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Inflation scares and bond market signals: Is the Fed's policy action asymmetric?

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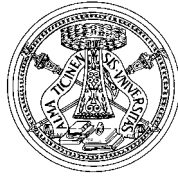
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Quaderni di Dipartimento

**Inflation scares and bond market signals:
Is the Fed's policy action asymmetric?**

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Inflation scares and bond market signals: Is the Fed's policy action asymmetric?

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March 8, 2007

Abstract

This paper is concerned with the question whether, over the last two decades, the priority to maintain the credibility as inflation 'fighter' induced the Fed to conduct a monetary policy that can be considered asymmetric in nature. Considering that the longer-term bond interest rate is the most reliable indicator used to gauge Fed's credibility to low inflation, we investigate the equilibrium-type relationship between the monetary policy instrument, the Federal Funds rate, and 10-year government bond rate within a threshold-cointegration framework in a similar fashion of Enders and Siklos (2001). Our empirical findings indicate that, during the period 1980-2005, both the direction and magnitude of changes in long-term expected inflation, reflected in 10-year bond rate movements, influenced the timing of the Fed's action. Therefore, we conclude that asymmetries in the term-spread dynamics are the result of the Fed's behaviour that can be considered representative of a monetary policy essentially asymmetric.

1 Introduction

Since the early 1980s, controlling inflation become a key feature of the monetary policy for the most important central banks. Numerous disinflations have taught central bankers around the world that credibility -having a reputation for pursuing price level stability consistently and persistently- is the foundation of an effective anti-inflationary monetary policy. Even though in United States the low-inflation objective has nearly been achieved, the acquisition and maintenance of credibility for the commitment to low inflation remains one of the hallmarks of the Fed's monetary policy. According to the theory, we support the idea that the continuous interaction between central bank and the public, involved by the monetary policy, is the essence of the fragility in the low-inflation equilibrium. The public is very sensitive to any central bank departure from the behavior it has come to anticipate: the absence of an implicit mutual understanding between the public and the central bank about future policy actions

may result in an inflation bias. Potential for future inflation revealed by the public can be thought of as the result of a monetary policy which tends to generate expectations of excessively expansionary policy in the future. Such expectations encourage workers and firms to ask for wage and prices increases to protect themselves from higher expected costs as long as the central bank demonstrates its credibility to low inflation by tightening its policy.

Starting from these latter premises, we can assume that the market is able to 'discipline' a central bank and that the policymaker's reaction is oriented to maintain the public's confidence that controlling inflation is a permanent feature of the longer-term monetary strategy.

Historical experiences, over the last two decades of U.S. monetary policy, seem to confirm this view. In many circumstances, the Fed's policy action was directed at resisting the so-called "inflation scares" episodes. These latter are defined by Goodfriend (1993) as episodes signaling sharp changes in inflation expectations as evidence of worsening credibility on the Fed's anti-inflationary policy. Inflation scares can be captured by different economic indicators. However, the most reliable indicator is represented by the very long-term nominal bond interest rates (i.e. ten- or thirty-year nominal government bond rates). According to the Fisherian theory and the empirical findings on the information content of the U.S. term structure provided by Mishkin (1990a, 1990b) and Fama (1984), the long-end of the term structure is likely to reveal information about expected future inflation because the volatility of expected inflation outweighs (in the long-term) that of the real interest rate. Recently, Merha (1994, 1996, 1998), Ireland (1996) and Bordo (2001) provide empirical evidence supporting this view.

Goodfriend (1993, 1998, 2002) interprets sharp movements in the long-term bond rate as the forward-looking response of the bond market with regard to doubts about the future of central bank policy regarding price stability. The numerous episodes of inflation scares, discussed by the author, provide evidence that the low-inflation equilibrium sustained by the Fed's credibility is fragile and that for much time the Fed's action can be interpreted as repeated attempts to contain inflationary expectations.

In this paper, we go into Goodfriend's innovative study thoroughly. We analyze formally how long-term bond rate movements influenced the *timing* of the U.S. monetary policy action over the last two decades. We expect that both the *magnitude* and the *direction* of changes in inflation expectations, reflected in 10-year bond rate movements, involve a different degree of policy inertia. In fact, the problem of maintenance of credibility as inflation 'fighter' might have induced the Fed to be aggressive in trying to prevent intolerable rising inflationary pressures via higher Fed Funds rates and cautious about lowering them in the presence of decreasing inflation expectations. If this the case, U.S. monetary policy can be considered asymmetric in nature.

In order to test this latter hypothesis, we examine the equilibrium-type relationship between Fed Funds rate and 10-year bond rate. After defining the term-spread as the residuals of the estimated cointegrating equation between interest rates, we test the hypothesis that the term-spread dynamics may result

asymmetric due to the nature of the monetary policy action. The approach used to detect the existence of asymmetries in the term-spread adjustment process is based on a testing procedure introduced, in the univariate context, by Enders and Granger (1998) and generalized to the cointegration context by Enders and Siklos (2001). It involves to estimate a particular class of models, the so-called threshold autoregressive models, introduced by Tong (1983), Enders and Granger (1998) and Canen and Hansen (1998).

The remainder of this paper is organized as follows. In section 2, we discuss the information content of long-term bond rates and look at how bond yields measure up as a guide to inflation expectations and monetary policy. In section 3, we present the econometric formulation of the Threshold Autoregressive (TAR) and Momentum Threshold Autoregressive (M-TAR) models and their application to performing cointegration tests allowing for asymmetric adjustment process. Section 4 explains how to investigate the relationship between Fed Funds rate and 10-year bond rate within the M-TAR framework. Section 5 presents our empirical findings and section 6 concludes.

