University of Wisconsin Milwaukee
UWM Digital Commons

Theses and Dissertations

May 2013

Essays on Monetary Policy and Financial Markets

Hardik Arvind Marfatia
University of Wisconsin-Milwaukee

Follow this and additional works at: https://dc.uwm.edu/etd

Part of the Economics Commons

Recommended Citation
Marfatia, Hardik Arvind, "Essays on Monetary Policy and Financial Markets" (2013). *Theses and Dissertations*. 135.
https://dc.uwm.edu/etd/135

This Dissertation is brought to you for free and open access by UWM Digital Commons. It has been accepted for inclusion in Theses and Dissertations by an authorized administrator of UWM Digital Commons. For more information, please contact open-access@uwm.edu.

ESSAYS ON MONETARY POLICY AND FINANCIAL MARKETS

by

Hardik A. Marfatia

A Dissertation Submitted in

Partial Fulfillment of the

Requirements for the Degree of

Doctor of Philosophy

in

Economics

at

University of Wisconsin Milwaukee

May 2013

ABSTRACT

ESSAYS ON MONETARY POLICY AND FINANCIAL MARKETS

by

Hardik A. Marfatia
The University of Wisconsin Milwaukee, 2013.
Under the Supervision of Dr. N. Kundan Kishor.

My dissertation utilizes the valuable information present in forward looking financial securities to understand important aspects of monetary policy analysis. In the first chapter, I attempt to address the long standing empirical challenge of estimating the forward-looking component of the New Keynesian Phillips Curve (NKPC). Since future inflation expectations are unobservable, I use the information in the inflation-indexed bond market to estimate the NKPC for the U.K. In order to account for any possible measurement error present in the inflation-indexed bond market proxy, the unobserved component model is used. This approach has the advantage of being able to extract the unobserved inflation expectations using the Kalman filter and to jointly estimate the parameters of the NKPC. Results show that the estimated inflation expectations from the model play a significant role

in explaining the inflation dynamics in the U.K. Evidence also suggests in favor of the Phillips curve tradeoff between inflation and output.

The second chapter of my dissertation uses the information in the federal funds futures market and conduct an event study to evaluate the impact of Fed's policy actions on 35 leading stock price indices across the world. Using a time-varying parameter model, the study finds significant time-variation in the response of global equity markets to U.S. monetary policy surprises with a greater response during the crisis periods. Interestingly, in the recent financial crisis stock markets in Europe and the U.S. responded negatively to unanticipated interest rate cuts by the Fed. Further, the response of the Asia-pacific region and Latin American stock markets is at least as strong as the response of the U.S. and the European equity markets.

The final chapter of my dissertation asks the question whether the federal funds futures rate contains information about the Treasury bill rate. Using high frequency daily data, I examine the dynamic relationship between these two interest rates. The results show that the one month federal funds futures rate move together with the 3-month T-bill rate in the long-run. More importantly, in contrast to the existing literature, any deviation from this long-term equilibrium is corrected by subsequent movements in both the T-bill rate and the federal funds futures rate.

©Copyright by Hardik A. Marfatia 2013

All Rights Reserved.

I dedicate this thesis to my loving parents and my beloved wife. I owe you everything and wish I had enough words to express my deep love and gratitude for you.

TABLE OF CONTENTS

LIST OF FIGURES

LIST OF TABLES

ACKNOWLEDGEMENTS

All great achievements become possible *only* with the help and support of the people around you. I would like to express my heartfelt gratitude to all the teachers for their guidance, my family and relatives for their love and support, and numerous friends for all the fun and exciting experiences in my path to Ph.D.

I would like to express my special appreciation and sincere gratitude to my advisor N. Kundan Kishor for the continuous support and guidance throughout my Ph.D study and research. I greatly appreciate his contributions of time, ideas, and efforts to make my Ph.D. experience productive and stimulating. His patience, motivation, enthusiasm, and deep insights makes him a great mentor and an ideal advisor. His advice on both research as well as on my career have been invaluable.

I would also like to thank the members of my PhD committee, Mohsen Bahmani, Rebecca Neumann, Suyong Song, and Filip Vesely for their support and helpful suggestions. Their guidance has served me well and I owe them my heartfelt appreciation. Special thanks to Antonio Galvao and Hamid Mohtadi for their helpful career advice and suggestions. My sincere thanks also goes to Mohsen Bahmani for offering

me diverse teaching opportunities. It has been my great privilege to be a part of UWM Econ dept, and all my professors, colleagues, and friends here will always remain dear to me.

I will forever be thankful to my former Professor Romer Correa. He has been helpful in providing advice many times in my career path to the graduate school.

This achivement would have not been possible without the love and inspiration of my parents, Arvind Marfatia and Charu Marfatia and my sister Dr. Khushali Ratra. I greatly owe to my wife, Rajvi Marfatia, whose love, encouragement, and patience has been truly instrumental in finishing this journey. I gratefully acknowledge the support of my uncle Dr. Vijay Dave and my aunt Ranjan Dave. I also want to thank my in-laws for their unconditional support.

CHAPTER 1
Overview

Expectations of economic agents plays a vital role in both theoretical and empirical models. Measuring the expectations are also necessary to evaluate the effectiveness of monetary policy. While explicit incorporation of expectations in many theoretical models is driven almost by necessity, the empirical estimation of expectations is notoriously challenging. It is common to use survey based measures or econometric techniques to capture future expectations of economic and policy variables. In recent times, the prices of certain types of forward-looking financial market instruments have been found to contain significant information that embeds the expectations of economic agents. These instruments act as proxies for future expectations of economic variables and have several appealing properties. They provide the advantages of being a natural market based metric, are entirely forward-looking, are available at high-frequency and under certain conditions are highly efficient. My thesis attempts at utilizing this superior forward-looking information content of financial markets to understand some of the key issues involved in monetary policy analysis.

Estimating inflation dynamics is one of the central topics of discission among monetary policy makers and macroeconomist. The New Keynesian Phillips Curve

(NKPC) has been found to provide a reasonable micro-founded explanation to how the macroeconomic inflation process is driven by the real economy, past inflation and expected future inflation. since inflation expectations are not directly observable one of the empirical challenges in the estimation of NKPC comes from the measurement of the forward looking component of the model. However, measuring these expectations are imperative for researchers interested in understanding the inflation dynamics of the economy.

The first essay attempts at addressing this challenge of measuring inflation expectations and estimate the inflation dynamics in the U.K.. I utilize the information in the inflation-indexed bond markets. The inflation premium - the yield spread between inflation-indexed bonds and nominal bonds - is found as an effecient proxy for expected future inflation. However, due to considerations of liquidity and other risk factors, it becomes necessary to account for an unobservable measurement error in the inflation premium. This issue is addressed by using unobserved component approach developed by Harvey (1985) and others. The main advantage of this approach comes from using the Kalman filter to explicitly estimate the unobserved expected inflation from the observed inflation premium. Our results show that estimated inflation expectations from our model plays a significant role in explaining the inflation dynamics in the U.K.. The evidence also suggests that the inflation expectations from the model are better able to capture the actual evolution of inflation process as compared to the inflation premium as a proxy for expected inflation. Further, mild but statistically significant tardeoff

is found between inflation and the real economy.

It is true that estimating the inflation dynamics is central to monetary policy making. However, the most immediate impact of monetary policy actions are felt on the financial markets. It is of considerable interest to monetary policy makers to assess the impact of policy changes on asset prices, particularly stock prices. The analysis is particularly complicated by the problems of endogenity and identification of monetary policy shocks. In analyzing the impact of monetary policy, the event study methodology has been found to greatly reduce the problem of endogenity (Bernanke and Kuttner 2005; Cook and Hahn 1989; Gürkaynak et al. 2005). For identifcation of policy shocks it is imperative to capture the market expectation of future policy actions. The federal funds futures market is found to possess significant information that can "efficiently" predict the future monetary policy actions (Gürkaynak et al. 2007; Krueger and Kuttner 1996; Robertson and Thornton 1997). The second essay, I utilize the information in the federal funds futures market to analyze the impact of U.S. monetary policy shocks on the 35 leading stock markets of the world. The unexpected component of FOMC policy announcements is extracted from federal funds futures data using Kuttner (2001) methodology. An event study is then undertaken to measure the high-frequency response of stock returns to unexpected changes the Fed's policy announcements. Commonly, similar event study is done using the fixed coefficient approach (Bernanke and Kuttner 2005; Ehrmann and Fratzscher 2009; Hausman and Wongswan 2011; Wongswan 2009). This is in contrast to anecdotal and em-

pirical evidence which suggest that the impact of monetary policy on stock market may vary with time (Andersen et al. 2007; Bekaert and Harvey 1995; Campbell et al. 1997).

To account for the possible time variation in the response of stock markets the second essay adopts a time-varying coefficient framework. Following the pioneering work of Cooley and Prescott (1976), the time-variation is modeled as driftless random walks, and is estimated using the maximum likelihood via the Kalman filter. To account for possible heteroscedasticity present in the error term the paper follows Harvey et al. (1992) and allow conditional heteroskedasticity in the error term. The results show significant time-variation in the response of the global equity markets to U.S. monetary policy surprises, where an unanticipated interest rate cut leads to an increase in stock returns. The findings also suggest that the foreign stock markets respond more to U.S. monetary policy surprises during the crisis periods. The paper also find that unlike previous episodes of crisis, the stock markets in Europe and the U.S. responded negatively to unanticipated interest rate cuts by the Fed during the recent financial crisis.

Understanding the interest rate movements is another topic that interests monetary policy makers, macroeconomist and financial market participants. Given the significance of the information content in the federal funds futures rate, it would be interesting to investigate how the fed funds futures rate link with the short term interest rates in the U.S.. This question is investigated in the third essay of my thesis. Using high frequency daily data I examine the dynamic relationship

between the federal funds futures rate and the 3-month T-bill rate. The results show that the one month federal funds futures rate is cointegrated with the 3-month T-bill rate, and thus move together in the long-run. Further, any deviation of the one month federal funds futures rate and the T-bill rate from their long run equilibrium is corrected by subsequent movements in both the federal funds futures rate and the T-bill rate. Using this long term relationship the federal funds futures rate and the T-bill rate is decomposed into a trend and cycle using the multivariate Beveridge-Nelson methodology. The results show a big positive cycle in the federal funds futures rate before 2008 implying a future downward movement in federal funds futures rate.

CHAPTER 2

Estimating New Keynesian Phillips Curve for the U.K. : Evidence from Inflation-Indexed Bonds Market

2.1 Introduction

It is widely accepted that inflationary expectations are not directly observable. However, measuring these expectations become necessary for researchers interested in understanding the inflation dynamics of the economy. The New Keynesian Phillips Curve (NKPC) has been found to provide a reasonable micro-founded explanation to how the macroeconomic inflation process is driven by the real economy, the past inflation and the expected future inflation. One of the empirical challenges in estimation of the NKPC come from the measurement of the forward-looking component. Usually based on rational expectations assumption, researchers have used realized inflation or survey based forecasts as a proxy for expected inflation and estimated NKPC using Generalized Method of Moments (GMM). However, Bårdsen et al. (2004) and many other studies show that using GMM for NKPC estimation may not be the best strategy since the model is likely to be either weakly identified or dynamically mis-specified.

In the present study, I use the valuable information in the inflation-indexed bond markets to estimate the NKPC for the U.K.. Using the U.K. data provides the advantage of the longest sample period that is available for the inflation indexed bond markets. The inflation premium rates – the yield difference between inflation-indexed and conventional treasury bonds – are generally considered to embed the future inflation expectations. These rates have the appealing characteristics of being a market based metric of inflation expectations that gets updated on daily basis and is available for several horizons.[1] However, due to consideration of liquidity and other risk factors, the inflation premium is related to, but not equal to the expectations of future inflation (Gürkaynak et al. 2010; Shen and Corning 2002). In light of these evidences, Gulyàs and Startz (2006) uses the inflation premium as a proxy for the forward-looking component of NKPC and use the spot and forward inflation premium rates as instruments. However, this approach is likely to generate inconsistent estimates since it can easily be contemplated that the shocks to actual inflation are correlated with the inflation premium rate. Further, it does not take into account the liquidity and other risk factors that are highlighted in the literature.

In this paper, instead of using the inflation premium as a direct proxy I argue that neither the underlying inflation expectations nor the measurement errors are directly observable. Hence, I use the unobserved component (UC) approach proposed by Harvey (1985); Watson (1986) and Clark (1987) to address the issue

[1] Scholtes 2005 shows the relative merit of inflation expectation extracted from indexed linked bonds as compared to the survey based proxies.

of measurement error. Using this approach, the observed inflation premium is decomposed into two unobserved components namely, the inflation expectations and the measurement error. Gaining insights from Morley et al. (2003) and Lee and Nelson (2007), I model inflation expectations as a random walk process and the measurement error as an autoregressive process. One of the benefits of such specification is that it opens the possibility to estimate the correlation between shocks to expected inflation and shocks to measurement error (Morley et al. 2003). I also consider other correlation structures of the model as a robustness test. The distinctive feature of this approach is that the Kalman algorithm can be used to extract the unobserved inflation expectations from the inflation premium and jointly estimate the parameters of the NKPC model. Since this strategy exploits the dependence structure between different components of the model it is expected to provide more reliable and accurate estimates.

The results show that future inflation expectations play a statistically significant role in explaining the actual inflation dynamics in the U.K.. The coefficient of forward-looking component extracted from the unobserved component model is estimated around 0.34 and is statistically significant. The estimates also imply the presence of substantial degree of price stickiness. Importantly, evidence suggests a significant relationship between inflation and the stance of real economy. The slope of Phillips curve is flat in the range of 0.15. These results are robust to various alternative maturities of the inflation premium rates considered in the paper.

I find that inflation expectations estimated from the unobserved component model adequately accounts for measurement error present in the inflation premium and very closely tracks the actual inflation process. The estimated inflation expectations offers a higher explanatory power relative to the inflation premium as proxy of the actual inflation rate. Further, the correlation between unobserved inflation expectations and measurement error is significant around -0.86. The measurement errors that are attributable to the liquidity premium and other risk factors are also significant, highly persistent, and possibly volatile.

The remainder of the paper is organized as follows. Section II presents some fundamental features of the inflation indexed bond markets and its link to the expected future inflation. A brief literature overview about NKPC is discussed in section III. Section IV specifies the unobserved component model with the discussion of the results in section V. A variety of robustness check is conducted in section VI, followed by concluding remarks in the last section.

2.2 Inflation-Indexed Securities and Inflation Expectations

The U.K. government first issued inflation-indexed financial securities (also called Index-linked gilts) in the early 1980's. Initially, this market was thin but given the appealing properties of these securities its market has grown steadily over time. In 2009, more than 25% of £353 billion total outstanding debt stock was in the form of index-linked bonds.

The underlying intent of these bonds is to protect the bond holders from future

rise in inflation by adjusting the coupons and principal to the evolution of inflation price index. Hence, these bonds are also called real return bonds. For example, assume that an investor buys a 4-year conventional Treasury bond at par with a 6 percent coupon rate and also buys a 4-year inflation-indexed bond at par with a coupon rate of 4.5 percent. Now, suppose the inflation rate measured by the retail price index (RPI) turns out to average 2.5 percent in the next 4 years. In this case the investor earns a 6 percent nominal return but the real return on this investment would only be 3.5 percent. On the other hand, the real return on the inflation-indexed bond would be 4.5 percent while the nominal return on it would be 7 percent. Thus, the real return of the investors are protected by the risk of unexpected inflation over the span of these 4 years.

The difference between the yields on conventional Treasury bonds (quoted in nominal terms) and inflation-indexed bonds (quoted in real terms) of comparable maturity is called the *inflation premium*. In a world where the investors are risk-neutral and only concerned with real returns, ideally the inflation premium would be equal to the expected future inflation rate. This is because any information that the market believes is going to affect the inflation in the future would be quickly reflected in the prices of these two types of securities so as to equalize the real return. Thus, if the 10-year real return bonds are currently trading at 4.5 percent and comparable maturity nominal government bonds are trading at 6.5 percent yield, then the yield difference of 2 percent is attributed to expected future inflation. In real world, in order to derive inflation expectations from the inflation

premium further adjustment is needed for two principle reasons. Since nominal bonds are exposed to uncertain future real returns the risk averse investors would demand a *inflation risk premium*. On the other hand, relatively small market size of inflation indexed bonds leads to the presence of *liquidity risk premium* on these bonds.

The inflation premium as a measure of future inflation expectations offers significant merit over survey based methods. Survey based measures usually cover a small portion of the population, are updated infrequently, and the accuracy may be compromised by survey respondents that answers the questions casually. In contrast, the inflation premium is based on wide investors base, is entirely forward-looking, timely, and updated every working day for a wide range of maturities. Further, given the amount of investors money in the index-linked gilts market the price discovery is expected to be far more accurate. Thus, it provides a superior measure of inflation expectations as compared to other measures like the survey based measures.

Cedric Scholte's of the Bank of England discusses the dual benefits of the inflation premium as a proxy for expected inflation (Scholtes 2005). At the shorter end of the forecast horizon, the study shows the information superiority of the 2-year inflation premium as compared to Barclays Basix surveys of inflation expectations. Meanwhile, the longer-term inflation premium rates serve as a measure of the monetary policy maker to assess the inflation credibility. Breedon and Chadha (1997) also show that the forecasting performance of inflation expectations derived from

the index-linked bond market is atleast as good as the expectations derived using either the nominal term structure alone or the forecasts derived from macroeconometric models.

The choice of curve-fitting technique is important in obtaining the inflation premium rates. The U.K. index-linked gilt market has the advantage of having securities outstanding over wide range of maturities up to 25 years. This leads to a reasonably well-specified fitting of a relatively smooth yield curve. The Bank of England uses a modified cubic smoothing splines methodology to fit nominal and real yield curves.[2] The daily data for the inflation premium across various maturities for both spot and forward rates is published by the Bank of England. The spot inflation premium rates can be looked upon as an average rate of inflation expected to rule over a given period. Similarly, the forward implied inflation premium rates can be interpreted as the rate of inflation expected to rule over a given period which begins at some future date.[3]

In this study, I use the 4-year spot inflation premium for the baseline NKPC estimation. As a robustness check, I also use 4-year forward, 10-year spot and 10-year forward inflation premium rates. Even though it would be ideal to use shorter maturity yields as a measure of inflation expectations, the limitation of data availability dictates us to choose 4-year maturity as it is the shortest maturity available without any missing values.[4] The dataset spans across 21 years from the

[2]See Anderson and Sleath (2001) and Deacon and Derry (1994) for details.

[3]In the limit, one can calculate instantaneous forward implied inflation rates just as with real and nominal rates. See http://www.bankofengland.co.uk/statistics/Pages/yieldcurve/default.aspx for more details.

[4]While much shorter maturities exists, BOE states that "we only provide data at maturities

first quarter of 1988 to the last quarter of 2009.

2.3 Literature Review

In an influential work, Gali and Gertler (1999) developed and estimated a hybrid specification of the NKPC. This specification has been widely used as a description of macroeconomic inflation dynamics derived from micro-foundations. It shows how the current inflation rate is affected by past inflation, future inflation expectations and aggregate demand pressure. Specifically,

$$\pi_t = \alpha + \gamma^b \pi_{t-1} + \delta x_t + \gamma^f E(\pi_{t+1}|I_t) + \epsilon_{\pi,t} \tag{2.1}$$

where π_t and π_{t-1} is the actual and the lagged inflation rate and $E(\cdot|I_t)$ is the expectations of inflation in period $t+1$ conditional upon the information set available in period t. The variable x_t is the 'driving variable' that captures the aggregate demand side pressures. This may be represented by Hodrik Presscot filter de-trended output or real marginal costs (Gali and Gertler 1999; Gali et al. 2001 among others) or model based output gap (Neiss and Nelson 2005).

Barring the differences as to what constitutes the driving variable, this specification nests other forms of Phillips curve developed in the literature. For example, if $\gamma^b = 1$ and $\gamma^f = 0$ then it leads to the basic adaptive expectations Phillips curve with dominant inflation stickiness (Fuhrer 1997; Fuhrer and Moore 1995; Lindé

where we think the curve can be fitted so that it is stable and meaningful" (Notes on the Bank of England Yield Curves). Gulyàs and Startz (2006) also use 4-year and ahead maturities to capture inflation expectations in the estimation of the NKPC.

2005; Roberts 2005; Rudebusch 2002). Alternatively, $\gamma^b = 0$ and $\gamma^f = 1$ leads to a pure forward-looking NKPC formulation. Studies have found this specification to a more appropriate model for inflation dynamics (Cogley and Sbordone 2008; Gali et al. 2005; Sbordone 2002, 2005).

Given the importance of these parameters especially in the context of monetary policy analysis (see Clarida et al. 1999, 2000) it is not surprising that empirical estimations of the hybrid NKPC parameters has received widespread attention. However, rationalizing the empirical findings across studies is not straight forward due to differences in the exact specification, the proxy used for the driving variable x_t, the econometric technique used and the country and the period under scrutiny.

One of the key challenges involved in estimation of the hybrid NKPC is to measure the expectations of future inflation which are not directly observable. Most existing studies use realized inflation or survey based inflation forecast in conjunction with Generalized Method of Moments (GMM) estimation or other instrumental variable methods.[5] However, Bårdsen et al. (2004); Mavroeidis (2005); Nason and Smith (2008) discuss the limitations of GMM in forward-looking models that use large number of instruments and a general correction of the covariance matrix. These studies recommend against indiscriminate application of GMM in estimating NKPC, since it is very likely to be either mis-specified or spuriously identified. Lindé (2005) also find that single equation methods, e.g. GMM, are likely to produce imprecise and biased estimates for NKPC and argue in favor of

[5]It is common to use four to six lags of inflation together with other instruments like interest rates, exchange rates, stance of real economy etc to correct for high-order serial correlation in the covariance matrix.

the model with full information maximum likelihood (FIML).

As an alternative measure of future inflation rate, the inflation-indexed bonds market is found to contain useful information about investors forecasts of inflation expectations. The price of these financial securities provides natural market-based proxy of inflation expectations. They provide "the advantages of being available for a wide range of maturities, entirely forward-looking, timely, and updated every working day"(Scholtes 2005). Even while studies have found presence of significant information content in this market about inflation expectations, it is imperative to extract the inflation risk premium and liquidity premium components from the yields of these securities to estimate the underlying inflation expectations (Barra and Campbell 1997; Gürkaynak et al. 2010; Joyce et al. 2010; Pflueger and Viceira 2011; Shen and Corning 2002).

In other but related strand of literature focus on exploiting the Kalman filter technique to estimate the unobserved state variables within the state-space framework (Clark 1987; Harvey 1985; Watson 1986). Basistha and Nelson (2007) estimates a forward-looking NKPC for the U.S. by treating the output gap as unobservable and using Michigan survey of inflation expectations. In another study Kim et al. (2011) estimate unobserved trend-cycle models of the U.S. inflation that is consistent with NKPC. In a pure forward-looking specification ($\gamma^b = 0$), Lee and Nelson (2007) show the slope of the Phillips curve depends critically on the horizon of the forward-looking inflation expectations, provided the cyclical component of unemployment (a proxy for x_t) is highly persistent.

In this paper, I draw support from these different but related strands of literature. I use the superior information in the inflation-index bond markets to estimate a hybrid version of NKPC. I treat inflation expectations as unobserved and also allow it to be correlated with the unobserved measurement error present in the inflation premium.

2.4 Model Specification

2.4.1 Baseline Model

Following Gali and Gertler (1999), I specify the following hybrid Phillips curve that is based on optimizing forward-looking and backward-looking monopolistically competitive firms:

$$\pi_t = \alpha + \gamma^b \pi_{t-1} + \delta g_t + \gamma^f \pi_t^e + \epsilon_{\pi,t} \quad \epsilon_{\pi,t} \sim i.i.d.N(0, \sigma_\pi^2) \tag{2.2}$$

The above specification closely represents equation (2.1) except that I use π_t^e to represent the unobserved expected future inflation. To measure these expectations, I use current 4-year and 10-year spot and forward inflation premium rates that are derived from inflation-indexed bonds market. In the above equation the actual inflation rate π_t is calculated as annual percentage change in Retail Price Index (RPI). Specifically, $\pi_t = [ln(RPI_t) - ln(RPI_{t-4})]$x100. The seasonally adjusted RPI is sourced from Office for National Statistics, UK. The variable g_t represents the output gap computed as the deviation of real GDP from trend or

potential output based on standard Hodrik Prescott (HP) filter (Gali 2002; Jondeau and Bihan 2005; Mihailov et al. 2011). As a measure of output, I use log of seasonally adjusted quarterly real GDP from International Financial Statistics dataset. Any differences in the frequency of the data availability are matched to make it quarterly averages. The theory suggests the value of δ to be positive implying that any increase in the output above its trend will lead to upward pressure on inflation. Further, theoretically the sum of γ^b and γ^f is expected to be equal to one since there is no long run Phillips curve tradeoff.[6]

The basic OLS regression results from estimating the above equation are presented in Table 2.1. The results show the value of δ to be positive and significant signifying the presence of short run tradeoff between inflation and output (or unemployment). Evidence also suggests that backward looking component play a dominant role in explaining the actual inflation behavior in the U.K. The coefficient of future expectations of inflation (γ^f) is significant but much smaller in value around 0.13. Interestingly, the sum of the weights on backward-looking component (γ^b) and forward-looking component (γ^f) is nearly one.[7] Thus, in the long-run actual inflation is affected only by past and future inflation but not by the demand side pressures. This is consistent with the theory according to which there is no Phillips curve tradeoff in the long run.

[6] The underlying structural parameter of the model include the degree of price stickiness, the degree of "backwardness", the inter-temporal discount factor etc. See Gali and Gertler (1999) for detailed derivation.

[7] The Wald test statistic also show the null hypothesis of $\gamma^b + \gamma^f = 1$ cannot be rejected at even one percent confidence level. These results are not presented to maintain brevity but available upon request.

These findings hold regardless of the proxy used to measure inflation expectations. However, the results must be read with an important caveat that the only adjustment is in the form of Newey-West HAC Standard Errors & Covariance. No explicit treatment of measurement error present in the inflation premium is undertaken and which is absolutely necessary. I elaborate on this issue in the following section.

2.4.2 The Unobserved Component Model

The information in the inflation-indexed bonds is often the best proxy that captures future inflation expectations (Scholtes 2005). The inflation premium is derived as the yield difference between inflation-indexed bonds and the conventional treasury bonds. However, presence of liquidity premium on conventional treasury bonds and the differences in the risk profiles of these two securities opens the possibility of an unknown measurement error. Specifically, the inflation premium equals inflation expectations plus inflation risk premium on the conventional treasury bonds minus liquidity premium on the inflation-indexed bonds (Gürkaynak et al. 2010; Pflueger and Viceira 2011; Shen and Corning 2002). My model acknowledges that the inflation premium measures the true underlying expectations of inflation only with an error. Further, neither the inflation expectation nor the measurement errors are directly observable. Hence, rather than using the observed proxies of inflation expectations like survey based measures, I treat inflation expectations and measurement error as unobserved state variables in the state-space

representation. Particularly,

$$\pi_t^B = \pi_t^e + V_t \tag{2.3}$$

where, π_t^B is the inflation premium obtained from the inflation-indexed bond yields. This is decomposed into inflation expectations π_t^e and the measurement error V_t, both components being unobserved. The state-space approach allows us to extract these two components from the observed inflation premium rate.

