STOCK MARKET VOLATILITY AND MONETARY POLICY

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Abstract

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This thesis comprises three essays. The first essay examines the effect of federal funds rate surprises on implied stock market volatility using U.S. data. While volatility is measured using two popular implied volatility indices (VIX and VXO indexes), different techniques are employed to measure federal funds rate surprises from federal funds futures data at the daily and monthly frequencies. We find that the surprises significantly increase volatility, even when timing uncertainty is accounted for. Consistent with the efficient markets hypothesis, we find that the expected component of a target rate change; as well as the target rate change itself, do not significantly affect volatility. Nonlinearities and asymmetries are explored in the response of volatility to the direction of the rate change and the sign of the surprise. The evidence of asymmetries and nonlinearities is found to be weak.

The second essay investigates the dynamic response of U.S. stock market variables to monetary policy shocks and the transmission of monetary policy shocks to the stock market using vector autoregressive models. We find that volatility is increased and excess returns are decreased contemporaneously due to a monetary policy shock but

that the persistence of the effect depends on the model used. A daily analysis using conditional heteroskedasticity models confirms the results found with vector autoregressive models.

The third essay uses Canadian data to examine risk premiums and predictability in futures contracts (BAX futures) on short-term Canadian interest rates (Bankers' Acceptances). While evidence for a constant risk premium is found, the predictive regressions employed only uncover weak signs of predictability (and time-varying risk premiums) in returns on BAX futures. This result is confirmed by forecast efficiency regressions. Lastly, out-of-sample forecasting of Bankers' Acceptances returns is undertaken. Forecasting results reveal the superior predictive ability of the model exploiting the restrictions of economic theory in comparison to random walk, autoregressive and error correction models.

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Chapter 1 Introduction

This thesis centers on the response of stock market returns and volatility to monetary policy surprises (shocks). While the main subject of this thesis is the interaction between the stock market and monetary policy, the three essays that comprise it closely examine the properties and uses of futures contracts. In the first two essays, the response of various measures (implied, realized and conditional) of stock market volatility to monetary policy shocks (or surprises) is studied. The response of volatility to monetary policy shocks is studied within a regression framework (first essay) and a dynamic framework (second essay) that sheds more light on the joint dynamics of several stock market variables.

A recurring theme throughout the thesis revolves around gauging monetary policy expectations from the stock market via financial futures. Futures market efficiency is the basic requirement allowing for the use of financial futures contracts to measure monetary policy expectations. An efficient market incorporates all relevant and publicly available information into futures prices. Efficiency, in turn, allows for measuring interest rate and monetary policy expectations in an effective manner. The first two essays employ a popular short-term U.S. interest rate future contract, known as 30 day interest rate futures or simply as federal funds futures, written on the federal funds rate (the monetary policy rate set by the Federal Reserve) to measure and identify monetary policy shocks.

The third essay centers on examining efficiency and unbiasdness in one of the most actively traded Canadian short-term interest rate futures contracts: BAX futures. The results indicate that efficiency holds for BAX futures. Naturally, this allows for the

BAX contract's use as a market measure of short-term interest rate expectations in Canada analogously to the use of federal funds futures as a market gauge of U.S. interest rate and monetary policy expectations. Hence, the three essays clearly relate through their interest in the properties and uses of financial futures for measuring short-term interest rate and monetary policy expectations.

The following provides a more detailed outline of the methodology and results found in this thesis. In the second chapter (first essay), the effects of expected and surprise elements in federal funds target rate changes on implied volatility are examined. While volatility is measured using the VIX and VXO indexes (two popular U.S. implied volatility indices), we use different techniques to measure federal funds rate surprises from federal funds futures data at the daily and monthly frequencies. We find that the surprises significantly increase volatility, even when timing uncertainty is accounted for. Consistent with the efficient market hypothesis, we find that the expected component of a target rate change; as well as the target rate change itself, do not significantly affect volatility. Monthly results reveal that an increase in volatility is due to surprises, although the significance of this effect decreases once account is taken of macroeconomic factors. Macroeconomic factors, such as industrial production growth and inflation, are found to affect volatility. Nonlinearities and asymmetries are explored in the response of volatility to the direction of the rate change and the sign of the surprise. The evidence of asymmetries and nonlinearities is found to be weak.

In the third chapter (second essay), we investigate the effect of monetary policy shocks on stock market volatility. In the first stage, a monthly recursive vector autoregression (VAR) is used to identify monetary policy shocks and measure their effect on stock market volatility and excess returns. Secondly, federal funds futures data are incorporated directly into VARs to assess any changes in the dynamic responses of stock market volatility and excess returns to a monetary policy shock. Thirdly, monthly monetary policy shocks, identified using federal funds futures data, are used to examine the transmission mechanism through which monetary policy shocks affect stock market volatility. Namely, monetary policy shocks derived from federal funds futures are introduced as an exogenous variable in a vector autoregression including five financial variables: excess returns, the real interest rate, the change in the bill rate, the dividend yield and volatility. It is found that volatility is increased and excess returns are decreased contemporaneously due to a monetary policy shock, but that the persistence of the effect depends on the model used. Using an EGARCH model, daily monetary policy shocks derived from the term structure of federal funds futures data are found to decrease returns and increase volatility. We relate the effect of monetary policy shocks on stock market volatility to the leverage effect.

