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The Impact of Monetary Policy Transparency on Risk and Volatility of Interest Rates: Evidence from the United States

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Abstract

Because of using unsatisfactory measures of the monetary policy transparency the existing literature found mixed empirical results for the relationship between the monetary policy transparency and risk as well as volatility. This paper extends the literature by using a recently developed monetary transparency index [Kia’s (2011) index] which is dynamic and continuous. It was found that the more transparent the monetary policy is the less risky and volatile the money market will be.

Keywords: Monetary policy transparency, risk, volatility, money market

JEL Codes: E43, E51, E52, E58
I. Introduction

The empirical findings of the impact of monetary policy transparency on the volatility of interest rates and/or forecast error/risk are mixed. Tabellini (1987), for example, shows that when market participants face parameter uncertainty (or multiplicative uncertainty) and learn over time, this learning process, using Bayes’ rule, is the source of additional volatility in asset prices. In this case, more transparency tends to reduce market volatility. Furthermore, Thornton (1996), using Fed Funds futures and Mean Squared Error, finds the consequences of the Fed’s policy shift toward immediate disclosure on the federal funds rate in 1994 resulted in a lower forecast error for all interest rates. Blinder (1998) argues, theoretically, that a more open public disclosure of central bank policies may enhance the efficiency of markets. Haldane and Read (2000), using dummy variables for the change in the monetary policy transparency, show that a higher monetary policy transparency leads to a lower conditional variance in the yield curve in the United Kingdom and the United States. Their cross-country (Italy, Germany, UK and US) empirical results also confirm their finding.

Blinder, et al. (2001), using descriptive accounts of transparency (do’s and don’ts of the central bankers’ actions), find a higher market transparency leads to a lower forecast error. Geraats (2002) theoretically shows the transparency reduces the variance of private sector forecast errors. Rafferty and Tomljanovich (2002) study whether the Federal Reserve System’s 1994 policy shift toward more open disclosure improved or worsened the predictability of financial markets. They find that since 1994, the forecasting error has decreased for interest rates on U.S. bonds of most maturity lengths, and that the expectations hypothesis has performed better at the low end of the yield curve.
Lasaosa (2005) analyzes the impact of the announcements on market activities and concludes that the increase in transparency facilitates the prediction of monetary policy in the UK once the latest macroeconomic data are known. She investigates the impact of the announcement five, fifteen and sixty minutes after an announcement. Then she compares the result with those days with no announcement at the same times. Swanson (2006) argues that increases in the monetary policy transparency in the US have played a significant impact on the private sector’s forecast improvement. He uses Fed futures and Eurodollar options rate to estimate and compare forecast errors as well as implied volatility, respectively, in different periods. Blinder, et al. (2008) conclude that central bank communication leads to improvement in the ability of the market to predict monetary policy.

Ehrmann and Fratzscher (2009), using both dummy variable and coding approach (i.e., the value of −1 for statements indicating an easing inclination, 1 for statements indicating a tightening, and 0 otherwise), find that the timing of statements is important. Specifically, statements in the pre-FOMC (Federal Open Market Committee) purdah resulted in a rise of market volatility, but statements in the post-FOMC purdah period and outside the purdah led to significantly lower market volatility. Biefang-Frisancho Mariscal and Howells (2010), using dummy variables, show that the various forms of Fed communications are highly significant on the volatility and market’s expectations. Finally, Jansen (2011) shows that the clarity in central banks’ communications reduces the volatility of the medium-term interest rates.

However, there are also studies which found monetary policy transparency does not affect the volatility, but actually contributes to more volatility in the financial markets. In fact, in a 1976 Freedom of Information Act filing, the Fed argued in favor of secrecy motivated by its desire to reduce interest rate variability [see Goodfriend (1986)]. Dotsey (1987) argues that the
cleaner and more frequent the “signal” (or the more transparent monetary policy is) the larger the responsiveness of interest rates to news, and thus the greater their volatility.

Biefang-Frisancho Mariscal and Howells (2007), using dummy variables, update Haldane and Read’s (2000) study on the UK by investigating the changes in anticipation on either side of the two major reforms in UK policy making in 1992 and 1997. They conclude that the Bank of England independence and its associated reforms have not added to the understanding of monetary policy by market participants. Chadha and Nolan (2001) use dummy variables accounting for days on and before the announcement day to investigate the impact of the announcement on the volatility of interest rate over selected periods in the UK. They concluded that transparency did not contribute to the volatility of interest rate.

The main problem with the current literature, in general, is that central bank communications are difficult to measure. Furthermore, using dummy variables, as it was also mentioned by Lasaosa (2005) and Kia (2011), can result in a misleading conclusion as some announcements may have already been taken into account by the market participants before the announcement. Furthermore, as Blinder, et al. (2008) argue, there are unobserved factors which affect asset prices. A rise in the volatility as a result of central bank communications may be due to the reaction of financial markets to shocks other than central bank communication. Moreover, the communication of the central bank may be due to a sudden change in economic outlook or some other news which also increase the volatility in asset prices. Therefore, the higher volatility is not due to the central bank communications, but to the shocks which caused the communications.