2 Bond market and long-term expected inflation

Fisher's theory represents perhaps the most famous theory of nominal interest rate determination. According to Fisher's theory of interest, movements in nominal bond yields originate in two sources: changes in real interest rates and changes in expected inflation. Unfortunately, neither ex-ante real interest rate nor inflation expectations over the life of the bond can be directly observed and then measuring their contribution in accounting for movements in bond yields is not straightforward. Despite the numerous empirical studies on the Fisher's effect (see Mishkin (1990a, 1990b), Fama (1984), Bernanke (1990) and Tkacz (2004)), and the widely accepted assumption that, at longer maturities, the term structure can be used to assess future market's inflationary prospects, a stable relationship between nominal interest rate and its embodied expected inflation component has proven extremely difficult to establish. Following Bordo (2001) and Bordo and Deward (2003), we can use a simple old-fashioned approach to demonstrate that 10-year US Treasury bond rate primarily reflects changes in long-term inflationary expectations. The approach is based on the *Fisher effect* that an increase in longer-term inflation expectations would be reflected in longer-term interest rates and the *Golden Rule effect* that in longer-term equilibrium the real rates of interest would equal the real growth trend of real output. The *Fisher-Golden Rule hypothesis* is that the difference between the bond rate and trend real growth is a crude measure of implicit bond market inflation expectations over the maturity of a bond. Assuming that 10-year moving average of real GDP growth rate is a proxy of the real growth trend, figure 1 shows that a stable relationship between 10-year bond rate and long-term expected inflation exists.

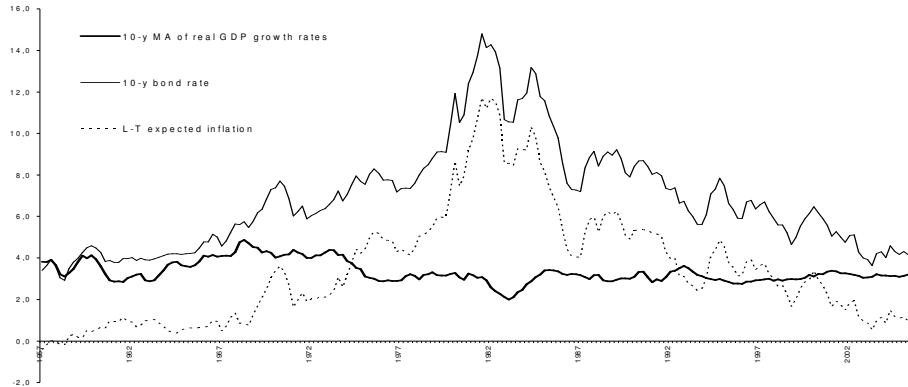


Figure 1: Long-term expected inflation is obtained as the difference between 10-year nominal bond rate and 10-year moving average of Real GDP (at 2000 prices) growth rates over the period 1957Q1 to 2005Q3.

However, deciphering signals from the bond market is not always straightforward. Complicating the debate on the information content of the bond rates is the fact that, besides an improvement in real growth prospects, long-term interest rates can sometimes move for a change in the risk premium. In addition, a more uncertain economic environment is generally thought to put upward pressure on long-term yields relative to short-term yields. Ireland (1996) uses a Lucas's model (1978) that generalizes the Fisherian theory accounting for the effects of uncertainty, and he demonstrates that, over the sample 1969-1995, the 10-year real interest rates are relatively stable and the size of the risk premium embodied in the 10-year bond rates is small. In particular, the author estimates bounds on the plausible size of the risk premium. Results indicate that these bound are very tight (28 basis point band). Therefore, considering that real interest rates are stable, the author concludes that any change in the 10-year bond rate in excess of the bound almost certainly signals a change in inflationary expectations.

In our analysis, we account for the uncertainty about the relationship between 10-year bond rate and expected inflation. We estimate a *data-determined* super-consistent threshold value, together with M-TAR model coefficients, that can be used to interpret the extent to which movements in the term-spread (caused by changes in the bond rate) most likely reflect changes in long-term inflationary expectations. In practice, as discussed in empirical section, we assume that the Fed interprets any increase of the term-spread in excess of the threshold sufficient enough to fear of current inflation scares will materialize in the future.

2.1 Inflation scares and the Fed's reaction: historical evidence

The above discussed empirical findings seem to justify the Federal Open Market Commission's common practice for using long-term bond yields as indicators of expected inflation. In this section, according to Goodfriend's approach, we review market's inflation scares episodes that, over the last two decades, severely tried the Fed's credibility regarding price stability and influenced the timing of the Fed's policy action.