The next step is to specify the laws of motion for these two latent state variables. Gaining insights from Lee and Nelson (2007), I assume inflation expectations to be considerably persistent and thus specify inflation expectations to follow a random walk process. To model measurement error, I follow Morley et al. (2003) and specify it as an AR(2) process. It is necessary to specify sufficient dynamics in the measurement error in order to achieve meaningful decomposition of the inflation premium. The other advantage of allowing measurement error to follow an AR(2) is that it allows us to estimate correlation between shocks to inflation expectations and the measurement error.

$$\pi_t^e = \pi_{t-1}^e + \epsilon_{\pi^e,t} \qquad \epsilon_{\pi^e,t} \sim i.i.d.N(0, \sigma_{\pi^e}^2) \tag{2.4}$$

$$V_t = \phi_1 V_{t-1} + \phi_2 V_{t-2} + \epsilon_{v,t} \qquad \epsilon_{v,t} \sim i.i.d.N(0, \sigma_v^2) \tag{2.5}$$

The main NKPC equation (2.2) together with the inflation premium equation (2.3) constitutes the measurement equation of the state-space representation. The transition equation takes the form specified in equations (2.4) and (2.5). The state-

space representation of the model is given by:

Measurement Equation

$$\begin{bmatrix} \pi_t \\ \pi_t^B \end{bmatrix} = \begin{bmatrix} \alpha \\ 0 \end{bmatrix} + \begin{bmatrix} \gamma^b & \delta \\ 0 & 0 \end{bmatrix} \begin{bmatrix} \pi_{t-1} \\ g_t \end{bmatrix} + \begin{bmatrix} \gamma^f & 1 & 0 & 0 \\ 1 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \pi_t^e \\ \epsilon_{\pi,t} \\ V_t \\ V_{t-1} \end{bmatrix}$$

$$y_t = \mu + A z_t + H\beta_t$$

Transition Equation

$$\begin{bmatrix} \pi_t^e \\ \epsilon_{\pi,t} \\ V_t \\ V_{t-1} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & \phi_1 & \phi_2 \\ 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \pi_{t-1}^e \\ \epsilon_{\pi,t-1} \\ V_{t-1} \\ V_{t-2} \end{bmatrix} + \begin{bmatrix} \epsilon_{\pi^e,t} \\ \epsilon_{\pi,t} \\ \epsilon_{v,t} \\ 0 \end{bmatrix}$$

$$\beta_t = F\beta_{t-1} + U_t, \quad U_t \sim i.i.d.N(0, Q)$$

Allowing for the correlation between inflation expectations and the measurement error, we have

$$\begin{bmatrix} \epsilon_{\pi^e,t} \\ \epsilon_{\pi,t} \\ \epsilon_{v,t} \\ 0 \end{bmatrix} \sim i.i.d.N \left(\begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma^2_{\pi^e} & 0 & \rho_{\pi^e v}\sigma_{\pi^e}\sigma_v & 0 \\ 0 & \sigma^2_{\pi} & 0 & 0 \\ \rho_{\pi^e v}\sigma_{\pi^e}\sigma_v & 0 & \sigma^2_v & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix} \right)$$

The above state-space specification involves three shocks in the system, namely the shocks to inflation ($\epsilon_{\pi,t}$), the shocks to inflation expectations ($\epsilon_{\pi^e,t}$) and the shocks to the measurement error ($\epsilon_{v,t}$). These three shocks have among themselves three covariance terms in the variance-covariance matrix (Q). I assume the two unobserved components to be non-orthogonal, and thus allow for the correlation between inflation expectations and measurement error (i.e. $\rho_{\pi^e,v} \neq 0$). The actual inflation rate is assumed to be orthogonal to the other two shocks in the system.[8] These restrictions are adequate to identify the parameters of the model. The parameters of interests are mainly δ since it captures the relative trade off between inflation and the real economy, and γ^f (consequently γ^b) as it measures the forward-looking (and backward-looking) component of NKPC model. All the parameters are estimated using the maximum likelihood method and then the unobserved components are estimated using Kalman filter technique.

Using the unobserved component model within the state-space framework to estimate NKPC provides with some attractive features. First, it allows us to extract the two unobserved component using the Kalman filter technique. Second, it

[8]Relaxing these restrictions and testing for alternative specifications is included as a part of the robustness check in the paper.

jointly estimates the NKPC equation along with the unobserved inflation expectations and measurement error. This strategy is expected to provide more accurate and reliable parameter estimates because it exploits the dependence structure between different components of the model.

2.5 Estimation Results

Table 2.2 presents the results from the unobserved component model represented by equation (2.2)-(2.5). Based on theoretical foundations (Gali and Gertler 1999) and common practice in empirical studies, I first assume that there is no long run trade off between inflation and aggregate demand side pressures. The implication is that the joint effect of lagged inflation and the future inflation on current inflation is equal to unity. Hence the estimation is undertaken by restricting $\gamma^b + \gamma^f = 1$. The results from OLS estimation suggest that this assumption is not very restrictive.[9] However, I also estimate the benchmark model without imposing this restriction.

The results in Table 2.2 show estimates from both the restricted (column 2-3) and the unrestricted (column 4-5) model. Evidence suggests existence of weak but statistically significant relationship between inflation and the stance of real economy. As theory suggests the coefficient of output gap (δ) is positive and statistically significant. The estimated slope of Phillips curve is relatively flat at

[9]The sum of the two components are driven by deep parameters which takes the form of the discount factor (See Gali and Gertler (1999)). Consequently, it is common in literature to restrict the value of to 0.99 (Neiss and Nelson 2005; Nelson and Nikolov 2004; Sbordone 2003; Smets and Wouters 2003).

0.15 for the restricted model and 0.19 for the unrestricted model. Thus, a 100 basis point increase in the actual output relative to the trend output leads to about 0.15-0.19 basis point acceleration in the current inflation rate.

A flat slope of Phillips curve is well known in studies that incorporate the forward-looking inflation component. Gulyàs and Startz (2006) use the inflation premium as a measure of expected inflation and find similar results.[10] In contrast, Balakrishnan and Lopez-Salido (2002) and Kara and Nelson (2003) find that the relationship between marginal cost and inflation disappears in the mid-1980s in case of the U.K.[11]

Evidence in Table 2.2 further reveals statistically significant role of future inflation in explaining the U.K. inflation dynamics. The coefficient on the forward-looking component of NKPC (γ^f) is 0.34 and statistically significant. Consequently under one interpretation, about 34 percent and 66 percent of economic agents have a forward-looking and backward-looking pricing setting behavior respectively. I also find that shocks to inflation expectations as measured by the standard deviation (σ_{π^e}) is 0.66 and statistically significant. Thus, these shocks play an important role the evolution of actual inflation process.

Most of the studies that use GMM find a much bigger role of forward-looking component. However, GMM estimation is most likely to be biased in favor of

[10] This paper uses two stage least square with instruments like the spot and forward inflation premium rates to address the measurement error present in the inflation premium.

[11] In the context of the U.S. Basistha and Nelson (2007) estimate NKPC and model the output gap as unobserved. They also find the significant role of forward-looking component but flat Phillips curve in the range of 0.18-0.27. Lee and Nelson (2007) also find dominance of forward-looking component but find CBO measure of output gap to be a small driving factor in inflation dynamics of the U.S.

apparently dominant forward-looking behavior, irrespective of the true nature of the forward and backward-looking dynamics of inflation (Bårdsen et al. 2004; Jondeau and Bihan 2005; Mavroeidis 2005; Nason and Smith 2008). Also, Jondeau and Bihan (2005) find the empirical choice of the forcing variable has negligible impact on the estimated degree of forward-lookingness.

Since the underlying issue in the estimation process involves measuring the expected inflation in the NKPC equation it becomes important to analyze the behavior of the two unobserved components of my model. Results in Table 2.2 show that significant and strong negative correlation exists between the inflation expectations and the measurement error. This correlation coefficient ($\rho_{\pi^e v}$) is estimated at -0.84 in the restricted model. Thus, any shock to the expectations of future inflation will immediately shift the path of inflation expectations, leaving the the prices of inflation linked bonds and consequently on the inflation premium temporarily away from the true inflation expectations. This implies a contemporaneous negative correlation between the measurement error and the inflation expectations.

Both the AR coefficients that describe the transition process of the measurement error is statistically significant. This evidence provides support in favor of the AR(2) specification for measurement error. However, I acknowledge that this is not the standard procedure since the model is not identified if the null hypothesis of $\phi_2 = 0$ is true. The sum of these AR estimates (ϕ_1, ϕ_2) is 0.94 reflecting a highly persistent measurement error in the inflation premium rates as a

measure of the true underlying expected inflation. Also, the standard deviation of the measurement error (σ_v) is statistically significant at 0.43 suggesting that shocks to measurement errors are important in explaining the observed inflation premium. In accordance with the existing literature findings (Barra and Campbell 1997; Gürkaynak et al. 2010; Joyce et al. 2010; Pflueger and Viceira 2011; Shen and Corning 2002), this indicates a persistent and possibly volatile presence of the liquidity premium and other risk factors that contribute to the measurement error.

The next question is how well does inflation expectations extracted from my approach capture the evolution process of the actual inflation in the U.K.? The preliminary evidence to this question can be found in Figure 2.1. This figure plots the actual inflation rate, the inflation premium and the extracted inflation expectations of the unobserved component model. A visual inspection clearly reveals that inflation expectations from my model tracks actual inflation strikingly better than inflation expectations as measured by the inflation premium rates. For most part of the sample period before recent financial turmoil the inflation premium over predicts the actual inflation. Scholtes (2005) find that both the (quarterly) Basix survey and the inflation premium generally overpredicted the two-year-ahead inflation after 1991. Though the inflation premium rate tracks two-year-ahead RPI inflation better than survey forecasts. Thus, inflation expectations estimated from the UC model seems to ranked better then both survey based measure and the inflation premium.

To further support this finding, I also run a simple OLS of actual inflation rate on these two alternative proxies. If my model is better able to capture the true inflation dynamics as compared to the inflation premium then at the minimum it should be able to offer higher explanatory power. The results of this test are presented in Table 2.6. As expected from the graphs, inflation expectations from the unobserved model (column 2-3) offer higher explanatory power as compared to the inflation premium (column 4-5). The R-square in case of extracted inflation expectations is 0.96 as compared to 0.91 in case of 4-year spot inflation premium. Further, the coefficient on expected inflation is higher in column 2 (0.56) as compared to column 4 (0.14). This implies significantly greater impact of inflation expectations from the UC model on actual inflation as compared to the expected inflation from the inflation-indexed bonds market.

These results should not be interpreted as case of lack of information in the inflation-indexed bond markets. In fact it means just the opposite. The inflation-indexed bond markets contain significant information of future inflation rate but only after appropriate accounting of the (unobserved) measurement errors.

2.6 Robustness Check

To check for robustness of my results, I undertake two exercises. First exercise involves considering alternative proxies in the form of different maturities for both the spot and forward inflation premium rates. In the second robustness test, I relax the zero restrictions imposed on the correlation of other components of

the state-space system presented in equations (2.2)-(2.5). For the rest of the analysis, I assume the sum of the coefficient of backward-looking and forward-looking component equal to theoretically recommended and empirically supported value of one.

The results presented in Table 2.3 show the estimation from using 4-year forward (column 2-3), 10-year spot (column 4-5) and 10-year forward (column 6-7) inflation premium rates as alternative proxy of inflation expectation in NKPC equation. Clearly, the key coefficients of the NKPC equation is largely unaffected. The coefficient of expected inflation ($\gamma^f = 1 - \gamma^b$) is 0.49 for 4-year forward, 0.46 for 10-year spot and 0.26 rates for 10-year forward inflation premium rates. Thus, using these alternative proxies, I find an even greater role of the forward-looking component in explaining the inflation dynamics in the U.K.. The slope of Phillips curve is statistically significant and still relatively flat about 0.14 to 0.18. The correlation between the two unobserved components and the standard deviations of all the shocks in the system are significant and close to the benchmark case.

Next, I consider all the possible correlation structures between various shocks of the state-space represented in equations (2.2)-(2.5). Specifically, I allow for correlation between actual inflation and inflation expectations ($\rho_{\pi\pi^e} \neq 0$) and then between actual inflation and the measurement error ($\rho_{\pi v} \neq 0$). Further, since there are three shocks in the system, I also relax the zero restriction condition for the two correlations one at a time. This leads to three possible correlation combination between actual inflation shocks, inflation expectations shocks, and measurement

error shocks.

Table 2.4 reveals the results of relaxing the zero covariance between the three shocks in the system one at a time. Columns 4-5 show that the slope coefficient of the Phillips curve is statistically significant at 0.15 when the correlation between the actual inflation and unobserved expectation of inflation is allowed to be non-zero. The statistical significance also holds in case of a non-zero covariance between actual inflation and measurement error (Columns 2-3), though with a much flatter slope coefficient. The coefficient of forward-looking component is smaller in magnitude (γ^f is around 0.06 to 0.20) its effect is still not negligible and has statistically significant impact on inflation dynamics. When I restrict only one covariance to be equal to zero the results presented in the Table 2.5 show that the coefficient of future inflation rates (γ^f) is around 0.3 and the slope of the Phillips curve (δ) is between 0.11-0.16. The statistically significant negative relationship between unobserved inflation expectations and measurement error as well as the importance of the standard deviations of the shocks is not affected qualitatively as well.

Further, Figure 2.2 as well as Figures 2.3 and 2.4 plot the comparison between the actual inflation, the inflation premium and the expected inflation extracted from the different specification discussed above. All the plots of inflation expectations estimated using unobserved component model fit the actual inflation considerably better or at least as well as the inflation premium rates as a measure of future inflation.

Overall, whether I consider alternative maturity spectrum of inflation indexed bonds or spot and forward inflation premium rates or alternative covariance structures in the unobserved component model, the main results are unaffected qualitatively and to most extent even quantitatively. These findings reinforce results found earlier in the benchmark case of using the 4-year spot inflation premium as a measure of expected inflation in NKPC estimation.

2.7 Conclusions

The expectations of the future inflation pose a significant challenge in estimating the forward-looking component of the NKPC model since the expectations are not directly measurable. I address this issue by using the information in the inflation-index bond markets. However, due to liquidity and other risk factors, the inflation premium rates derived from these markets are not exactly equal to the future expected inflation. To account for this, I allow for the possible measurement error in my model. Using an unobserved component approach, I extract inflation expectations from the inflation premium and jointly estimate a hybrid NKPC for the U.K. The results show that the estimated inflation expectations from unobserved component model play a significant role in explaining the actual inflation dynamics in the U.K. Evidence also suggest in favor of the possible trade-off between inflation and the real economy. The main results hold regardless of the maturity of inflation premium rate (both spot and forward) or different correlation structures of the state-space model.

Table 2.1: OLS Estimate of NKPC for the U.K.

	4-Year Spot		4-Year Forward		10-Year Spot		10-Year Forward	
Parameter	Estimate	SE	Estimate	SE	Estimate	SE	Estimate	SE
α	0.044	0.109	0.026	0.138	-0.050	0.134	-0.085	0.147
γ^b	0.880	0.041	0.907	0.043	0.879	0.043	0.907	0.032
δ	0.124	0.052	0.139	0.047	0.130	0.048	0.153	0.043
γ^f	0.086	0.041	0.065	0.054	0.111	0.056	0.099	0.050
R-Sq	0.937		0.936		0.938		0.936	
Obs	87		87		87		87	

The Table shows the OLS estimation of the following hybrid version of NKPC equation: $\pi_t = \alpha + \gamma^b \pi_{t-1} + \delta g_t + \gamma^f \pi_t^e + \epsilon_{\pi,t}$. The π_t^e is based on 4-year and 10-year spot and forward inflation premium rates. g_t represents output gap based on the Hodrick Prescott (HP) filter. The standard errors (SE) are Newey-West HAC Standard Errors and Covariance adjusted.

Table 2.2: NKPC Parameter Estimates of Benchmark Unobserved Component Model

Parameter	Estimate	SE	Estimate	SE
The Phillips Curve Coefficients and AR Coefficients of Measurement Error				
α	0.035	0.187	-0.328	0.191
γ^b	0.686	0.099	0.814	0.081
δ	0.145	0.069	0.186	0.064
γ^f	-	-	0.283	0.090
ϕ_1	1.457	0.123	1.530	0.088
ϕ_2	-0.516	0.123	-0.607	0.082
The Standard Deviations and Correlation of the Shocks				
σ_{π^e}	0.661	0.112	0.665	0.127
σ_v	0.439	0.122	0.428	0.134
σ_π	0.325	0.045	0.352	0.036
$\rho_{\pi^e v}$	-0.861	0.085	-0.932	0.073
Log Likelihood Value				
	-15.0107		-13.841	

The Table reports the estimates from the unobserved component model for the NKPC equation based 4-year spot inflation premium rates. The parameter estimates and standard errors (SE) in columns 2 and 3 are from the restricted model ($\gamma^b+\gamma^f = 1$) and the estimates in columns 4 and 5 are from unrestricted model ($\gamma^b+\gamma^f \neq 1$). Both the estimations allow for the correlation between the two unobserved components of the model namely, inflation expectations and measurement error ($\rho_{\pi^e v} \neq 0$).

Table 2.3: NKPC: Alternative Measures of Inflation Premium Based on Maturity

	4-Year Forward		10-Year Spot		10-Year Forward	
Parameter	Estimate	SE	Estimate	SE	Estimate	SE
The Phillips Curve Coefficients and AR Coefficients of Measurement Error						
α	0.060	0.468	0.306	0.687	-0.153	0.100
γ^b	0.511	0.162	0.557	0.120	0.775	0.083
δ	0.184	0.076	0.154	0.070	0.141	0.039
ϕ_1	1.284	0.130	1.360	0.131	1.722	0.050
ϕ_2	-0.330	0.133	-0.386	0.132	-0.741	0.043
The Standard Deviations and Correlation of the Shocks						
σ_{π^e}	0.633	0.109	0.585	0.106	0.385	0.039
σ_v	0.579	0.124	0.454	0.140	0.083	0.029
σ_π	0.244	0.087	0.280	0.060	0.385	0.035
$\rho_{\pi^e v}$	-0.856	0.059	-0.903	0.044	-0.737	0.156
Log Likelihood Value						
	-3.662		11.897		5.450	

The Table shows the NKPC estimates from the unobserved component model based alternative measures of the inflation premium which includes 4-year froward (columns 2-3), 10-year spot (columns 4-5) and forward rate (columns 6-7). All the three estimations allow for correlation between the unobserved inflation expectations and measurement error.

Table 2.4: NKPC: Alternative Covariance Structures I

Parameter	Estimate	SE	Estimate	SE
The Phillips Curve Coefficients and AR Coefficients of Measurement Error				
α	-0.018	0.194	0.002	0.045
γ^b	0.796	0.074	0.942	0.038
δ	0.088	0.042	0.146	0.036
ϕ_1	1.507	0.297	0.901	0.134
ϕ_2	-0.521	0.292	-0.203	0.060
The Standard Deviations and Correlation of the Shocks				
σ_{π^e}	0.429	0.042	0.225	0.036
σ_v	0.118	0.075	0.344	0.031
σ_π	0.387	0.033	0.427	0.035
$\rho_{\pi^e v}$				
$\rho_{\pi^e \pi}$			0.621	0.159
$\rho_{\pi v}$	-0.436	0.533		
Log Likelihood Value				
	-17.201		-12.851	

The Table reports NKPC estimates by placing the zero restrictions on two out of the three shocks of the unobserved component model. In all the estimation 4-year spot inflation premium is used as a proxy of inflation expectations. Columns 2 and 3 is estimated assuming non-zero correlation between actual inflation and the unobserved inflation expectations (i.e. $\rho_{\pi^e v} = 0$ and $\rho_{\pi^e \pi} = 0$). The last two columns is estimated assuming that there is a non-zero correlation between actual inflation and the unobserved measurement error (i.e. $\rho_{\pi^e v} = 0$ and $\rho_{\pi v} = 0$).

Table 2.5: NKPC: Alternative Covariance Structures II

Parameter	Estimate	SE	Estimate	SE	Estimate	SE
The Phillips Curve Coefficients and AR Coefficients of Measurement Error						
α	0.035	0.177	0.018	0.143	-0.061	0.052
γ^b	0.694	0.098	0.735	0.133	0.933	0.040
γ	0.155	0.073	0.164	0.070	0.112	0.037
ρ_1	1.464	0.116	1.464	0.106	0.116	0.315
ρ_2	-0.523	0.115	-0.526	0.107	-0.003	0.025
The Standard Deviations and Correlation of the Shocks						
σ_{π^e}	0.678	0.128	0.694	0.145	0.422	0.041
σ_v	0.455	0.136	0.474	0.157	0.125	0.068
σ_π	0.325	0.043	0.335	0.049	0.428	0.034
$\rho_{\pi^e v}$	-0.878	0.092	-0.893	0.091		
$\rho_{\pi\pi^e}$			0.061	0.108	-0.085	0.191
$\rho_{\pi v}$	0.043	0.142			1.000	0.003
Log Likelihood Value						
	-14.968		-14.880		-16.9182	

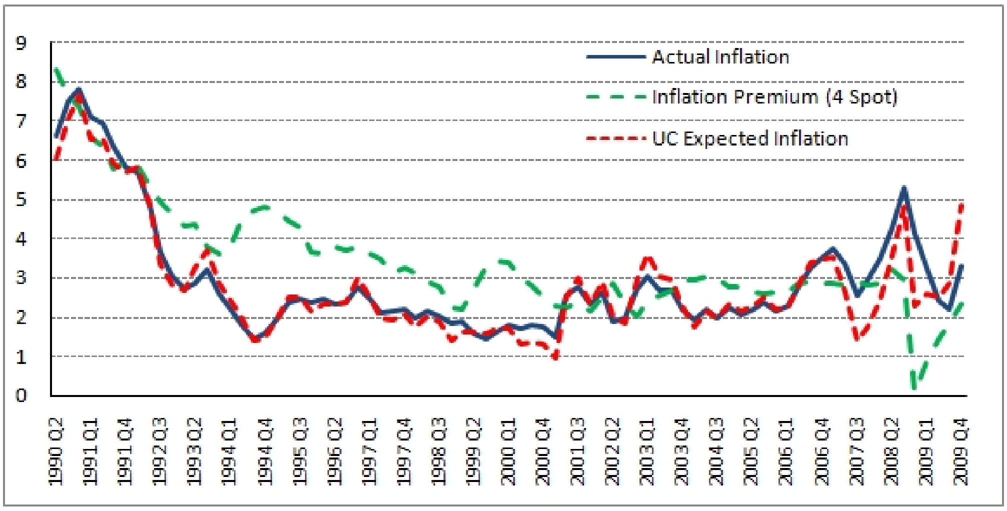
The Table reports NKPC estimates by placing the zero restrictions on only one of the three shocks of the unobserved component model. In all the estimation 4-year spot inflation premium is used as a proxy of inflation expectations. The shocks to actual inflation ($\epsilon_{\pi,t}$) and unobserved inflation expectations ($\epsilon_{\pi^e,t}$) are assumed to orthogonal in Columns 2 and 3 (i.e. $\rho_{\pi^e\pi} = 0$). The shocks to actual inflation ($\epsilon_{\pi,t}$) and unobserved measurement error ($\epsilon_{v,t}$) are assumed to orthogonal in Columns 4-5 (i.e. $\rho_{\pi v} = 0$). The shocks to the unobserved inflation expectations ($\epsilon_{\pi^e,t}$) and measurement error ($\epsilon_{v,t}$) is orthogonal in the last two columns (i.e. $\rho_{\pi^e v} = 0$).

Table 2.6: Explaining the Actual Inflation Using the Two Alternative Proxies of Inflation Expectations

Parameter	Estimate	SE	Estimate	SE
Constant	0.015	0.084	0.042	0.133
Lagged Inflation	0.475	0.052	0.860	0.043
UC Expected Inflation	0.535	0.054		
Inflation Premium			0.098	0.046
R-Sq	0.97		0.91	
Obs	79		79	

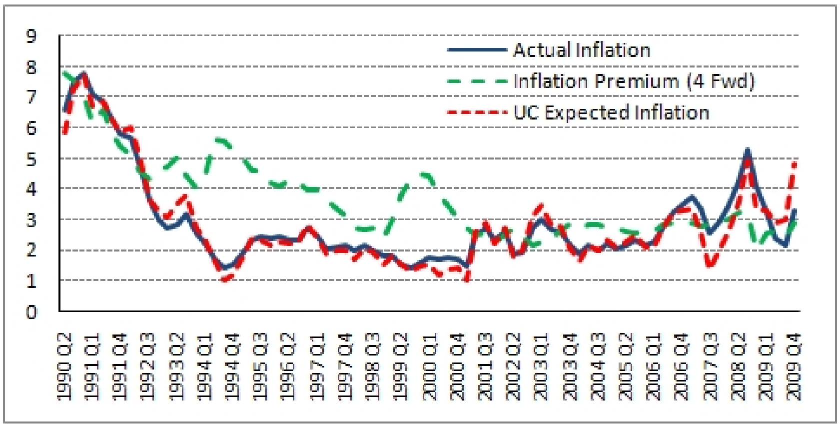
The Table shows the relative superiority of the inflation expectation from the estimated unobserved component model and inflation expectations as measured by the inflation premium from the inflation-indexed bond market. Particularly, in the regression estimation actual inflation is the dependent variable and the list of regressors include a constant, lagged inflation and expected inflation which comes from unobserved component (columns 2-3) and 4-year spot inflation premium rate (columns 4-5). The standard errors (SE) are Newey-West HAC Standard Errors and Covariance adjusted.

Figure 2.1: Comparison of Actual Inflation, 4-Year Spot Inflation Premium Rate and Expected Inflation from Benchmark Unobserved Component Model

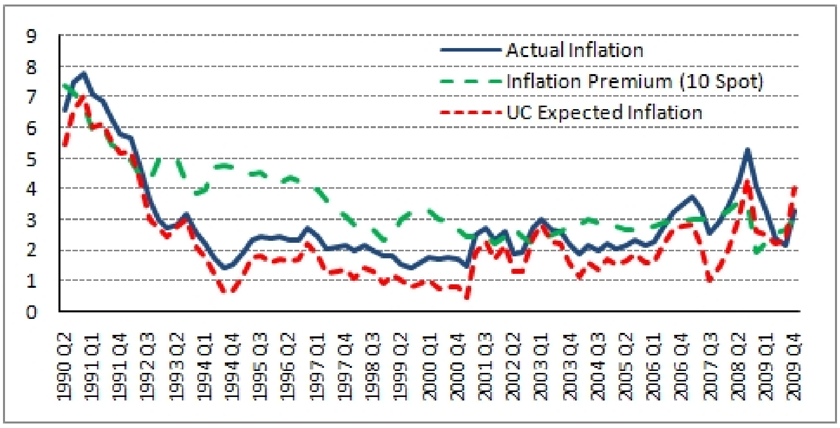


(a) Inflation Premium Based on 4-Year Spot Rate

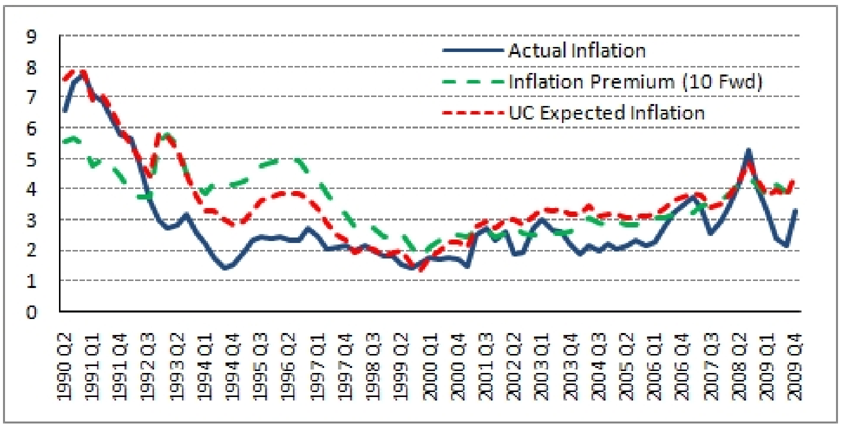
Figure 2.2: Comparison of Actual Inflation, Inflation Premium across Different Maturities and Expected Inflation from Unobserved Component Model



(a) Inflation Premium Based on 4-Year Forward Rate

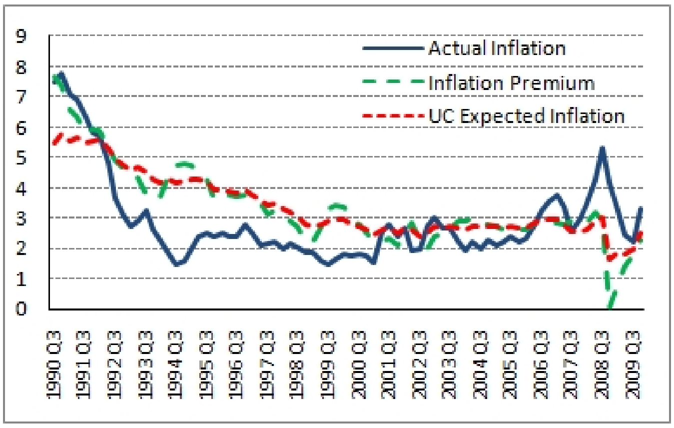


(b) Inflation Premium Based on 10-Year Spot Rate

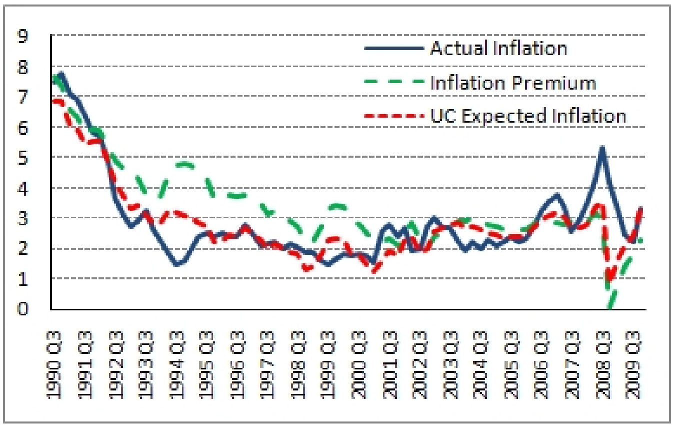


(c) Inflation Premium Based on 10-Year Forward Rate

Figure 2.3: Alternative Covariance Structures I: Comparison of Actual Inflation, 4-Year Spot Inflation Premium and Expected Inflation from Unobserved Component Model

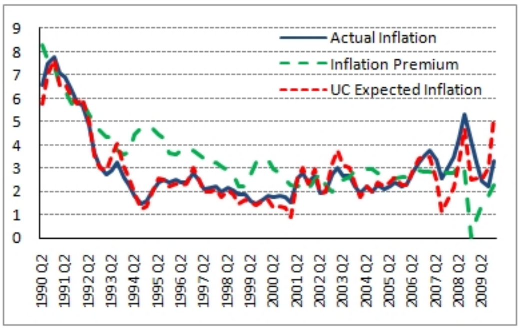


(a) Correlation Between Actual Inflation and Inflation Expectations.