The fourth chapter (third essay) provides an analysis of risk premiums in the Canadian Bankers' Acceptances futures (BAX) market. While evidence for a constant risk premium is found, the predictive regression setting employed only uncovers weak signs of predictability in excess and holding-period returns on BAX futures. Lack of predictability in futures returns indicates the absence of time-varying risk premiums. Forecast efficiency regressions are employed to study unbiasdness and efficiency in the BAX market and it is found that efficiency cannot be rejected. Out-of-sample forecasting of spot (Bankers' Acceptances) returns demonstrates the superior predictive ability of the models that exploit the unbiasdness restriction in comparison to random walk, autoregressive and error correction models.

The thesis makes the following contributions to the literature. The first essay extends the literature by being the first to study the effect of federal funds surprises on implied volatility at both the daily and monthly frequencies. Furthermore, it is the first to provide an account of nonlinearity and asymmetry in the response of volatility to federal funds rate surprises. The second essay makes a contribution to the literature by assessing the dynamic effect of monetary policy shocks on realized stock market volatility and studying the transmission mechanism of monetary policy shocks to the stock market. Moreover, it provides an original daily analysis of the effect of surprises computed from the term structure of federal funds futures data on conditional stock market volatility. The third essay extends the literature by providing a detailed account of the presence, magnitude and determinants of risk premiums in Canadian BAX contracts. The out-of-sample forecasting results also add to the literature by providing strong evidence of predictability in financial returns.

Chapter 2 Stock market volatility, federal funds rate surprises and economic factors: what drives volatility?

The effect of monetary policy on the securities' markets has been of central importance to investors, policymakers, the financial press and academics. Central bankers have long debated the effects of monetary policy on financial variables. While most policymakers argued that stock markets do respond to monetary policy announcements and actions, opinions differed with regards to whether central banks should respond to asset price fluctuations, market turbulence or perceived stock market bubbles. Analysts and pundits point to the pronounced reaction of the stock market to news regarding monetary policy. Stock markets respond, according to analysts and the financial press, to a multitude of monetary policy related announcements such as regular meetings of the Federal Open Market Committee (FOMC), wording of FOMC statements, changes in the Fed leadership or to changes in the stance and direction of monetary policy.

While the financial press allocates significant resources to the collection and analysis of news relating to monetary policy, financial institutions assign analysts, referred to as "Fed watchers", whose role centers on inferring and forecasting the stance of monetary policy. According to Greenspan (2007), the financial press's interest in news regarding monetary policy prompted a major financial news network to devise a gimmick called the "briefcase indicator" in order to gain insight into possible monetary policy actions. More recently, the change in the Fed leadership from Greenspan to Bernanke led to extensive reporting and analysis by major financial magazines regarding the perceived differences and merits of each Chairman's monetary policy emphasis. The direction and magnitude of the Fed's next move or news about changing monetary policy goals or leadership is consequently regarded as essential information to rational investors and is closely monitored by financial institutions and the press. Underlying such attention is a maintained belief that monetary policy does affect the return on various securities and hence has effects on investors' portfolio returns.

The inherent relationship among short term interest rates, on the one hand, and fixed income securities' returns, on the other, led early studies to examine the effects of monetary policy actions on bond returns. Although a voluminous literature examines the effects of monetary policy on real activity or the effect of macroeconomic variables on the equity premium, financial economists only recently examined the direct effect of monetary policy actions and announcements on stock market returns and volatilities. This relatively recent academic interest led to a rapidly expanding literature that reached interesting conclusions from a practitioner, academic and policy making perspective.

Monetary policy exerts an effect on the stock market through various channels. First, a change in the federal funds rate is closely associated with changes in various short-term interest rates. This, in turn, influences the discount rate used to value the cash flows of different equities and may thus increase or decrease returns and volatility. A second channel through which monetary policy exerts an effect on the stock market is through financial leverage: each rate move by the Fed varies the cost for firms to finance their activities through issuing debt. Both of these channels can impact stock market returns and volatilities. Bernanke and Kuttner (2005) describe these channels while Mishkin (2007) gives a textbook account of both the role of financial markets in the transmission of monetary policy and the effect of monetary policy on stock markets.

In this paper, we examine the effects of the expected and surprise elements of federal fund changes on implied volatility, as measured by the VIX and VXO indexes, using different techniques to measure surprises at the daily and monthly frequency. We capitalize on Gospodinov, Gavala and Jiang's (2006) insight and the availability of implied volatility indices to treat volatility as an observable rather than a latent process. This insight allows the estimation and forecasting of conditional mean models for volatility. Further, nonlinearities and asymmetries in the response of volatility to target rate surprises are investigated at the daily frequency. In addition to federal funds rate surprises, the paper also considers the effect of macroeconomic factors on implied volatility at the monthly frequency.

The paper makes several contributions to the literature. Firstly, at the time of writing, this is the first paper to examine the effects of expected and surprise components of federal funds target moves on implied volatility. Examining the effect of federal funds rate surprises on volatility is a recent topic, with a single manuscript by Chulia-Soler, Martens and Van Dijk (2007). Secondly, this paper examines nonlinearities in the response of implied volatility target rate changes, a subject that has not been developed in earlier studies. Third, the paper measures the effect of timing surprises on volatility, a subject that remained unexplored until now. Fourthly, this is the first paper to use monthly data to investigate the robustness, magnitude and direction of the relationship between implied volatility and target rate surprises.