Consequently, the dummy variables may not actually reflect the impact of the announcements or the change in the policy. Moreover, as also mentioned, among others, by Blinder, et al. (2008), the coding approach is subjective and there is always a possibility of
misclassifications. To investigate clearly the impact of monetary policy transparency on forecast error, risk or volatility we need to have an objective market-based monetary policy transparency index. Such an index should also be dynamic so that it can be used for a long range of time series data. Kia (2011) has developed such an index for the United States which can be used to measure monetary policy transparency for a country or, simultaneously, a series of countries, using time-series as well as cross-sectional data. Using Kia’s index, this study investigates the impact of monetary policy transparency on risk (forecast error) and volatility in the money market in the United States.

The next section briefly explains Kia’s index and will be followed with a section on the impact of the index on the risk in the money market. Section IV is devoted to the analysis of the impact of monetary policy transparency on the volatility in the money market in the United States. The final section provides some concluding remarks. The description of the data and the definitions of dummy variables are given in the Appendix.

II. Brief Description of Kia’s (2011) Index

Kia postulates that when there is no expected change in the structure of the interbank market and/or the credibility of the United States government to pay its debt at equilibrium we will have \( Dif_t = FF_t - TB_t \) = risk premium - maturity risk premium = Tdif_t, where FF is the Fed Fund rate, TB is the three-month Treasury Bill rate and Tdif is the trend value or the equilibrium level of Dif. Under full transparency Dif is equal to Tdif and under opacity he finds Dif as:

\[
Odif_t = FF_{t-1} + \alpha (OER_t - ER_{t-1}) + \alpha \tau_{t-1} - TB_{t-1} - \gamma_1 (FF_t - TFF) \\
- \gamma_2 \left\{ \frac{\sigma_{T+1}^2}{\sigma_{T+1}^2 + \sigma_\tau^2} (\alpha \tau_t - TFF_{T+1}) + [TFF_{T+1} - TFF_T] \right\},
\]

where Odif is Dif under opacity, OER is the banking system optimal excess reserve, ER is the actual excess reserve, \( \tau \) is the shock to the economy, TFF is the target Fed Funds rate, and \( TFF \) is the mean of the target
rate. \( \sigma_{T+1}^2 \) is the variance of the target rate at the next FOMC meeting day, \( \sigma_{\tau}^2 \) is the variance of \( \tau \), and \( \alpha \) and \( \gamma \)'s are coefficients, see Kia (2011), Equation (6). Kia found the deviation between Dif under opacity and full transparency (D) to be:

\[
D_t = \text{Odif}_t - \text{Tdif}_t = + \gamma_2 \left\{ \left[ \frac{\sigma_{T+1}^2}{\sigma_{T+1}^2 + \sigma_{\tau}^2} \right] \left( \alpha \tau_t \right) - \text{TFF}_{T+1} \right\}.
\]

He considered the maximum/minimum of \(|D_t|\), at the event (FOMC meetings) dates in the sample period, to be the least/most transparent monetary policy over the period, and calculated the index (defined basic index) \( T_t = \text{transparency index} = \frac{100}{e^{\hat{D}_t}} \). If \(|D| = 0\%, we will have \( T = 100\%. Clearly the higher is \(|D_t|\), the lower will be the transparency index. In this formula, Kia calculated Tdif as the average of Dif between two event days. Since event days are at irregular intervals, Kia (2011) takes the quarterly average of the basic index to construct quarterly observations of the index.

He also developed the high frequency data of the index \( \hat{T}_t = 100/e^{\hat{D}_t} \), where \( \hat{D}_t = |Dif_t - Adif_t| \), where Dif\(_t\) is defined as before (= FF\(_t\) - TB\(_t\)) and Adif\(_t\) = \( \frac{1}{n} \sum_{j=1}^{j} \text{Dif}\_t \). Here \( j \) is the last event day and \( n \) is the number of days since the last event day. Using \( \hat{D}_t \), we can calculate an index for non-event days \( \hat{T}_t \). Kia’s index is dynamic and can be used to calculate an index for each minute of the day using the intra-day observations on FF and TB.

### III. Risk in the Money Market and the Monetary Policy Transparency

In this section we will investigate, using Kia’s index, whether a higher monetary policy transparency leads to a higher or lower forecastability of the market participants. Consequently, they will ask for a lower or a higher risk premium as the Fed conveys more of its private
information to the market. See Data Appendix for the sample period and sources of the data. Note that Kia’s index is market based, i.e., it reflects what market participants perceive from hints, actions or reactions (to exogenous shocks) of the monetary authorities and not what these authorities intend to convey to the market. In other words, the public availability of the data does not suffice to achieve transparency, see Kia (2011). As he stressed, market participants may observe a different norm/direction in the policy during the day or within a month or a period than what the central bank actually follows.