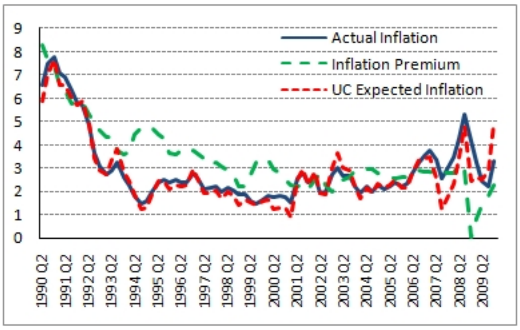


(b) Correlation Between Actual Inflation Expectations and Measurement.

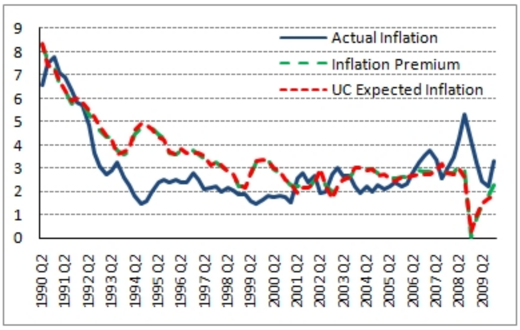
Figure 2.4: Alternative Covariance Structures II: Comparison of Actual Inflation, 4-Year Spot Inflation Premium and Expected Inflation from Unobserved Component Model



(a) Correlation Between Inflation Expectations and Measurement Error and Also Between Inflation Expectations and Actual Inflation.



(b) Correlation Between Inflation Expectations and Measurement Error and Also Between Actual Inflation and Measurement Error.



(c) Correlation Between Actual Inflation and Inflation Expectations and Also Between Actual Inflation and Measurement Error.

CHAPTER 3

The Time-Varying Response of Foreign Stock Markets to U.S. Monetary Policy Surprises: Evidence from the Federal Funds Futures Market

3.1 Introduction

The impact of U.S. monetary policy shocks on asset prices, especially equity prices, has been one of the topics of great interest to policymakers and market participants. It enables policymakers to assess the wealth channel in the monetary policy transmission process, and it affects the portfolio valuations of the market participants. Different methods have been proposed to estimate monetary policy shocks in the literature. Among these methodologies, monetary policy shocks based on the Federal funds futures rate have received widespread attention recently.[1] Using this approach, Bernanke and Kuttner (2005) provide evidence that a typical unanticipated Fed rate cut of 25 basis points leads to roughly 1 percent increase in the S&P 500 index returns. According to them, such a policy action elicits a positive response because it favorably impacts the future dividend streams, re-

[1] See Bernanke and Kuttner (2005); Kuttner (2001) among others.

duces the discount rate and increases the equity market premium. This line of study has been extended to foreign stock markets, as globalization and technological revolution have made the global markets much more tightly interlinked with each other. In the international context, Ehrmann and Fratzscher (2009); Hausman and Wongswan (2011) find that foreign equity returns respond positively to an unanticipated interest rate cut by the Fed. They attribute the cross country variation in responses to the level of financial market integration and the degree of exchange rate flexibility of the country.[2]

Most of the existing studies on the high frequency response of stock returns to monetary policy shocks however, use a fixed-coefficient approach. The underlying assumption is that the response of stock returns to monetary policy shocks remain unchanged over time. This is in contrast to the anecdotal and formal evidence that suggest that the response of stock returns varies over time. For example, Andersen et al. (2007) find that the equity market's response to macroeconomic news depends on the stage of the business cycle. Similarly, equity risk premia, which explain the response of the stock returns to monetary policy surprises according to Bernanke and Kuttner (2005) are found in the literature to vary over time. Campbell et al. (1997) also find that equity market premia vary over time and this variation is quite large relative to the variation in expected real interest rate. Similarly, Bekaert and Harvey (1995) have found that a number of emerging markets exhibit time-varying global market integration. All these factors sug-

[2]Ehrmann and Fratzscher (2009); Wongswan (2009) reach similar conclusions.

gest that modeling the stock market reaction to monetary policy surprise using a fixed-coefficient approach is not appropriate.

The study contributes to the existing literature by taking into account the possible time-variation in the foreign equity market's responses to U.S. monetary policy shocks. This paper is an attempt to examine the empirical regularity in how the foreign stock markets respond over time to U.S. monetary policy surprises that have been derived from the high frequency federal funds futures market. The paper do not test for the explicit linkage between country characteristics and its response over time to U.S. monetary policy surprise. While this may be an important issue to explore, this is beyond the scope of this paper. Following the pioneering work of Cooley and Prescott (1976), the time-variation is modeled as a driftless random walk, and is estimated using the maximum likelihood via the Kalman filter. The argument is that this is an appealing and flexible way of uncovering changes in the responsiveness of stock returns to monetary policy shocks. The framework adopted also allows for heteroscedasticity present in the error term that is typical in a high frequency stock returns data. To do so, we follow Harvey et al. (1992) and allow conditional heteroskedasticity in the error term.

The results show significant time-variation in the stock market response to U.S. monetary policy surprises for all 35 countries under this study. The study finds substantial comovement in the response of the European equity markets to U.S. monetary policy surprises for the whole sample period. This is not surprising since there is a high degree of financial integration among the European coun-

tries. The response of the stock markets of emerging market economies and Latin American economies to U.S. monetary policy shocks is found to be much more divergent as compared to the European stock markets. The results also suggest that the emerging markets' stock markets are more sensitive to the Fed's surprises, especially during the recessions and the crisis periods.

Evidence also suggest a noticeable common feature in the stock markets' reaction to monetary policy surprise. During abnormal periods that include recessions and different crisis episodes[3], U.S. monetary policy surprise has a much larger impact on the equity markets of almost all the 35 countries in the sample. For example, a hypothetical 25 basis points unexpected rate cut by the Fed during a crisis period could elicit a positive response in the stock returns up to as high as 2.5 percent in case of the U.S., Canada and European countries, 7.5 percent in case of Korea, and 5 percent in case of Hong Kong, Singapore and the Latin American economies. This is not a surprising result since the stock market's response to macroeconomic news have been found to depend upon different states of the business cycles according to Andersen et al. (2007). One of the interesting findings of the study relates to the response of the stock markets to monetary policy surprise during the 2008 financial crisis. The results suggest that the surprise cut in the interest rate by the Fed during the 2008 crisis was followed by a decline of the stock markets in the U.S. and Europe. One possible explanation of this puzzling behavior of the stock market is that the market viewed the succes-

[3]In the sample, there are three abnormal events: the LTCM crisis in 1997, 2001 recession and the financial crisis of 2008.

sive and aggressive rate cuts by the Fed as a sign of an increasingly deteriorating economic conditions.

The remainder of the paper is organized as follows. Section 2 provides brief literature review; Section 3 describes the data and methodology adopted in the study. The baselines estimates are presented in Section 4. Section 5 reports the results of the TVP-GARCH model. Section 6 presents robustness results and section 7 concludes.

3.2 Brief Literature Review

Addressing the endogeneity problem and choosing an appropriate proxy for the identification of monetary policy shocks have been the two main challenging tasks in the study of the impact of monetary policy shocks on stock returns. In order to control for the influence of other variables that affect the stock markets, researchers have often used the event study methodology in the analysis of the impact of monetary policy (Bernanke and Kuttner 2005; Bomfim 2003; Cook and Hahn 1989; Gürkaynak et al. 2005). Under this methodology, by looking at a narrow window around the event in question and combining it with high frequency data, one can greatly reduce the problem of endogeneity, if not completely eliminate it.

The identification of monetary policy shocks has generated widespread interest in macroeconomics. In order to measure the policy shock one needs to capture the market expectation. Several methods have been utilized by researchers to mea-

sure unexpected changes in monetary policy.[4] One relatively recent and popular method of estimating monetary policy shocks uses information from the Federal funds futures market. This has been used by Krueger and Kuttner (1996) among others, who find that the Fed funds futures rate is an unbiased predictor of the Fed funds rate and is an 'efficient' measure of Fed funds rate. Gürkaynak et al. (2007) also show the superiority of the Fed funds futures price among different market based measures of monetary policy expectations. Given the superiority of the Fed funds futures data to measure policy expectations, Bernanke and Kuttner (2005) find that a typical unanticipated 25 basis point rate cut has been associated with a 1.3 percent increase in the S&P 500 index.[5] They attribute such negative equity returns response more to the change in expected future excess returns (the equity premium) and less due to revision of the expectations of discounted future dividends streams and the change in real interest rate.

This line of study has also been extended to foreign stock markets. In the context of the U.K. equity markets, Bredin et al. (2007b) follow Bernanke and Kuttner (2005) and use Sterling LIBOR futures contract as a proxy for the U.K. monetary policy shocks. They find considerably smaller response of the FTSE stock index to the U.K. monetary policy shocks than the one found in the U.S. markets. Like the U.S., the equity premium is the key driving factor explaining the U.K. stock returns, but only in case of traditional sectors, like auto parts,

[4]See Cochrane and Piazzesi (2002); Ellingsen and Soderstrom (2001); Faust et al. (2004); Poole and Rasche (2000); Rigobon and Sack (2004) among others.

[5]Other studies that have looked further into the response of individual sector stock returns and the impact on volatility include Bredin et al. (2007a) for the impact on REIT industry and Chuliá et al. (2007) that covers the S&P 500 index constituents.

chemicals, oil and gas and steel but not at an aggregate level.

There are significant number of studies that focus on the impact of the U.S. monetary policy shocks on global markets.[6] For example, Wongswan (2009) has found that a hypothetical 25 basis points Fed rate cut elicits a response of 0.5 to 2.5 percent increase in foreign equity price indexes, whereas Ehrmann and Fratzscher (2009) found the impact to be around 0.67 percent. Hausman and Wongswan (2011) found that similar cut leads to an increase of about 1 percent in foreign stock markets. Among the various economies of the world, the equity markets of Canada, Hong Kong, Korea, Singapore and Mexico are found to be the most sensitive to the U.S. monetary policy rate surprise. Kim and Nguyen (2009) also analyze the spillover effects of interest rate news from the Fed and the ECB on the Asia-Pacific stock markets. They find the speed of adjustment for the Fed's news to be mixed across these markets but the ECB news was absorbed slowly in general.

In terms of the factors leading to the response of foreign stock markets to monetary policy shocks, Wongswan (2009) attributes the variation in the response to be more a function of degree of financial integration with the U.S. than the proxies for real integration or exchange rate flexibility. However, Ehrmann and Fratzscher (2009) attribute the cross country variation in response to the degree of global integration of the country - and not the country's bilateral integration with the U.S. Also, equity indexes in countries with a less flexible exchange rate regime

[6]Studies that focus on the impact of other macroeconomic news (other than monetary policy news) on the stock markets include Flannery and Protopapadakis (2002).

respond more to the U.S. monetary policy surprises (Hausman and Wongswan 2011).

While all these studies adopt the fixed-coefficient approach in assessing the impact of the U.S. monetary policy on stock returns, in this study the responses are allowed to vary across time. This is important since most of the factors that explain such stock market response are found to vary with time. Thus, a TVP model would allow us to uncover richer dynamics of the relationship between the U.S. monetary policy and the foreign stock markets.

3.3 Data and Methodology

3.3.1 Data

The sample period for the study uses daily data from May 17, 1994 through June 25, 2008. The study examines all the Federal Open Market Committee (FOMC) meetings for the whole sample period. The extension of the sample period to 2008 is also one of the contributions of the study. Since the FOMC meets eight times a year (approximately every six weeks), the sample contains 114 scheduled meetings decisions and nine inter-meeting decisions. The 50 bps rate cut announcement made on September 17, 2001 meeting is excluded since it is an instance when several central banks across the globe (including the Fed) and the financial markets responded jointly to September 11, 2001 terrorist attack.

The high-frequency daily data used in the study includes the price of the

Federal funds futures interest rates and the closing stock price index of leading economies of the world. The first data set consists of the 30-day Federal funds futures market contracts that most closely track the effective overnight Federal funds rate for a specific month.[7] The Chicago Board of Trade (CBOT) has been offering Federal funds futures contracts since October 1988 for several different deliveries going from the current month to five months ahead. Even while contracts with longer deliveries exist, liquidity in those contracts is significantly lower. These contracts not only allow the market participants to hedge interest rate risk, but also serves the important role of revealing the expectations of the market about the future monetary policy actions.

In the present study the most liquid spot months contract is used for the purpose of extracting out the monetary policy surprises at each FOMC meeting. Among the variety of market-based measures of monetary policy expectations the Federal funds futures rate is found to dominate all other instruments in predicting the future path of monetary policy at horizons out over several months (Gürkaynak et al. 2007).[8]

The second data set contains the daily data on the closing stock price index of the U.S. and thirty-five other leading stock markets of the world, which includes fifteen European economies, thirteen economies from the Asia-Pacific region, four

[7]Federal funds futures prices are expressed as 100 minus the expected average effective federal funds rate for the delivery month. For example, if a January contract has a price of 92.75, it reflects an anticipated average rate of 7.25 percent for that month.

[8]Alternative measures used includes term Eurodollar interest rate (Cochrane and Piazzesi 2002), the Eurodollar futures interest rate (Rigobon and Sack 2004), the three-month Sterling LIBOR futures rate in case of the U.K. monetary policy surprise (Bredin et al. 2007b)

Latin American economies, two Middle Eastern economies and Canada, the details of which are presented in Table 1.[9] The daily stock returns are then calculated from this data set.

Table 1 reports the descriptive statistics of each country's stock market returns that are realized on the FOMC meeting dates. The average stock return during the FOMC meeting dates in most of the Asian countries is higher than the U.S. and the European stock markets. Further, the U.S. monetary policy surprise leads to a much greater degree of variation in the stock market movements in the Asia-Pacific region as compared to the European economies and the U.S. The standard deviation of stock returns in Asian economies is considerably higher than most of the European countries. For example, the standard deviation of the stock markets of Hong Kong, Korea, Indonesia and Philippines is nearly twice that of the U.S. and the U.K. Overall, non-OECD economies' stock markets exhibit higher volatility on the FOMC announcement dates.

3.3.2 Methodology

This section briefly discusses the methodology to calculate monetary policy shocks from the Federal funds futures market, and also outline the event study approach that shows how to estimate the impact of monetary policy shocks on stock returns.

[9]Entire data for all the countries stock price index is obtained from publicly available source, yahoo finance and the Federal funds futures contract rate has been provided by Kenneth N. Kuttner.

Measurement of Monetary Policy Surprise

As stated before, the Federal funds futures price is used as a natural market based proxy for the otherwise unobserved market expectations. The main premise of using the future price is that all the future expectations about interest rates would be embedded in the futures price today and any change in the futures rate post-FOMC meeting is because the announced rate change (or no change) was unexpected before the FOMC meeting.[10] The key advantage of such a measure to gauge policy surprise is that it is free of model selection and the 'generated regressors' problems. In order to compute the unexpected changes in monetary policy, the methodology proposed by Kuttner (2001) is adopted, according to which the one day surprise is computed as:

$$\Delta r_{\tau}^{u} = \frac{m_s}{m_s - \tau}(f_{s,\tau}^{0} - f_{s,\tau-1}^{0}) \tag{3.1}$$

where, Δr_{τ}^{u} is the monetary policy surprise, m_s is the number of days in month s, $f_{s,\tau}^{0}$ is the current-month futures rate on day τ of the month s and $f_{s,\tau-1}^{0}$ is the current-month futures rate on day $\tau - 1$.[11] This holds true for all the days in the

[10]Though there is body of evidence in the finance literature (e.g. Campbell and Shiller 1991; Cochrane and Piazzesi 2005; Fama and Bliss 1987 that shows evidence against expectations hypothesis and thus the possible presence of 'risk premia' (i.e. the predictable returns in excess of the risk-free rate), Piazzesi and Swanson (2008) find no such evidence in case of current months Fed futures price that is used in this study, though the risk premia exists at 3 months to one year horizon.

[11]As suggested by Kuttner (2001), the adjustments as in (3.1) are necessary to address two main challenges. First, since the Fed funds futures settlement price is based on the average of the relevant month's effective overnight Fed funds rate and not based on the rate on any specific day of the month the time averaging must somehow be undone. The second complication arises, especially in case of daily frequency, is that the futures rate is based on the effective Fed funds rate and not on the target Fed funds rate.

month except for the first and the last day. When the change comes on the first day of the month, its expectations would have been reflected in the prior months spot rate, so the one-month futures rate on the last day of the previous month, $f^1_{s-1,\tau-1}$ is used instead of $f^0_{s,\tau-1}$. Similarly, when the change comes on the last day of the month, the difference in the one-month futures rate is used. This is because when the target rate change comes on the last day of the month it would have no impact on the spot month's rate, as the market Fed funds rate does not change until the day following the target change. Furthermore in order to avoid amplifying the month-end noise, no adjustments in terms of scaling is made when the announcement is made within the last three days of the month.

This method enables one to purge out the unexpected changes from the rate change decisions. Only the unexpected changes is considered in this study, since the expected changes have been found to have no statistical impact on the financial markets.[12] This is consistent under the *efficient market hypothesis* under which the security price incorporates all information that is available at any point in time. Hence, in order to evaluate the impact of the monetary policy only the surprise changes in policy matters.

The issue of the timing of the FOMC announcements is particularly relevant in the high frequency data analysis. Usually the FOMC announcements are made at 2:15 p.m. Eastern time and since the Fed funds futures market closes at 3:00 p.m. Eastern time, the closing futures prices would have typically incorporated the news

[12]Bernanke and Kuttner (2005); Hausman and Wongswan (2011); Kuttner (2001).

of the FOMC's decision. In these instances the target rate changes are assigned to the dates of the announcements. However, in instances where the announcement came only after 3:00 p.m. Eastern time, the policy surprise is measured as the difference between the closing futures rate on the announcement date and the opening rate on the following date. This is consistent with Bernanke and Kuttner (2005); Kuttner (2001)

Event Study Approach

Researchers have often adopted the event study methodology using high frequency data in order to control for the endogeneity problem and also control for the possible joint response of the policy and the stock markets to new information (Bernanke and Kuttner 2005. Following the popular conventional Cook and Hahn (1989) style analysis, the study analyzes the impact of the policy surprises on the equity markets across the world. A brief outline of the *event study* performed in the current study is presented below.

- *Event definition*: The event of interest in this paper is the FOMC meetings, and the period over which the stock returns are examined is defined as an *event window*. Following Bernanke and Kuttner (2005) and Hausman and Wongswan (2011), one day is selected as the size of the event window.[13]

- *Normal and abnormal returns*: To examine the impact of the event, abnor-

[13] However qualitatively, just like Hausman and Wongswan (2011), the results too remain robust when we re-estimate for the sample period covered in Wongswan (2009) study which uses a much shorter window size than one day that is used in the present study.

mal returns are measured as the actual *ex post* return *(event returns)* of the stock market over the event window over and above the normal return i.e. the returns that would have been expected if the event did not take place. Adopting a variation of *constant-mean-return model*,[14] the normal returns are calculated based on the previous twenty one trading day's average daily return, which is roughly equivalent to one-month average before the event (FOMC meeting). Appropriate adjustments are made to the *normal returns* in case the meeting dates fall within the span of 21 trading days of each other. This is especially true in case of the inter-meeting moves.

Calculating the stock market returns in this manner is novel as compared to the previous studies.[15] This approach of measuring abnormal returns allows us to model the country's own stock market dynamics in normal times. One can argue that the methodology implicitly assumes that the FOMC announcements are exogenous to the changes in stock markets. However, evidence in Bernanke and Gertler (1999) suggests that the Fed's reaction has generally failed to find a direct response of monetary policy to stock market fluctuations. If this holds true for US market then it is not difficult to accept the same for stock markets of other countries. Thus, for example, if the average return of the U.K. equity markets in June is 2 percent, and the June-end FOMC announcement of an unexpected rate

[14]Under the constant mean model, the normal return is simply a fixed average number. However, it would be unreasonable to assume that the average returns for 14 years for any country is fixed. Hence, the average return is allowed to vary with time. Further even while the period of one month may seem arbitrary, it is reasonably long and has been chosen keeping the frequency of the FOMC meetings in mind.

[15]For the sake of simplicity the term *stock returns* in rest of the paper would mean these *abnormal returns.*

cut leads to an increase in the stock returns by 3 percent, then it is reasonable to conclude that the impact attributable just to this rate cut decision is 1 percent. The underlying assumption is that 2 percent returns in the U.K.'s stock market would have taken place even in the absence of this event.

Due to the difference in the time zones for various countries, getting the dates right for the event impact is also very important. The study follows Hausman and Wongswan (2011) in order to pin down the relevant closing stock index price for the country. In case of European countries, Asia Pacific countries and Middle Eastern countries the stock returns of the following day of the FOMC announcement is taken. Since these markets are closed at the time of scheduled FOMC announcements, the impact of FOMC announcements on these markets occurs only on the following day.[16] For Latin American equity markets, the event impact is calculated as the two average returns over the FOMC meeting date and the following date. This ensures evenness of comparison and also because these markets are relatively slower to incorporate the arrival of new information. However, the relevant closing stock index price in case of the U.S. and Canada is the same day as the FOMC announcement since these markets are not only relatively more advanced but also fall in nearly the same time zone.

[16]In case of FOMC announcement that occur before 1 p.m. eastern time, we take the same day returns for the European countries as these equity markets are still open at the time of FOMC announcement.

3.4 The Response of Foreign Stock Markets to U.S. Monetary Policy Surprise

3.4.1 Baseline Estimates

In this section, the estimated response of foreign stock markets to U.S. monetary policy shocks is discussed. To do so, monetary policy surprise is estimated using Kuttner's (2001) approach, as outlined in the previous section. Figure 1 shows the unexpected changes in monetary policy actions computed from the spot month Federal funds futures rate. The estimates of surprise measure in this paper matches closely with Kuttner (2001) study that uses data up to 2000.[17] Using the same methodology, the analysis is extended to include seven extra years of data. The sample period in this study ends on June 25, 2008.

After extracting the surprise element from the policy action the next step is to estimate the impact of the U.S. monetary policy surprise on the abnormal returns of country's stock price index. Following Bernanke and Kuttner (2005), and use the regression specification in equation 3.2 to estimate the impact of U.S. monetary policy surprise on the foreign stock returns.

$$R_t^i = \alpha^i + \beta^i \Delta r_t^u + e_t^i \tag{3.2}$$

where, R_t^i represents the abnormal return of country i at the event date t as defined earlier and Δr_τ^u is monetary policy surprise. The coefficient β^i shows how much

[17] Any deviation may possibly be due to the rounding off errors.

the abnormal return of a country i's stock market responds to 100 basis point unanticipated interest rate increase in the U.S. Given the exceptional nature of recent financial crisis, the above regression is estimated for both the full sample period from 1994-2008 and also for the 1994 - 2006 sample period. Excluding the period of recent financial crisis is especially important since the impact of conventional monetary policy measures is severely limited when the short-term interest rate hits the zero lower bound.

The estimated results are shown in Table 3.2. Based on the results from the full sample, except for the European equity markets, an unexpected rate cut by the Fed boosts the stock market indexes across the global economy. A hypothetical surprise rate cut (negative surprise) of 25 basis points by the Fed triggers a jump in Hang Seng by 2.5 percent, Jakarta Composite index by 1.7 percent, Bombay Stock Exchange index by 1.5 percent, Toronto stock exchange index by 1.5 percent and both Argentinian and Mexican stock market index by 1.3 percent. Interestingly, these are even higher than the 1 percent response of the S&P composite index itself. Furthermore, if the recent crisis period is excluded, then the results become even stronger and shows the strong impact of the U.S. monetary policy shocks during the normal times. The full sample estimates are qualitatively similar to what other researchers have found in the literature. The estimates are slightly different in magnitude because of the longer sample size and also the dependent variable is abnormal returns in this study, as compared to normal returns in the existing literature.

The results presented above assume that the response of the global stock returns to the U.S. monetary policy shock has remained constant over time, but there is a strong anecdotal and formal evidence which suggest that the response of stock returns vary over time. To capture these time-varying effects, an ad-hoc approach is adopted first by performing the split-sample estimation. This helps us in finding out whether preliminary evidence exists for time-variation in response of stock returns to monetary policy shocks. The whole sample is divided into three sub-samples: 1994-2001, 2001-2006, and 2006-2008. These results are further supported by Quandt-Andrews breakpoint test for one or more unknown structural breakpoints in the whole sample.[18]

The results shown in Table 3.3 suggest that the estimates of response coefficients are highly sensitive to the sample period chosen. Even though the sample-period split is arbitrary, it clearly shows that the estimates are highly sensitive to the choice of the sample period.[19] Even the results from Quandt-Andrews parameter stability test presented in Table 3.3 are largely consistent with these split sample results. The test reveals strong evidence for parametric instability since the null of no structural change is rejected for most of the countries in the study. During the 1994-2001 period, according to the estimates the stock markets across the world were much more sensitive to U.S. monetary policy shocks with an important exception of Latin American stock markets. The impact of U.S.

[18]A standard trimming level of 15% is used in the test for almost all the countries.

[19]Even when we look at just the event returns, as most of the earlier studies, there is still an evidence of sample selectivity issue. This is not presented in the study to maintain brevity. Furthermore, we do not suggest that these are the precise dates that are associated with the changes in the response of the foreign equity returns to the U.S. monetary policy shocks.

monetary policy shocks was significant for all the Latin American stock markets during 2001-2006 period. In case of the U.S. and Canada the R-squared rises from 0.20 and 0.41 to 0.48 and 0.46 from 1994-2001 to 2001-2006 period respectively. This is not case with European and Asia-Pacific stock markets. Interestingly the 2006-2008 sub-sample which nearly overlaps with the recent global financial crisis the Fed's policy surprise led to an unusually positive response of the stock markets Europe but usual negative response of the stock markets in rest of the world.[20]

The preliminary evidence presented above suggests that there is significant time-variation in the response of global stock market returns to monetary policy shocks. Therefore, it is more appropriate to adopt a model that is able to capture the dynamic response of the stock market returns over time.