The first inflation scare episode occurred in the fall of 1979. The second oil shock, the Soviet Union's invasion of Afghanistan and Fed's hesitation to rise Fed Funds rate in the face of a weakening economy convinced private sector to increase expectations of future inflation. The 10-year bond yield jumped 2 percentage points in the first quarter of 1980 signaling the sudden collapse of confidence in the Fed's commitment to low inflation. The 3 percentage points increase of the Fed Funds rate, in March 1980, is interpreted by many commentators as the Fed's attempt to restore its credibility as inflation fighter. The fragility of the Fed's credibility was apparent again in 1984 when the bond rate, after falling to about 10 percent in late 1982, registered another inflation scare by rising to around 14 percent, even though the Fed had by then brought actual inflation down from over 10 percent to around 4 percent. The Fed reacted by rising the nominal Fed Funds rate from the 8 percent range to the 11 percent range. Inflation remained low, so the tightening took the real short-term interest rate up by about 3 percentage points to around 7 percent briefly in mid-1984. The high real short rate needed to contain the inflation scare bringing real GDP growth down to a sustainable 2 to 3 percent range in the second half of 1984. The 6 percentage point drop in the bond rate from its June 1984 peak to the 7 percent range in early 1986 indicated that the Fed acquired enormous additional credibility for low inflation, in large part no doubt due to the aggressive inflation-fighting action taken. A third inflation scare episode was faced during 1987. In the period between March and October of 1987, 10-year bond yield jumped 2 percentage points for two reasons. First, a change in central bank governor affected financial markets: Volcker was near the end of his term as chairman and there were doubts about whether the Fed, under Volcker's successor, would have continued to place a high priority on low inflation target. Second, the October 1987 stock market crash forced the Fed to ease monetary policy and put off raising interest rates until the spring of 1988. Thus, Fed policy action created some doubt about its credibility regarding price stability and produced the effect that long bond rates didn't fall until late 1992.

Despite years of sustained progress in bringing the inflation rate down, a number of episodes in the bond market, during the Greenspan era, continued to evidence that the Fed's anti-inflationary credibility remained exceedingly brittle. The swings in the 10-year bond rate over the period 1993-94 was less dramatic than in the early 1980s, but nonetheless substantial. Rising from a low of about 5.3 percent in October 1993, the bond rate peaked at around 8 percent in November 1994 signaling a change in the market's inflation prospects. During

the January 2001, despite cuts in the Fed's short-term interest rate target, long-term interest rates drifted upward. Many commentators agree to consider this lack of response signaled that bond investors' long-term inflation expectations were on the rise and, hence, that the Fed eased too much. Finally, many of the 'rallies' of the American bond market, since 2001 recession, are interpreted as the result of the market's inflation outlook shifts. In particular, the 2002 super-rally, following the disruption to the economy from the September 11 attacks, reflected, in part, expectations of further monetary ease from already low rate levels. Contrary, the early 2003 deflation-scare rally, ended by cumulative declines in 10-year yields of at least 50 basis points, had the distinctive characteristic to reflect the fall of market's participant inflation expectations and encouraged the Fed to provide unprecedented monetary accommodation even at official interest rates close to zero.

3 Testing for cointegration using threshold models

In this section we introduce the methodology developed by Enders and Siklos (2001) to test for threshold cointegration. These authors generalize to the cointegration context the testing procedure suggested by Enders and Granger (1998) that can be used to test the null hypothesis of a unit-root against the alternative of stationarity with asymmetric adjustment. In particular, Enders and Siklos (2001) propose a residual-based test of cointegration allowing for asymmetric error-correction mechanism. Let $\{x_{kt}\}_1^T$ denote observable variables integrated of order one, $I(1)$. Suppose that the long-run equilibrium relationship is given by:

$$x_{1t} = \hat{\beta}_0 + \hat{\beta}_2 x_{2t} + \dots + \hat{\beta}_n x_{nt} + u_t \quad (1)$$

where $\hat{\beta}_k$ (with $k = 1, \dots, n$) are estimated parameters, $\hat{\beta}_0$ is the intercept term and u_t is the disturbance term. Following the Engle-Granger (1987) two-step procedure, the existence of cointegration involves stationary u_t , $u_t \sim I(0)$. In order to test residuals for stationarity, after estimating regression residuals, in the second step we have estimate the auxiliary equation used as the basis to perform traditional Augmented Dickey-Fuller (ADF) test:

$$\Delta \hat{u}_t = \rho \hat{u}_{t-1} + \sum_{i=1}^p \gamma_i \Delta \hat{u}_{t-i} + v_t \quad (2)$$

where \hat{u}_t are the estimated residuals from equation (1) while lagged differences, $\Delta \hat{u}_{t-i}$, allow for autocorrelation in the residuals v_t . Under the null, $H_0 : \rho = 0$, u_t is a nonstationary series and then we reject the presence of cointegration. If we accept the alternative hypothesis, $H_1 : -2 < \rho < 0$, residuals, u_t , are stationary and long-run equilibrium relationship (1) is accepted. In this case, the

Granger's representation theorem guarantees the existence of an error-correction representation of the variables in the form:

$$\Delta x_{kt} = \alpha_k \left(x_{1t-1} - \hat{\beta}_0 - \hat{\beta}_2 x_{2t-1} - \dots - \hat{\beta}_n x_{nt-1} \right) + \sum_{j=1}^n \sum_{i=1}^{p-1} \gamma_{ji} \Delta x_{jt-i} + \varepsilon_{kt} \quad (3)$$