Examining and modeling implied volatility could be more insightful than examining realized volatility for several reasons. First, implied volatility as measured by the VIX or VXO index, can be interpreted as market uncertainty as noted in Whaley

(2000). It thus carries more informational content than other measures of volatility (including realized volatility used by Chulia-Soler, Martens and Van Dijk (2007)) as noted in the survey article of Granger and Poon (2003). Second, implied volatility measured by the VIX or VXO index is observable and market traded, while realized volatility is not. This distinguishing feature of implied volatility makes it of natural practical interest for investors seeking to treat volatility as an underlying asset rather than an unobservable measure. Third, the steady growth of trading volume in options and futures written on implied volatility indices makes examining the factors affecting implied volatility even more timely. For example, given a certain (even directional) forecast of implied volatility, Hull (2008) provides a textbook treatment of the different trading strategies involving put and call options (for example, straddles and strips) which can be used by investors. Since options written on the S&P 500 and S&P 100 are widely available, such trading strategies can obviously be used by profit seeking investors.

In contrast to the recent manuscript of Chulia-Soler, Marterns and Van Dijk (2007) which examines the effect of surprise and expected components of target rate moves on realized volatility, this paper uses a different and arguably more relevant measure of volatility, different data frequencies, different specifications of shocks (including timing shocks not discussed in Chulia-Soler, Marterns and Van Dijk (2007)), as well as a considerably larger sample. We can thus investigate more fully the robustness of the conclusions relating federal funds target rate surprises to volatility. Moreover, Chulia-Soler, Marterns and Van Dijk (2007) use high frequency data to discern the effect of surprises on volatility. Although interesting and useful in many respects, the use of high frequency data only considers the effects of shocks on implied volatility in a very narrow time window (typically of five or ten minutes). This paper finds that the response of volatility is longer lived and argues that such a response is more informative from a trading and policy making perspective. Lastly, we offer an interpretation of the results which draw upon, and are in line with, earlier findings in the literature (Bernanke and Kuttner (2005) among others) regarding the negative effect of federal funds rate surprises on stock market returns.

In comparison to the literature measuring the effect of macroeconomic and monetary policy announcements on implied volatility (also reviewed below), this paper uses the informational content of Fed target rate announcements and captures the qualitative and quantitative effects of such announcements. The essay proceeds as follows: Section 2.2 describes the data and data sources. Section 2.3 discusses the econometric methodology used. Section 2.4 presents the results and section 2.5 summarizes the conclusions and proposes avenues for future research.

2.1 Literature Review

The natural relationship between fixed income securities and interest rates led researchers to examine this relationship. Leading among the attempts to examine the effects of monetary policy actions on securities returns was the event study regression approach of Cook and Hahn (1989). Cook and Hahn (1989) study the effect of federal funds target changes on bond returns and find a positive, statistically significant effect of target increases on bond returns of all maturities. The relationship between monetary policy actions and long-term interest rates has in turn been investigated by Roley and Sellon (1995) and Thornton (1997) without achieving a clear conclusion. The literature subsequently evolved towards studying the effects of monetary policy on stock returns and various techniques were proposed to uncover the strength and direction of the response of stock market returns to monetary policy. Thorbecke (1997) employs vector autoregressive, narrative and factor analysis techniques to conclude that monetary policy influences stock returns significantly. On the other hand, Patelis (1997) uses Fama-French long horizon regressions and vector autoregressive techniques to conclude that monetary policy variables have significant predictive power for forecasting excess stock returns. Crowder (2006) and Goto and Valkanov (2002) utilize structural vector autoregressions to uncover a significant response of excess returns to monetary policy shocks.

Researchers have also made use of the increasing availability of high frequency financial data to re-examine the significance of the reaction of stock market returns to variations in monetary policy. In the US context, D'Amico and Farka (2002) use structural vector autoregressive techniques to find that stock market returns respond negatively to monetary shocks. In a closely related paper, D'Amico and Farka (2006) use high-frequency data on returns of the S&P 500 and S&P 500 futures to identify a monetary vector autoregression and reiterate their previous findings with regard to the effect of monetary policy shocks on stock returns. Cochrane and Piazzesi (2002) examine the relationship between monetary policy shocks and interest rates while Bohl, Skilos and Sondermann (2007) analyze the effects of monetary policy surprises on European stock market returns. The central feature of the papers reviewed above is their emphasis on the effect of surprises in target rate changes (whether measured as the residual from a monetary VAR or in a high frequency setting) on returns from financial assets.

A parallel literature which examines the effects of monetary policy on stock market volatility has evolved. This literature differs from the before-mentioned papers in two respects: it examines second moments and focuses almost entirely on detecting the effects of monetary and macroeconomic policy announcement, rather than examining the effects of actual rate moves on volatility. That is, the aim of this literature has been to check whether the mere presence of a scheduled or unscheduled announcement significantly impacts volatility, with no regard to the content of the announcement which is unknown *a priori.* Ederington and Lee (1993, 1996) launched this literature by examining the effects of information releases on the implied volatility of options. Ederington and Lee (2006) distinguish between scheduled and unscheduled news announcements and maintain that since the timing, but not the informational content, of scheduled news announcements is known at the outset, implied volatility should increase prior to major economic and financial announcements and decrease thereafter. This is dubbed the "pre-announcement" effect. Donders and Vorst (1996) examine a similar announcement hypothesis using Dutch data on firm-specific implied volatility and reach very similar conclusions. In a similar spirit, Nikkinen and Salhlstrom (2004) and Chen and Clements (2007) use the Chicago Board of Exchange's (CBOE) S&P500 implied volatility index (VIX) to investigate the effect of scheduled monetary and macroeconomic news announcements on implied volatility. Both papers employ dummy variables corresponding to the days of release of major macroeconomic announcement, ranging from the release of the employment, producer or consumer price indices reports, but differ with regard to the effect of announcements on implied volatility. While Nikkinen and Sahlstrom (2004) report an increase in implied volatility prior to the announcement date and decrease following the announcement, Chen and Clements (2006) maintain that implied volatility falls on the days of the FOMC meeting with no significant movement on the days preceding and succeeding the FOMC meeting. Carr and Wu (2006) examine the behaviour of average implied volatility ten days prior to and succeeding a scheduled FOMC meeting. The authors find that average volatility is considerably higher on the ten days preceding the FOMC meeting and drops afterwards.