3.5 The Time-Variation in the Response of Global Stock Markets

3.5.1 Model Specification

Since the fixed-coefficient estimates fail to capture the dynamics of the equity market reaction to the U.S. monetary policy shocks, the study adopts a time-varying parameter (TVP) framework to model the changes in response of stock returns to U.S. monetary policy surprises. This allows us to model the changing responses of

[20] However, one has to be cognizant of relatively small sample size and the fact that in this period the central banks across the world took extraordinary measures to combat the crisis. Also, in this period some of the monetary policy surprises coincided with bad economic news leading to the perverse response of the American and European equity markets. For example, the collapse of Bear Stearns in March 2008 and the associated 75 basis point rate cut.

world equity markets across time to the Fed's policy actions. There are alternative approaches of modeling time-variation that includes structural break, as well as Markov switching in the response coefficients. Usual test of time-variation has a low power against the alternative, that is, it is difficult to distinguish between different forms of time-variation. As in Boivin (2006), it is important to note that structural break models are very special cases of time-variation and do not allow for the gradual evolution of how stock markets respond to monetary policy shocks. Moreover, time-varying parameter model may also be used as a good approximation of multiple breaks in the response coefficients.[21] Thus, for each country i the following state-space TVP model is considered:

$$R_t^i = \alpha_t^i + \beta_t^i \Delta r_t^u + e_t^i \tag{3.3}$$

The coefficient β_t^i measures the time-varying response of the foreign stock market returns to U.S. monetary policy surprise. Based on the fixed-coefficient model presented before and the findings in the existing literature, this coefficient is expected to be negative implying that an unexpected increase in the interest rate by the Fed would be associated with a drop in stock market returns.[22] To allow time-variation, β_t^i is assumed to follows a random walk.

$$\beta_t^i = \beta_{t-1}^i + v_t^i \quad where, v_t^i \sim N(0, Q) \tag{3.4}$$

[21]Stock and Watson (2002) and Boivin (2006) discuss merits of the TVP model over other forms of structural break.

[22]Bernanke and Kuttner (2005); Ehrmann and Fratzscher (2009); Hausman and Wongswan (2011); Wongswan (2009).

Furthermore, equity market movements have often witnessed to possess the feature of volatility clustering, therefore to model volatility we follow Sims (1999) and Sims and Zha (2002) and allow for heteroskedasticity in the disturbance term e_t. To model heteroskedasticity, the error term is allowed to follow ARCH(p) or GARCH (1,1) process.[23] The baseline model is GARCH (1,1) and is represented as:

$$e_t^i|\psi_{t-1} \sim N(0, h_t^i) \tag{3.5}$$

$$h_t^i = \phi_0^i + \phi_1^i(e_{t-1}^i)^2 + \phi_2^i h_{t-1}^i \tag{3.6}$$

The study also explicitly take into account the possibility of spurious inference in the GARCH (1,1) model as highlighted in Ma et al. (2007). Using Monte Carlo experiments Ma et al. (2007) show that when the ARCH parameter is small or insignificant the Zero-Information-Limit Condition formulated by Nelson and Startz (2007) holds and the GARCH(1,1) model is weakly identified. This leads to spurious GARCH effect that is possibly strong when in fact it is absent. Hence, to avoid spurious inference, the empirical strategy recommended by Ma et al. (2007) is followed and estimate an ARCH(p) model instead of the GARCH (1,1) model for the countries were the ARCH parameter (ϕ_1^i) is insignificant .

In order to predict and update the parameter estimates through the Kalman filter process, we need to calculate equation (6), which itself is a function of the past unobserved shocks and the square of the past unobserved shocks. Hence, it

[23] See Harvey et al. (1992) and Kim and Nelson (1998, 2006).

becomes necessary to modify the model before applying the Kalman filter estimation procedure. A relatively straightforward modification of the model involves augmenting the error term into the state vector of the measurement equation as stated below:

$$R_t^i = \begin{bmatrix} 1 & \Delta r_t^u & 1 \end{bmatrix} \begin{bmatrix} \alpha_t^i \\ \beta_t^i \\ e_t^i \end{bmatrix} \tag{3.7}$$

$$(R_t = X_t^* \beta_t^*)$$

$$\begin{bmatrix} \alpha_t^i \\ \beta_t^i \\ e_t^i \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \end{bmatrix} \cdot \begin{bmatrix} \alpha_{t-1}^i \\ \beta_{t-1}^i \\ e_{t-1}^i \end{bmatrix} + \begin{bmatrix} \sigma_{0,t}^i \\ \sigma_{1,t}^i \\ e_t^i \end{bmatrix} \tag{3.8}$$

$$\left(\beta_t^* = F^* \beta_{t-1}^* + v_t^* \right)$$

$$\& \; Q^* = E(v_t^* v_t^{*'}) = \begin{bmatrix} Q & 0 \\ 0 & h_t \end{bmatrix} \tag{3.9}$$

3.5.2 Empirical Results

Basic Findings

The state-space model presented in the previous section is estimated by maximum likelihood via the Kalman filter. The filtered time-varying responses of different

countries' stock returns to U.S. monetary policy surprise are shown in Figures 3.3-3.7. The baseline model for heteroskedasticity in the error term is GARCH (1,1), but the study also take into account the spurious inference problem in GARCH model when the the ARCH coefficient is insignificant and zero information limit condition applies, as pointed out by Ma et al. (2007). In cases where Ma et al. (2007) critique applies, the paper estimates a time-varying parameter model with ARCH(p) error terms. In the sample considered, ARCH model is estimated for 22 out of 36 countries.[24] It should also be mentioned however, that time-varying response coefficients are qualitatively similar for both GARCH and ARCH models except for Argentina and India.

The results suggest interesting pattern for different groups of countries. First, the response of the U.S. and the global stock markets to unexpected changes in the Fed's policy actions does in fact vary significantly over time. Second, European equity markets exhibit a very high degree of comovement in their response to the U.S. monetary policy surprise. Third, the impact of U.S. policy surprise on the developing world is relatively more volatile. Fourth, in spite of the difference in the response pattern across countries, important similarities exists during the periods of crisis and between similar economies. Fifth, less surprisingly, the response of the world equity markets to the Fed's policy action in the recent financial turmoil differs significantly for developed markets and the emerging market economies.

[24]In particular, the GARCH model is estimated for Australia, Hong Kong, Malaysia, and all the European stock markets except Greece, Ireland, Italy, Spain and Switzerland. The ARCH model is estimated for stock markets of the U.S., Canada, China, India, Indonesia, Japan, Korea, New Zealand, Pakistan, Philippines, Singapore, Taiwan and Latin American stock markets.

The stock market response during the current period is also significantly different as compared to the earlier periods.

The results suggest significant time-variation in the response of global equity markets to U.S. monetary policy surprises. There are several economic reasons why the response should vary over time. For example, in a fixed-coefficient framework, Bernanke and Kuttner (2005) suggest that the equity market premium and expected future dividends explain why stock market reacts negatively to a positive surprise in interest rates. Equity market premium, however, has been found to vary with time as in Campbell et al. (1997). This suggests that stock market response to monetary policy shocks may also vary with time. The time-variation is also consistent with Andersen et al. (2007) who find that the equity market's response to macroeconomics news depends on the stage of the business cycle. In the international context, over and above such factors, additional channels like exchange rate exists. This channel transmits the impact of the U.S. monetary policy action on the foreign stock prices.[25] Since the response of exchange rate to monetary policy shocks itself evolves over time, it is not surprising to find time-varying response of foreign stock markets to U.S. monetary policy surprise.

For the whole sample period, the impact of the Fed's policy surprise on the U.S. and the foreign stock markets [Figures 3.3-3.7] was maximum during the period of 1998-2001, and minimum during the 2002-2006 period. Further the response pattern varied considerably during 1998-2001, but was far more stable in

[25] See Wongswan (2009) for more details.

the 2002-2006 period. This holds true for almost all the stock markets across the world with few notable exceptions like Indonesia [Figure 5(b)], Philippines [Figure 5(e)], China [Figure 5(h)] and Egypt [Figure 7(a)]. Between 2001-2006, the Fed's policy actions have not only become more predictable, but also the response of stock markets to Fed's policy surprises has declined.

Looking at the TVP estimates, it is found that a hypothetical 25 basis point rate cut during the 1998-2001 period leads to a 2.5 percent jump in S&P 500, and more than 5 percent increase in the stock market indexes of Korea, Hong Kong and Singapore. The Latin American and European stock markets behaved similarly, though the magnitude of response was slightly lower as compared to these economies. On the other hand, during the 2002-2006 period, the response to such a hypothetical rate cut declined to less than 1 percent for most of the foreign stock markets and less than 1.25 percent in case of the S&P 500 stock index returns.

The evidence presented above suggests that the response of foreign stock markets to Fed's policy surprises is time dependent. It would also be interesting to investigate whether these time varying responses depend on the size of the monetary policy intervention. Hence, the correlation between the monetary policy surprise and the TVP estimates of each country is computed. The results suggest that except for the current financial crisis period, the correlation coefficient for most of the stock markets is positive but weak. In general this means that the

impact of monetary policy shock may be related to the size of the shock.[26] However, the correlation gets weakened if the recent financial crisis period is included in the calculations. This is understandable given the *unusual* response of foreign stock markets in the recent crisis period the details of which are discussed in the following section.

The results also show a significant degree of comovement in the responses across different regions.[27] Figures 2(a) & 2(b) suggests a clear pattern in how the European economies' stock markets respond to the U.S. monetary policy surprise. First, the variation across time in the impact of the U.S. policy shock on most of the European markets is very similar, even though the magnitude differs by a small margin. This result is not very surprising since the European financial markets are much more tightly linked than the financial markets in any other region of the world. This is also confirmed by the evidence presented in the Table 3.4. This table shows the correlation between the time-varying response coefficients of different European stock markets. The correlation between the response coefficient of S&P 500 and the FTSE, DAX and CAC is 0.71, 0.68 and 0.68 respectively. Similarly, the correlation for other European economies lie in the range of 0.8-0.9. One possible explanation for the high degree of correlation could be the progressive integration of the international stock markets (Morana and Beltratti 2008).

[26]The possibility of a nonlinearity in the relationship of the stock markets response to some threshold level of monetary policy shocks based on the Hansen (2000) threshold test is also considered. The preliminary evidence for big economies like the U.S., U.K., Canada and Germany among others suggest that threshold effect doesn't exist i.e. Hansen's test doesn't reject the null of no threshold effect. To maintain brevity the results for threshold analysis are not reported.

[27]Kizys and Pierdzioch (2009) study the changing patterns of the international comovement of stock returns and how these are systematically linked to asymmetric macroeconomic shocks.

Similar behavior is found in the responses by stock markets of the Latin American economies. This can be observed by looking at Figures 6(a)-6(d). In case of the Asian markets, it is found that these markets don't fit into one single group in how it responds to U.S. monetary policy shocks [Figures 5(a)-5(m)]. Formal support for this is presented in Table 3.5. The correlation in the response of the stock markets of China, India, Philippines and Indonesia with the response of other stock markets in the region is around 0.2 and even negative in some of the cases.

Overall, the correlation between the response coefficients of different OECD economies is mostly above 0.6 [Table 3.4], while for the non-OECD economies' stock markets it is far below 0.3 and even negative [Table 3.5]. Hence, based on the response pattern [Figures 3.3-3.7] and the evidence presented in Table 3.4 and Table 3.5 we can conclude that the time-varying impact of the U.S. monetary policy shocks on the stock market of OECD economies is more tightly linked to each other. However, this is clearly not the case with the stock markets of the emerging market economies.

As shown in Figures 5(a)-5(m), most of the equity markets in the emerging economies of the Asia-Pacific region like Korea [Figure 5(c)], Hong Kong [Figure 5(a)], Singapore [Figure 5(f)], China [Figure 5(h)], Indonesia [Figure 5(b)] and Latin America [Figures 6(a)-6(d)] respond much more aggressively to the Fed's policy surprise as compared to the U.S. and the European markets. Thus, a 25 bps unexpected rate cut by the Fed could elicit a positive response up to as high

as 2.5 percent in case of the U.S., 7.5 percent in case of Korea, 5 percent in case of Hong Kong, Singapore and the Latin American economies. These responses are significantly higher than the baseline fixed-coefficient point estimates of 0.5 percent for the U.S., 1 percent in case of Korea, 2.5 percent in case of Hong Kong, 1.25 percent in case of Singapore and around 1 percent in case of Latin American stock markets. The differences could possibly be explained by the fact that point estimates are only able to capture the average response for the whole sample period and it masks the significant time-variation that is clearly evident from the TVP model adopted in this study.

These findings highlights the importance of modeling the stock return response in a time-varying parameter setting. The time-varying estimates provide a much richer dynamics of the relationship of the foreign stock markets to the U.S. monetary policy shocks. Further, the U.S. monetary policy plays at least as important role in emerging economies equity market, as it does in the European equity markets. Also, less surprisingly, the impact of the Fed's policy surprise on the European equity markets is very similar.

The Response During the Crisis Periods

There is widespread evidence in the literature that suggests that the behavior of the stock markets during recessions and crisis episodes is very different than the normal times.[28] In addition to confirming these existing findings, the study

[28] Andersen et al. (2007); Bomfim (2003); McQueen and Roley (1993); Schwert (1989) among others.

also report that the usual response pattern was reversed in the 2008 crisis period in case of the OECD economies. The results are shown in Figures 3.3-3.7. The sample period in this study covers three crisis episodes: the LTCM crisis of 1997, the economic downturn of 2001, and the 2008 financial crisis.

The results suggest that during the crisis periods, the foreign equity markets become more sensitive to U.S. monetary policy shocks. This is consistent with the theoretical model of Ribeiro and Veronesi (2002) and the empirical findings of Gulen et al. (2011). According to these studies, the news is more informative about the true state of the economy in contractions, resulting in higher cross-market correlations. This implies that global financial markets become much more sensitive to U.S. monetary policy shocks during economic downturns. To the extent that investors across the globe perceives the Fed's policy actions to reveal important information about the state of the U.S. economic activity and that this news has increased relevance in the periods of contractions, any surprise move by the Fed during such period will mean increased response of the equity markets to FOMC announcements.

Furthermore, even though the behavior of the European stock market was similar to that of other equity markets across the world during the first two crisis episodes, its response during the recent financial crisis was remarkably different as shown in Figures 2(a) & 2(b) and also Figures 4(a)-4(o). While in the first two crisis periods the Fed rate cuts were associated with a rise in the European equities, the sign of equity response in the recent crisis was reversed. This shows that during

the recent financial turmoil the panic among investors was so severe that it greatly reduced the ability of the monetary authorities to lift the stock market sentiments via interest rates cuts. One possible explanation of this response reversal is that this policy shock was perceived by market participants as a signal of a deteriorating economic outlook. This is certainly plausible under the scenario where the Fed has informational advantage over the private market participants (Romer and Romer 2000).

For the Asia-pacific region and Latin America, however, we do not find a reversal in the response of stock markets to U.S. monetary policy surprise during the recent crisis. The results are shown in Figures 5(a)-5(m). The results suggest that the overall response of the stock markets in emerging countries in 2008 was similar to the responses during the LTCM crisis and the 2001 recession. The limited exposure of these economies to 'sophisticated' derivative products and lower losses on these products may have been one of the reasons why these countries did not respond the way the advanced economies did to U.S. monetary policy surprise in 2008. Another possible explanation could arise from the interest rate differential channel and international portfolio adjustments. An increased interest rate differential between the U.S. and the emerging market economies due to such steep rate cuts lead to carry trade, thereby leading to an increased sensitivity of these markets to the U.S. policy actions. It can also be argued that the impact of monetary policy shock on stock returns may depend upon the level of uncertainty in the economy. The different response during the current financial crisis certainly

provide some support to this view. However, it should also be noted that the increase in uncertainty may also arise due to the actions of central banks. The relationship between the effectiveness of monetary policy shock and uncertainty is an interesting topic in itself and it warrants a careful examination.

3.6 Robustness Check

There is widespread evidence that suggests that stock market movements in the U.S. leads the world stock markets, or there is a significant degree of comovement with the U.S. stock markets. (for example, Eun and Shim (1989) for the developed markets and Bekaert and Harvey (1997) for the emerging markets). Hence, in order to check for the robustness of the results, the daily returns of the S&P 500 are included as an additional variable in the baseline time-varying regression model as represented by equation (3). In particular, the following regression is estimated:

$$R_t^i = \alpha_t^i + \beta_t^i \Delta r_t^u + \gamma_t^i R_t^{US} + e_t^i \tag{3.10}$$

where, R_t^{US} is the daily returns on the S&P index. The coefficient γ_t^i measures the extent of the influence of the U.S. stock market movements on the countries i stock returns. This coefficient measuring the comovement of countries stock returns with the U.S. stock returns is also allowed to vary over time.

It is evident from the plots in Figures 3.3-3.7 that the earlier results from the TVP model remain robust even after controlling for the S&P returns for

almost every country under this study. This implies that the Fed's surprises have a direct impact on the world equity markets apart from its indirect influence via the U.S. equity market movements, and that these responses vary over time. Not surprisingly, the magnitude of the response shrinks slightly for most of the countries. This is consistent with the literature that suggests that the movement in the U.S. markets does influence the world equities through varying extent even apart from the impact of the Fed's policy actions.

The results are consistent with the fixed-coefficient framework of Hausman and Wongswan (2011) who find evidence that foreign equity markets respond directly to FOMC announcements even after controlling for the U.S. stock returns. In the paper, the time-varying response plots represented by the dotted lines in Figures 3.3-3.7 are better able to uncover the direct impact of the U.S. monetary policy shocks after controlling for the influence of the U.S. stock returns on the foreign stock markets.

In case of the European stock markets, Figures 4(a)-4(o) reveal that once the influence of the U.S. stock market is controlled for, the impact of the FOMC announcements on the European stock returns is reduced to some extent. However, this is not the case for most of the non-European stock markets. The stock market response of these countries to FOMC announcements does not alter significantly even after controlling for the U.S. stock returns. This suggests that the extent of comovement between the U.S. and the European stock markets is much more stronger than comovement between the U.S. and the Asian, Latin American and

Middle Eastern stock markets. Japanese and Australian markets are exception to this, as the response coefficients change significantly after for S&P returns in these two countries. This reflects the strong interlinkage between these economies and the U.S. financial markets.

Therefore, the results in this study suggest that the time-varying response of foreign stock market returns to U.S. monetary policy shocks remain qualitatively unchanged even after controlling for the direct impact of S&P 500 returns in the regression equation. In terms of the actual estimates, it is found that the stock market response of the countries that have developed financial markets and have strong financial linkages with the U.S. witness significant reduction after one controls for the S&P 500 returns. However, it is also found that the response of most of the non-European stock markets remain almost unchanged even after the inclusion of S&P 500 returns in the model specification.

3.7 Conclusions

This essay performs an estimation of the time-varying response of the foreign stock markets to U.S. monetary policy surprises derived from the Federal funds futures market. The study notes that a fixed-coefficient approach of estimating the equity return response to U.S. monetary policy surprise is not able to capture the gradual evolution of the global stock market sensitivity to U.S. monetary policy changes. These responses may change over time because of factors like time-varying financial integration, the state of the business cycles and the time-variation in equity risk

premia itself. The results show significant time-variation in the response of global equity markets to U.S. monetary policy surprises. Further, the impact of the Fed policy shocks on the Asia-pacific region and Latin American economies is at least as strong as the impact on the U.S. and the European equity markets.

The results also show that the stock markets across the world respond more to U.S. monetary policy surprises during the crisis periods. In the sample, these crisis periods include the LTCM crisis of 1997, the 2001 recession and the financial crisis of 2008. The findings suggest an interesting pattern in how the European and the U.S. market responded to a surprise cut in Federal funds rate during the 2008 financial crisis. Rather than boosting the stock market, the study finds that the surprise cut in the interest rate during the 2008 financial crisis led to a fall in the stock returns across Europe and the North American markets. This reflects the abnormal nature of the current recession and the surprise cut in the interest rate may have revealed information about deteriorating economic outlook.

Table 3.1: Summary Statistics

Country	Code	Period	Obs	Mean	Median	Max	Min	S.D.
Argentina	MERV	8-Oct-96	103	0.20	0.22	7.36	-4.53	1.67
Australia	AORD	17-May-94	123	0.38	0.22	6.27	-1.52	1.21
Austria	ATX	17-May-94	123	0.40	0.38	4.96	-2.55	1.23
Belgium	BFX	17-May-94	123	0.34	0.28	4.81	-4.25	1.22
Brazil	BVSP	17-May-94	123	0.56	0.44	6.76	-4.67	1.76
Canada	TSE	3-Jan-00	83	0.39	0.28	5.58	-1.88	1.28
Chile	IGPA	3-Jan-00	76	0.06	-0.04	3.07	-1.88	0.66
China	SSEC	4-Jan-00	76	0.05	0.09	5.80	-4.51	1.58
Denmark	OMXC	3-Jan-00	76	0.19	0.06	4.59	-4.24	1.51
Egypt	CCSI	2-Jul-97	97	0.24	0.17	5.43	-3.01	1.06
France	CAC	17-May-94	123	0.25	0.31	4.86	-4.56	1.55
Germany	DAX	17-May-94	123	0.27	0.52	4.76	-5.50	1.64
Greece	ATH	3-Jan-00	74	0.22	0.32	3.91	-4.65	1.36
Hongkong	HSI	17-May-94	123	0.64	0.19	13.43	-3.10	2.21
India	BSE	1-Jul-97	97	0.59	0.46	7.69	-2.22	1.66
Indonesia	JKSE	1-Jul-97	96	0.81	0.51	11.73	-3.80	2.42
Ireland	ISEQ	7-Apr-98	78	0.25	0.24	5.12	-2.86	1.52
Israel	TASE	1-Jul-97	96	0.12	-0.01	4.35	-3.05	1.36
Italy	ITY	2-Jun-03	46	0.20	0.30	3.54	-2.95	1.16
Japan	NIK	17-May-94	123	0.40	0.28	7.84	-3.49	1.74
Korea	KOSPI	1-Jul-97	97	0.66	0.36	9.63	-5.74	2.28
Malayasia	KLSE	17-May-94	123	0.62	0.34	10.41	-2.55	1.70
Mexico	MXX	17-May-94	123	0.29	0.24	6.49	-2.88	1.36
Netherland	AEX	17-May-94	123	0.36	0.55	5.00	-4.64	1.46
New Zealand	NZSE	30-Apr-04	39	0.25	0.13	3.36	-1.63	0.91
Norway	NOR	7-Feb-01	66	0.23	0.13	4.71	-4.25	1.47
Pakistan	PAK	2-Jul-97	97	-0.31	-0.11	2.85	-6.89	1.64
Philippines	PSEI	3-Jan-00	76	0.60	0.22	12.86	-2.53	1.97
Portugal	POR	24-Jan-00	75	0.08	0.05	2.92	-3.41	1.26
Singapore	STIS	17-May-94	122	0.52	0.33	8.43	-2.18	1.65
Spain	IBEX	17-May-94	123	0.18	0.29	4.35	-4.77	1.55
Sweden	SWD	8-Jan-01	66	0.39	0.08	5.67	-2.94	1.68
Switzerland	SSMI	9-Nov-90	123	0.32	0.18	3.67	-3.85	1.12
Taiwan	TWI	2-Jul-97	95	0.26	0.07	7.71	-5.04	1.86
UK	FTSE	17-May-94	123	0.28	0.35	5.81	-2.72	1.17
USA	SNP	17-May-94	123	0.41	0.22	5.21	-2.81	1.28

The following table shows the summary statistics of the event returns of the global stock market indexes. The column titled ‘Period’ refers to the beginning period for various countries when the data is available. The column titled ‘Obs’ refers to the number of FOMC meeting announcements from the start date till June 25, 2008.

Table 3.2: Response of Global Equity Indexes to the Fed Surprise

	Full Sample			Excluding Recent crisis		
Country	β^i	R-Sq	Obs	β^i	R-Sq	Obs
U.S. and Canada:						
U.S.	-0.02*	0.02	123	-0.07***	0.24	104
	(0.01)			(0.01)		
Canada	-0.06***	0.32	83	-0.08***	0.35	64
	(0.01)			(0.01)		
Europe:						
Austria	-0.00	0.00	123	-0.03***	0.07	104
	(0.01)			(0.01)		
Belgium	0.01	0.01	123	-0.00	0.00	104
	(0.01)			(0.01)		
Denmark	0.00	0.00	76	-0.03	0.02	57
	(0.01)			(0.02)		
France	-0.00	0.00	123	-0.04**	0.05	104
	(0.01)			(0.02)		
Germany	0.01	0.00	123	-0.02	0.02	104
	(0.01)			(0.02)		
Greece	0.01	0.00	74	-0.05**	0.12	55
	(0.01)			(0.02)		
Ireland	0.01	0.00	78	0.01	0.00	59
	(0.01)			(0.02)		
Italy	0.03**	0.12	46	-0.23*	0.13	27
	(0.01)			(0.11)		
Netherland	0.00	0.00	123	-0.03*	0.03	104
	(0.01)			(0.02)		
Norway	0.01	0.00	66	-0.00	0.00	47
	(0.02)			(0.02)		
Portugal	-0.00	0.00	75	-0.03*	0.06	56
	(0.01)			(0.02)		
Spain	-0.02	0.01	123	-0.07***	0.14	104
	(0.01)			(0.02)		
Sweden	0.01	0.01	66	-0.05*	0.07	47
	(0.02)			(0.03)		
Switzerland	0.01	0.00	123	-0.00	0.00	104
	(0.01)			(0.01)		
UK	-0.01	0.01	123	-0.04***	0.11	104
	(0.01)			(0.01)		

continued on next page

	Full Sample			Excluding Recent crisis		
	β^i	R-Sq	Obs	β^i	R-Sq	Obs
Latin America:						
Argentina	-0.05***	0.09	103	-0.10***	0.26	84
	(0.01)			(0.02)		
Brazil	-0.03*	0.02	123	-0.03	0.02	104
	(0.02)			(0.02)		
Chile	-0.02***	0.14	76	-0.01**	0.07	57
	(0.01)			(0.01)		
Mexico	-0.05***	0.13	123	-0.04**	0.06	104
	(0.01)			(0.01)		
Asia-Pacific:						
Australia	-0.03***	0.06	123	-0.02**	0.04	104
	(0.01)			(0.01)		
China	-0.01	0.01	76	0.02	0.02	57
	(0.02)			(0.02)		
Hongkong	-0.10***	0.21	123	-0.08***	0.14	104
	(0.02)			(0.02)		
India	-0.06***	0.13	97	-0.04**	0.05	78
	(0.01)			(0.02)		
Indonesia	-0.07***	0.10	96	-0.05*	0.04	77
	(0.02)			(0.03)		
Japan	-0.01	0.00	123	-0.00	0.00	104
	(0.02)			(0.02)		
Korea	-0.04*	0.04	97	-0.09***	0.12	78
	(0.02)			(0.03)		
Malayasia	-0.01	0.00	123	-0.01	0.00	104
	(0.01)			(0.02)		
New Zealand	0.00	0.00	39	0.16	0.12	20
	(0.01)			(0.10)		
Pakistan	0.03**	0.04	97	0.05*	0.05	78
	(0.02)			(0.02)		
Philippines	-0.01	0.00	76	-0.03	0.03	57
	(0.02)			(0.02)		
Singapore	-0.05***	0.09	122	-0.07***	0.14	104
	(0.01)			(0.02)		
Taiwan	0.00	0.00	95	-0.06**	0.07	76
	(0.02)			(0.02)		

continued on next page

	Full Sample			Excluding Recent crisis		
	β^i	R-Sq	Obs	β^i	R-Sq	Obs
Middle East:						
Egypt	-0.03***	0.09	97	0.00	0.00	78
	(0.01)			(0.01)		
Israel	0.01	0.01	96	-0.03*	0.05	78
	(0.01)			(0.02)		

The table shows the estimates from the regression equation for the full sample (1994-2008) and for the (1994-2006) sample period that excludes the recent financial crisis episode. The estimated regression equation is: $R_t^i = \alpha^i + \beta^i \Delta r_t^u + e_t^i$ where, R_t^i represents the abnormal return of country i at the event date t and Δr_t^u captures the monetary policy surprise calculated form the Fed funds futures data. Standard errors are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1%, respectively.