The implicit assumption underlying the error-correction model (ECM) is that, the system exhibits *symmetric* adjustment around the equilibrium value. However, in the presence of asymmetric adjustment, the above discussed standard cointegration framework is misspecified. As showed by Pippenger and Goering (1993), Balke and Fomby (1997) and Enders and Granger (1998), in the presence of asymmetric adjustment, standard tests for cointegration, such as Engle-Granger (1987) and Johansen's (1988, 1991) tests, all suffer of misspecification problem and have low power properties. Therefore, Enders and Siklos (2001) propose to test for cointegration using a class of models presented by Enders and Granger (1998) in an univariate context as the basis of unit-root tests in the presence of potential asymmetric adjustment. They consider two alternative specifications of the auxiliary equation (2):

$$\Delta \hat{u}_t = I_t \rho_1 \hat{u}_{t-1} + (1 - I_t) \rho_2 \hat{u}_{t-1} + \sum_{i=1}^p \gamma_i \Delta \hat{u}_{t-i} + v_t \quad (4)$$

where the Heaveside index I_t is defined such that:

$$I_t = \begin{cases} 1 & \text{if } \hat{u}_{t-1} \geq \tau \\ 0 & \text{if } \hat{u}_{t-1} < \tau \end{cases} \quad (5)$$

in the Threshold Autoregressive (TAR) model or:

$$I_t = \begin{cases} 1 & \text{if } \Delta \hat{u}_{t-1} \geq \tau \\ 0 & \text{if } \Delta \hat{u}_{t-1} < \tau \end{cases} \quad (6)$$

in the Momentum-Threshold Autoregressive (M-TAR) model, where τ is the unknown threshold value, that is the value above or below the regime of the adjustment process changes. According to Petrucelli and Woolford (1984) a necessary and sufficient conditions for the stationarity of u_t is the condition that $\rho_1 < 0$, $\rho_2 < 0$ and $(1 + \rho_1)(1 + \rho_2) < 1$. If these conditions are met, regression residuals are found stationary and series x_{kt} (for $k = 1, \dots, n$) can be said to be threshold-cointegrated.

Interpreting the properties of the TAR model is straightforward. If \hat{u}_{t-1} is above its long-run equilibrium value, the adjustment is $\rho_1 \hat{u}_{t-1}$, and if \hat{u}_{t-1} is below long-run equilibrium, the adjustment is $\rho_2 \hat{u}_{t-1}$. The TAR model captures the key aspects of any "deep" movements in the series. In fact, if we suppose to have $-1 < \rho_1 < \rho_2 < 0$, then the negative phase of the $\{u_t\}$ sequence will tend to be more persistent than the positive phase. Contrary, the adjustment process in the M-TAR model depends on the previous period's change in \hat{u}_t . This

implies that the momentum model can capture the possibility of asymmetrically "sharp" movements in a series. If, for example, we assume $|\rho_1| < |\rho_2|$ then the M-TAR model exhibits little decay for positive $\Delta\hat{u}_{t-1}$ but substantial decay for negative $\Delta\hat{u}_{t-1}$. Thus, increases tend to persist but decreases tend to revert quickly toward the equilibrium value.

Enders and Siklos (2001) suggest to test for threshold cointegration using a standard F statistic for the null hypothesis of no cointegration, $H_0 : \rho_1 = \rho_2 = 0$, that implies residuals u_t are not stationary, against the alternative that series are cointegrated and the system exhibits a potential asymmetric adjustment toward the equilibrium. The empirical test statistic is compared with the appropriate critical values reported in their paper at the 10%, 5% and 1% significance levels and obtained via Monte Carlo experiments. If the null of no cointegration is rejected, it is possible to test for symmetric adjustment in the series \hat{u}_t , $H_0 : \rho_1 = \rho_2$, against the alternative hypothesis of asymmetric adjustment, $H_1 : \rho_1 \neq \rho_2$, performing a standard F -test because the system is stationary. Therefore, if we accept the null of symmetric adjustment, the Engle-Granger's cointegration test emerges as a special case. Contrary, if we ascertain the presence of cointegration and asymmetric adjustment process towards equilibrium, we can write the corresponding Threshold Vector Error-Correction model (T-VECM) as follows:

$$\Delta x_{kt} = \alpha_k^+ \hat{u}_{t-1}^+ + \alpha_k^- \hat{u}_{t-1}^- + \sum_{k=1}^n \sum_{i=1}^{p-1} \gamma_{ki} \Delta x_{kt-i} + \varepsilon_{kt} \quad (7)$$

where $\hat{u}_{t-1} = (x_{1t-1} - \hat{\beta}_0 - \hat{\beta}_2 x_{2t-1} - \dots - \hat{\beta}_n x_{nt-1})$ are the residuals of long-run equation (1) while \hat{u}_{t-1}^+ and \hat{u}_{t-1}^- are defined as:

$$\begin{cases} \hat{u}_{t-1}^+ = I_t \hat{u}_{t-1} \\ \hat{u}_{t-1}^- = (1 - I_t) \hat{u}_{t-1} \end{cases} \quad (8)$$

where the Heaviside index values, I_t , depends on the selected specification (TAR or M-TAR).

3.1 Estimation of the threshold models

Before testing for cointegration, the threshold model (4) has to be estimated. The estimation procedure involves to estimate via OLS the value of the threshold, τ , along with the parameters values of the model, ρ_1, ρ_2 and γ_i for $i = 1, \dots, p$. Chan (1993) showed that searching over all values of possible attractor points so as to minimize the sum of squared errors from the fitted model (4) yields a super-consistent estimate of the threshold. The method involves sorting potential thresholds values into ascending order discarding $\pi = 15\%$ of the largest and smallest values. Then, the optimal threshold value is estimated consistently by means of a direct search:

$$\begin{aligned}\hat{\tau} &= \arg \min_{\tau \in T} \hat{\sigma}^2(\tau) \\ T &= \{\tau | y_{([\tau m])} \leq \tau \leq y_{([1-0.15m])}\}\end{aligned}\tag{9}$$

where $y_{(\cdot)}$ denotes the order statistic (\hat{u}_t or $\Delta\hat{u}_t$ depending on the threshold's specification). $m = T$ in the TAR model while $m = (T - 1)$ in the M-TAR model, being T the number of observations. $\hat{\sigma}^2(\tau)$ denotes the error-term variance of the regression (4) for a given τ .