In this study, risk is measured by using the pure (rational) expectations model of the term structure (RE). According to RE the term premia are set identically to zero, which implies that at any moment in time, the expected TB, for example, prevailing at the beginning of three months from now \((1 + 3 \text{TB}_t^e)\) should be equal to the implied forward three-month Treasury bill rate \((\text{FTB}_t)\) in the absence of term premium or any other risk. From the first statement of the theory [e.g., Van Horne (1965)], we know that \(\text{FTB}_t = \left(1 + \frac{\text{TB}_6}{4}\right)^2/(1 + \frac{\text{TB}_3}{4})\) – 1. Here \(\text{TB}_6\) is the six-month spot rate and we assume both six- and three-month spot Treasury bill rates are at the annual rate. Specifically, we can write:

\[
1 + 3 \text{TB}_t^e = \text{FTB}_t. \tag{1}
\]

If this equality is violated, investors and speculators, trade three- and six-month Treasury bills, to capture potential speculative profits, until Equation (1) is restored. For example, if \(1 + 3 \text{TB}_t^e > \text{FTB}_t\), speculators sell their six-month bills and buy three-month bills, pushing the price of six-month bills down (\(\text{TB}_6\) will go up) and increasing the price of three-month bills up (\(\text{TB}_3\) will go down). This speculative activity continues until the potential for speculative profits is eliminated, i.e., \(1 + 3 \text{TB}_t^e\) is again equal to \(\text{FTB}_t\). Furthermore, by orthogonal decomposition at any given time \(t\), we have:

\[
\text{TB}_t = \text{TB}_t^e + V_t \tag{2}
\]
where $V_t$ is the agents’ forecast error in the absence of transaction costs, risk and other premia (including term premium, liquidity premium and reinvestment premium). Substituting (1) in (2) yields:

$$TB_{t+1} = FTB_t + V_{t+1}. \tag{3}$$

If the market is efficient, $TB_{t+1} - FTB_t = V_{t+1}$ is stationary [e.g., Campbell and Shiller (1987)] and, in the absence of risk premia and transaction costs, has a zero mean. The error term ($V_t$) is stationary as both Dickey-Fuller and Phillips-Perron tests reject the null hypothesis that $V_t$ is not stationary. The absolute value of the augmented Dickey Fuller $t$ was estimated to be 6.97 and the absolute value of the Phillips-Perron non-parametric $t$ for the lag length of 4 was estimated to be 7.23, both $t$ statistics results are higher than 3.43 (1% critical value). However, the mean of $V_t$ over our sample period was found to be -0.31%, at the annual rate, with an autocorrelated-heteroskedastic adjusted $t$ statistic of -20.64. The mean of the absolute value of $V$ was found to be 0.41%, at the annual rate, with an autocorrelated-heteroskedastic adjusted $t$ statistic of 32.90. Both of these means are far from being zero, indicating term premium or other risk premia exist, assuming a trivial transaction cost. Although a completely different approach was used, this result (i.e., on average, the RE hypothesis is valid in the United States money market, and risk premia exist) is consistent, among many others, with the finding of Van Horne (1965), Mankiw and Miron (1986) and Taylor (1992).

We will, consequently, modify Equation (3) to

$$TB_{t+1} = FTB_t + RP_{t+1} + V_{t+1} = FTB_t + W_{t+1}, \tag{4}$$

---

1 The lag length in augmented Dickey-Fuller or Phillips-Perron nonparametric tests was obtained according to AIC and SC criteria for a global lag of 20 days.

2 Autocorrelation is due to the overlapping observations. I used Newey and West’s (1987) robust error for 5-order moving average to correct the standard error.
where RP is risk premia and $W_t = RP_t + V_t$. Note that RP includes term, liquidity, interest exposure and reinvestment risk premia where reinvestment risk premium has a negative effect on RP. To test the impact of monetary policy transparency on the forecast errors or risk in the money market we need to test the relationship between $W_t$ in Equation (4) and the transparency index. For arguments and econometric tests on the relationship between transparency and forecast errors of market participants, see, e.g., Thornton (1996), Haldane and Read (2000) and Blinder et al. (2001).

We estimate the following equation:

$$|W_t| = \xi_0 + \xi_1LT_{t-1} + DUM_{t-1}' \varsigma + \epsilon_t,$$  \hspace{1cm} (5)

where $|W_t|$ is the absolute value of the forecast error from Equation (4), $LT_t$ is the logarithm of $\hat{T}_t$, $\xi$’s are constant parameters, $\varsigma$ is a vector of constant parameters and $\epsilon_t$ is the white noise disturbance term. Vector DUM is defined as:

$$DUM = (M_t, T_t, WED_t, TH_t, D851231_t, D861231_t, GREEN_t, Bernanke_t, OCT87_t, ASIA_t, TA_t, TAF_t, SWED_t, REMA_t, minutes_t, transcripts_t, state_t, D940418_t, D970819_t, lrr_t, D981015_t, D99518_t, D000202_t, D010103_t, D010418_t, D020319_t, EDAY_t, TARATE_t, USCrisis_t).$$

For the definition of these dummy variables see the Appendix. DUM is included in the equation in order to capture the impact of monetary policy regime changes as well as other shocks on the risk premia.