Table 3.3: Response of Global Equity Indexes to the Fed Surprise for Different Sample Periods

Country	May 94 - Jan 01 β^i	R^2	Obs	Jan 01 - Sept 06 β^i	R^2	Obs	Sept 06 - Jun 08 β^i	R^2	Obs	F-Test [P-Val]
U.S. and Canada:										
USA	-0.06***	0.2	56	-0.09***	0.48	49	0.04	0.14	20	14.06
	(0.02)			(0.01)			(0.02)			[0.00]
Canada	-0.09***	0.41	16	-0.08***	0.46	49	-0.05***	0.32	20	3.26
	(0.03)			(0.01)			(0.02)			[0.14]
Europe:										
Austria	-0.04***	0.14	56	-0.02	0.03	49	0.03	0.05	20	5.26
	(0.01)			(0.02)			(0.03)			[0.03]
Belgium	-0.01	0.0	56	0.01	0.01	49	0.02	0.05	20	1.85
	(0.02)			(0.02)			(0.02)			[0.43]
Denmark	-0.05	0.18	9	-0.02	0.01	49	0.02	0.04	20	2.80
	(0.04)			(0.02)			(0.02)			[0.21]
France	-0.05**	0.1	56	-0.03	0.04	49	0.04*	0.17	20	6.57
	(0.02)			(0.02)			(0.02)			[0.01]
Germany	-0.02	0.02	56	-0.01	0	49	0.04**	0.24	20	4.92
	(0.02)			(0.03)			(0.02)			[0.04]
Greece	-0.11	0.14	8	-0.05**	0.12	47	0.03	0.11	20	10.28
	(0.11)			(0.02)			(0.02)			[0.00]
Ireland	0.01	0.0	11	0.01	0.0	49	0.01	0.01	20	1.73
	(0.04)			(0.02)			(0.03)			[0.47]
Italy	-	-	-	-0.23*	0.13	27	0.04*	0.17	20	2.53
	-			(0.11)			(0.02)			[0.26]
Netherland	-0.04**	0.09	56	-0.02	0.01	49	0.04*	0.16	20	6.15
	(0.02)			(0.03)			(0.02)			[0.02]
Norway	-	-	-	0.00	0.0	47	0.01	0.01	20	0.59
	-			(0.02)			(0.03)			[0.96]
Portugal	-0.10**	0.65	8	-0.03*	0.07	49	0.02	0.08	20	9.79
	(0.03)			(0.02)			(0.02)			[0.00]
Spain	-0.10***	0.29	56	-0.04**	0.08	49	0.04*	0.15	20	13.77
	(0.02)			(0.02)			(0.02)			[0.00]
Sweden	-	-	-	-0.05*	0.07	47	0.04	0.12	20	5.62
	-			(0.03)			(0.02)			[0.02]
Switzerland	-0.01	0.02	56	0.01	0.01	49	0.02	0.05	20	1.95
	(0.01)			(0.02)			(0.02)			[0.40]
UK	-0.04***	0.15	56	-0.05***	0.15	49	0.02	0.05	20	5.74

continued on next page

	May 94 - Jan 01			Jan 01 - Sept 06			Sept 06 - Jun 08			F-Test
	β^i	R^2	Obs	β^i	R^2	Obs	β^i	R^2	Obs	[P-Val]
	(0.01)			(0.02)			(0.02)			[0.02]
Latin America:										
Argentina	-0.11***	0.35	36	-0.10***	0.29	49	0.00	0	20	9.28
	(0.03)			(0.02)			(0.02)			[0.00]
Brazil	-0.01	0	56	-0.07***	0.18	49	-0.02	0.03	20	3.36
	(0.03)			(0.02)			(0.02)			[0.13]
Chile	-0.02	0.27	9	-0.01*	0.06	49	-0.03*	0.18	20	1.63
	(0.01)			(0.01)			(0.01)			[0.51]
Mexico	-0.05**	0.09	56	-0.04***	0.15	49	-0.06***	0.34	20	4.55
	(0.02)			(0.02)			(0.02)			[0.05]
Asia-Pacific:										
Australia	-0.02*	0.06	56	-0.02*	0.08	49	-0.03	0.08	20	5.74
	(0.01)			(0.01)			(0.03)			[0.02]
China	0.02	0.19	9	0.02	0.01	49	-0.03	0.07	20	4.24
	(0.02)			(0.02)			(0.03)			[0.07]
Hongkong	-0.08**	0.11	56	-0.09***	0.41	49	-0.11**	0.3	20	4.18
	(0.03)			(0.02)			(0.04)			[0.07]
India	-0.03	0.04	30	-0.04*	0.07	49	-0.06**	0.29	20	5.19
	(0.03)			(0.02)			(0.02)			[0.03]
Indonesia	-0.13**	0.18	30	0.00	0	48	-0.08*	0.19	20	4.84
	(0.05)			(0.02)			(0.04)			[0.04]
Japan	-0.02	0.01	56	0.02	0.01	49	-0.02	0.02	20	1.68
	(0.02)			(0.03)			(0.02)			[0.49]
Korea	-0.14**	0.19	30	-0.09***	0.21	49	0.00	0	20	5.78
	(0.06)			(0.03)			(0.03)			[0.02]
Malayasia	0.00	0	56	-0.02*	0.06	49	-0.01	0.01	20	1.22
	(0.03)			(0.01)			(0.02)			[0.68]
New Zealand	-	-	-	0.16	0.12	20	0.01	0.01	20	5.82
	-			(0.10)			(0.01)			[0.02]
Pakistan	0.09**	0.14	30	0.01	0.01	49	0.02	0.11	20	4.16
	(0.04)			(0.02)			(0.01)			[0.07]
Philippines	-0.02	0.02	9	-0.03	0.05	49	0.00	0	20	4.75
	(0.04)			(0.02)			(0.04)			[0.05]
Singapore	-0.09***	0.22	56	-0.05***	0.14	49	-0.02	0.03	19	4.02
	(0.02)			(0.02)			(0.03)			[0.08]
Taiwan	-0.08*	0.12	29	-0.07***	0.15	48	0.05**	0.22	20	8.28
	(0.04)			(0.02)			(0.02)			[0.00]

continued on next page

	May 94 - Jan 01			Jan 01 - Sept 06			Sept 06 - Jun 08			F-Test
	β^i	R^2	Obs	β^i	R^2	Obs	β^i	R^2	Obs	[P-Val]
Middle East:										
Egypt	0.00	0	30	0.00	0	49	-0.05**	0.24	20	5.82
	(0.01)			(0.01)			(0.02)			[0.02]
Israel	-0.08***	0.33	30	(0.02)	0.02	49	0.04*	0.18	19	8.47
	(0.02)			(0.02)			(0.02)			[0.00]

The table shows the estimates from the following regression equation for the three sub-sample periods. $R_t^i = \alpha^i + \beta^i \Delta r_t^u + e_t^i$ where, R_t^i and Δr_t^u represent the abnormal return and the policy surprises respectively. Standard errors are reported in parentheses. *, ** and *** denote significance at 10%, 5% and 1%, respectively.

Table 3.4: Correlation Matrix: Europe and Canada

The following table shows the correlation between the response coefficient of each country with the other country in the region and the U.S. The response coefficient are obtained from the TVP-GARCH model:

$R_t^i = \alpha_t^i + \beta_t^i \Delta r_t^u + e_t^i$

Thus, the matrix below presents correlation of the time varying response (β_t) for any country i and j.

Europe-Canada	Austria	Belgium	Denmark	France	Germany	Greece	Ireland	Italy	Netherlands	Norway	Portugal	Spain	Sweden	Switzerland	UK	Canada	USA
Austria	1																
Belgium	0.74	1															
Denmark	0.65	0.52	1														
France	0.87	0.52	0.61	1													
Germany	0.84	0.42	0.59	0.94	1												
Greece	0.87	0.74	0.91	0.81	0.83	1											
Ireland	0.10	0.19	0.41	0.18	0.06	0.19	1										
Italy	0.59	0.67	0.82	0.61	0.60	0.61	-0.35	1									
Netherlands	0.81	0.57	0.58	0.93	0.87	0.80	0.11	0.65	1								
Norway	0.44	0.61	0.61	0.49	0.49	0.57	0.14	0.97	0.53	1							
Portugal	0.68	0.62	0.93	0.69	0.67	0.90	0.37	0.69	0.66	0.57	1						
Spain	0.77	0.81	0.59	0.58	0.48	0.56	0.77	0.12	0.48	0.34	0.56	1					
Sweden	0.91	0.93	0.75	0.93	0.94	0.93	0.04	0.65	0.93	0.64	0.93	0.61	1				
Switzerland	0.91	0.83	0.81	0.72	0.67	0.86	0.53	0.41	0.63	0.41	0.85	0.91	0.88	1			
UK	0.51	-0.01	0.55	0.81	0.85	0.76	0.13	0.60	0.75	0.45	0.64	0.08	0.90	0.27	1		
Canada	0.16	-0.11	0.68	0.16	0.08	0.52	0.17	0.31	0.12	0.39	0.52	0.09	0.15	0.02	0.15	1	
USA	0.61	0.11	0.69	0.69	0.69	0.72	-0.02	0.77	0.54	0.58	0.63	0.27	0.58	0.40	0.65	0.76	1

Table 3.5: Correlation Matrix: Asia-Pacific Region and Latin America

The following table shows the correlation between the response coefficient of each country with the other country in the region and the U.S. . The response coefficient are obtained from the TVP-GARCH model:

$R_t^i = \alpha_t^i + \beta_t^i \Delta r_t^u + e_t^i$

Thus, the matrix below presents correlation of the time varying response (β_t) for any country i and j.

Asia-Pacific	Australia	China	Hong Kong	India	Indonesia	Japan	Korea	Malaysia	N. Zealand	Pakistan	Philippines	Singapore	Taiwan	USA
Australia	1													
China	0.15	1												
Hong Kong	0.46	0.05	1											
India	0.53	-0.20	0.65	1										
Indonesia	0.06	0.21	0.20	0.05	1									
Japan	-0.11	-0.01	0.56	0.64	0.09	1								
Korea	0.26	0.32	0.77	0.37	0.20	0.35	1							
Malayasia	0.68	0.43	0.03	-0.22	-0.10	-0.69	0.14	1						
N. Zealand	0.01	-0.81	-0.23	0.29	-0.46	0.27	-0.53	-0.57	1					
Pakistan	0.35	0.36	-0.24	-0.20	-0.09	-0.74	0.12	0.92	-0.58	1				
Philippines	0.03	-0.31	0.30	0.10	0.31	0.07	0.34	0.12	0.28	0.19	1			
Singapore	0.21	0.49	0.60	0.41	0.45	0.66	0.45	-0.25	-0.45	-0.48	0.33	1		
Taiwan	0.26	0.25	0.49	0.15	0.13	-0.03	0.56	0.41	-0.66	0.37	0.26	0.28	1	
USA	0.36	0.08	0.58	0.22	0.15	0.13	0.57	0.35	-0.68	-0.06	0.36	0.51	0.85	1

Latin America	Argentina	Brazil	Chile	Mexico	USA
Argentina	1				
Brazil	0.60	1			
Chile	0.46	0.65	1		
Mexico	0.53	0.51	0.43	1	
USA	0.75	0.82	0.61	0.47	1

Figure 3.1: Monetary Policy Surprise Computed from the Federal Funds Futures Interest Rate

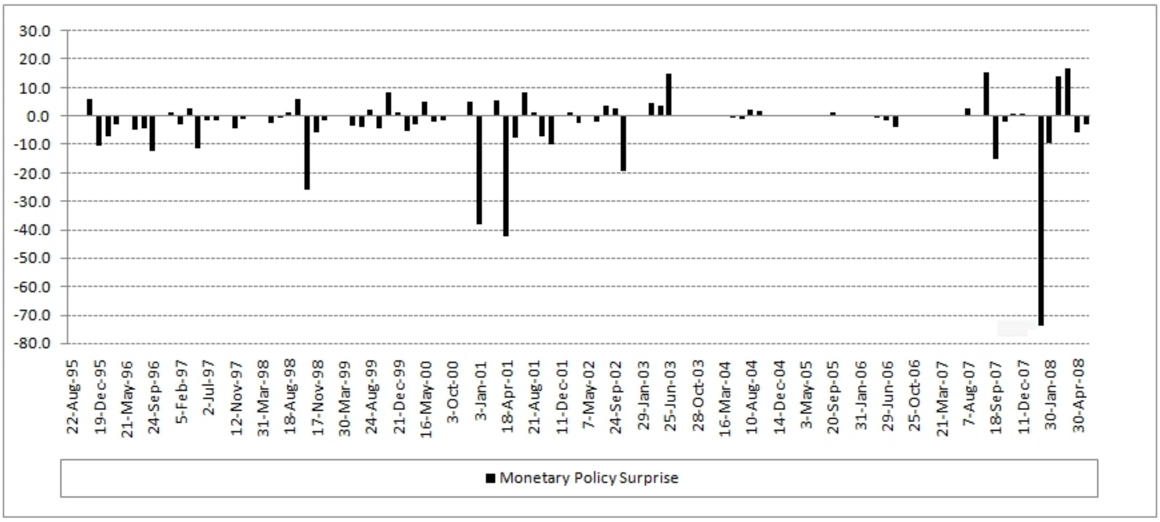
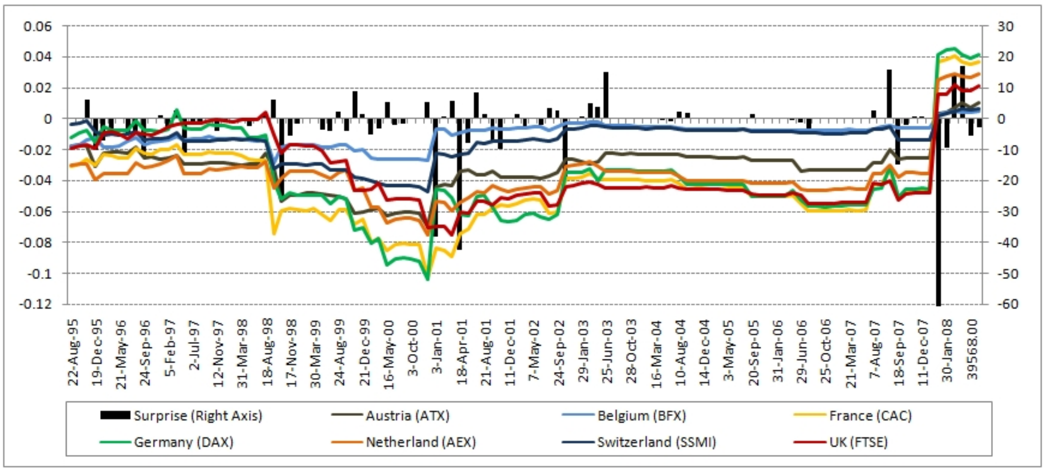
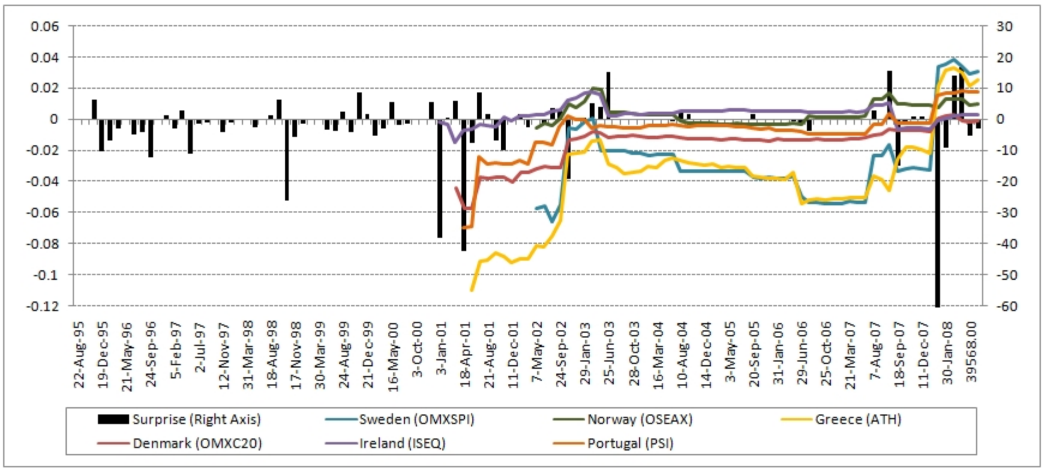


Figure 3.2: The Filtered Time Varying Response of the European Stock Markets to FOMC Announcements.

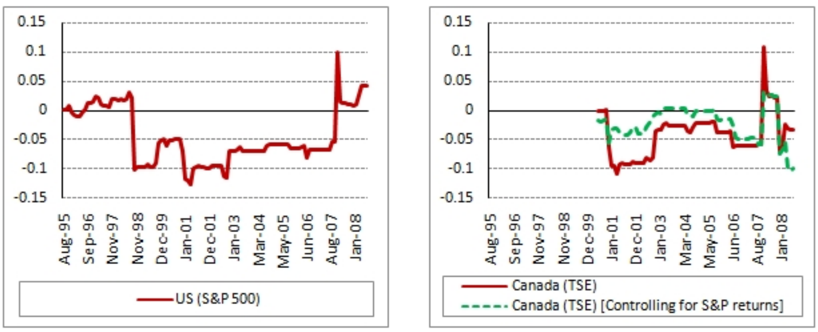


(a)



(b)

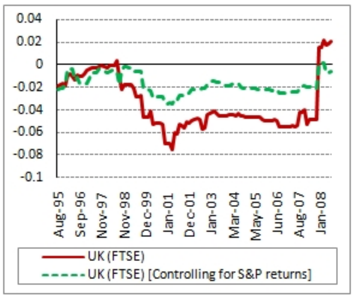
Figure 3.3: The Filtered Time Varying Response of the U.S. and Canadian Stock Markets to FOMC Announcements.



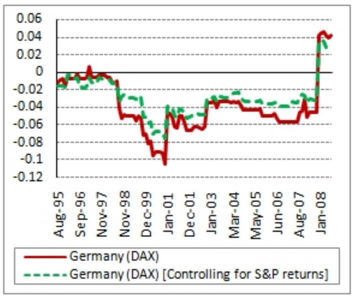
(a) The U.S.

(b) Canada

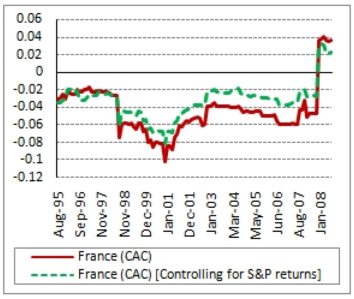
Figure 3.4: The Filtered Time Varying Response of the European Stock Markets to FOMC Announcements.



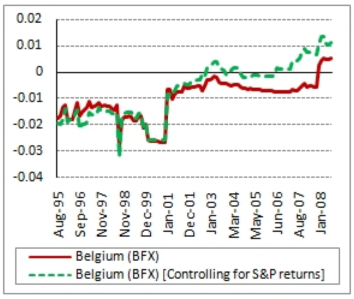
(a) UK



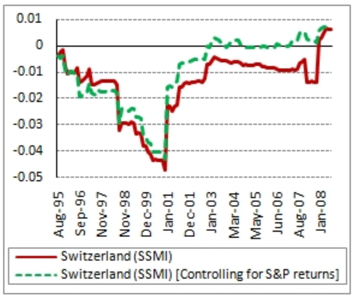
(b) Germany



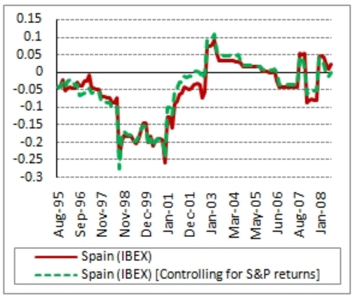
(c) France



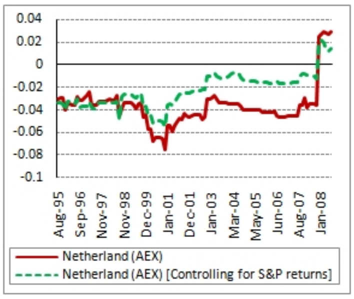
(d) Belgium



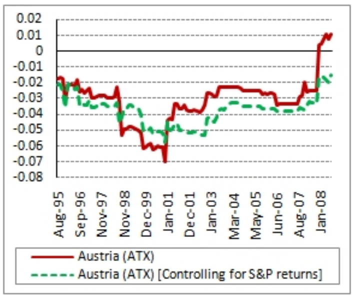
(e) Switzerland



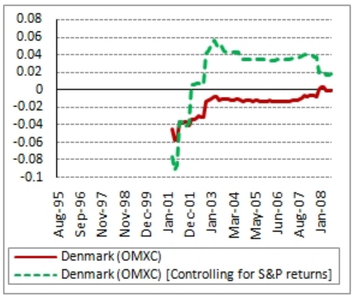
(f) Spain



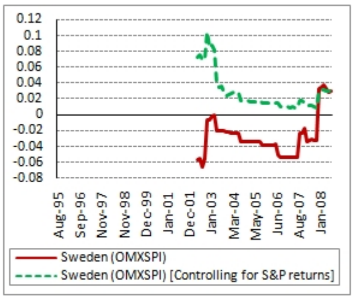
(g) Netherland



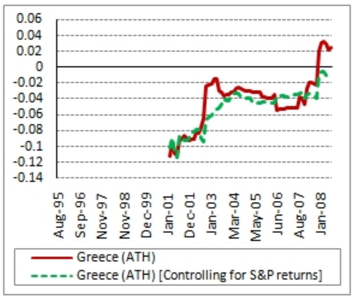
(h) Austria



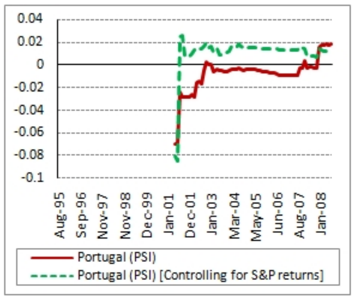
(i) Denmark



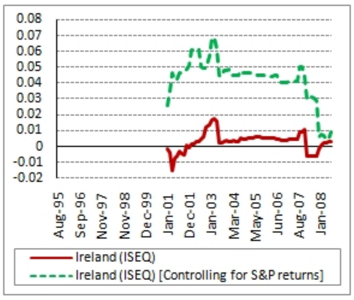
(j) Sweden



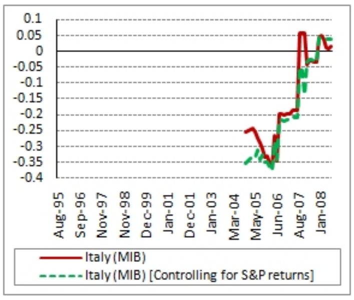
(k) Greece



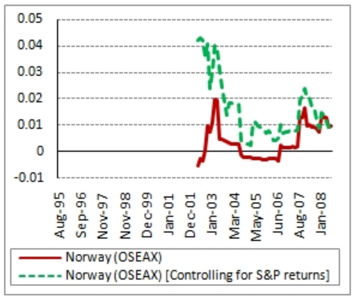
(l) Portugal



(m) Ireland

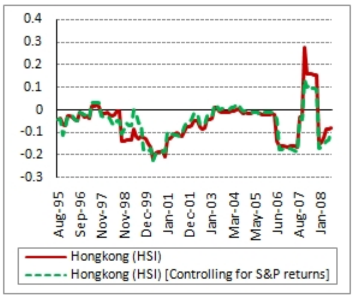


(n) Italy

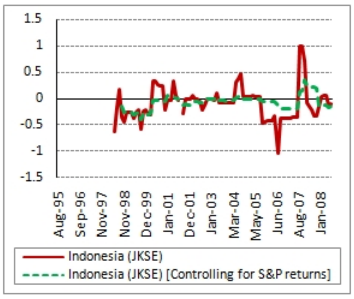


(o) Norway

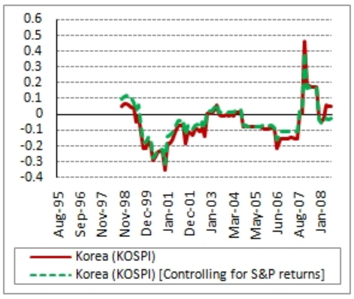
Figure 3.5: The Filtered Time Varying Response of Stock Markets in Asia-Pacific Region to FOMC Announcements.



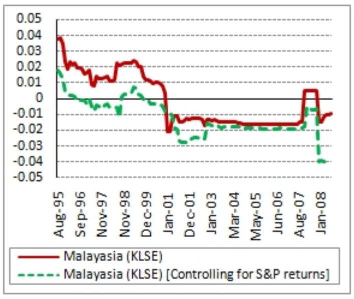
(a) Hong Kong



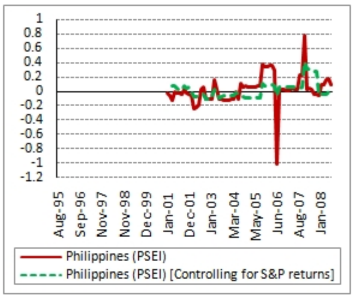
(b) Indonesia



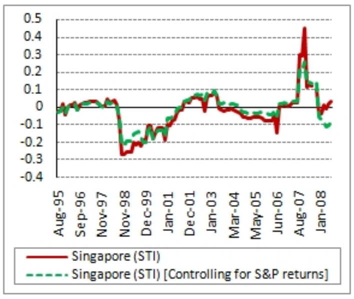
(c) Korea



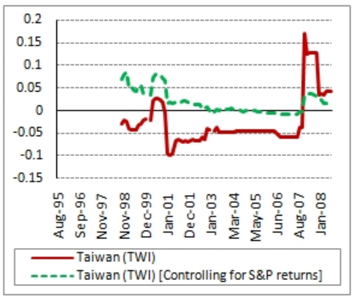
(d) Malaysia



(e) Philippines

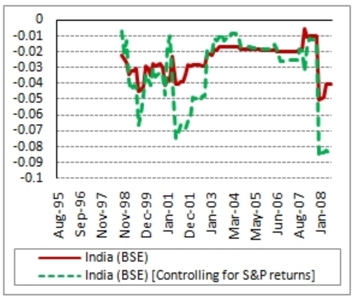


(f) Singapore

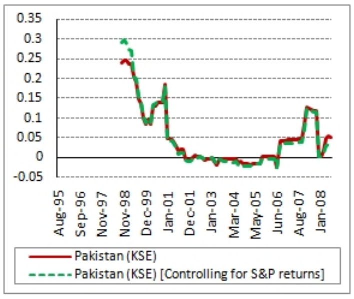


(g) Taiwan

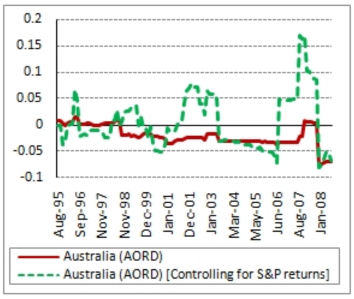
(h) China



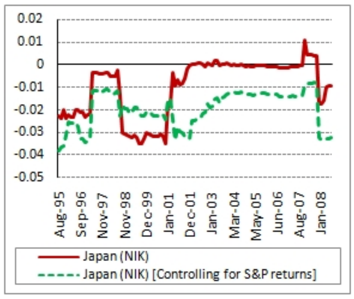
(i) India



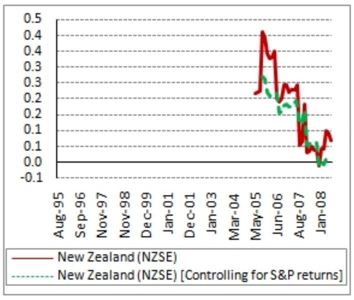
(j) Pakistan



(k) Australia

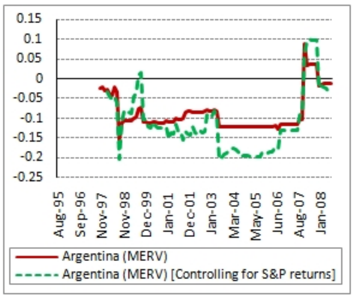


(l) Japan

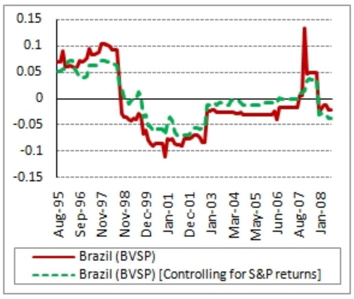


(m) New Zealand

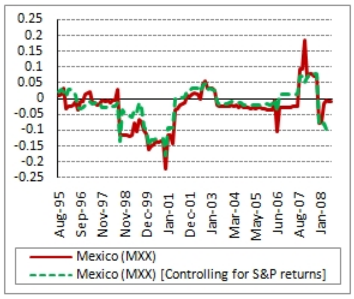
Figure 3.6: The Filtered Time Varying Response of the Latin American Stock Markets to FOMC Announcements.