4 Testing for asymmetric policy action

In order to investigate how bond rates movements influenced the *timing* of monetary policy action over the last two decades, we examine the equilibrium-type relationship between Fed Funds rate and 10-year bond rate employing a M-TAR model. We prefer a "momentum" model to a simple TAR specification because of interpreting the regime switch in the M-TAR can be particularly useful when policymakers are viewed as attempting to smooth out any large changes in the series. In our analysis, after defining the term-spread as the residuals of the estimated cointegrating equation between the above considered nominal interest rates, we are interested to demonstrate that the Fed might have taken strong measures to offset shocks producing significant *changes* in the term-spread if such shocks were deemed to indicate increases rather than decreases in inflationary expectations. This implies to test the hypothesis that Fed Funds rate adjusts to the 10-year bond rate producing asymmetries in the term-spread dynamics depending on the *magnitude* and the *direction* of the shocks.

Following the two-step cointegration testing procedure presented in section 3, in the first step, we estimate via OLS the following long-run relationship:

$$i_{FF,t} = c + \beta i_{10y,t} + u_t\tag{10}$$

where $i_{FF,t}$ is the Federal Funds rate and $i_{10y,t}$ is the 10-year nominal bond rate. c and β represent the parameters to be estimated while u_t denotes regression residuals and can be interpreted as the (*reversal*) term-spread of our interest. In contrast with expectations hypothesis theory (EHT), we consider the possibility of a reversal relationship between interest rates: short-term rate may responds to the long-term interest rate. As we have discussed in section 2, this assumption is consistent with a policymaker who sets short-rate including movements of the long-term bond rates in the decision-making process. Moreover, we don't assume a priori that short-term and long-term rates are linearly cointegrated with cointegrating vector $[1, -1]$ as postulated by EHT. The slope coefficient in equation (10), may differ from one, $\beta \neq 1$. In the second step, we estimate the following M-TAR model:

$$\Delta \widehat{u}_t = I_t \rho_1 \widehat{u}_{t-1} + (1 - I_t) \rho_2 \widehat{u}_{t-1} + \sum_{i=1}^p \gamma_i \Delta \widehat{u}_{t-i} + v_t \quad (11)$$

where \widehat{u}_t are the residuals of the long-run equation (10) and the Heaviside indicator function I_t is defined according to (6) as:

$$I_t = \begin{cases} 1 & \text{if } \Delta \widehat{u}_{t-1} \geq \tau \\ 0 & \text{if } \Delta \widehat{u}_{t-1} < \tau \end{cases}$$

After estimating the parameters of the model via OLS method and using the *search-grid* methodology proposed by Chan (1993), we perform the sequential testing procedure suggested by Enders and Siklos (2001). If the series $i_{FF,t}$ and $i_{10y,t}$ are found to be threshold-cointegrated then, in contrast with EHT, the adjustment process of the term-spread u_t cannot be considered symmetric irrespective the extent by which it deviates from its attractor point τ . In order to obtain a confirmation of the M-TAR empirical findings and identify the sources of fluctuations in the term-spread, we estimate the following T-VECM :

$$\begin{cases} \Delta i_{FF,t} = c_1 + \lambda_{FF}^+ \widehat{u}_{t-1}^+ + \lambda_{FF}^- \widehat{u}_{t-1}^- + \sum_{i=1}^{p-1} \gamma_{1,i} \Delta i_{FF,t-i} + \sum_{i=1}^{p-1} \delta_{1,i} \Delta i_{10y,t-i} + \varepsilon_{FF,t} \\ \Delta i_{10y,t} = c_2 + \lambda_{10y}^+ \widehat{u}_{t-1}^+ + \lambda_{10y}^- \widehat{u}_{t-1}^- + \sum_{i=1}^{p-1} \gamma_{2,i} \Delta i_{FF,t-i} + \sum_{i=1}^{p-1} \delta_{2,i} \Delta i_{10y,t-i} + \varepsilon_{10y,t} \end{cases} \quad (12)$$

where λ^+ and λ^- indicate the different speed of convergence depending on whether equilibrium deviations are rising or falling. $\widehat{u}_{t-1}^+ = I_t \widehat{u}_{t-1}$ while $\widehat{u}_{t-1}^- = (1 - I_t) \widehat{u}_{t-1}$. The economic interpretation of the short-run dynamics is straightforward. Assuming $i_{10y,t}$ is found to be an exogenous variable¹, $\Delta \widehat{u} < \widehat{\tau}$ is indicative of a sharp rise in the 10-year bond rate. In particular, given some degree of uncertainty about the relationship between 10-year bond rate and expected inflation as discussed in section 2, the estimated threshold value $\widehat{\tau}$ can be used to interpret the extent to which movements in the (*reversal*) term-spread most likely reflect changes in long-term inflationary expectations. We assume that only increases in the 10-year bond rate producing changes in the term-spread below the estimated threshold value, such that $\Delta \widehat{u} < \widehat{\tau}$, almost certainly signal a change in market's expected inflation. Therefore, such a change produces a response by the Fed to dampen inflationary expectations. This implies an upward adjustment of the Fed Funds rate, $\Delta i_{FF,t} > 0$, until the term-spread reverts to its equilibrium value with a speed of convergence equals to λ_{FF}^- . Analogously, a change in the term-spread, such that $\Delta \widehat{u} \geq \widehat{\tau}$, caused by a permanent negative shock on the bond rates (i.e. due to a deflation scare episode) entails a downward movements of the Fed Funds rates and the term-spread adjusts back to its attractor point with a speed of convergence equals to λ_{FF}^+ . Therefore, in order investigate the influence of the bond rate movements on the *timing* of monetary policy, we are interested to compare the magnitude of the decay