Note that variables LT and DUM enter in Equation (5) with one lag length (three months earlier) since the implied forward rate was used three months before (at the time of forecast) the actual rate was realized. Since our sample is daily observations, LT is our extended index and the lag length is 90 days. Note that the index and the calculated forward rate [an element of $|W_t| (= TB_t - FTB_{t-1})$] have the same lag length in Equation (5). However, there is no theoretical reason to believe that the index can be influenced by the calculated forward rate. Furthermore, we can
investigate the causality between these two stationary variables by estimating each variable by its 20 lagged values as well as the lagged values of the other variable.\textsuperscript{3}

By doing so, we found the Wald test on the coefficients of twenty lagged values of $LT_t$ in a regression of $LT_t$ on its twenty lagged values as well as twenty lagged values of $|W_t|$ is 177.76 ($p$-value=0.00), while the Wald test on twenty lagged values of $|W_t|$ is 17.77 ($p$-value=0.60). At the same time, the Wald test on the coefficients of twenty lagged values of $|W_t|$ in a regression of $|W_t|$ on its twenty lagged values as well as twenty lagged values of $LT_t$ is 5916.18 ($p$-value=0.00), while the Wald test on twenty lagged values of $LT_t$ is 42.45 ($p$-value=0.00). This result implies that $LT_t$ Granger causes $|W_t|$ while $|W_t|$ does not Granger cause $LT_t$. Specifically, we conclude $LT_t$ is strongly exogenous in Equation (5).\textsuperscript{4}

Equation (6) is the parsimonious estimated result of Equation (5), where the figures in brackets are standard errors adjusted for autocorrelation and heteroskedasticity.

$$
|W_t| = 1.11 (0.21) - 0.14 (0.05) LT_{t-1} - 0.16 (0.02) TA_{t-1} + 0.15 (0.7) REMA_{t-1} \\
-0.18 (0.06) GREEN_{t-1} + 0.53 (0.03) D010917_{t-1} + 0.43 (0.01) D010103_{t-1} \\
+ 0.05 (0.01) D010418_{t-1} + 0.58 (0.16) OCT87_{t-1} - 0.63 (0.25) D851231_{t-1} \\
-1.73 (0.38) D861231_{t-1} - 0.22 (0.07) BERNANKE_{t-1}. \quad (6)
$$

$R^2 = 0.12$, $\sigma = 0.37$, Godfrey(5) = 10029 (significance level=0.00), White=664 (significance level=0.00), ARCH(5)=3467 (significance level=0.00), RESET=0.57 (significance level=0.64), where $\sigma$ is the standard error of the estimate, Godfrey is the five-order Godfrey’s (1978) test, White is the White’s (1980) general test for heteroskedasticity, ARCH is the five-order Engle’s (1982) test and REST is the Ramsey (1969) misspecification test. According to Godfrey’s test

\textsuperscript{3} For the stationary test result on the index see Kia (2011).

\textsuperscript{4} All Wald test results are adjusted for autocorrelation and heteroskedasticity.
result, the error is autocorrelated and both White and ARCH tests results confirm the existence of heteroskedasticity. Because of overlapping observations these results would be expected.

The estimated coefficient of LT is negative and statistically significant implying that as the monetary policy is more transparent the forecast errors and risk premia will fall. This result confirms the finding of Thornton (1996), Haldane and Read (2000), Blinder, et al. (2001) and Swanson (2006). According to the estimated coefficient of dummy variable TA, the Fed policy of announcing policy decision (Target rate) at the conclusion of each FOMC meeting, since February 4, 1994, resulted in a lower risk and forecast error and risk in the money market in the United States.

The positive and statistically significant estimated coefficient of REMA implies that modifying the reserve maintenance period from one week (for most large institutions) to two weeks (for all institutions) in February 1984 resulted in a higher forecast error, while the negative and statistically significant coefficient of the dummy variables GREEN and BERNANKE means the forecast error and risk in the money market fell during the tenure of Chairman Greenspan and continued to fall during the tenure of Chairman Bernanke. The estimated coefficient of D010917, as one would expect, is positive and statistically significant, which reflects a higher risk environment associated with September 2001.

Furthermore, as the positive and statistically significant estimated coefficient of D010103, D010418 and OCT87 indicates, the unexpected change in the target rate on January 3, 2001 and April 18, 2001 as well as during the October 87 stock crisis resulted in a higher forecast error. The surprising result is the estimated coefficients of dummy variables D851231 and D861231. Both are negative and statistically significant implying a high volatility of FF on December 31, 1985 and December 30 and 31, 1986 resulted in a lower forecast error on those days.
I also used quarterly averages of the daily observations to create a quarterly sample to test the impact of monetary policy transparency on forecast errors/risk. For this test I used Kia’s (2011) basic index T. For quarterly observations of T, following Kia (2011), I also took the average of the basic index in each quarter. Since the quarterly constructed index is highly correlated with the set of variables in DUM, most of the estimated coefficients, including the estimated coefficient of the index, while having the correct sign were statistically insignificant (the estimated result is available upon request). I, therefore, report the estimated Equation (5) without DUM. The estimated Equation (5) with LT being the logarithm of the quarterly index is as follows, where the figures in brackets are standard errors adjusted for autocorrelation:

$$|W_t| = 3.15 (1.26) - 0.62 (0.28) LT_{t-1}. \quad (7)$$

$$\bar{R}^2=0.04, \sigma=0.38, DW=1.32, \text{Godfrey}(5)=3.53 \text{ (significance level}=0.003), \text{White}=1.20 \text{ (significance level}=0.94), \text{ARCH}(5)=6.67 \text{ (significance level}=0.25), \text{RESET}=0.02 \text{ (significance level}=0.99).$$

According to Godfrey’s test result the error is autocorrelated.