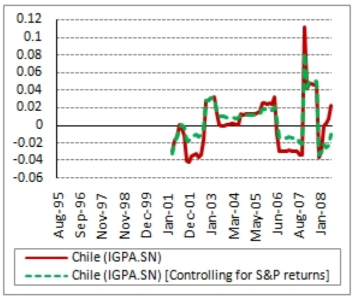
(a) Argentina



(b) Brazil

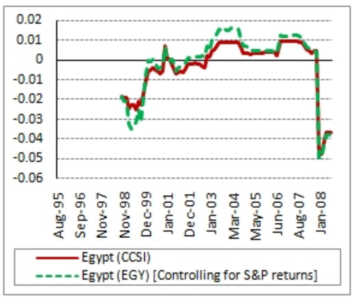


(c) Mexico

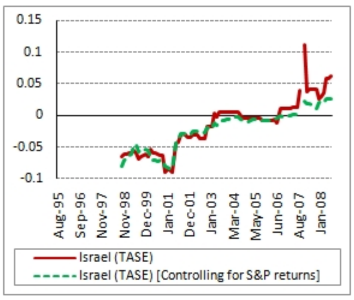


(d) Chile

Figure 3.7: The Filtered Time Varying Response of the Middle Eastern Stock Markets to FOMC Announcements.



(a) Egypt



(b) Israel

CHAPTER 4

Does Federal Funds Futures Rate Contain Information about the Treasury Bill Rate?

4.1 Introduction

One of the channels of monetary policy transmission is through its impact on the interest rates at different maturities along the yield curve. The explicit theoretical link between interest rates across different maturities is provided by the expectations hypothesis (EH hereafter). Though the empirical support for the EH is generally weak, it is well established that interest rates across maturities tend to co-move (Bradley and Lumpkin. 1992; Engsted and Tanggaard 1994; Hall et al. 1992).[1] Following expectations hypothesis, the federal funds rate and the yield on bonds especially at the shorter horizon tend to move together. However, the dynamic relationship between the effective federal funds rate and the Treasury bill rate (T-bill hereafter) is constrained by the fact that the effective federal funds rate always tend to revert back to the target federal funds rate, whereas the yield on the T-bill reflects current and expected future short-term interest rates (Sarno and Thornton 2003). To gain insight into the dynamic relationship between the T-bill market and expected monetary policy stance, we need a measure of the expected federal

[1]Even though comovement of the interest of different maturities is a necessary condition for EH to hold, it is by no means sufficient. See for example Miron (1991).

funds rate. Even though the expectations of Fed's policy actions are not directly observable, the federal funds futures market is found to possess significant information that can *efficiently* predict future monetary policy actions (Gürkaynak et al. 2007; Krueger and Kuttner 1996; Robertson and Thornton 1997).

Recognizing that the federal funds futures market contains valuable information about future movements in federal funds rate, this paper examines the dynamic relationship between the federal funds futures market and the T-bill market. The federal funds futures market data allows us to include all the relevant information of the market participants' expectations of the monetary policy actions that will impact the yield curve. This paper is certainly not the first study that links the federal funds rate with the T-bill rate. Sarno and Thornton (2003) also examine the dynamic relationship between the daily *effective* federal funds rate and the T-bill rate and find that they move together in the long-run. They find that if there is a disequilibrium in the short-run, only the *effective* federal funds rate moves to correct for the disequilibrium. The results in Sarno and Thornton (2003) are not very surprising since the *effective* federal funds rate always tends to revert back to the target federal funds rate. The *effective* federal funds rate does not incorporate any changes in federal funds rate that the market expects to take place in the future. The use of federal funds *futures* rate allows us to get around this problem, as it incorporates information about market's expectation about the future federal funds rate.

Using daily data from 1989 to 2008 the study finds that one month federal funds (FF hereafter) futures rate and the 3-month T-bill rate are cointegrated, and hence move together in the long run. Since one month FF futures rate co-move with the T-bill rate in the long run, any short run disequilibrium needs to be corrected. It turns out that if

there is disequilibrium in the cointegrating relationship between one month FF futures rate and the T-bill rate, both these rates error-correct. It is important to note that this evidences does not explicitly test the expectation hypothesis nor does this alone prove the existence of it. The results are in contrast to what Sarno and Thornton (2003) found, where only the effective federal funds rate move to correct for the short-run disequilibrium. This is not surprising since both the federal funds futures rate and the T-bill rate incorporate expectation about future changes in the federal funds rate, but the effective federal funds rate is constrained by the current federal funds rate target. The effective funds rate is constrained because the Fed ensures that it gets close to the target rate.

Given the finding that both the one month FF futures rate and the T-bill rate move to correct for the disequilibrium, we can utilize this information to decompose the movements in these two rates into trend and cycle using multivariate Beveridge-Nelson decomposition. Decomposing the FF futures rate and the T-bill rate into the trend and cycle, the study finds that there was a big positive cycle in the federal funds futures rate before 2008 implying a future downward movement in the federal funds futures rate. The study also find a negative cycle in T-bill market during the financial crisis implying the yield on T-bill was below the long-run trend.

The remainder of the paper is structured as follows. Section 2 provides a brief literature review followed by the conceptual framework in section 3. Section 4 presents the data used in the study. Empirical analysis in section 5 contain the results from error correction model and the multivariate Beveridge-Nelson trend-cycle decomposition. Robustness of results is checked in Section 6 and conclusions are presented in Section 7.

4.2 Related Literature

A big portion of the literature focuses on testing the theoretically appealing expectations hypothesis of interest rates according to which the long-term rate is determined by the market expectations for the short-term rate plus a constant risk premium.[2] Following Engle and Granger (1987) seminal work, most empirical studies have preferred cointegration and error correction technique to formally test the expectations hypothesis (Booth 1991; Campbell and Shiller 1991; Engsted and Tanggaard 1994; Hall et al. 1992; Lardic and Mignon 2004; Zhang 1993, among others). The cointegrating relationship between interest rates of different maturities implies that a single non-stationary common factor underlies the time series behavior of each of the yield to maturity and that the term risk premium is stationary.

In its essence the Expectations Hypothesis suggests that the forward rates should have significant predictive power for future short-term interest rates. Dominguez and Cinca (2002) uses data on Eurodeposits for number of currencies and studies whether forward rates can be used to improve interest rate forecasts. They find forward rates are able to produce better forecasts than those obtained from the own past of interest rates, especially at the shorter end of the yield curve.

The other strand of literature has focused on the money markets, both of the US and the world. For example, Zhou (2007) study the dynamic relationship between the federal funds rate and short-term Eurodollar deposits rates. The study finds strong evidence of co-movement between these interest rates and the adjustment of these rates

[2] The empirical evidence supporting the EH is generally weak. Campbell et al. (1997) and Sarno et al. (2007) find evidence against EH while weak support is found in Campbell and Shiller (1991); Hall et al. (1992); Zhang (1993) and Lardic and Mignon (2004).

toward long-run equilibrium appears to be related to the targeting procedural changes. Ajayi and Serletis (2009) examine the dynamic relation between daily Eurodollar and US Certificate of Deposit (CD) interest rates. They find linear causality only from the US CD rates to the Eurodollar interest rates. However, they also find significant bidirectional nonlinear causality between these two interest rates. Several researchers has also investigated the interest rate linkages at the short and long end of the yield curve, both within U.S. markets (Berument and Froyen 2008; Swiston 2007) and between the U.S. and other major economy's financial markets (Bryant and Joyeux 2010; Ehrmann et al. 2011; Lindenberg and Westermann 2012).

In the context of the U.S. money markets, Simon (1990) finds that during the federal funds targeting regime the spread between the 3-month T-bill rate and the federal funds rate have no predictive power for the funds rate itself. Barnes (2012) studies the target federal funds rate-CD rate linkage. The study particularly focuses on comparing the time series behavior of these two interest rates between pre-and post-crisis of 2007. Barnes (2012) finds a breakdown in the typical target-CD rate linkage in August 2007. During the 2-year period following 2007, even while the target and other short-term rates were falling banks kept the CD rates firm as a way to attract funds in a period of growing uncertainty.

Sarno and Thornton (2003) examine the dynamic relationship between the effective federal funds rate and the 3 month T-bill rate. The study employs nonlinear asymmetric VECM and find that the long-run equilibrium relationship between these two rates remain stable across monetary policy regimes. Importantly, the asymmetric adjustment that they find to be present in the error correction process is mainly done by the funds rate. The 3-month T-bill do not adjust in a significant way and thus not predictable in

that sense.[3] They stress on the role of expectation by the market about future monetary policy actions in explaining such interesting behavior.

In a different but related strand of literature Krueger and Kuttner (1996) show that the FF futures market incorporates virtually all the publicly available quantitative information that can efficiently predict short run movements in the monetary policy. Similar conclusions have also been found in Robertson and Thornton (1997) and Gürkaynak et al. (2007). Kuttner (2001) finds that unexpected policy action derived from the FF futures rate has a significant impact on the bill, note, and bond yields.

These evidence in the literature motivates the question raised in the paper. Does federal funds futures rate contain information about the Treasury bill rate? Our study is closest to Sarno and Thornton (2003) in flavor. However, in our study we use the federal funds *futures* rate that has been found to contain valuable information about future monetary policy stance and is an efficient predictor of federal funds rate in future. Unlike Sarno and Thornton (2003) study which finds that the T-bill rate do not adjust to correct for its long-run disequilibrium with the effective federal funds rate, our study shows that the T-bill rate plays a significant role in the short-run disequilibrium, and has much more important role in the daily dynamic relationship between the federal funds futures rate and the T-bill rate. Using this short-run error correction property, our study also undertakes a multivariate Beveridge-Nelson trend-cycle decomposition of the two interest rates.

[3] Similar evidence of weak predictability of T-bill rate is found by Campbell and Shiller (1991), Hardouvelis (1988) and Simon (1990).

4.3 Conceptual Framework

The expectations hypothesis provides the theoretical link between interest rates across different maturities. Number of studies have noted that cointegration and error correction techniques are natural ways of testing the implications of the EH (Campbell and Shiller 1991; Engle and Granger 1987; Hall et al. 1992; Stock and Watson 1988). Even though the objective of the present study is not to formally test the expectations hypothesis *per se* it would be informative to present a brief conceptual framework.

Let $R_{k,t}$ and $F_{k,t}$ denote the yield to maturity of a k-period pure discount bond and the forward rate respectively. The relationship linking $R_{k,t}$ and $F_{k,t}$ may be described according to the Fisher-Hicks recursive formula, as $R_{k,t} = \frac{1}{k}\left[\sum_{j=1}^{k} F_{j,t}\right]$ for k = 1, 2, ... (Campbell et al. 1997; Hall et al. 1992). It is well accepted that the forward rate differ from the expected future yield to maturity due to presence of the term premia that arise out of investors risk considerations and preferences for liquidity. Hence, the forward rate could be characterized as $F_{j,t} = E_t(R_{k,t+j-1}) + \phi_{j,t}$ where, $E_t(R_{k,t+j-1})$ represents the expected rates conditioned on information available at time t and $\phi_{j,t}$ is the term premium. We can then re-write the Fisher-Hicks formula as follows:

$$R_{k,t} = \frac{1}{k}\left[\sum_{j=1}^{k} E_t(R_{1,t+j-1})\right] + \delta_{j,t} \tag{4.1}$$

where $\delta_{j,t} = \sum_{j=1}^{k} \phi_{j,t}$ captures the effects of term premia components. The pure expectations hypothesis asserts that the term premia are all identically equal to zero, $\delta_{j,t} \equiv 0$. A milder version of the EH asserts a less stringent proposition that the term premia is constant over time.

Assuming that yield to maturity across maturities are integrated I(1) processes one can obtain the equation that links the yields at different maturities.

$$R_{k,t} - R_{1,t} = \frac{1}{k}\left[\sum_{i=1}^{k-1}\sum_{j=1}^{i} E_t \Delta R_{1,t+j} + \sum_{j=1}^{k}\phi_{j,t}\right] \tag{4.2}$$

This equation essentially shows that the yield on assets of different maturities tend to move together. Assuming the risk premia is constant over time the right hand side of equation (4.2) is stationary. Following from this, the left hand side is stationary as well i.e. $(R_{k,t} - R_{1,t}) \sim I(0)$ and that $(1, -1)'$ is the cointegrating vector for any series $X_t = [R_{k,t}, R_{1,t}]'$. Empirically, according to the Granger Representation Theorem (Engle and Granger 1987), cointegration between a set of variables implies the existence of a vector equilibrium correction model (VECM). This provides the rationale for modeling the dynamic inter-relationship between interest rates using a VECM approach that is adopted in the present study.

The implication of the dynamic structure of VECM for the Fisher-Hicks formula is that the movements in the long term rates are affected by market expectations of the movements in the short term rates. Since the federal funds futures rate is found to efficiently incorporate all the information about market expectations of monetary policy stance (Gürkaynak et al. 2007; Krueger and Kuttner 1996; Robertson and Thornton 1997), expectations hypothesis would suggest a co-movement in the long run between the 3-month T-bill rate and the Federal funds futures rate. It is important to note that given the institutional features of the futures markets, the present study uses the expectations hypothesis as the underlying framework but it is not a straightforward

application of equation (4.2).[4]

4.4 Data

This study uses 19 years of daily data from 05/17/1989 through 06/25/2008 for the one month federal funds futures rate and the U.S. 3-month constant maturity T-bill rate. The sample period ends in 2008 to avoid the problem of zero lower bound in the federal funds rate during the current financial crisis. The study also use the spot month federal funds futures rate and the target federal funds futures rate to support the study. The daily data for the U.S. 3-month T-bill rate is obtained from the Federal Reserve Bank of St. Louis's FRED database. The Federal funds futures contract rate as well as the target Federal funds has been provided by Kenneth N. Kuttner.[5]

The 30-day federal funds futures contract most closely track the average daily effective federal funds rate for a given calendar month as calculated and reported by the Federal Reserve Bank of New York. Since the Federal funds futures contract is based on the effective overnight Federal funds rate for a given month, it tends to be highly correlated with other short-term interest rates. The Chicago Board of Trade (CBOT) has been offering Federal funds futures contracts since October 1988 for several different deliveries going from the current month to five months ahead.[6] Even though contracts with longer deliveries exist, liquidity in those contracts is significantly lower.

[4]This is because the federal funds futures contracts are traded and settled against the average daily effective federal funds rate for the delivery month, which makes direct application of the basic Fisher-Hicks formula less appropriate.

[5] Typically, single time series of futures prices is spliced together from individual futures contract prices based on liquidity considerations (See for example, Ahn et al. (2002)).

[6]Federal funds futures contracts have a nominal value of $ 5 million and the settlement price is expressed as 100 minus the expected average effective federal funds rate for the delivery month. For example, if a January contract has a price of 95.75, it reflects an anticipated average federal funds rate of 4.25 percent for that month.

These contracts serve two important purposes. First, it allows the market participants to hedge interest rate risk. Second, it also serves the role of tracking expectations of the market about the future monetary policy actions. Among the variety of market-based measures of monetary policy expectations Gürkaynak et al. (2007) finds that federal funds futures dominate all the other securities in forecasting monetary policy at horizons out to six months.

The descriptive statistics of the data used is presented in Table 4.1. A usual yield curve is expected to slope upward. However, contrary to that the one month FF futures rate is greater than the 3 month T-bill rate by 20 basis points on an average. This is not very surprising given the fact that the T-bill rate is free of default risk whereas the funds rate is not.[7] Further, given the fact that FF futures rate captures the expectations of the market about future monetary policy actions it seems reasonable that the variability of the FF futures rate is marginally higher than the T-bill rate. For almost all the variable the third and forth moment suggests that the underlying distribution may be normal.

4.5 Empirical Analysis

4.5.1 Common Trends in Federal Funds Futures Rate and T-Bill Rate

It is well established in the literature that interest rates across different maturities move together in the long run. In order to study the dynamic relationship between the federal funds futures rate and the T-bill rate, we first examine the existence of the long-run relationship between these two rates.

[7]Sarno and Thornton (2003) note that the possible explanation to the higher effective Federal funds rate and the T-bill rate could be attributed to the difference in the tax treatment on the interest income earned from the two instruments.

Let f_t and r_t denote the one month FF futures rate and the 3-month T-bill rate respectively. As a preliminary step, non-stationarity in these two variables needs to be established. The results in Table 4.2 clearly suggest that both ADF and PP test do not reject the null of unit root for both the rates at conventional levels of significance. At the same time, the test results of a difference series does induce stationarity in each case. This clearly establishes that the two variables are integrated of order one. Since the two variables are non-stationary the equation representing the long run relationship between the FF futures rate and the T-bill rate is given by:

$$f_t = \beta_0 + \beta_1 r_t + \epsilon_t \tag{4.3}$$

In order to support the expectations hypothesis the necessary (though not sufficient) condition that needs to be satisfied is that these variables are cointegrated. Thus, f_t and r_t would share a common trend in the long run in which case the estimated cointegrating residual $\hat{\epsilon}_t = f_t - \hat{\beta}_0 - \hat{\beta}_1 r_t$ should be stationary.[8] As a first step, it is important to test for the stationarity of the cointegrating residual and perform the Johansen cointegration test for the number of cointegrating vector. To estimate the cointegrating vector the Stock and Watson (1988) dynamic ordinary least squares (DOLS) methodology is adopted. Newey-West heteroskedastic autocorrelation consistent standard errors are estimated since there is a significant degree of serial correlation in the residuals if just OLS is used.[9] In the DOLS estimation of the cointegrating vector [Eq. 4.4] fifteen lags based

[8]The "hats" represent estimated values in rest of the paper.

[9]No adjustment is necessary for the generated regressors problem since Stock and Watson (1988) have shown that the estimates of the $\beta's$ are superconsistent i.e. the true parameter converges to the true values at rate T rather than $\sqrt{T}$ as in OLS. This is in spite of the fact that the explanatory variable and error terms are correlated.

on the Schwartz criterion are selected.

$$f_t = \beta_0 + \beta_1 r_t + \sum_{j=-15}^{+15} \gamma' \Delta r_{t-j} + e_t \tag{4.4}$$

The null of the unit root in the residual is rejected at 1 percent level of significance. The stationarity of the residuals means that the two rates share a common long term trend. The results in Table 4.3 show the DOLS estimate of the cointegrating parameter to be -1.05 which is very close to -1 that is implied by the expectations hypothesis. The cointegrating residual obtained from the DOLS procedure is plotted in Figure 4.1.

The estimated cointegrating residual $\hat{\epsilon}_t > 0$ ($\hat{\epsilon}_t < 0$) essentially implies that the FF futures rate is too high (low) or that the T-bill rate is too low (high) relative to the equilibrium relationship. Though the cointegrating residual has a mean zero (by construction) it is evident from Figure 4.1 that with the onset of the recent financial crisis the residual has remained significantly positive. The FF futures rate has remain significantly above the long-run value implied by the cointegrating relationship. The long-run cointegrating relationship implies that either the federal funds rate is expected to fall or the T-bill rate is expected to rise in future.

For robustness check, Johansen cointegration test is also performed. Both the trace statistic and the maximum eigenvalue statistic indicate the presence of one cointegrating relationship. Null of no cointegration is rejected, whereas the null of one cointegrating relationship is not rejected for the maximum eigenvalue statistic as well as for the trace statistic.[10]

[10] The detailed test results for both the unit uoot test and the Johansen's cointegration test are available upon request.

4.5.2 The Dynamic Relationship

The EH presented before postulates that expectations of future short interest rates shape the term structure of longer interest rates. Especially this should be true in case where the short term rate is the Federal Reserve policy variable. In that case one can expect that the 3-month T-bill rate should adjust to the expected changes in the Fed funds rate. Even though it is difficult to measure market expectations, strong evidence exists that suggests that the FF futures market contracts provides an "efficient" forecast about the future course of Fed's policy actions.[11] Thus, the movement in the T-bill rate should be predictable by using the joint dynamics between the FF futures rate and the T-bill rate that is proved to exist in the preceding section.

More specifically, we know that the FF futures rate and the T-bill rate share a common trend in the long run. This implies that either the FF futures rate or the T-bill rate or both should be predictable since at least one of the variables would correct for any short term disequilibrium. The Engel and Granger representation theorem provides the VECM representation of the cointegrated system as follows:

$$\Delta Y_t = \nu + \alpha\hat{\beta}' Y_{t-1} + \Gamma(L)\Delta Y_{t-1} + e_t \tag{4.5}$$

where $Y_t = (1, f_t, r_t)'$, ΔY_t is the vector of the first differences, $(\Delta f_t, \Delta r_t)'$, and $\Gamma(L)$ is a finite-order distributed lag operator, $\alpha = (\alpha^f, \alpha^r)'$ represents the vector of adjustment parameters. The following VECM has been estimated

$$\Delta f_t = \gamma_{10} + \gamma^f_{11}\Delta f_{t-1} + \gamma^r_{12}\Delta r_{t-1} + \cdots + \gamma^f_{15}\Delta f_{t-3} + \gamma^r_{16}\Delta r_{t-3} + \alpha^f\hat{\beta}' Y_{t-1} + e_{ft} \tag{4.6}$$

[11] See Gürkaynak et al. (2007); Krueger and Kuttner (1996); Robertson and Thornton (1997).

$$\Delta r_t = \gamma_{20} + \gamma_{21}^f \Delta f_{t-1} + \gamma_{22}^r \Delta r_{t-1} + \cdots + \gamma_{25}^f \Delta f_{t-3} + \gamma_{26}^r \Delta r_{t-3} + \alpha^r \hat{\beta}' Y_{t-1} + e_{rt} \quad (4.7)$$

where $\hat{\beta}' Y_{t-1} = f_{t-1} - \hat{\beta}_0 - \hat{\beta}_1 r_{t-1}$ is the disequilibrium error or cointegrating residual from the last period. Three lags are chosen based on the Schwartz information criterion criteria. Even though the Akaike information criterion or Hannan-Quinn criterion suggests longer lags the results do not change qualitatively (and to a large extent even quantitatively) if more lags were included instead. If $\alpha's$ are statistically significant then it implies that the variable in the current period would adjust to restore the long-run equilibrium subsequent to any shock that distorts the equilibrium that occurred in the last period. Thus, if the Y_t is cointegrated then at least one of the $\alpha's$ must be significantly different from zero.

The results presented in Table 4.4 show that both α^f and α^r are significant at 5-percent level. Any deviation of either of the rates in the current period from their shared long term trend would lead to a correction in both of these rates in coming trading session in order to restore to their long run equilibrium path. Thus, both the one day ahead 3-month T-bill rate and the FF futures rate could be predicted by exploiting the long run joint dynamics between these two interest rates.[12] This result is in contrast to Sarno and Thornton (2003), where they find that only effective federal funds rate moved to correct for the disequilibrium.

To examine how the results would change if the spot month FF funds rate is used in place of one month FF futures rate, we perform the cointegration and the VECM analysis with spot FF futures rate. The hypothesis is that the one month ahead FF

[12] In fact even one week lagged value of cointegrating residual still posses the forecasting power of both the FF futures rate as well as the T-bill rate. Both α^f and α^r are significant for the seven day lagged value of the error correcting residual. The results are not presented here to conserve space but are available upon request.

futures contract would be expected to contain more information in predicting the 3-month T-bill rate as compared to the spot months FF futures contract. If the hypothesis were true then one month federal funds futures rate would have stronger predictive power for movement in T-bill rates than the spot federal funds futures rate.

The DOLS procedure with spot month futures contract indicate the cointegrating vector to be -1.05 which is the same as what we find when the one month futures contract is used. The VECM estimates from the spot month Fed futures contract presented in Table 4.6 provides evidence in support of the expectations hypothesis. When the spot month FF futures contract is used the α^r is statistically not different from zero.[13] Also, the adjusted R-squared drop by a very small margin as well. Even though the spot month FF futures and the T-bill rate are cointegrated it is only the spot month FF futures rate that adjusts to restore any disequilibrium error that occurred in the last period, but not the T-bill rate. This implies that the information content in the cointegrating residual obtained by using the spot month FF futures is significantly less than when the one month FF futures rate is used. It is important to note that while this may be a necessary conditions it is not a sufficient condition to confirm the expectations hypothesis of interest rates.

Following Sarno and Thornton (2003), I further check for robustness of the results by including $\Delta(FF^E - FF^T)_{t-1}$ as a RHS variable in the VECM model, where FF^E is the *effective* Fed funds rate and FF^T is the targeted funds rate. This is important because the past deviation of the *effective* Fed funds rate from the target Fed funds rate may have implications on current movements in the 3-month T-bill rate and the one

[13]Three lags are chosen based on the Schwartz criterion. Further, it also keeps the comparison even. However, results do not change in any significant way if more lags are included as implied by other lag length criterion's.

month FF futures rate. If the source of the current changes in the T-bill rate is mainly due to this factor, then it would undermine the result that the one month FF futures rate has significant information content for the movements in the T-bill rate.

The results of this robustness check is presented in Table 4.7. The results clearly remain robust to this specification. This is in contrast to Sarno and Thornton (2003) results. In their study the adjustment coefficient of the effective Fed funds rate reduces once $\Delta(FF^E - FF^T)_{t-1}$ is included as a RHS variable in the VECM. The coefficient of the error correction for both the one month FF futures rate as well as the T-bill rate do not change both quantitatively and in significance. This highlights the point that the long run relationship of one month FF *futures* market with the T-bill rate that is used in this paper has significant information content in predicting the T-bill rate which is over an above than the information implied by the Fed's targeting procedure.

Given the information advantage of the one month FF futures and its relationship with the 3-month T-bill rate, the analysis is now extended to decompose these two interest rates into the trend and the cycle.

Multivariate Beveridge-Nelson Decomposition

Using the VECM results, the study performs trend-cycle decomposition using the Beveridge-Nelson (1981) (BN hereafter) methodology. Under this methodology the trend in each variable is simply the long-run forecast of the variable that is furnished by the trend component of the multivariate BN decomposition for the cointegrated system $(f_t, r_t)'$. The results obtained from the BN methodology would reinforce the findings that the FF futures rate would have relatively more deviation from the trend as compared to the T-bill rate.

The BN methodology decomposes a non-stationary series into a random walk component and a stationary component which is the cycle of the non-stationary series. Applying the Engel-Granger theorem the BN decomposition of Y_t has the following representation:

$$Y_t = y_0 + \mu t + \Psi(1) \sum_{j=1}^{t} \epsilon_j + \breve{\epsilon}_t - \breve{\epsilon}_0 \tag{4.8}$$

where

$$\Psi(1) = \beta_{\perp}(\alpha'_{\perp}\Gamma(1)\beta_{\perp})^{-1}\alpha'_{\perp} \tag{4.9}$$

and the deterministic trend is given by $TD_t = y_0 + \mu t$, the BN trend is $TS_t = \Psi(1)\sum_{j=1}^{t}\epsilon_j$ and the cycle is $C_t = \breve{\epsilon}_t - \breve{\epsilon}_0$. Also, $\breve{\epsilon}_t = \breve{\Psi}(L)\epsilon_j$ and $\Psi(L) = \Psi(1) + (1-L)\breve{\Psi}(L)$.

The practical implementation of this trend-cycle decomposition for cointegrated system has been done by using Morley's (2002) state space technique.[14] The state space representation of the above model and decomposition of the variables into a trend and a cycle using above technique is presented in Appendix 1. Using this methodology the FF futures rate and the T-bill rate have been decomposed into the trend and the cycle.

Figures 4.2 and 4.3 presents the respective trend and cycle for the FF futures and the T-bill rate. Figure 4.2 shows that FF futures rate tends to have large deviations from its trend and thus a large cyclical component. In case of the T-bill rate Figure 4.3 shows that the cyclical component is also significant. These results are consistent with the VECM results where it is found that both FF futures and T-bill rate move to correct for the short-run disequilibrium implying the presence of significant cyclical

[14] Morley (2002) methodology provides a simple technique of trend-cycle decomposition of cointegrated variables using state-space method.

component in both of these time series.