¹In the empirical section, we test this hypothesis.

factors λ_{FF}^- and λ_{FF}^+ . Only in the case these latter differ from each other, we conclude that Fed's policy action can be considered asymmetric in nature. In particular, the case $\lambda_{FF}^+ < \lambda_{FF}^-$ would confirm that maintaining price stability remains a priority for the U.S. monetary policy.

5 Empirical results

In this section we present the data used in our analysis and the estimation results.

5.1 Data

We use monthly values of the Federal Funds rate (i_{FF}) and 10-year U.S. government bond rates (i_{10y}) for the sample period October 1980 to September 2005. The data are obtained from the Federal Reserve Bank of St. Louis economic database (FRED[®]). The sample covers a period well-characterized by an (implicit) form of inflation-targeting monetary policy. This period motivates our interest in investigating the effects of the Fed's policy action on the term-spread dynamics. The analysis is implemented using GAUSS 5.0 programming language.

5.2 Integration and cointegration analysis

Before employing the M-TAR testing procedure, we perform the standard augmented unit-root test proposed by Dickey and Fuller (ADF test) for the individual interest rate series and Johansen's cointegration test. Table 1 displays results of ADF tests.

	i_{FF}	Δi_{FF}	i_{10y}	Δi_{10y}
ADF	-1.99	-4.17***	-1.21	-12.11***

Table 1: Unit-root tests. The ADF test equation includes a constant term while the number of lag lengths considered are selected according to Schwartz information criterion. *, **, *** denote statistical significance at the 10%, 5% and 1% levels, respectively. Critical values are tabulated in Dickey and Fuller (1979) and MacKinnon (1991).

We cannot reject the null of unit-root hypothesis for both i_{FF} and i_{10y} at conventional level of significance. However, both the two series are integrated process of order one. This result is consistent with the general agreement (see Stock and Watson (1988)) that short-term and long-term interest rates are $I(1)$ process. Table 2 reports results of standard cointegration test between i_{FF} and i_{10y} . Unfortunately, the λ -trace statistics cannot reject the null hypothesis of no cointegration. Therefore, the term-spread is not stationary and then the traditional EHT theoretical ground is no more consistent.

$No.CE(s)$	$\lambda - Trace$	5% _{cv}	10% _{cv}
$r = 0$	14.24	15.41	20.04
$r \leq 1$	1.95	3.76	6.65

Table 2: Trace test. The results are obtained after estimating a bivariate VAR(3) model and including an intercept in the potential cointegrating vector. r denotes the number of cointegrating vectors under the null and the last two columns reports the critical values at 5% and 1% levels of significance.

In contrast, testing for cointegration using the M-TAR approach provides evidence for the existence of threshold-cointegration between i_{FF} and i_{10y} . Following the two-step estimation procedure illustrated in section 3, first we estimate the long-run relationship (10) and then we use the regression residuals to estimate the parameters of the M-TAR model (11). Results are presented in table 3.

Panel A		
i_{FF}	i_{10y}	c
1	1.17	-2.63
	[0.03]	[0.26]

Panel B						
$\hat{\tau}$	$\hat{\rho}_1$	$\hat{\rho}_2$	F_c	F_a	p_{AIC}	p_{SC}
-0.43	-0.03	-0.22	14.16***	16.17***	5	5
	[0.02]	[0.04]				

Table 3: M-TAR estimates. Standard errors figures in square brackets. Critical values for the F-tests are tabulated in Enders and Siklos (2001). *, **, *** denote statistical significance at the 10%, 5% and 1% levels, respectively.

The panel A of table 3 reports the OLS estimates of the cointegrating vector $[1 \ \beta \ c]'$. Panel B summarizes the parameters estimates of the M-TAR model. Its specification is augmented with lagged changes in the residual series. The lag-length, p , is determined according to Akaike (p_{AIC}) and Schwarz (p_{SC}) information criterion. F_c is the value of the F -statistics used to test the hypothesis of no cointegration, $H_0 : \rho_1 = \rho_2 = 0$ against the alternative hypothesis of cointegration and possible asymmetric adjustment dynamics. F_a indicates the value of the F -statistics used to test the hypothesis of symmetric adjustment, $H_0 : \rho_1 = \rho_2$, against the alternative of asymmetric adjustment, $H_1 : \rho_1 \neq \rho_2$. The M-TAR based cointegration tests strongly suggest that i_{FF} and i_{10y} are threshold-cointegrated with cointegrating vector $[1, -1.17]'$. The estimated parameters of the M-TAR model indicate that conditions for the stationarity of $\{u_t\}$, $-2 < (\hat{\rho}_1, \hat{\rho}_2) < 0$ and $(1 + \hat{\rho}_1)(1 + \hat{\rho}_2) < 1$, are satisfied. F_a -test indicates that the behavior of the term-spread between Fed Funds rate and 10-year bond rate is asymmetric. In particular, discrepancies from the equilibrium such

that $\Delta \hat{u} < -0.43$ are eliminated relatively quickly, whereas other changes, such that $\Delta \hat{u} \geq -0.43$, display a large amount of persistence. Figure 2 depicts the theoretical (*reverse*) term-spread, obtained as simple difference between Fed Funds rate and 10-year treasury bond rate, $(i_{FF,t} - i_{10y,t})$, and residuals from cointegrating regression (10), \hat{u}_t .