According to Equation (7) the estimated coefficient of LT is negative and statistically significant confirming the result reported above. In sum, we conclude that the more the monetary policy is transparent the less will be forecast errors and risk premia.

**IV. Volatility in the Money Market and the Monetary Policy Transparency**

In this section I will examine the relationship between monetary policy transparency, using Kia’s (2011) index, and risk, measured by the volatility, in the money market. We know that FF and TB are cointegrated and the adjustment toward the long-run equilibrium largely

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5 Again I found LT to be strongly exogenous in Equation (5) as the Wald test on the coefficients of four lagged (incorporating a year) values of $|W_t|$ in a regression of $|W_t|$ on its four lagged values as well as four lagged values of LT, is 57.12 ($p$-value=0.00), while the Wald test on four lagged values of LT is 16.29 ($p$-value=0.00). Alternatively, the Wald test on the coefficients of four lagged values of LT in a regression of LT on its four lagged values as well as four lagged values of $|W_t|$ is 27.43 ($p$-value=0.00), while the Wald test on four lagged values of $|W_t|$ is 6.14 ($p$-value=0.19). This result implies that LT, is strongly exogenous in Equation (5).
occurs through the movements in the FF rate rather than the TB rate \cite{Sarno and Thornton 2003 and Kia 2010}. Hence, we would expect the volatility of the daily movements in FF, say VFF, to affect the volatility of the daily movements in TB, say VTB, (i.e., the risk in the money market). Let us assume the volatility of TB is a function of the volatility of FF and policy regime changes as well as other exogenous shocks specified in EDUM defined below. Furthermore, assume such a relationship has a linear approximation as specified by Equation (8):

$$V_{TB_t} = \Gamma_0 \gamma LT_t + \sum_{i=1}^{k} \Gamma_i VFF_{t-i} + \sum_{i=1}^{k} \Phi_i VTB_{t-i} + EDUM_t' \Gamma + \epsilon_t,$$

(8)

where $\Gamma_0, \ldots, \Gamma_k$, $\Phi_0, \ldots, \Phi_k$, $\gamma$ and $\Gamma$ are constant parameters. Dummy vector EDUM = (DUM, STU, TUE1, HB1, HA1, HB3, HA3, LDY, LDBYA, LQBA). Dummy variables included in EDUM are defined in Table 1. To capture the possible volatility in the money market created by other factors, like window dressing, holidays and other seasonality, following Hamilton (1996), beside dummy variables in vector DUM, I included dummy variables STU, TUE1, HB1, HA1, HB3, HA3, LDY, LDBYA, LQBA and LQ, see Table 1 in Data Appendix.

Note that in Equation (8) VFF is predetermined and if $\gamma$ is negative/positive then the higher the monetary policy transparency (LT) is, the lower/higher the volatility of the three-month Treasury bill rate will be. Following, among many, Schwert (1989) and Kia (2003), the methodology developed by Davidian and Carroll (1978) was used. Let $x$ be any variable in column vector $x_t = (\Delta TB_t, \Delta FF_t,)'$ and estimate Equation (9) for $\Delta TB_t$ and $\Delta FF_t$.

$$x_t = \sum_{i=1}^{20} \alpha^x x_{t-i} + EDUM_t' \mu^x + u_xt, \text{ where } u_xt \sim \text{niid}(0, \Sigma).$$

(9)

The parameters $\alpha^x$'s and vector $\mu^x$ are assumed to be constant. I assume a lag length of 20 days (reflecting a month) is sufficient for the market participants to learn from the past movements in
the TB rate. The dummy variables included in vector EDUM capture the shocks on the rate during our sample period. Furthermore, a 20th-order autoregression for the absolute values of errors from Equation (9), including dummy variables in vector EDUM that allow for different daily standard deviations, should be estimated:

$$|\hat{u}_{xt}| = \sigma^x_t = \sum_{i=1}^{20} \delta^x_i \sigma^x_{i-1} + \text{EDUM}_t' \eta^x + v_t,$$  

(10)

where $\delta^x_i$, for $i = 1$ to 20 and the column vector $\eta^x$ are constant parameters. The absolute value of the fitted value of $u_{xt}$ (i.e., $|\hat{u}_{xt}|$) is the standard deviation (adjusted for heteroskedasticity and autocorrelation due to overlapping observations) of $x_t$, for $x_t = \Delta TB_t$ and $\Delta FF_t$. However, since the expected error is lower than the standard deviation from a normal distribution, following Schwert (1989), all absolute errors are multiplied by the constant 1.2533.

As it was also mentioned by Kia (2003), the conditional volatility in Equation (10) represents a generalization of the 20-day rolling standard estimator used by Officer (1973), Fama (1976) and Merton (1980). This is due to the fact that the conditional volatility estimated by Equation (10) allows the conditional mean to vary over time in Equation (9), while it also allows different weights to be applied to the lagged absolute unpredicted changes in Treasury bills and Fed funds rates.