By using dummy variables to measure the response of implied volatility to macroeconomic and monetary announcements, the "announcement" effect literature does not fully exploit the informational content of such announcements. Bomfim (2003) notes that interest in examining the effect of news contained in an announcement, and not the mere presence of the announcement itself, should take into account the informational content of the announcement and distinguish between the surprise and expected components of each announcement. As opposed to earlier studies such as Cook and Hahn (1989), the basic premise is that in an *efficient stock market,* an actual rate move by the Fed should be decomposed into an expected and an unexpected component to uncover the effect of each. Cook and Hahn (1989) simply use the actual rate move and make no such distinction. Lombra and Kearney (2004) take into consideration Bomfim's (2003) recommendation and measure the response of the changes in the VIX index to surprise elements of employment and producer price index announcements. The authors gauge market expectations on employment and the producer price index using median forecasts of market professionals reported by Money Market Services International. They report a positive and statistically significant response of VIX to unanticipated changes in employment.

In three papers of considerable interest to this research, Kuttner (2001), Bernake and Kuttner (2005) and Chulia-Soler, Martens and Van Dijk (2007) examine the effect of expected and surprise components of federal funds target rate changes, respectively, on bond returns, stock returns and stock market volatility. In order to measure market expectations, the authors use scaled one-day changes in federal funds futures rates as first proposed by Kuttner (2001). Measuring federal funds rate surprises using futures contracts became a subject of wide research and implementation. We will detail the use of federal funds futures contracts and review the relevant literature in sections 2.2.4 and 2.2.6 of this chapter. All three papers state that the surprise component of a target rate change affects the financial variables of interest, while the expected and actual target rate changes do not. Specifically, Bernanke and Kuttner (2005) find that stock returns respond negatively and significantly to surprises while responding positively but weakly to the expected component of target rate moves. Chulia-Soler, Marten and van Dijk (2007) extend the analysis of the effect of expected monetary policy and monetary policy surprises on volatility in a high-frequency data setting. The authors use realized volatility and realized correlations at market and sector levels and conclude that surprise movements in the funds rate affect volatility and returns, while actual rate changes do not. They further develop threshold autoregressive models to examine nonlinearities in high frequency returns and volatilities and deduce, contrary to Bernanke and Kuttner (2005), that there is evidence of nonlinearity in returns, but not in volatility.

2.2 Data

2.2.1 Daily event-study data

The daily data used in this paper consist of observations on (i) the level of VIX (VXO) (ii) the level of the VIX (VXO) for the previous day, (iii) the actual target rate change, and (iv) the expected component of a target rate change as well as the surprise component of a target rate change for all dates on which there is a meeting of the Federal Open Market Committe (FOMC)[1] for the period 08/02/1990 to 11/12/2007. Thus, we define "an event" as a meeting of the FOMC. Our data contain a total of 165 meetings of the FOMC of which 31 involved target rate increases, 39 involved target rate decreases and 95 saw no target rate change. The changes comprise 25 basis points, 50 basis points or 75 basis points federal funds target rate changes. We omit the observation relating to the 17/09/2001 FOMC intermeeting from our sample, as in Bernanke and Kuttner (2005) and others, as this is the first day of trading following the September 11th attacks.

The surprise and expected components of a target rate change are computed, as discussed at length in section 2.2.4, using daily data on federal funds futures prices for different maturities obtained from the website of the Commodity Research Bureau (CRB)[2]. The daily federal funds futures price data spans the 06/10/1988 to 31/12/2007 period. Daily data on the level of the VIX (VXO) index is made available from the

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[1] The dates of FOMC meetings as well as FOMC minutes and press statements can be found at the Board of Governors' website: http://www.federalreserve.gov/monetarypolicy/fomc.htm

[2] I would like to thank my supervisor for providing me with this data.

website of the Chicago Board of Exchange (CBOE)[3]. While daily data for the VXO starts on 02/01/1986, the VIX is only available starting 02/01/1990. Due to the availability of the VIX and the importance of correctly dating a target rate change (as discussed shortly), our data are confined to the 08/02/1990 to 11/12/2007 period. Summary statistics of the implied volatility data is found in table 2.1 while time series plots of our implied volatility data are reported in figures 2.1 and 2.2.

In a setting such as ours, the correct timing of the target rate move by the Fed is essential. Accordingly, a swift review of the FOMC targeting procedures is warranted. Prior to 1994, the FOMC did neither announce nor set a specific intended funds rate target. Further, the 1990-1994 period saw numerous unscheduled intermeetings of the FOMC. Stock market participants were not easily capable of discerning whether the FOMC changed its monetary policy stance. Given that the FOMC's directives are carried out by the trading desk of the New York Fed, market participants had to examine the trading desk's actions to infer the stance and direction of the FOMC move, if any. The directives of the FOMC usually became apparent to market participants when carried out the day following the FOMC decision. This pre-1994 lack of transparency in FOMC actions complicates the dating of target rate moves in our sample, especially since the trading desk sometimes implemented the FOMC's directives with a time lag. To circumvent problems associated with the dating of target rate changes, we rely on the dating widely agreed upon and reported in the literature. Specifically, Gurkaynak, Sack and Swanson (2004) list all the FOMC meeting dates and actions from 1990 to 2004 in their data appendix. Their dating scheme is nearly identical to that reported in the

[3] Daily data on the VIX (VXO) can be downloaded from: http://www.cboe.com/micro/vix/historical.aspx

literature such as Kuttner (2001), Poole and Rasche (2000), Poole, Rasche and Thornton (2002) and Chulia-Soler, Martens and Van Dijk (2007). The uncertainty relating to the FOMC target change and the stock market's participants knowledge of it is resolved in the post-1994 period when the FOMC began announcing its actions upon making them. This drive towards greater transparency by the Fed culminated in a decision to announce a numeric value for their intended target rate starting August 1997.