During the recent financial turmoil Figure 4.2 reveals that the federal funds rate cycle was above the trend. At the same time, Figure 4.3 shows that the cycle of the T-bill rate was negative during the recent crisis. In the financial crises, since the risk perception of market participants remained at significantly elevated levels, this converted into increased demand for the default-risk free Treasury security. This meant that the T-bill rate remain below the long term trend and at the same time the FF futures rate remain significantly above the long term trend. The positive cycle in the FF futures rate indicated a future reduction in federal funds rate and this was validated by the actual change in the federal funds rate during 2008-09 period.

4.6 Robustness Results

Given the exceptional behavior of the relationship between the FF futures rate and the T-bill rate in the recent financial turmoil it becomes necessary to see how would the results change if the period of recent financial crisis is excluded. Thus, in order to check robustness of the results the study performs the same analysis for the period from 05/17/1989 to 07/31/2007.

The DOLS estimate of the cointegrating vector for this is period is estimated to be -1.06 which is close to the full sample estimate of -1.05 and also the theoretically implied estimate of -1. Thus, the estimate of the cointegrating vector remains unaffected by exclusion of the financial crisis. Following from this, the VECM model is estimated for the period excluding recent crisis period. These results are presented in Table 4.8. To maintain consistence Schwartz information criterion is used which suggests five lags

in the VECM model.[15]

The VECM results presented in Table 4.8 suggests that the findings in the paper clearly remain robust if the recent crisis period is excluded. Both the adjustment coefficients, α^r and α^f still remain highly significant. In fact the absolute value of α^r increases from 0.01 to 0.02 and absolute value of α^f increases marginally as well. This is implies a greater role of the information content in FF futures rate in predicting the movement in the T-bill rate as indicated by these speed of adjustment coefficients.[16] Thus, a much stronger relationship exists between the FF futures rate and the T-bill in the 1989-2007 period.

The trend-cycle analysis for the sample from 1989 to 2007 is also investigated, the results of which can be found in Figures 4.4 and 4.5. Comparing the cycle in Figure 4.2 for the period before 2007 with Figure 4.4 it is evident that there is no significant change in FF futures cycle if one excludes the recent period of financial distress. In case of the T-bill cycle also, Figures 4.3 and 4.5 provide evidence of robustness of the earlier results. The results suggest a smaller cyclical component for both the FF futures rate and the T-bill rate at the end of the sample period in 2007 as compared to the full sample period when the sample ended in 2008. This is intuitive since the bigger cyclical component for the full sample period in 2008 was reflecting the stress in the financial market

Overall the analysis remains robust to the exclusion of recent period financial turmoil. If anything, the result suggests a stronger relationship between the FF futures

[15]Again the results do not change if more lags are included as suggested by other criterion's. The other detailed results that includes DOLS output, unit root tests and cointegration tests for the truncated sample are not presented to conserve space.

[16]The adjustment coefficients do not change if $\Delta(FF^E - FF^T)_{t-1}$ is included as additional RHS vraible in the VECM.

rate and the T-bill rate.

4.7 Conclusions

In this paper, using high frequency daily data the study examines the dynamic relationship between the federal funds futures rate and the 3-month T-bill rate. The results show that the 1-month federal funds futures rate is cointegrated with the 3-month T-bill rate and thus move together in the long run. More importantly, any deviation of the one month federal funds futures rate and the T-bill rate from their long run equilibrium forces both the rates to error correct in order to restore to the equilibrium path. However, in a cointegrated system between the spot month federal funds futures rate and the T-bill rate, only the spot month FF futures rate does the error correction, but the T-bill rate does not adjust to correct for the short term disequilibrium. These findings are consistent with the implications of the expectations hypothesis of the term structure of interest rates.

Since the one month FF futures rate and the T-bill rate move to correct for the disequilibrium we can exploit this long run property to decompose the federal funds futures rate and the T-bill rate into the trend and cycle using multivariate Beveridge-Nelson methodology. The trend-cycle results show that there was a big positive cycle in the FF futures rate before 2008 implying a future downward movement in the FF futures rate. Further, a negative cycle in T-bill market during the financial crisis implying the yield on T-bill was below the long-run trend. The results are also robust to the exclusion of the recent financial crisis sample period.

Table 4.1: Descriptive Statistics

Variable	Mean	Median	Max	Min	Std.Dev.	Skewness	Kurtosis
FF Target	4.48	5.00	9.75	1.00	1.98	0.07	2.70
FF Effective	4.51	4.95	10.48	0.86	2.00	0.08	2.70
FF Futures (Spot)	4.50	4.96	9.83	0.99	1.99	0.07	2.70
FF Futures (1 Mth)	4.50	4.99	9.72	0.84	1.97	0.00	2.61
3-Mth T-bill	4.29	4.70	8.92	0.61	1.87	-0.04	2.53

Note: The table reports the summary statistics for Federal funds (FF) rates and the 3-Month constant maturity T-bill rate.

Table 4.2: Unit Root Test Results

Variable	ADF		PP	
	t-Stat	P-Value	t-Stat	P-Value
FF Futures (1 Mth)	-2.06	0.57	-2.01	0.59
3-Mth T-bill	-1.88	0.67	-1.69	0.76
Δ FF Futures (1 Mth)	-9.26	0.00	-63.06	0.00
Δ 3-Mth T-bill	-22.51	0.00	-62.94	0.00

Note: The table reports the results from the two unit root tests, namely Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. The last two row presents the test statistics for both the variables in their first difference (Δ).

Table 4.3: DOLS Estimate of Cointegrating Vector

Coefficient	Value	Std.Error	P-Value
β_0	0.002	0.0217	0.914
β_1	1.046	0.0047	0.000

Note: The table reports the DOLS estimate of the cointegrating vector. Standard errors are Newey-West HAC errors.

Table 4.4: Estimates from Cointegrated VAR Using One Month FF Futures Rate

Dependent Variable	Δf_t		Δr_t	
Δf_{t-1}	0.070	(4.437)	0.109	(5.056)
Δf_{t-2}	-0.001	(-0.040)	0.109	(5.048)
Δf_{t-3}	-0.009	(-0.560)	0.126	(5.819)
Δr_{t-1}	0.042	(3.595)	0.058	(3.625)
Δr_{t-2}	-0.003	(-0.235)	-0.120	(-7.624)
Δr_{t-3}	0.001	(0.069)	-0.120	(-7.591)
$\hat{\beta}' Y_{t-1}$	-0.017	(-6.052)	0.011	(2.947)
$\bar{R}^2$	0.020		0.042	

Note: The table reports the result from VECM estimation by using the one month FF futures contract. The t-statistic are in parenthesis. Second last row shows the adjustment coefficient of the one day lagged value of the estimated cointegrating residual $\hat{\beta}' Y_{t-1} = f_{t-1} - \hat{\beta}_0 - \hat{\beta}_1 r_{t-1}$.

Table 4.5: Estimates from Johansen's Method Using One Month FF Futures Rate

	Δf_t	Std.Error	Δr_t	Std.Error
Normalized cointegrating coefficients	1.000	-	-1.038	0.012
Adjustment coefficients	-0.019	0.003	0.011	0.004
1 Cointegrating Equation: Log likelihood 16343.95				

Note: The table reports the result from Johansen maximum likelihood method by using the one month FF futures contract. The t-statistic are in parenthesis.

Table 4.6: Estimates from Cointegrated VAR Using Spot Month FF Futures Rate

Dependent Variable	Δf_t		Δr_t	
Δf_{t-1}	0.019	(1.283)	0.049	(2.652)
Δf_{t-2}	-0.015	(-1.056)	0.078	(4.253)
Δf_{t-3}	-0.025	(-1.701)	0.053	(2.900)
Δr_{t-1}	0.024	(2.032)	0.078	(5.228)
Δr_{t-2}	-0.002	(-0.146)	-0.104	(-7.011)
Δr_{t-3}	0.006	(0.463)	-0.097	(-6.489)
$\hat{\beta}'Y_{t-1}$	-0.027	(-12.256)	-0.005	(-1.795)
$\bar{R}^2$	0.036		0.030	

Note: The table reports the result from VECM estimation by using the spot month FF futures contract. The t-statistic are in parenthesis. Second last row shows the adjustment coefficient of the one day lagged value of the estimated cointegrating residual $\hat{\beta}'Y_{t-1} = f_{t-1} - \hat{\beta}_0 - \hat{\beta}_1 r_{t-1}$.

Table 4.7: Robustness Check: Cointegrated VAR Using One Month FF Futures Rate

Dependent Variable	Δf_t		Δr_t	
Δf_{t-1}	0.072	(4.522)	0.108	(5.004)
Δf_{t-2}	0.000	(0.025)	0.108	(5.009)
Δf_{t-3}	-0.009	(-0.614)	0.126	(5.850)
Δr_{t-1}	0.039	(3.380)	0.059	(3.736)
Δr_{t-2}	-0.002	(-0.202)	-0.121	(-7.644)
Δr_{t-3}	0.000	(0.056)	-0.120	(-7.584)
$\hat{\beta}'Y_{t-1}$	-0.017	(-6.057)	0.011	(2.948)
$\Delta(FF^E - FF^T)_{t-1}$	0.000	(2.597)	-0.005	(-1.522)
$\bar{R}^2$	0.021		0.042	

Note: The table reports the result from robustness check by including the lagged deviation of the effective fed funds rate FF^E from the target Fed funds rate FF^T. The t-statistic are in parenthesis. The table shows that the adjustment coefficient of the one day lagged value of the estimated cointegrating residual $\hat{\beta}'Y_{t-1} = f_{t-1} - \hat{\beta}_0 - \hat{\beta}_1 r_{t-1}$ remain robust even after the inclusion of $\Delta(FF^E - FF^T)_{t-1}$.

Table 4.8: Robustness Check: Estimates from Cointegrated VAR Using One Month FF Futures Rate (Excluding Recent Crisis)

Dependent Variable	Δf_t		Δr_t	
Δf_{t-1}	0.072	(4.283)	0.125	(6.162)
Δf_{t-2}	0.021	(1.241)	0.104	(5.137)
Δf_{t-3}	0.000	(-0.013)	0.053	(2.616)
Δf_{t-4}	0.035	(2.059)	0.062	(3.059)
Δf_{t-5}	0.014	(0.815)	-0.002	(-0.080)
Δr_{t-1}	0.037	(2.622)	0.018	(1.087)
Δr_{t-2}	-0.011	(-0.771)	-0.114	(-6.792)
Δr_{t-3}	-0.008	(-0.553)	-0.091	(-5.418)
Δr_{t-4}	-0.053	(-3.801)	0.019	(1.155)
Δr_{t-5}	-0.014	(-0.995)	0.078	(4.691)
$\hat{\beta}'Y_{t-1}$	-0.018	(-4.827)	0.021	(4.817)
$\bar{R}^2$	0.017		0.047	

Note: The table reports the robustness of results by excluding the recent financial crisis period. The VECM estimates are obtained by using the one month FF and T-bill rate. The t-statistic are in parenthesis. Second last row shows the adjustment coefficient of the one day lagged value of the estimated cointegrating residual $\hat{\beta}'Y_{t-1} = f_{t-1} - \hat{\beta}_0 - \hat{\beta}_1 r_{t-1}$.

Figure 4.1: Estimated Cointegrating Residual

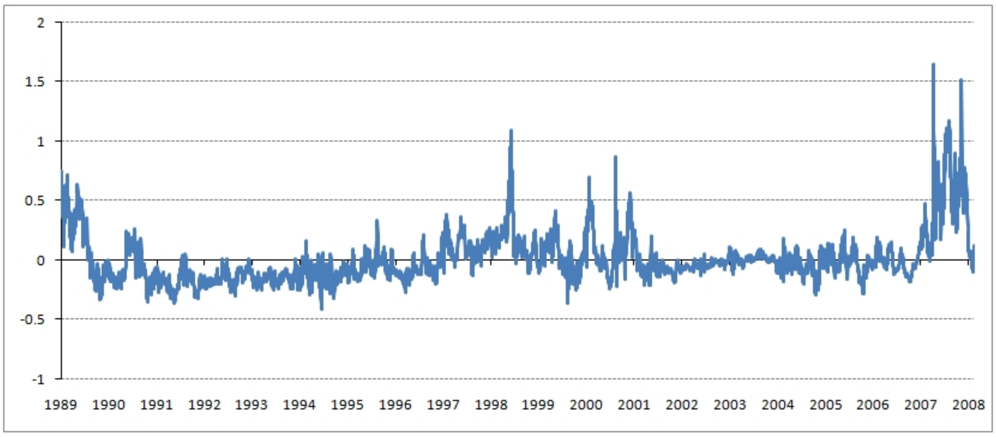
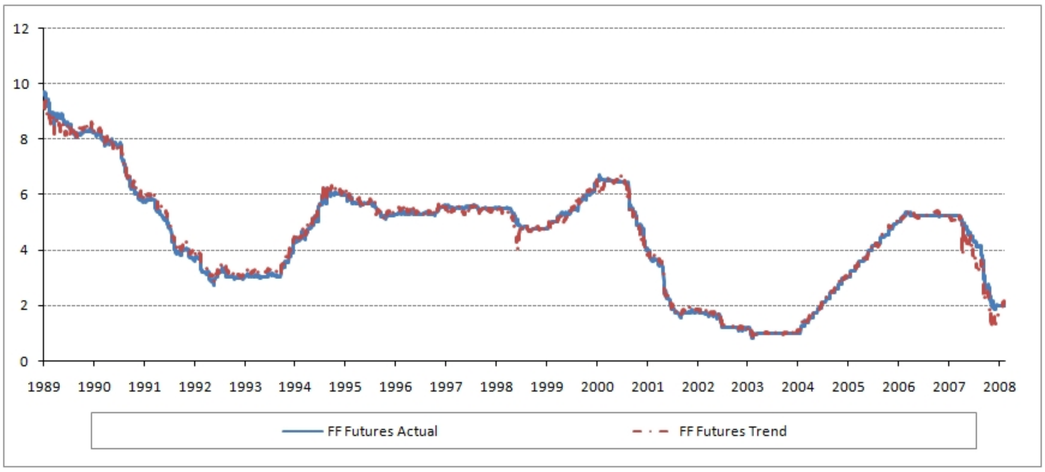
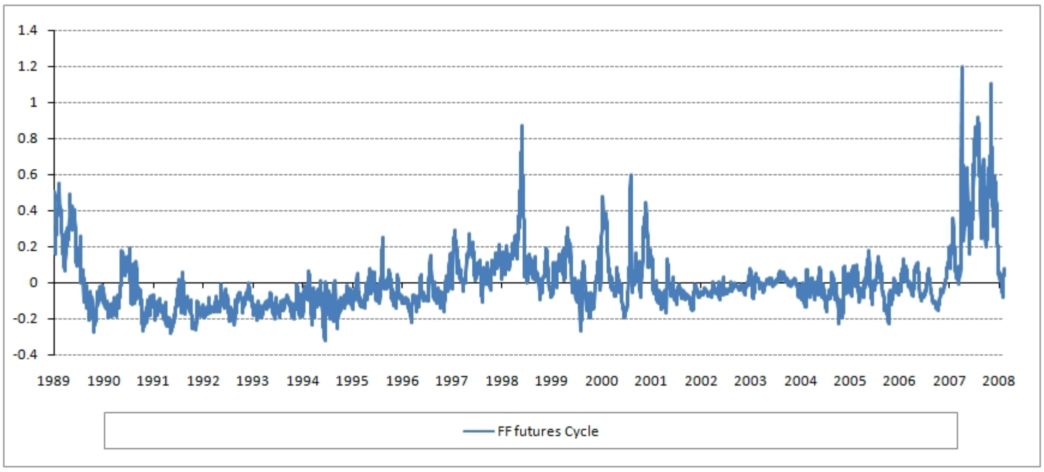


Figure 4.2: Estimates of the Trends and Cycle in FF Futures Rate.

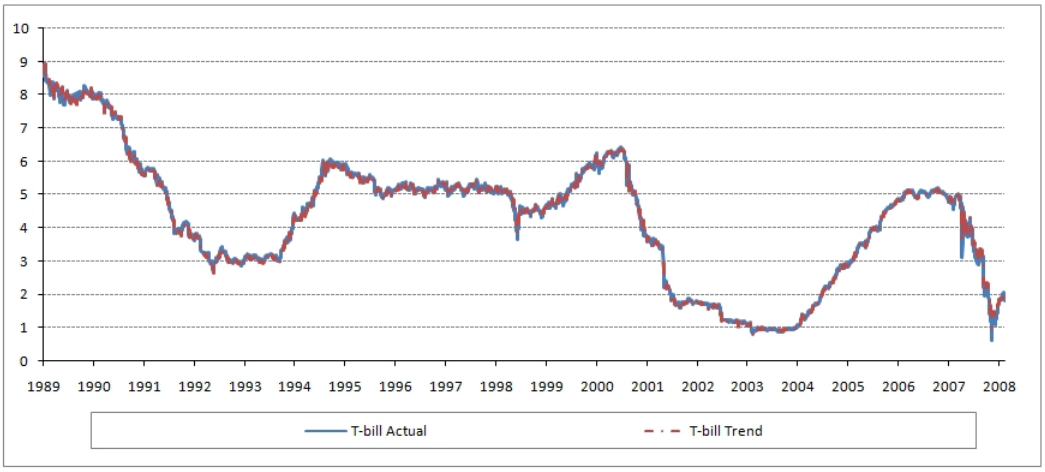


(a) FF Futures Rate Actual and Trend

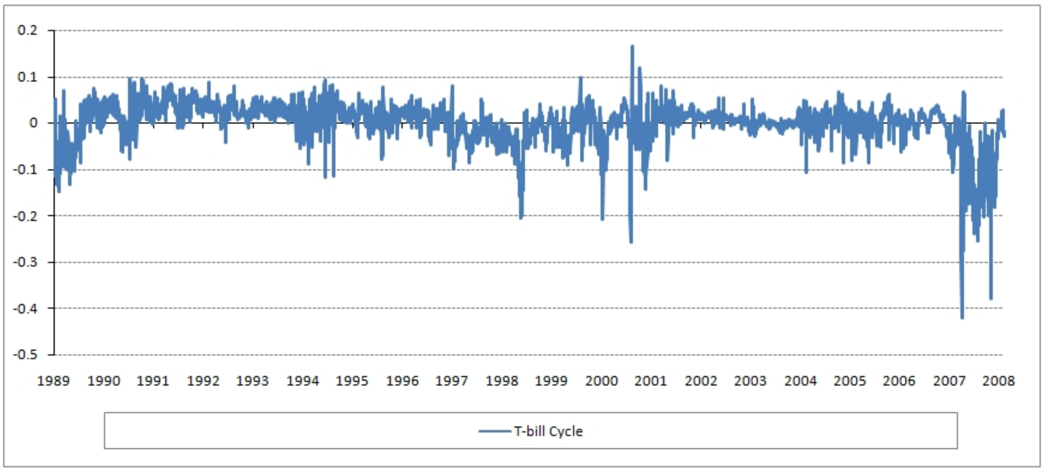


(b) FF Futures Rate Cycle

Figure 4.3: Estimates of the Trends and Cycle in T-bill Rate.

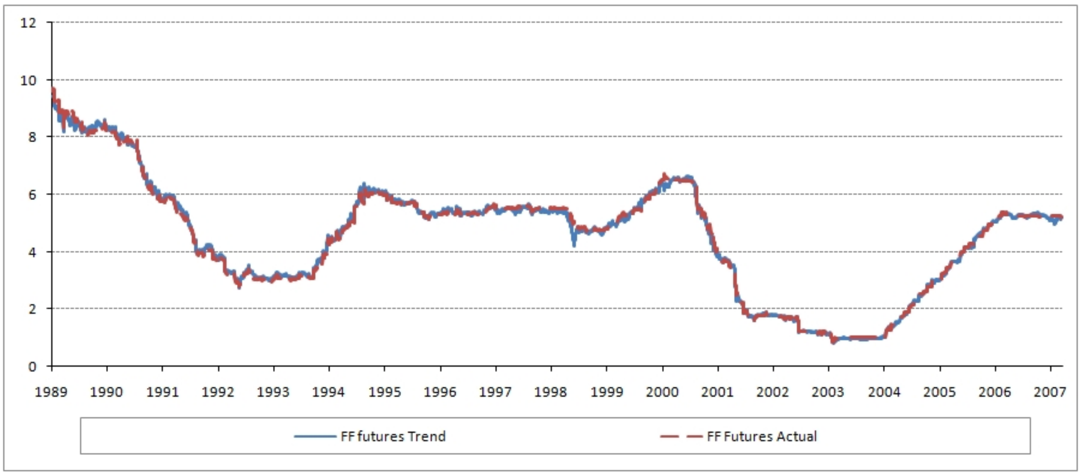


(a) T-bill Rate Actual and Trend

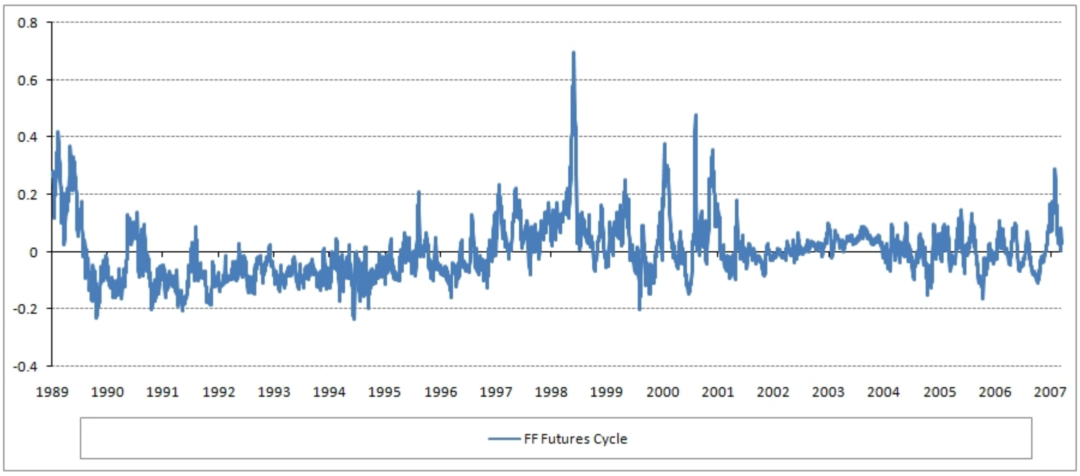


(b) T-bill Rate Cycle

Figure 4.4: Estimates of the Trends and Cycle in FF Futures Rate (Excluding Recent Crisis).

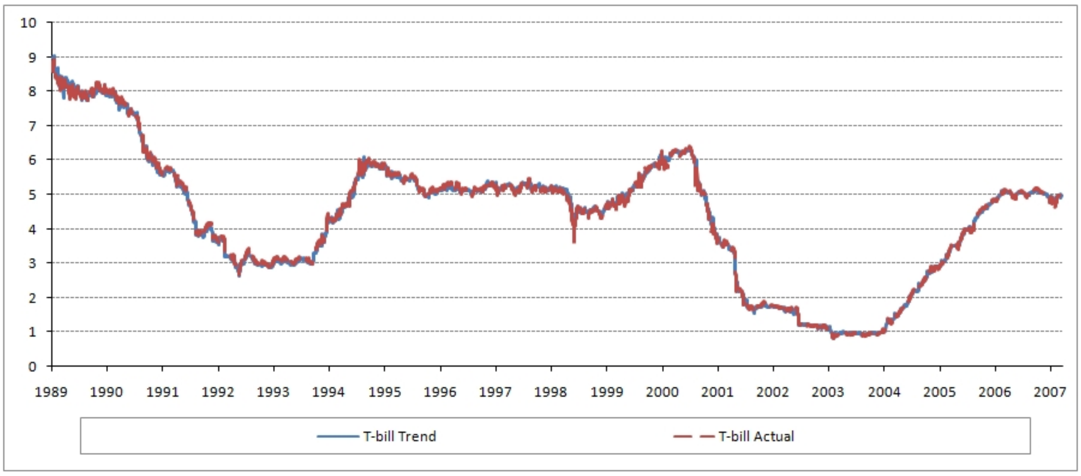


(a) FF Futures Rate Actual and Trend

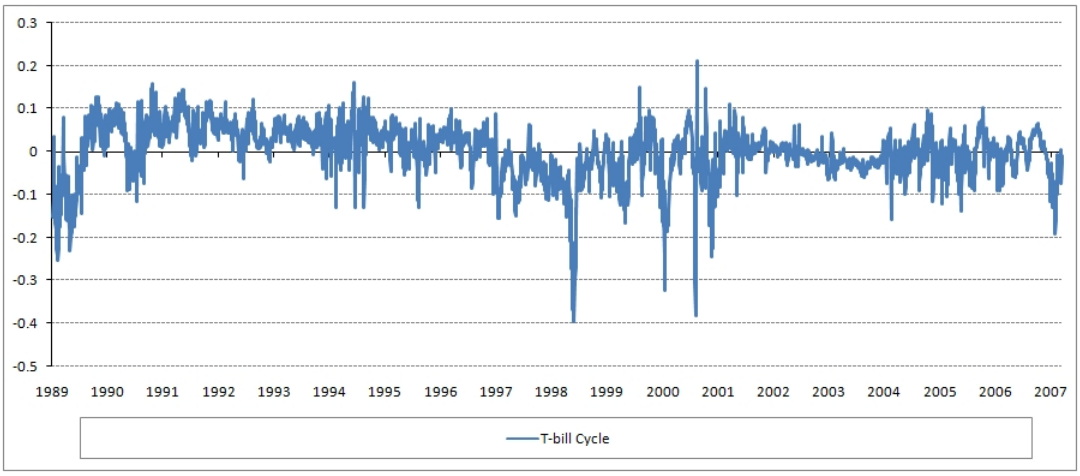


(b) FF Futures Rate Cycle

Figure 4.5: Estimates of the Trends and Cycle in T-bill Rate (Excluding Recent Crisis).



(a) T-bill Rate Actual and Trend



(b) T-bill Rate Cycle

CHAPTER 5
Concluding Remarks

Expectations of economic agents play a central role in monetary policy analysis. However, empirical estimation is significantly challenged since the true underlying expectations are not directly observable. In a set of three essays, I addresses this challenge to study topics relating to monetary policy. To this end, I use the information in various forward-looking financial market instruments. These instruments capture expectations of the agents and have several appealing features of being market determined, forward looking, available at high frequency, and mostly efficient.

One of the central topics relating to monetary policy is understanding the inflation-output relationship in the economy. In the first essay, I study the trade off between inflation and output within the hybrid New Keynesian Phillips Curve framework for the U.K.. Under this framework inflation expectations is one of the important factors driving current inflation. To measure these unobserved inflation expectations, I take advantage of information contained in the inflation-indexed bonds market. More importantly, since the true inflation expectations are not directly observable the estimation is performed using an unobserved component model. Results suggest that inflation expectations estimated from this model play a statistically significant role in driving inflation dynamics in the U.K.. Further, there is evidence of a mild but statistically significant

trade-off between inflation and output.

Another topic that interests the monetary policy maker is assessing the asset price channel of monetary policy. This is because the most immediate impact of monetary policy actions is felt on the financial markets. The second essay evaluates the impact of federal reserve policy surprises on the stock markets across the world. These surprises are derived from high frequency federal funds futures market contracts. Several researchers have found these instruments to efficiently capture the expected monetary policy actions by the Fed. One of the main contributions of the study is that it accounts for the possible time variation in the response of stock returns across the world to U.S. monetary policy surprises. Evidence suggest existence of significant time variation in the response of world stock markets to fed policy surprises, where an unanticipated interest rate cut leads to an increase in stock returns. The time varying pattern show that the foreign stock markets respond more to U.S. monetary policy surprises during the crisis periods. Further, responses to the fed policy surprise was remarkably different during current financial crisis which is not found in earlier episodes of crisis.