Note that here the conditional mean of these rates was also allowed to vary with the shocks represented by dummy variables included in vector EDUM. Furthermore, Engle (1993) reviews the merit of this measure of volatility, among others. This measure of volatility is similar to the autoregressive conditional heteroskedasticity (ARCH) model of Engle (1982), which, in its various forms, has been widely used in the finance literature. Davidian and Carroll (1978) argue that the specification in Equation (10) based on the absolute value of the prediction errors is more robust than those based on the squared residuals in Equation (9).
However, it should be noted that the variables in equations (8) and (10), excluding dummy variables, are generated regressors. Consequently, when these equations are estimated, their \( t \) statistic should be interpreted with caution. To cope with this problem, following, among many, Kia (2003), the equation for the conditional volatility [i.e., Equation (8)] is estimated jointly with the equations determining the conditional volatilities of \( \Delta TB \) and \( \Delta FF \) using the generalized Least Squares (GLS) estimation procedure (SUR).\(^6\)

In the GLS system, two equations are generated by Equation (9), two equations are generated by Equation (10) and including Equation (8) a system of five equations with 7303 observations (with a final sum of 6658 usable observations) is estimated. In the GLS estimation, for each equation and the system of equations, I used Newey and West’s (1987) robust error for 5-order moving average to correct for heteroskedasticity and autocorrelation. The GLS estimator incorporates the possibility of cross-equation correlation among the error terms. The final parsimonious GLS estimation result of Equation (8) is given by Equation (11), where standard errors appear in brackets.\(^7\)

\[
\begin{align*}
VTB_t &= 0.06 (0.004) - 0.008 (0.001) LT_t + 0.009 (0.003) VFF_{t-1} + 0.42 (0.01) VTB_{t-1} \\
&
+ 0.21 (0.01) VTB_{t-3} - 0.006 (0.001) TAF - 0.002 (0.001) GREEN_t \\
&- 0.01 (0.002) USCrisis_t
\end{align*}
\]

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\(^6\) See Kia (2003), Footnote 4, for a full explanation on why in our case the GLS estimation technique should be used.

\(^7\) The stationarity test results for VTB are as follows: The absolute value of the augmented Dickey Fuller \( t \) for a lag length of 5 is 18.05 and the absolute value of the Phillips-Perron non-parametric \( t \) test for the lag length of 3 is 57.75, both \( t \) statistics are higher than 3.43 (1% critical value) indicating the conditional volatility VTB is stationary. The stationarity test results for VFF are as follows: The absolute value of the augmented Dickey Fuller \( t \), for a lag length of 9 is 16.04 and the absolute value of the Phillips-Perron non-parametric \( t \)-test for the lag length of 4 is 53.26, both \( t \) statistics are higher than 3.43 (1% critical value) indicating the conditional volatility VFF is stationary. Note that the lag length for these tests was determined in such a way that, for a global lag of 20 days, the AIC and SC criteria are at their minimum.
$\bar{R}^2=0.52$ and $\sigma=$ the standard error of the regression $=0.01$, Godfrey(5)=$604.61$
(significance level=$0.00$), White=$782.84$ (significance level=$0.00$), ARCH(5)=$3608.87$
(significance level=$0.00$), RESET=$33.31$ (significance level=$0.00$). As it would be expected
specification test results indicate both the existence of autocorrelation and heteroskedasticity.

The estimated coefficient of our monetary policy index (LT) is negative, indicating that a
more transparent monetary policy leads to a lower volatile money market. This result confirms
the finding of Tabellini (1987), Haldane and Read (2000), Swanson (2006), Ehrmann and
Fratzscher (2009) and Biefang-Frisanco Mariscal and Howells (2010), among many, that a
higher degree of transparency tends to lower market volatility. Among all dummy variables
included in EDUM, the coefficients of dummy variables GREEN, TAF and USCrisis were found
to be statistically significant. As the negative coefficient of dummy variable GREEN indicates,
the volatility and risk in the money market fell during the tenure of Chairman Greenspan. As it
would be expected, the estimated coefficient of TAF is negative, implying that the policy regime
change of announcing policy decisions after each FOMC meeting, since October 1989, led to a
lower volatility in the money market in the United States. The estimated sign of dummy variable
USCrisis is negative reflecting that the current US financial crisis led to a lower volatility of
interest rate.

I also repeated the above exercise with our quarterly data explained above and Kia’s
basic index. I found the basic index has a negative effect on volatility, but the estimated
coefficient was statistically significant only at 93% level (the full result is available upon
request). This could be due to a lack of observations on the index in each quarter. In sum, it was
shown in this section that a higher monetary policy transparency leads to a lower volatility in the
money market in the United States.
V. Concluding Remarks

The empirical findings on the impact of monetary policy transparency on risk and volatility are mixed. One possible explanation for these findings is the lack of a proper index to measure monetary policy transparency. The literature so far has been using dummy variables constructed according to the central banks’ announcements and actions or coding approach. However, these measures for the monetary policy transparency have some limitations, such as a lack of an objectively designed index or indexes without time-series properties. Furthermore, dummy variables may not really reflect the impact of an announcement or changes in the policy. That is because market participants may have already incorporated the impact of a policy change before the central bank communication.

This is also possible when shocks which resulted in the central bank communications already affected the economy, say the volatility. Clearly using different measures for monetary policy transparency results in mix empirical findings. This is especially true when these different measures have different shortcomings. Kia (2011) developed an objective market-based monetary policy transparency index. The index is dynamic and continuous and can be used to measure monetary policy transparency for a country or, simultaneously, a series of countries, using time-series as well as cross-sectional data. In this paper I used Kia’s (2011) index to investigate the impact of monetary policy transparency on risk and volatility in the United States.