We thus rely on the dating scheme of the literature for the pre-1997 period and extend it (using the FOMC statements and transcripts published on the Board of Governors' website) until the end of 2007. Our dataset containing the dates of FOMC actions and the decomposition of each target rate move into an expected and surprise component forms an extension to the datasets reported in the literature to take into account recent policy moves. A more detailed account of the FOMC's operating procedure can be found in Kuttner (2001), Bernanke and Kuttner (2005), Poole, Rasche and Thornton (2002) and Bomfim (2003).

2.2.2 Monthly Data

A monthly analysis is undertaken using a dataset from November 1988 to December 2007. These data begin in October 1988 and one observation is lost due to lags. Daily and monthly macroeconomic data are obtained from the Federal Reserve Bank of St. Louis Economic Database (FRED)[4]. The dataset consists of the monthly level of implied

[4] Macroeconomic and interest rate data from FRED's website are available at: http://research.stlouisfed.org/fred2/

volatility, the monthly measure of the expected component of a target rate change, the monthly surprise measure of a target rate change as well as macroeconomic variables such as the industrial production index, employment and inflation. Monthly interest rate data, such as the three month T-Bill rate, the yield of AAA rated corporate bonds and the yield of BAA rated corporate bonds are also used.

The construction of the monthly expected and surprise components of a target rate change is detailed in section 2.2.6 below and involves using a daily time series for the federal funds target rate. The macroeconomic and interest rate data used, either directly or indirectly (to construct a monthly surprise measure), are the following: seasonally adjusted nonfarm payrolls for all employees, seasonally adjusted consumer price index for all urban consumers (all items), the three month T-Bill rate (secondary market rate), Moody's seasoned AAA corporate bond yield, Moody's seasoned BAA corporate bond yield as well as a daily time series for the federal funds rate target[5]. The monthly value for implied volatility is taken to be the observation on the last trading day of the month. A total of 231 observations are available.

2.2.3 Implied volatility indices

The two measures of implied volatility used throughout this paper are the Chicago Board of Exchange (CBOE) VIX and VXO implied volatility indices. While the former is constructed from options written on Standard and Poor's 500 (SPX) index, the latter's

[5] The daily time series for the federal funds rate target prior to 1994 is due to Thornton (2005). After 1994, the target funds rate series is derived from FOMC statements. A full daily dataset for the federal funds target rate starting in 1954 is available from the FRED website.

construction is based on options written on Standard and Poor's 100 (OEX) index[6]. A weighted average of four American puts and calls with strike prices around the at-the-money point go into the construction of the VXO index. The index's construction is such that it has a constant thirty calendar day expiry. Gospodinov, Gavala and Jiang (2006) point to the fact that such near-the-money, close-to-maturity options are the most informative about volatility since they maximize the first derivative of the option price with respect to volatility (the "vega" of the option). Underlying the construction of the VXO index is the option pricing formula of Merton (1973) and Black and Scholes (1973). Details about the construction of the VXO index could be found in Whaley (2000) and Fleming, Ostdiek and Whaley (1995). In fact, it was Whaley (1993) who introduced the VXO index into academic studies and subsequently referred to it as "the investor fear gauge". Since its construction is based on options prices, the VXO could indeed be thought of as market uncertainty or more accurately as investors' expectations of future volatility. In their extensive survey on forecasting stock market volatility, Poon and Granger (2003) cite the construction of the VXO as "good practice" and note its superior predictive ability.

Unlike the VXO, the other CBOE volatility index we employ, namely the VIX, does not only use the at-the-money options. Rather, it is a weighted average of out-the-money, European-style puts and calls written on the S&P 500 with a wide range of strikes. The VIX's construction is model independent. Although the VIX is constructed in

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[6] CBOE undertook a change relating to the construction and nomenclature of their indices. Previously, the VIX index used to refer to options written on the S&P 100 index. Starting September 22, 2003, the VIX refers to options written on the S&P500 index and the old VIX was renamed as VXO with its construction and dissemination unchanged.

a slightly different manner than the VXO, it still retains the important features captured by the VXO and also has a constant thirty calendar days to expiry. Full details on the construction of the VIX are provided from the CBOE website. An excellent review of construction, properties and interpretation of both indices can be found in Carr and Wu (2006) who argue that the VXO index is inflated upward by construction and needs to be adjusted. Such an adjustment is undertaken in Carr and Wu (2006) and Dotsis, Psychoyios and Skiadopulos (2007)[7]. The construction of the VIX index does not suffer from such a problem and hence there is no need to adjust it (see Carr and Wu (2006)).