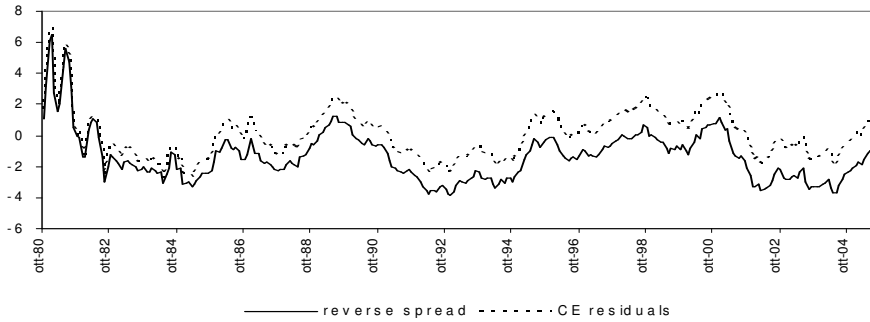


Figure 2: Simple reverse spread $(i_{FF,t} - i_{10y,t})$ and residuals from CE. Differences are mainly due to the introduction of a the intercept term in the long-run equation.

After ascertaining the presence of threshold-cointegration, we estimate the bivariate T-VECM to obtain confirmation of the M-TAR empirical findings. Moreover, dynamic representation of the system allow us to identify the sources of fluctuations in the term-spread and investigate the causality direction in the monetary policy transmission mechanism.

5.3 T-VECM estimation

According to Granger's representation theorem, we proceed to estimate the following T-VECM:

$$\begin{cases} \Delta i_{FF,t} = c_1 + \lambda_{FF}^+ \hat{u}_{t-1}^+ + \lambda_{FF}^- \hat{u}_{t-1}^- + \sum_{i=1}^{p-1} \gamma_{1,i} \Delta i_{FF,t-i} + \sum_{i=1}^{p-1} \delta_{1,i} \Delta i_{10y,t-i} + \varepsilon_{FF,t} \\ \Delta i_{10y,t} = c_2 + \lambda_{10y}^+ \hat{u}_{t-1}^+ + \lambda_{10y}^- \hat{u}_{t-1}^- + \sum_{i=1}^{p-1} \gamma_{2,i} \Delta i_{FF,t-i} + \sum_{i=1}^{p-1} \delta_{2,i} \Delta i_{10y,t-i} + \varepsilon_{10y,t} \end{cases}$$

Table 4 reports parameters estimates obtained via Full Information Maximum Likelihood (FIML) estimation method.

The signs of the estimated adjustment coefficients, $\hat{\lambda}^+$ and $\hat{\lambda}^-$, are coherent with the direction of the shocks. Both the Federal Funds rate and 10-year bond rate adjust in the "right" direction for positive and negative values of $\Delta \hat{u}$. Tests of significance provide additional confirmation of the M-TAR findings and

	\hat{c}	$\hat{\lambda}^+$	$\hat{\lambda}^-$	$\hat{\gamma}_{\cdot,1}$	$\hat{\gamma}_{\cdot,2}$	$\hat{\delta}_{\cdot,1}$	$\hat{\delta}_{\cdot,2}$
$\Delta i_{FF,t}$	-0.04 [0.02]	-0.03 [0.02]	-0.14 [0.04]	0.50 [0.05]	-0.26 [0.05]	0.20 [0.07]	-0.07 [0.07]
$\Delta i_{10y,t}$	-0.02 [0.02]	0.01 [0.01]	0.05 [0.03]	-0.08 [0.04]	0.13 [0.04]	0.35 [0.06]	-0.20 [0.06]
Tests:							
$H_0 : \lambda_{FF}^+ = \lambda_{FF}^-$	6.23 (0.01)						
$H_0 : \lambda_{10y}^+ = \lambda_{10y}^-$	1.27 (0.26)						
$H_0 : \lambda_{10y}^+ = \lambda_{10y}^- = 0$	3.20 (0.20)						
$H_0 : \gamma_{2,1} = \gamma_{2,2} = 0$	6.18 (0.06)						

Table 4: T-VECM estimates. Standard errors in square brackets. Significance levels in parenthesis. Lag length (p=3) is determined using both the Akaike and Schwartz information criteria.