It was found that the more transparent the monetary policy in the United States is the less risky and volatile the money market will be. Moreover, the rational expectations model of the term structure is valid in the United States money market, but risk premia in this market exist. It was also found that modifying the reserve maintenance period from one week (for most large institutions) to two weeks (for all institutions) in February 1984 resulted in a higher risk in the
money market, while the risk fell during the tenure of Chairman Greenspan and continued to fall during the tenure of Chairman Bernanke.

Furthermore, September 2001, the unexpected change in the target rate on January 3, 2001 and April 18, 2001 as well as during the October 87 stock crisis resulted in a higher risk in the United States. However, a high volatility of FF on December 31, 1985 and December 30 and 31, 1986 resulted in a lower risk on those days. It was also found that the policy regime change of announcing policy decisions after each FOMC meeting, since October 1989, and the current US financial crisis led to a lower volatility in the money market in the United States. Finally, we conclude that the practice of a more transparent monetary policy leads to more stability (lower volatility) and lower risk in the financial markets.

REFERENCES


Data Appendix

The daily data on the effective Fed funds rate and the three-month Treasury bill rate (secondary market) for the period 1982 (October 5)- 2010 (September 30) are used. The choice of the sample period is based on the availability of data for the construction of the transparency index which was developed by Kia (2011). The period covers more than 27 years with 7304 effective daily observations. The source of these data is the St. Louis Federal Reserve website. The effective Fed funds rate is a weighted average of the rates on Fed funds reported daily by a group of brokers to the Federal Reserve Bank of New York. Both Fed funds and three-month Treasury bill rates are expressed as bond equivalent yields on a 365-day basis. The definitions of the dummy variables in DUM and EDUM are given in the following table.