Time series plots and summary statistics of our implied volatility measures are given in table 2.1 and figures 2.1 and 2.2. From visual inspection as well as by observing the values of the first five autocorrelations for each of the implied volatility series, we note a high degree of persistence in implied volatility. In their examination of various implied volatility indices, Dotsis, Psychoyios and Skiadopulos (2007) find that the VIX and VXO indices are mean reverting. Because of the possible presence of a unit root in the implied volatility indices, we conduct augmented Dickey-Fuller (ADF) tests to test the null of a unit root in each of the volatility indices. Our results for these tests appear in table 2.1 and the null of a unit root is rejected for each series. Since ADF tests for a unit root are known to exhibit low power when the alternative hypothesis is near unit root behaviour, we follow Elliot, Rothenberg and Stock's (1996) testing procedure and demean the series using GLS demeaning. This procedure increases the power of the unit root test when the alternative is near-unit root behaviour. Again, test results reported in

[7] The adjustment consists of scaling down the VXO index as : $VXOA = VXO \times \sqrt{22/33}$

the last row of Table 2.1 reject the null of a unit root for both the VIX and adjusted VXO indices at the 1% level. The high persistence in the implied volatility processes has to be accounted for in our estimation methodology, however. We return to this issue in section 2.3 of the paper.

2.2.4 Gauging daily expectations from the market: federal funds futures

Federal funds futures, officially known as thirty day interest rate futures, are in essence futures contracts that settle on the average of the month's overnight federal funds rate. The contract is cash-settled daily (i.e. it is marked to market) and the initial contract size is five million dollars. As opposed to other interest rate futures contracts, default risk in federal funds futures is negligible due to cash settlement and collateral requirements. Federal funds futures contacts trade on the Chicago Board of Trade (CBOT) where contracts for several different deliveries exist, ranging from the current month to five months ahead. Contracts with even longer deliveries exist, but liquidity in those contracts drops sharply.

Krueger and Kuttner (1996) test for unbiasdness and rationality of federal funds futures in forecasting federal funds rates. The authors find that federal funds futures efficiently embody all publicly available information and conclude that the contracts can be effectively used to identify surprises in FOMC target rate moves. In two recent manuscripts, Hamilton (2008a, 2009) revisits the evidence on the efficiency of federal funds futures. The author reiterates Krueger and Kuttner's (2001) findings relating to the excellent predictive ability of federal funds futures in forecasting the federal funds rate at the daily frequency.

Measuring federal funds rate surprises (at the daily frequency) from federal funds futures contracts was first formalized in a contribution by Kuttner (2001). Although Kuttner (2001) details the construction of expected and surprise components of a target rate change using futures data, the idea of market based measures of monetary policy expectations predate his contribution. Rudebusch (1998), Brunner (2000), Carlson, McIntire and Thomson (1998), Krueger and Kuttner (1997, 1998), Evans and Kuttner (1998), Robertson and Thornton (1997) and Soderstrom (2001) all investigate the possibility of measuring monthly federal funds rate surprises from federal funds futures data but do not describe a methodology that adequately accounts for the intricacies of federal funds futures contracts such as: (i) federal funds futures contracts settlement prices are based on the average of the month's federal funds rate; (ii) federal funds futures use the effective federal funds rate rather than the target rate; (iii) the possible presence of a risk premium in the futures price.

Kuttner's (2001) procedure for extracting the expected and surprise elements of a target rate change attempts to deal with such intricacies. The basic premise of measuring the surprise element of target rate changes from futures prices is the following: since federal funds futures incorporate all relevant available information, market expectations are embodied in the current month futures rate. In line with federal funds futures market efficiency, all available information prior to the FOMC meeting would have been factored into the previous or current day's price. Thus, the one-day change in current (spot) month's futures implied rate on FOMC meeting days measures the unexpected (surprise) component of a target rate move. More formally, denote by $f_{t,d}^{0}$ the spot futures rate on day d of month t. Kuttner's (2001) proposed surprise component on day d of month t[8] (for each target rate move) based on the one-day change in the spot futures rate is given by:

$$\Delta i^{u,0} = \frac{D}{D-d}(f_{t,d}^{0} - f_{t,d-1}^{0}) \tag{2.1}$$

where D denotes the number of days in month t with d=1,...,D and t=1,2,...,12. The scaling factor preceding the difference in futures rates is used to adjust for the nature of the federal funds futures which involve averaging of the overnight funds rate. This scaling adjusts the surprise component proportionally to the number of days affected by the target rate change. Further details on the properties of federal funds futures can be found in Evans and Kuttner (1998), Kuttner (2001). We refer to this surprise measure as the current month surprise.

Kuttner's (2001) proposed method has become widely used in the literature for measuring the surprise element of a target rate move. In fact, Bernanke and Kuttner (2005), Chulia-Soler, Martens and van Dijk (2007), Gurkaynak (2005), Poole, Rasche and Thornton (2002), Faust, Swanson and Wright (2004) and Gurkaynak, Sack and Swanson (2007) among many others, all use Kuttner's (2001) proposed method.

[8] Throughout this section, we abstract from indexing the surprise and expected measures of rate changes with the day d and month t for notational simplicity. For our daily, "event-type" analysis introduced in this section, it is implied that the surprise and expected measures are for day d of month t so that d and t are omitted from the subscript.

However, measuring the surprise component of a target rate move using (2.1) still suffers from shortcomings. Namely, Hamilton (2008a) notes that when the target rate move occurs towards the end of the month, the scaling factor in (2.1) becomes excessively large and cautions against using this measure towards the end of the month. In fact, many authors [including Bernanke and Kuttner (2005), Chulia-Soler, Martens and Van Dijk (2007) and Gurkaynak, Sack and Swanson (2007) and in particular Kuttner (2001) himself] use the unscaled change in the one-month ahead futures when the target rate change occurs towards the end of the month. We construct the surprise component of a target rate change using equation (2.1), but use the unscaled change in the one-month ahead futures rate when the target change occurs during the last seven calendar days of the month. When the change in the target rate occurs on the first day of the month, the one-month futures rate from the last day of the previous month is employed instead of $f^{0}_{t,d-1}$.