The information advantage of federal funds futures rate at capturing monetary policy actions raises another interesting question - does federal funds futures rate contain information about the Treasury bill rate? The third essay studies the dynamic relationship of the one month federal funds futures rate with the 3-month T-bill rate. Evidence suggests the two rate are cointegrated, and thus share a common long run trend. The key finding is that any deviation from the long run equilibrium leads to a correction in both the one month federal funds futures rate and the T-bill rate to restore the long run

trend. The study further extends the analysis by decomposition of the two variables into a trend and cycle using multivariate Beveridge-Nelson methodology.

Overall, my dissertation attempts at addressing empirical challenge of measuring expectations involved in monetary policy models. This is done by taking advantage advantage of the information in the forward looking financial market securities. In a set of three essays, I analyze various important topics related to monetary policy. Particularly, estimating inflation dynamics, assessing the asset price of monetary policy, and understanding the dynamic relationship of the money markets.

References

Ahn, Dong-Hyun, Matthew Richardson Jacob Boudoukh, and Robert F. Whitelaw. 2002. "Partial Adjustment or Stale Prices? Implications from Stock Index and Futures Return Autocorrelations." *Review of Financial Studies*, 15 (2): 655–689.

Ajayi, Richard A. and Apostolos Serletis. 2009. "Testing for causality in the transmission of Eurodollar and US interest rates." *Applied Financial Economics*, 19 (6): 439–443.

Andersen, Torben, Tim Bollerslev, Francis X. Diebold, and Clara Vega. 2007. "Real-Time Price Discovery in Global Stock Bond and Foreign Exchange Markets." *Journal of International Economics*, 73 (2): 251–277.

Anderson, Nicola and John Sleath. 2001. "New Estimates of the UK Real and Nominal Yield Curves." Bank of England Working paper No. 126.

Balakrishnan, Ravi and J. David Lopez-Salido. 2002. "Understanding UK Inflation: The Role of Openness." Bank of England Working Paper No. 186.

Barnes, Bob J. 2012. "The Change in the CD RateTarget Federal Funds Rate Linkage During the Financial Crisis." *Applied Economics Letters*, 19 (12): 1207–1211.

Barra, David G. and John Y. Campbell. 1997. "Inflation, Real Interest Rates, and the Bond market: A Study of UK Nominal and Index-Linked Government Bond Prices." *Journal of Monetary Economics*, 39 (3): 361–383.

Basistha, Arabinda and Charles R. Nelson. 2007. "New Measures of the Output Gap Based on the Forward-Looking New Keynesian Phillips Curve." *Journal of Monetary Economics*, 54 (2): 498–511.

Bekaert, Geert and Campbell Harvey. 1995. "Time-varying world market integration." *Journal of Finance*, 50 (2): 403–444.

Bekaert, Geert and Campbell Harvey. 1997. "Emerging Equity Market Volatility." *Journal of Financial Economics*, 43 (1): 29–78.

Bernanke, Ben and Kenneth Kuttner. 2005. "What Explains the Stock Market's Reaction to Federal Reserve Policy?" *Journal of Finance*, 60 (3): 1221–1257.

Bernanke, Ben and Mark Gertler. 1999. "Monetary policy and Asset Price Volatility." *Federal Reserve Bank of Kansas City Economic Review*, 84, 17–51.

Berument, Hakan and Richard Froyen. 2008. "Monetary Policy and U.S. Long-Term Interest Rates: How Close are the Linkages?" *Journal of Economics and Business*, 61 (1): 34–50.

Boivin, Jean. 2006. "Has U.S. Monetary Policy Changed? Evidence from Drifting Coefficients and Real-Time Data." *Journal of Money, Credit and Banking*, 38 (5): 1149–1174.

Bomfim, Antulio. 2003. "Pre – Announcement Effects News Effects and Volatility: Monetary Policy and the Stock Market." *Journal of Banking & Finance*, 27 (1): 133–151.

Booth, Paul. 1991. "Interest Parity, Cointegration, and the Term Structure in Canada and the United States." *Canadian Journal of Economics*, 24 (3): 595–603.

Bradley, Michael and Stephen A. Lumpkin. 1992. "The Treasury Yield Curve as a Cointegrated System." *Journal of Financial and Quantitative Analysis*, 27 (3): 449–463.

Bårdsen, Gunnar, Eilev S. Jansen, and Ragnar Nymoen. 2004. "Econometric Evaluation of the New Keynesian Phillips Curve." *Oxford Bulletin of Economics and Statistics*, 66 (s1): 671–686.

Bredin, Don, Gerard O'Reilly, and Simon Stevenson. 2007. "Monetary Shocks and REIT Returns." *Journal of Real Estate Finance and Economics*, 35 (3): 315–331.

Bredin, Don, Stuart Hyde, Dirk Nitzsche, and Gerard O'reilly. 2007. "UK Stock Returns and the Impact of Domestic Monetary Policy Shocks." *Journal of Business Finance & Accounting*, 34 (5-6): 872–888.

Breedon, Francis and Jagjit S. Chadha. 1997. "The Information Content of the Inflation Term Structure." Bank of England Working Paper No. 75.

Bryant, William D. and Roselyne Joyeux. 2010. "Interest Linkages Between the US, UK and German Interest Rates: Should the UK Join the European Monetary Union?" *International Review of Applied Economics*, 24 (6): 633–647.

Campbell, John and Robert Shiller. 1991. "Yield Spread and Interest Rates: A Bird's Eye View." *Review of Economic Studies*, 58 (3): 495–514.

Campbell, John, Andrew Lo, and Craig MacKinlay. 1997. *The Econometrics of Financial Markets*, Princeton University Press Princeton NJ.

Chuliá, Helena, Martin Martens, and Dick Van Dijk. 2007. "The Effects of Federal Funds Target Rate Changes on S&P100 Stock Returns Volatilities and Correlations." ERIM Report Series Reference No. ERS-2007-066-F&A.

Clarida, Richard, Jordi Gali, and Mark Gertler. 1999. "The Science of Monetary Policy: A New Keynesian Perspective." *Journal of Economic Literature*, 37 (4): 1661–1707.

Clarida, Richard, Jordi Gali, and Mark Gertler. 2000. "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory." *Quarterly Journal of Economics*, 115 (1): 147–180.

Clark, Peter K. 1987. "The Cyclical Component of U.S. Economic Activity." *The Quarterly Journal of Economics*, 102 (4): 797–814.

Cochrane, John and Monika Piazzesi. 2002. "The Fed and the Interest Rates − A High Frequency Identification." *The American Economic Review*, 92 (2): 90–95.

Cochrane, John and Monika Piazzesi. 2005. "Bond Risk Premia." *The American Economic Review*, 95 (1): 138–160.

Cogley, Timothy and Argia M. Sbordone. 2008. "Trend Inflation, Indexation, and Inflation Persistence in the New Keynesian Phillips Curve." *The American Economic Review*, 98 (5): 2101–2126.

Cook, Timothy and Thomas Hahn. 1989. "The effect of Changes in the Federal Funds Rate Target on Market Interest Rates in the 1970s." *Journal of Monetary Economics*, 24 (3): 331–351.

Cooley, Thomas and Edward Prescott. 1976. "Estimation in the Presence of Stochastic Parameter Variation." *Econometrica*, 44 (1): 167–184.

Deacon, Mark and Andrew Derry. 1994. "Deriving Estimates of Inflation Expectations from the Prices of UK Government Bonds." Bank of England Working Paper No. 23.

Dominguez, Emilio and Alfonso Novales Cinca. 2002. "Can Forward Rates be Used to Improve Interest Rate Forecasts?" *Applied Financial Economics*, 12 (7): 493–504.

Ehrmann, Michael, Marcel Fratzscher, and Roberto Rigobon. 2011. "Stocks, Bonds, Money Markets and Exchange Rates: Measuring International Financial Transmission." *Journal of Applied Econometrics*, 26 (6): 948–974.

Ehrmann, Micheal and Marcel Fratzscher. 2009. "Global Financial Transmission of Monetary Policy Shocks." *Oxford Bulletin of Economics and Statistics*, 71 (6): 739–759.

Ellingsen, Tore and Ulf Soderstrom. 2001. "Monetary Policy and Market Interest Rates." *The American Economic Review*, 91 (5): 1594–1607.

Engle, Robert and Clive W.J. Granger. 1987. "Cointegration and Error Correction: Representation, Estimation and Testing." *Econometrica*, 55 (2): 251–275.

Engsted, Tom and Carsten Tanggaard. 1994. "Cointegration and the US Term Structure." *Journal of Banking & Finance*, 18 (1): 167181.

Eun, Cheol and Sangdal Shim. 1989. "International Transmission of Stock Market Movements." *Journal of Financial and Quantitative Analysis*, 24 (2): 241–256.

Fama, Eugene F and Robert R Bliss. 1987. "The Information in Long-Maturity Forward Rates." *The American Economic Review*, 77 (4): 680–692.

Faust, Jon, John Rogers, Eric Swanson, and Jonathan Wright. 2004. "Identifying VARs Based on High-Frequency Futures Data." *Journal of Monetary Economics*, 51 (6): 1107–1131.

Flannery, Mark J. and Aris A. Protopapadakis. 2002. "Macroeconomic Factors Do Influence Aggregate Stock Returns." *Review of Financial Studies*, 15 (3): 751–782.

Fuhrer, Jeffrey C. 1997. "The (Un)importance of Forward-Looking Behavior in Price Specifications." *Journal of Money, Credit, and Banking*, 29 (3): 338–350.

Fuhrer, Jeffrey C. and George R. Moore. 1995. "Inflation Persistence." *Quarterly Journal of Economics*, 110 (1): 127–159.

Gali, Jordi. 2002. "New Perspectives on Monetary Policy, Inflation, and the Business Cycle." NBER Working Paper No. 8767.

Gali, Jordi and Mark Gertler. 1999. "Inflation Dynamics: A Structural Econometric Analysis." *Journal of Monetary Economics*, 44 (3): 195–222.

Gali, Jordi, Mark Gertler, and David López-Salido. 2001. "European Inflation Dynamics." *European Economic Review*, 45 (7): 1237–1270.

Gali, Jordi, Mark Gertler, and David López-Salido. 2005. "Robustness of the Estimates of the Hybrid New Keynesian Phillips Curve." *Journal of Monetary Economics*, 52 (6): 1107–1118.

Gulen, Huseyin, Yuhang Xing, and Lu Zhang. 2011. "Value Versus Growth: Time-Varying Expected Stock Returns." *Financial Management*, 40 (2): 381–407.

Gulyàs, Erika and Richard Startz. "The Phillips Curve: Forward-Looking Behavior and the Inflation Premium." 2006. Working Paper.

Gürkaynak, Refet, Brian Sack, and Eric Swanson. 2005. "Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements." *International Journal of Central Banking*, 1, 55–93.

Gürkaynak, Refet, Brian Sack, and Eric Swanson. 2007. "Market-Based Measures of Monetary Policy Expectations." *Journal of Business & Economic Statistics*, 15 (1): 45–58.

Gürkaynak, Refet, Brian Sack, and Jonathan H. Wright. 2010. "The TIPS Yield Curve and Inflation Compensation." *American Economic Journal: Macroeconomics*, 2 (1): 70–92.

Hall, Anthony, Heather M. Anderson, and Clive W.J. Granger. 1992. "A Cointegration Analysis of Treasury Bill Yields." *The Review of Economics and Statistics*, 74 (1): 116–126.

Hansen, Bruce E. 2000. "Sample Splitting and Threshold Estimation." *Econometrica*, 68 (3): 575–603.

Hardouvelis, Gikas A. 1988. "The Predictive Power of the Term Structure During Recent Monetary Regimes." *The Journal of Finance*, 43 (2): 339–356.

Harvey, Andrew C. 1985. "Trends and Cycles in Macroeconomic Time Series." *Journal of Business & Economic Statistics*, 3 (3): 216–227.

Harvey, Andrew C., Esther Ruiz, and Enrique Sentana. 1992. "Unobserved Component Time Series Models with ARCH Disturbances." *Journal of Econometrics*, 52 (1-2): 129–157.

Hausman, Joshua and Jon Wongswan. 2011. "Global Asset Prices and FOMC Announcements." *Journal of International Money and Finance*, 30 (3): 547–571.

Jondeau, Eric and Hervé Le Bihan. 2005. "Testing for the New Keynesian Phillips Curve. Additional international evidence." *Economic Modelling*, 22 (3): 521–550.

Joyce, Michael, Peter Lildholdt, and Steffen Sorensen. 2010. "Extracting Inflation Expectations and Inflation Risk Premia from the Term Structure: A Joint Model of the UK Nominal and Real Yield Curves." *Journal of Banking & Finance*, 34 (2): 281–294.

Kara, Amit and Edward Nelson. 2003. "The Exchange Rate and Inflation in the UK." *Scottish Journal of Political Economy*, 50 (5): 585–608.

Kim, Chang-Jin and Charles R. Nelson. 1998. *State-Space Models with Regime Switching*, MIT Press.

Kim, Chang-Jin and Charles R. Nelson. 2006. "Estimation of a Forward-Looking Monetary Policy Rule: A Time-Varying Parameter Model Using Ex Post Data." *Journal of Monetary Economics*, 53 (8): 1949–1966.

Kim, Chang-Jin, Pym Manopimoke, and Charles Nelson. "Trend Inflation and the New Keynesian Phillips Curve." 2011. Working Paper.

Kim, Suk-Joong and Do Quoc Tho Nguyen. 2009. "The Spillover Effects of Target Interest Rate News from the U.S. Fed and the European Central Bank on the Asia-Pacific Stock Markets." *Journal of International Financial Markets, Institutions & Money*, 19 (3): 415–431.

Kizys, Renatas and Christian Pierdzioch. 2009. "Changes in the International Comovement of Stock Returns and Asymmetric Macroeconomic Shocks." *Journal of International Financial Markets, Institutions & Money*, 19 (2): 289–305.

Krueger, Joel T. and Kenneth Kuttner. 1996. "The Fed funds futures rate as a predictor of Federal Reserve policy." *The Journal of Futures Markets*, 16 (8): 865–865.

Kuttner, Kenneth. 2001. "Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market." *Journal of Monetary Economics*, 47 (3): 523–544.

Lardic, Sandrine and Valerie Mignon. 2004. "Fractional cointegration and the Term Structure." *Empirical Economics*, 29 (4): 723–736.

Lee, Jaejoon and Charles R. Nelson. 2007. "Expectation Horizon and the Phillips Curve: The Solution to an Empirical Puzzle." *Journal of Applied Econometrics*, 22 (1): 161–178.

Lindé, Jesper. 2005. "Estimating New-Keynesian Phillips Curves: A Full Information Maximum Likelihood Approach." *Journal of Monetary Economics*, 52 (6): 1135–1149.

Lindenberg, Nannette and Frank Westermann. 2012. "Common Trends and Common Cycles Among Interest Rates of the G7-Countries." *Journal of Macroeconomics*, 34 (4): 1125–1140.

Ma, Jun, Charles Nelson, and Richard Startz. 2007. "Spurious Inference in the GARCH (1,1) Model When It Is Weakly Identified." *Studies in Nonlinear Dynamics & Econometrics*, 11 (1): 1–27.

Mavroeidis, Sophocles. 2005. "Identification Issues in Forward-Looking Models Estimated by GMM, with an Application to the Phillips Curve." *The Journal of Money, Credit and Banking*, 37 (3): 421–448.

McQueen, Grant and V. Vance Roley. 1993. "Stock Prices, News, and Business Conditions." *Review of Financial Studies*, 6 (3): 683–707.

Mihailov, Alexander, Fabio Rumler, and Johann Scharler. 2011. "Inflation Dynamics in the New EU Member States: How Relevant Are External Factors?" *Review of International Economics*, 19 (1): 65–76.

Miron, Jeffery A. 1991. "Comment on 'Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots' by John Y. Campbell and Pierre Perron." NBER Macroeconomics Annual 1991, Blanchard O. and Fischer S. (eds), 211-217.

Morana, Claudio and Andrea Beltratti. 2008. "Comovements in International Stock Markets." *Journal of International Financial Markets, Institutions & Money*, 18 (1): 31–45.

Morley, James. 2002. "A State-Space Approach to Calculating the Beveridge-Nelson Decomposition." *Economic Letters*, 75 (1): 123–127.

Morley, James, Charles R. Nelson, and Eric Zivot. 2003. "Why Are the Beveridge-Nelson and Unobserved-Components Decompositions of GDP So Different?" *The Review of Economics and Statistics*, 85 (2): 235–243.

Nason, James M. and Gregor W. Smith. 2008. "Identifying the New Keynesian Phillips Curve." *Journal of Applied Econometrics*, 23 (5): 525–551.

Neiss, Katharine S. and Edward Nelson. 2005. "Inflation Dynamics, Marginal Cost, and the Output Gap: Evidence from Three Countries." *Journal of Money, Credit and Banking*, 37 (6): 1019–1045.

Nelson, Charles R and Richard Startz. 2007. "The Zero-Information-Limitcondition and Spurious Inference in Weakly Identified Models." *Journal of Econometrics*, 138 (1): 47–62.

Nelson, Edward and Kalin Nikolov. 2004. "Monetary policy and Stagflation in the UK." *Journal of Money, Credit and Banking*, 36 (3): 293–318.

Pflueger, Carolin E. and Luis M. Viceira. 2011. "An Empirical Decomposition of Risk and Liquidity in Nominal and Inflation-Indexed Government Bonds." NBER Working Paper No. 16892.

Piazzesi, Monika and Eric Swanson. 2008. "Futures Prices as Risk-Adjusted Forecasts of Monetary Policy." *Journal of Monetary Economics*, 55 (4): 677–691.

Poole, William and Robert Rasche. 2000. "Perfecting the Market's Knowledge of Monetary Policy." *Journal of Financial Services Research*, 18 (2-3): 255–298.

Ribeiro, Ruy and Pietro Veronesi. 2002. "Excess Comovement of International Stock Markets in Bad Times: A Rational Expectations Equilibrium Model." Manuscript, University of Chicago.

Rigobon, Roberto and Brian Sack. 2004. "The Impact of Monetary Policy on Asset Prices." *Journal of Monetary Economics*, 51 (8): 1553–1573.

Roberts, John M. 2005. "How Well Does the New Keynesian Sticky-Price Model Fit the Data?" *Contributions to Macroeconomics*, 5 (1): 1–37.

Robertson, John C. and Daniel L. Thornton. 1997. "Using Federal Funds Futures Rates to Predict Federal Reserve Actions." Federal Reserve Bank of St. Louis Review, November/December, 45-53.

Romer, Christina D. and David H. Romer. 2000. "Federal Reserve Information and the Behavior of Interest Rates." *American Economic Review*, 90 (3): 429–457.

Rudebusch, Glenn D. 2002. "Assessing Nominal Income Rules for Monetary Policy with Model and Data Uncertainty." *The Economic Journal*, 112 (479): 402–432.

Sarno, Lucio and Daniel L. Thornton. 2003. "The Dynamic Relationship Between the Federal Funds Rate and the Treasury Bill Rate: An Empirical Investigation." *Journal of Banking & Finance*, 27 (6): 1079–1110.

Sarno, Lucio, Daniel L. Thornton, and Giorgio Valente. 2007. "The Empirical Failure of the Expectations Hypothesis of the Term Structure of Bond Yields." *Journal of Financial and Quantitative Analysis*, 42 (1): 81–100.

Sbordone, Argia M. 2002. "Prices and Unit Labor Costs: A New Test of Price Stickiness." *Journal of Monetary Economics*, 49 (2): 265–292.

Sbordone, Argia M. 2003. "A Limited Information Approach to the Simultaneous Estimation of Wage and Price Dynamics." Working Paper, Rutgers University.

Sbordone, Argia M. 2005. "Do Expected Future Marginal Costs Drive Inflation Dynamics?" *Journal of Monetary Economics*, 52 (6): 1183–1197.

Scholtes, Cedric. 2005. "On Market-Based Measures of Inflation Expectations." *Bank of England Quarterly Bulletin.*

Schwert, William. 1989. "Why Does Stock Market Volatility Change Over Time?" *The Journal of Finance*, 44 (5): 1115–1153.

Shen, Pu and Jonathan Corning. 2002. "Can TIPS Help Identify Long-Term Inflation Expectations?" *Economic Review*, Federal Reserve Bank of Kansas City, 61–87.

Simon, David. 1990. "Expectations and the Treasury Bill-Federal Funds Rate Spread over Recent Monetary Policy Regimes." *The Journal of Finance*, 45 (2): 567–577.

Sims, Christopher. 1999. "Drifts and Breaks in Monetary Policy." Manuscript, Princeton University.

Sims, Christopher and Tao Zha. 2002. "Macroeconomic Switching." Manuscript, Princeton University.

Smets, Frank and Raf Wouters. 2003. "An estimated dynamic stochastic general equilibrium model of the Euro area." *Journal of the European Economic Association*, 1 (5): 1123–1175.

Stock, James H. and Mark W. Watson. 1988. "Testing for Common Trends." *Journal of the American Statistical Association*, 83 (104): 1097–1107.

Stock, James H. and Mark W. Watson. 2002. "Has the Business Cycle Changed and Why?" Working Paper 9127, National Bureau of Economic Research.

Swiston, Andrew. 2007. "Where Have the Monetary Surprises Gone? The Effects of FOMC Statements." IMF Working paper.

Watson, Mark W. 1986. "Univariate Detrending Methods with Stochastic Trends." *Journal of Monetary Economics*, 18 (1): 49–75.

Wongswan, Jon. 2009. "The Response of Global Equity Indexes to U.S. Monetary policy Announcements." *Journal of International Money and Finance*, 28 (2): 344–365.

Zhang, Hua. 1993. "Treasury Yield Curves and Cointegration." *Applied Economics*, 25 (3): 361–367.

Zhou, Su. 2007. "The Dynamic Relationship Between the Federal Funds Rate and the Eurodollar Rates Under Interest-Rate Targeting." *Journal of Economic Studies*, 34 (2): 90–102.

Appendix A: State Space Representation of Unobserved Component Model

Measurement Equation

$$\begin{bmatrix} \pi_t \\ \pi_t^B \end{bmatrix} = \begin{bmatrix} \alpha \\ 0 \end{bmatrix} + \begin{bmatrix} \gamma^b & \delta \\ 0 & 0 \end{bmatrix} \begin{bmatrix} \pi_{t-1} \\ g_t \end{bmatrix} + \begin{bmatrix} \gamma^f & 1 & 0 & 0 \\ 1 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \pi_t^e \\ \epsilon_{\pi,t} \\ V_t \\ V_{t-1} \end{bmatrix}$$

$$y_t = \mu + Az_t + H\beta_t$$

Transition Equation

$$\begin{bmatrix} \pi_t^e \\ \epsilon_{\pi,t} \\ V_t \\ V_{t-1} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & \phi_1 & \phi_2 \\ 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \pi_{t-1}^e \\ \epsilon_{\pi,t-1} \\ V_{t-1} \\ V_{t-2} \end{bmatrix} + \begin{bmatrix} \epsilon_{\pi^e,t} \\ \epsilon_{\pi,t} \\ \epsilon_{v,t} \\ 0 \end{bmatrix}$$

$$\beta_t = F\beta_{t-1} + U_t, \quad U_t \sim i.i.d.N(0, Q)$$

$$\begin{bmatrix} \epsilon_{\pi^e,t} \\ \epsilon_{\pi,t} \\ \epsilon_{v,t} \\ 0 \end{bmatrix} \sim i.i.d.N \left(\begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_{\pi^e}^2 & 0 & \rho_{\pi^e v}\sigma_{\pi^e}\sigma_v & 0 \\ 0 & \sigma_\pi^2 & 0 & 0 \\ \rho_{\pi^e v}\sigma_{\pi^e}\sigma_v & 0 & \sigma_v^2 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix} \right)$$

Appendix B: State Space Representation of Multivariate Beveridge-Nelson Trend and Cycle

Based on Cochrane (1994) and Morley (2002) we present a state-space representation of the multivariate Beveridge-Nelson trend and cycle. We have the following VECM equations

$$\Delta f_t = \gamma_{10} + \gamma_{11}^f \Delta f_{t-1} + \gamma_{12}^r \Delta r_{t-1} + \cdots + \gamma_{15}^f \Delta f_{t-3} + \gamma_{16}^r \Delta r_{t-3} + \alpha^f \beta' Y_{t-1} + e_{ft}$$

$$\Delta r_t = \gamma_{20} + \gamma_{21}^f \Delta f_{t-1} + \gamma_{22}^r \Delta r_{t-1} + \cdots + \gamma_{25}^f \Delta f_{t-3} + \gamma_{26}^r \Delta r_{t-3} + \alpha^r \beta' Y_{t-1} + e_{rt}$$

The Beveridge-Nelson cycle is defined as

$$Y_t^c = -[E(\Delta Y_{t+1}|I_t) + E(\Delta Y_{t+2}|I_t) + ... + E(\Delta Y_{t+k}|I_t) + ...]$$

where I_t is the information available at time t. The state space representation of the

above model is as follows:

$$\begin{bmatrix} \Delta f_t^* \\ \Delta r_t^* \\ \Delta f_{t-1}^* \\ \Delta r_{t-1}^* \\ \Delta f_{t-2}^* \\ \Delta r_{t-2}^* \\ \beta' z_t \end{bmatrix} = \begin{bmatrix} \gamma_{11}^f & \gamma_{12}^r & \gamma_{13}^f & \gamma_{14}^r & \gamma_{15}^f & \gamma_{16}^r & \alpha^f \\ \gamma_{21}^f & \gamma_{22}^r & \gamma_{23}^f & \gamma_{24}^r & \gamma_{25}^f & \gamma_{26}^r & \alpha^r \\ 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ \gamma_{71}^f & \gamma_{72}^r & \gamma_{73}^f & \gamma_{74}^r & \gamma_{75}^f & \gamma_{76}^r & 1+\alpha^f-\alpha^r \end{bmatrix} \begin{bmatrix} \Delta f_{t-1}^* \\ \Delta r_{t-1}^* \\ \Delta f_{t-2}^* \\ \Delta r_{t-2}^* \\ \Delta f_{t-3}^* \\ \Delta r_{t-3}^* \\ \beta' z_{t-1} \end{bmatrix} + \begin{bmatrix} e_{ft} \\ e_{rt} \\ 0 \\ 0 \\ 0 \\ 0 \\ e_{zt} \end{bmatrix}$$

where $\gamma_{71}^f = \gamma_{11}^f - \beta\gamma_{21}^f, \gamma_{72}^r = \gamma_{12}^r - \beta\gamma_{22}^r, \gamma_{73}^f = \gamma_{13}^f - \beta\gamma_{23}^f, \gamma_{74}^r = \gamma_{14}^f - \beta\gamma_{24}^r, \gamma_{75}^f = \gamma_{15}^f - \beta\gamma_{25}^f, \gamma_{76}^f = \gamma_{16}^r - \beta\gamma_{26}^r$ and $e_{zt} = e_{ft} - \beta e_{rt}$ and the starred letters represents the mean adjusted variable. In matrix form the state space form can be written as

$$\Delta Y_t^* = F\Delta Y_{t-1}^* + e_t^*$$

where eigenvalues of the matrix F are less than unity in modulus. Then the cycle of the i^{th} component of vector Y_t^* can be written as $(i,i)^{th}$ element of the matrix $-(F + F^2 + F^3 + - - -) * \Delta Y_t^*$ which is equivalent of $(i,i)^{th}$ element of matrix $-F(I-F)^{-1} * \Delta Y_t^*$. The trend component is computed by substracting the cyclical component from the corresponding variable.

Curriculum Vitae

HARDIK A. MARFATIA

PLACE of BIRTH: Mumbai, India.

EDUCATION

PhD. Economics, University of Wisconsin-Milwaukee, 2013.
M.A. Economics, University of Wisconsin-Milwaukee, 2009.
M.A. Economics, University of Mumbai, India, 2005.
Accounts Tech., Institute of Chartered Accountants of India, New Delhi, 1998.
B. Commerce, University of Mumbai, India, 1998.

RESEARCH and TEACHING INTERESTS

Macroeconomics, Monetary Economics, Applied Econometrics, Financial Economics, Business Statistics and Managerial Economics.

AWARDS

Outstanding Tutor award – PASS (TARC), Univ. of Wisconsin-Milwaukee, USA.
Chancellor's Graduate School Award (Fellowship), 2009-2010.
First prize for top performance in Economics in First Year B.Commerce.
Award for best presentation on BASEL Core Principles in M.A.II. at University of Mumbai, India.