corroborate our assumption to consider 10-year bond rate as exogenous variable. The χ^2 statistics for the null hypothesis $H_0 : \lambda_{10y}^+ = \lambda_{10y}^- = 0$ indicate that 10-year bond rate is *weakly* exogenous variable. Moreover, we cannot reject the hypothesis that $i_{FF,t}$ does *not* Granger-cause $i_{10y,t}$ at 5% level of significance. Contrary, the adjustment coefficients in the first equation, $\hat{\lambda}_{FF}^+$ and $\hat{\lambda}_{FF}^-$, are different from zero suggesting that the Fed Funds rate adjusts to the 10-year bond rate. In particular, the χ^2 test for the null hypothesis $H_0 : \lambda_{FF}^+ = \lambda_{FF}^-$, indicates that the adjustment mechanism characterizing the short-run dynamics of the Fed Funds rate is asymmetric, meaning a varying strength of attraction to equilibrium. Consistent with the asymmetric behavior of the term-spread, we find that $\hat{\lambda}_{FF}^+ < \hat{\lambda}_{FF}^-$. This implies that a positive shock at time t on 10-year bond rate ($\Delta i_{10y,t} > 0$) producing a change in the (*reverse*) term-spread below the threshold, such that $\Delta \hat{u} < -0.43$, triggers an upward adjustment of the Fed Funds rate toward the equilibrium faster than in the case of negative shock when deviations of the term-spread are above the threshold. Therefore, according to M-TAR findings, increases in the term-spread tend to persist but decreases (below the threshold) tend to revert quickly toward the equilibrium as result of a quick adjustment of the Fed Funds rate.

Economic interpretation is straightforward. Rising inflation expectations, reflected in an increase in 10-year bond rate, leads the Fed to respond quickly by increasing the Fed Funds rate while the Fed is relatively less active when observing a fall in inflation expectations or tolerable volatility in the bond market due to some uncertain in economic environment. In particular, our estimates

indicate that the Fed’s response to dampen inflationary expectations is almost five times faster ($\lambda_{FF}^- = -0.14$) than in the case changes in the bond rate movements are not significant to evidence inflation scares episodes ($\lambda_{FF}^+ = -0.03$).

In order to evaluate the effects of an exogenous shock on 10-year bond rates, we perform a simulation exercise. After shocking the system (12) by increasing (decreasing) $i_{10y,t}$, we generate the adjustment time-paths for the Fed Funds rate under the hypothesized two regimes ($\Delta\hat{u} < -0.43$ and $\Delta\hat{u} \geq -0.43$). Figure 3 presents the simulated Fed Funds rate adjustment process to 10-year bond rate’s one percentage point shocks. It shows that asymmetries in the Fed’s responses strongly depend on both the *direction* and *magnitude* of the shocks. We conclude that the asymmetric Fed’s behavior, in response to bond market signals, is coherent with the (implicit) inflation targeting monetary policy pursued since the 1980s. The priority to maintain the commitment to low inflation induced the Fed to be aggressive in trying to prevent rising inflationary pressures via higher Fed Funds rates and cautious about lowering them in the presence of decreasing inflation expectations.

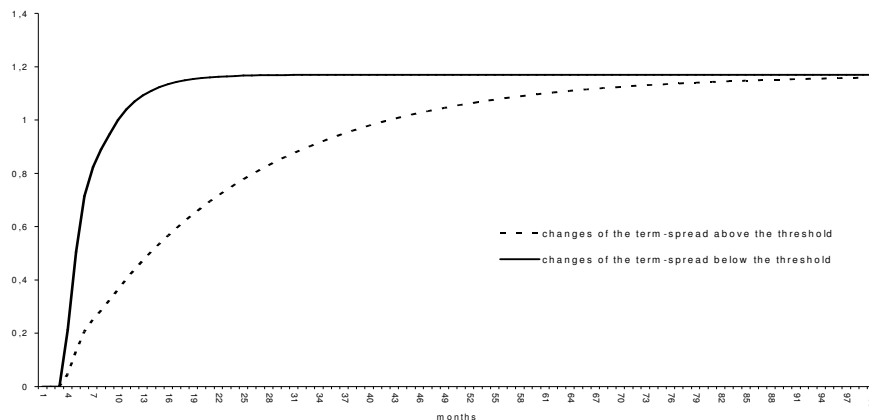


Figure 3: Adjustment paths of Fed Funds rate to changes in 10-year bond rate in both regimes ($\Delta\hat{u} < -0.43$ and $\Delta\hat{u} \geq -0.43$). The cumulative adjustment process is strongly asymmetric. Consistently with our findings that long-run elasticity between i_{FF} and i_{10y} is 1.17, i_{FF} approaches to this value faster when rising bond rates than in the case bond rates are falling.

6 Summary and conclusions

In this paper we have analyzed the influence of the Fed’s policy action on the behavior of the (*reverse*) term-spread defined as the residuals of the cointegrating relationship between Fed Funds rate and 10-year bond rate. After discussing

the importance of the 10-year bond rate as indicator of market's inflation expectations, we have investigated the nature of the equilibrium-type relationship between Fed Funds rate and 10-year bond rate using a Momentum Threshold Autoregressive Model (M-TAR). Estimates of the M-TAR model indicate that the term-spread dynamics is asymmetric. We suppose that this behavior is the effect of the asymmetric nature of the U.S. monetary policy over the last two decades. T-VECM estimates confirm our assumption: movements of the 10-year bond rate imply a different *timing* in the adjustment process of the Fed Funds rate depending on both the *magnitude* and the *direction* of such changes. In particular, given some uncertainty in deciphering bond rates signals and considering that resisting to inflation scares episodes is costly because induce the Fed to raise real short rates with potentially depressing effects on business condition, we have interpreted the estimated threshold value as a 'target' with respect to which changes of the term-spread in excess are assumed almost certainly due to a change in market's expected inflation. Therefore, we conclude that the priority to maintain the credibility as inflation 'fighter', induced the Fed to act quickly when changes in the term-spread reflect a rise in expected inflation and slowly when inflationary expectations decreases or upward movements in the term-spread are not so relevant to signal a crisis of confidence in the Fed's commitment to low inflation.

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