Other measures of the surprise component of a target rate change have been advanced in the literature. Poole and Rasche (2000) propose using the unscaled, one-day change in the one-month ahead futures rates. Let $f^{1}_{t,d}$ denote the one-month ahead futures rate for month t and day d, the Poole and Rasche (2000) surprise is computed from:

$$\Delta i^{u,1} = f^{1}_{t,d} - f^{1}_{t,d-1} \tag{2.2}$$

The Poole and Rasche (2000) measure avoids scaling, and arguably contains more information than (2.1). Indeed, Hamilton (2008b) argues that the one and two-month-ahead federal funds futures contracts contain significantly more information than the current month contract and supports their use. This surprise measure is referred to as the one-month ahead surprise.

In another recent paper, Gurkaynak, Sack and Swanson (2007) [see also Gurkaynak (2005)] also suggest incorporating information past the current month. To that effect, they propose using an appropriately weighted difference between the current month surprise in (2.1) and the one day change in the futures rate from the month containing the next FOMC meeting (in six to eight weeks time). Let $f_{t,d}^2$ denote the futures rate from the month containing the next FOMC meeting. Gurkaynak, Sack and Swanson (2007) advocate computing the surprise component in the target rate change as:

$$\Delta i^{u,2} = \frac{D_2}{D_2 - d_2}[(f_{t,d}^2 - f_{t,d-1}^2) - \frac{d_2}{D_2}\Delta i^{u,0}] \tag{2.3}$$

where d_2 and D_2 are, respectively, the day of the next of FOMC meeting and the number of days in the month containing the next FOMC meeting. The authors argue that their measure detailed in equation (2.3) has two main advantages: it captures the surprise component of a target rate move at a longer horizon than (2.1) and is not subject to the "timing" surprise (the surprise with regard to the *timing* of the Fed's next move) which is confounded in (2.1). We also employ Gurkaynak, Sack and Swanson (2007)'s surprise measure and refer to it as the two month ahead surprise (this label is not very accurate since the dates of the next FOMC meetings are not necessarily regularly spaced, but mostly occur within six to eight weeks after the current meeting). Due to the fact that this measure assumes knowledge of the date of the next FOMC meeting, it is computed only starting from 1994 as in Gurkaynak, Sack and Swanson (2007).

The current month and one-month-ahead (2000) surprise measures computed using (2.1) and (2.2) are very highly correlated (the correlation coefficient being 0.96) while the surprise measures in (2.1) and (2.3) display a weaker correlation of 0.51. Surprise measures (2.2) and (2.3) display a correlation coefficient of 0.63. Scatter diagrams illustrating the relationship between our three surprise measures are found in figures 2.3 and 2.4.

For each of the surprise components of a target rate move in (2.1), (2.2) and (2.3), the expected component of a policy rate move can be calculated as the difference between the actual rate move by the Fed and the unexpected part of the rate change:

$$\Delta i^{e,j} = \Delta i - \Delta i^{u,j}, j = 0,1,2, \tag{2.4}$$

where Δi denotes the actual (target) target rate change by the Fed and $\Delta i^{u,j}$ denotes one of the surprise measures detailed above.

Other interest rates futures can, in principle, be used to measure the surprise component of a target rate move. For instance, Rigobon and Sack (2002) use the three-month Eurodollar rate, Cochrane and Piazzesi (2002) employ the one-month Eurodollar deposit rate, Ellingsen and Soderstrom (2004) the three month Treasury bill rates. Gurkaynak, Sack and Swanson (2007) provide a review of the different possible market based measures of monetary policy expectations which include the term federal funds rate, term Eurodollar deposit rates, Eurodollar futures rates, Treasury Bill rates and commercial paper rates and test for the ability of each in forecasting the federal funds rate. They conclude that "federal funds futures rates clearly dominate other market-based measures of monetary policy expectations at horizons out to about five months" further encouraging their use.

2.2.5 Timing surprises

The spot month surprise measure in (2.1) is a useful tool for summarizing current surprises. Nevertheless, more information can be extracted from longer maturity federal funds futures contracts. This, in turn, permits the defining of several other surprise measures. For instance, macroeconomic news may well allow forward-looking rational investors to determine, with a high level of likelihood, that the monetary authority will undertake certain actions. Market participants, however, will be unsure as to the exact timing of such actions (for example, whether a target rate change will take place at the next FOMC meeting or the one after).

The surprise with regards to the timing of a policy action has been labelled a "timing" surprise in the literature[9]. However, the literature had slightly different methods of computing this surprise. Bernanke and Kuttner (2005) define the timing surprise as the difference between the current month surprise and the change in the three-month-ahead futures rate. We opt for using Bernanke and Kuttner's (2005) timing surprise definition,

[9] Other surprise measures were also proposed in the literature. Hamilton (2008c) proposes "level", "slope" and "curvature" surprises obtainable from daily changes in one-month, two-month and three-month ahead future rates while Gurkaynak (2005) suggests defining the "slope" surprise from five-month ahead futures rates. The "slope" surprise is thought to reflect the "expected pace of interest rate changes" according to Gurkaynak (2005) and is directly related and inferable from the phrasing of an FOMC policy statement.

while noting that all definitions of timing surprises proposed in the literature are very much the same. Specifically, we define a timing surprise as the difference between the current surprise and the change in the three-month-ahead futures rates. Let $f_{t,d}^{3}$ denote the three-month-ahead futures rate on day d of month t. Then, $time_t$ is the timing surprise in day d of month t, defined as:

$$time_t = \Delta i^{u,0} - (f_{t,d}^{3} - f_{t,d-1}^{3}) \tag{2.5}$$

Thus the timing surprise measures the change between current interest rate expectations and the change in longer-dated interest rate expectations.

