(2011\)](#), and opposed to those advanced by [Boudoukh and Richardson \(1993\)](#) early on.

5. Conclusions

In this paper, we use the Gregory-Hansen cointegration test to estimate the long-run relationship between stock prices, inflation level and inflation uncertainty in the U.S. The stock prices are estimated based on the DJ composite average index and there different sector indexes (industrial, transportation and utilities), during the period from 2006M1 to 2015M5. The inflation uncertainty is associated with the inflation gap and is computed based on the [Stock and Watson's \(2007\)](#) model with time-varying unbounded inflation trend, and based on the [Chan et al.' \(2013\)](#) approach, with a bounded inflation trend, the second approach being recommended in the case of inflation targeting regimes.

Our findings document the existence of the long-run relationship between variables. In particular, the cointegration relationship is shown by two Gregory-Hansen tests for DJCA, DJIA and DJUA (level shift with trend and regime shift with trend), and by one test for DJTA. The cointegration equations provide evidence in the favor of a negative impact of the inflation and its uncertainty on the stock prices, except for the DJTA index where the results are mixed. These results are opposed to the Fisher's effect but in agreement with recent findings advanced by the empirical literature and with the arguments advanced by [Fama \(1981\)](#), underlining the role of business cycles and the correlation existing between the real economy and equity markets.

However, the short-run equations show less consistent results, where the influence of the inflation and its uncertainty either do not exists, or it is influenced by the way uncertainty is defined. In general the inflation and its uncertainty have a similar impact on stock prices, meaning that they move together, indirectly supporting the Friedman – Ball and Cukierman – Meltzer hypotheses.

The present paper presents however some limits because we test the so called long-run relationship over a relatively reduced time-span (10 years). Therefore, a development of this work will be to address a longer time-horizon and more sector indexes to underline the differences which might exists in the perception of inflation and its uncertainty for different industries and economic sectors.

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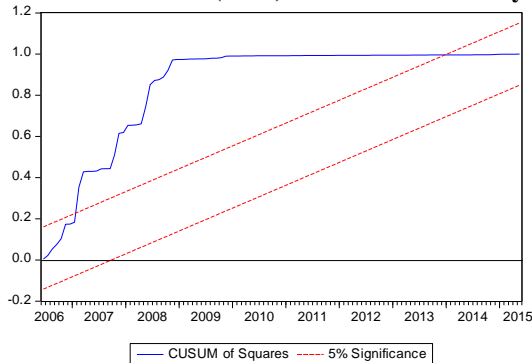
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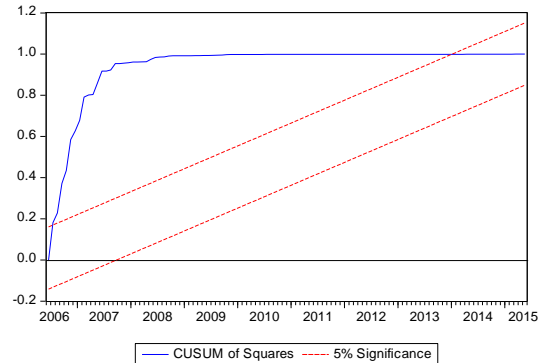
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Appendix - CUSUM SQAURES test

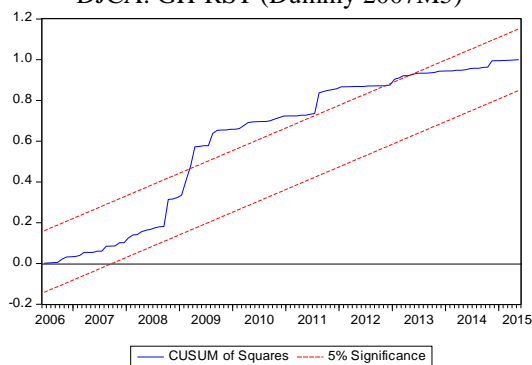
Stock and Watson's (2007) inflation uncertainty