<table>
<thead>
<tr>
<th>Dummy Variables</th>
<th>Definitions</th>
</tr>
</thead>
<tbody>
<tr>
<td>(M_t, T_t, WED_t, ) and (TH_t)</td>
<td>(M_t, T_t, WED_t, ) and (TH_t = 1) for Mondays, Tuesdays, Wednesdays and Thursdays, respectively, and are equal to zero, otherwise.</td>
</tr>
<tr>
<td>(D851231) and (D861231)</td>
<td>(D851231_t) and (D861231_t = 1) on December 30 and 31, 1985 and December 31, 1986, respectively, and are equal to zero, otherwise. These dummy variables are included to capture the high volatility of Fed funds rate on those days.</td>
</tr>
<tr>
<td>(GREEN_t)</td>
<td>(GREEN_t = 1) since August 11, 1987 when Alan Greenspan was appointed chair of the Fed and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(BERNANKE_t)</td>
<td>(BERNANKE_t = 1) since February 1, 2007 when Chairman Bernanke was appointed chair of the Fed and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(OCT87_t) and (ASIA_t)</td>
<td>(OCT87_t = 1) for October 19 to 30, 1987 and is equal to zero, otherwise, and (ASIA_t = 1) for October 17 to 30, 1997 and is equal to zero, otherwise. These dummy variables account for the October 87 and Asian crises, respectively. In both events, central banks in industrial countries flooded the money markets with liquidity to ease the downfall in the stock markets. The easing of the markets took at least until the end of October of the year the crisis took place.</td>
</tr>
<tr>
<td>(TA_t)</td>
<td>(TA_t = 1) since February 4, 1994 (when Fed started to announce target changes) and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(TAF_t)</td>
<td>(TAF_t = 1) since October 19, 1989 (when the Fed adopted the practice of changing the FF targets by 25 or 50 basis points) and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(SWED_t)</td>
<td>(SWED_t) accounts for settlement days on Wednesdays, i.e., it is equal to one on Wednesdays when it is a settlement day and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(REMA_t)</td>
<td>(REMA_t = 1) since February 2, 1984 when the reserve maintenance period was modified from one week (for most large institutions) to two weeks (for all institutions) and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(minutes_t)</td>
<td>(minutes_t = 1) since March 23, 1993, when the Fed began releasing the minutes of the FOMC meetings (with 6-8 week lag) and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(transcripts_t)</td>
<td>(transcripts_t = 1) since November 16, 1993 when the Fed began releasing the transcripts of the FOMC meetings (with 5-year lag) and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>(state_t)</td>
<td>(state_t = 1) since August 16, 1994 when the Fed began describing the state of the economy and further rationale for the policy action after FOMC decisions, and is equal to zero.</td>
</tr>
<tr>
<td>$\text{lrr}_t$</td>
<td>$\text{lrr}_t = 1$, since July 30, 1998 and is equal to zero, otherwise. On March 26, 1998 the Fed moved from contemporaneous reserve requirements back to lagged reserve requirements. This policy went into effect with the reserve maintenance period beginning July 30, 1998. Irr accounts for this policy regime change.</td>
</tr>
<tr>
<td>$\text{D970819}_t$</td>
<td>$\text{D970819}_t = 1$ since August 19, 1997, when the FOMC started to include a quantitative Fed funds target rate in its Directive to the New York Fed Trading Desk, and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{D99518}_t$</td>
<td>$\text{D99518}_t = 1$ since May 18, 1999, when the Fed extended its explanations regarding policy decisions, and started to include in press statements an indication of the FOMC’s view regarding prospective developments (or the policy bias), and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{D000202}_t$</td>
<td>$\text{D000202}_t = 1$ since February 2, 2000, when the FOMC started to include a balance-of-risks sentence in its statements replacing the previous bias statement, and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{D020319}_t$</td>
<td>$\text{D020319}_t = 1$ since March 19, 2002, when the Fed included in FOMC statements the vote on the directive and the name of dissenter members (if any), and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{D940418}, \text{D981015}, \text{D010103}, \text{D010418}$, and $\text{D010917}_t$</td>
<td>$\text{D940418}_t$, $\text{D981015}_t$, $\text{D010103}_t$, $\text{D010418}_t$, and $\text{D010917}_t$, are equal to one for April 18, 1994; October 15, 1998; January 3, 2001; April 18, 2001 and September 17, 2001 (when the Fed changed the FF target rate outside its regular meetings), respectively, and is equal to zero otherwise.</td>
</tr>
<tr>
<td>$\text{EDAY}_t$</td>
<td>$\text{EDAY}_t = 1$ for the days (“event”) when the Fed funds target rate was changed whether at a regularly scheduled FOMC meeting, or otherwise, and also for the days on which the FOMC met, but did not change the target rate. It is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{TARATE}_t$</td>
<td>$\text{TARATE}_t = 1$ for the days when the Federal funds target rate actually was changed and is equal to zero, otherwise. These days can be among the regularly scheduled FOMC meeting dates or other days. Note that TARATE is a subset of EDAY, as it excludes the days when FOMC met, but did not change the target.</td>
</tr>
<tr>
<td>$\text{USCrisis}_t$</td>
<td>$\text{USCrisis}_t = 1$ since September 2007, is equal to zero, otherwise. The housing bubble started to burst in 2006, and the decline accelerated in 2007 and 2008. Housing prices stopped increasing in 2006, started to decrease in 2007. The decline in prices meant that homeowners could no longer refinance when their mortgage rates were reset, which caused delinquencies and defaults of mortgages to increase sharply, especially among subprime borrowers. “It was in August 2007 when BNP Paribas, a large French bank, froze withdrawals in three investment funds that people began to panic. If a bank with zero obvious exposure to the U.S. mortgage sector could have this measure of difficulty, anyone could be hiding untold losses. This marked the official beginning of the credit crisis.” Harrison (2008).</td>
</tr>
<tr>
<td>$\text{STU}_t$</td>
<td>$\text{STU}_t = 1$ on Tuesdays before settlement Wednesdays and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{TUE1}_t$</td>
<td>$\text{TUE1}_t = 1$ on Tuesdays before settlement Wednesdays if Wednesday was a holiday, and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{HB1}_t$, $\text{HA1}_t$, $\text{HB3}_t$, and $\text{HA3}_t$</td>
<td>$\text{HB1}_t = 1$ for the day before a one-day holiday, and is equal to zero, otherwise. $\text{HA1}_t = 1$ for the day after a one-day holiday, and is equal to zero, otherwise. $\text{HB3}_t = 1$ for the day before a three-day holiday, and is equal to zero, otherwise. $\text{HA3}_t = 1$ for the day after a three-day holiday, and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{LDY}_t$ and $\text{LDBYA}_t$</td>
<td>$\text{LDY}_t = 1$ for the last day of the year, and is equal to zero, otherwise. $\text{LDBYA}_t = 1$ for 2 days before, 1 day before, on, 1 day after, or 2 days after the end of the year, and is equal to zero, otherwise.</td>
</tr>
<tr>
<td>$\text{LQBA}_t$ and $\text{LQ}_t$</td>
<td>$\text{LQBA}_t = 1$ for the day before, on, or after the last day of the first, second and third quarters, and is equal to zero, otherwise. $\text{LQ}_t = 1$ for the last day of the first, second, third and fourth quarters, and is equal to zero, otherwise.</td>
</tr>
</tbody>
</table>

* Most part of this table is taken from Kia (2010). See also Kia (2011).

$\text{DUM} = (\text{M}_t, \text{T}_t, \text{WED}_t, \text{TH}_t, \text{D851231}_t, \text{D861231}_t, \text{GREEN}_t, \text{BERNANKE}_t, \text{OCT87}_t, \text{ASIA}_t, \text{TA}_t, \text{TAF}_t, \text{SWED}_t, \text{REMA}_t, \text{minutes}_t, \text{transcripts}_t, \text{state}_t, \text{D940418}_t, \text{D970819}_t, \text{lrr}_t, \text{D981015}_t, \text{D99518}_t, \text{D000202}_t, \text{D010103}_t, \text{D010418}_t, \text{D010917}_t, \text{D020319}_t, \text{EDAY}_t, \text{TARATE}_t, \text{USCrisis}_t).$

$\text{EDUM} = (\text{DUM}_t, \text{STU}_t, \text{HB1}_t, \text{HA1}_t, \text{HB3}_t, \text{HA3}_t, \text{LDY}_t, \text{LDBYA}_t, \text{LQBA}_t, \text{LQ}_t).$