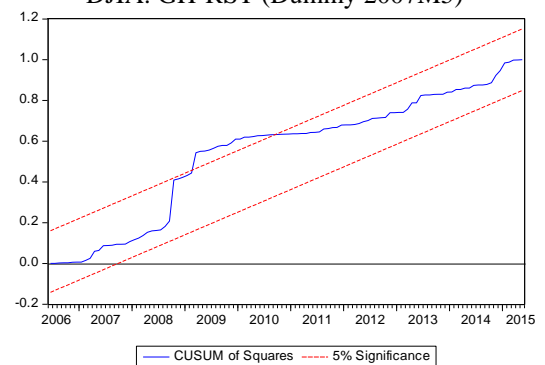
DJCA: GH-RST (Dummy 2007M5)



DJIA: GH-RST (Dummy 2007M5)

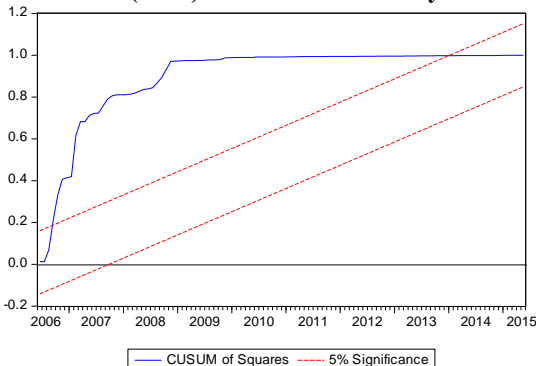


DJTA: GH-LS (Dummy 2013M3)

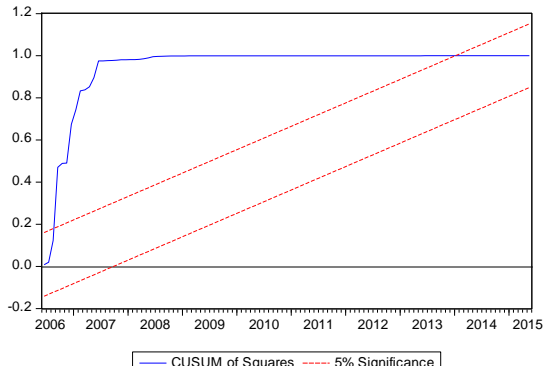


DJUA: GH-RST (Dummy 2008M10)

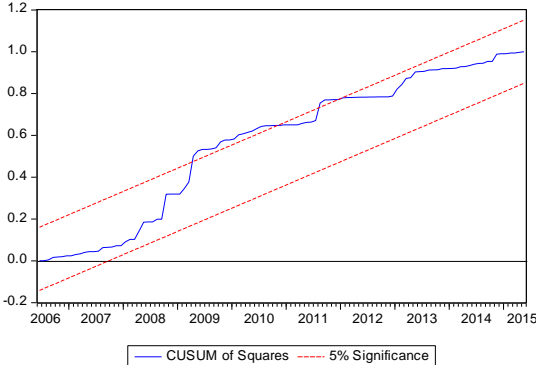
Chan et al.'s (2013) inflation uncertainty



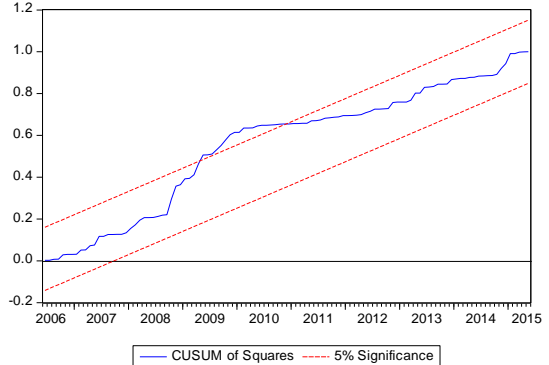
DJCA: GH-RST (Dummy 2007M7)



DJIA: GH-RST (Dummy 2007M5)



DJTA: GH-RS (Dummy 2013M5)



DJUA: GH-RST (Dummy 2009M4)

Note: The straight lines represent critical bounds at 5% significance level.