2.2.6 Measuring monthly expectations from federal funds futures

The analysis using surprise measures introduced above imposes the use of event-study analysis[10]. Rudebusch (1998), Brunner (2000) and more recently Bernanke and Kuttner (2005) suggest a monthly measure of the surprise and expected components of a rate change. Specifically, we use Bernanke and Kuttner's (2005) framework and define the

[10] Hamilton (2008a) develops a framework that allows for the use of daily changes in the futures implied rate series to measure surprises with no regard to the date of actual rate changes. This allows for the use of a regular time series approach but involves numerous parametric assumptions regarding the time series properties of federal funds futures contracts.

monthly surprise component as the difference between the average funds rate target for month t and the one-month-ahead futures rate on the last day of month (t-1)[11], that is:

$$\overline{\Delta i}_t^u = \frac{1}{D}\sum_{d=1}^{D} i_{t,d} - f_{t-1,D}^1 \qquad (2.6)$$

where $i_{t,d}$ denotes the federal funds target rate on day d of month t, $f_{t-1,D}^1$ the futures rate from the last day of month t-1 and D denotes the last day of month t and hence the number of days in month t. In turn, a measure of the monthly expected component of a target rate change can be computed as the difference between the futures rate on the last day of month (t-1) and the federal funds rate target on the last day on month (t-1), namely:

$$\overline{\Delta i}_t^e = f_{t-1,D}^1 - i_{t-1,D} \ . \qquad (2.7)$$

The measures suggested in (2.6) and (2.7) are constructed by averaging the federal funds target rate over all the days in a month. Such averaging is introduced because federal funds futures contracts settle on the average of the month's effective federal funds rate, and there exists no straightforward method of undoing this averaging. In contrast to daily surprises, the monthly surprises are regularly spaced and thus form typical time series data to which time series methods can be applied.

[11] Since our monthly surprises are a regular monthly time series, there is no need to index by the day of the target change.

2.2.7 The accuracy of surprise measures

Measuring federal funds rate surprises from futures data has been scrutinized in two ways. Firstly, the federal funds futures prices possibly include a risk premium that distorts the measured surprises in (2.1), (2.2) and (2.3). If the risk premium embedded in futures contracts is constant (or constant over the one-day interval of FOMC action), taking the one-day change as in (2.1), (2.2) and (2.3) would difference out any risk premium. However, Piazzesi and Swanson (2008) argue that long-horizon (four-month ahead contracts and longer) federal funds futures contracts contain time-varying risk premiums that depend on macroeconomic factors such as employment growth or Treasury yield spreads. These authors argue that such risk premiums are most pronounced at low frequencies (i.e. business-cycle frequencies).

Other research, such as the recent contributions by Hamilton (2008b) and the earlier findings of Sack (2004) and Durham (2004), point to a small (or nonexistent) time-varying risk premium in long-dated federal funds futures contracts. Sack (2004) notes that the impact of such a time-varying risk premium "is fairly limited for futures contracts with relatively short horizons, but increases as the horizon of the contract lengthens" while Durham (2004) finds a very small risk premium. Hamilton (2009) revisits Piazzesi and Swanson's (2008) evidence and notes a very small risk premium even for longer horizon (equal to or more than three month ahead) contracts. Robertson and Thornton (1997) and Poole, Rasche and Thornton (2002) also argue that the risk premium in futures contracts is very small and opt to ignore it in their analyses. Given that the measures advocated in (2.1), (2.2) and (2.3) use near-term futures contracts and a high frequency (one-day) difference in the implied futures rate, the existence of a slow moving risk premium would not impact our surprise measures (2.1), (2.2) and (2.3) as acknowledged by Piazzesi and Swanson (2008) themselves.

A second concern with using measures (2.1), (2.2) and (2.3) is the fact that federal funds futures contracts settle on the average of the month's effective federal funds rate, and not the federal funds target rate set by the FOMC. A large deviation between the effective and target federal funds rate can distort the surprise measures in (2.1), (2.2) and (2.3). Nevertheless, when the "targeting error" is a zero mean (but not necessarily i.i.d) random variable[12] and under efficiency of the futures market, Poole, Rasche and Thornton (2002) show that the surprise measures proposed in (2.1) and (2.2) are robust to deviations between the effective and the target funds rate.

The monthly surprise and expected component measures detailed in (2.7) and (2.8) above offer slight advantages over the daily surprise measured computed in (2.1) and (2.2). Bernanke and Kuttner (2005) remark that such advantages include the avoidance of any sample selection issues[13] since each month in the sample can contain a surprise policy action and any lack of action within a month can also surprise the markets. Nonetheless, the monthly shocks are more prone to the shortcomings of daily surprises discussed in this section and suffer from other limitations. On one hand, Evans and Kuttner (1998) note that such a definition of the surprises is subject to time

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[12] Previous research indicates that there is some persistence in the deviation between the effective and target funds rates. See Sarno, Thornton and Valente (2005) and Taylor (2001) for details.

[13] In our daily framework, we use all observations on which the FOMC meets regardless of whether a target rate change occurred or not. Thus, sample selection issues are avoided. Some authors condition their analyses on the presence of a target rate move, which arguably can lead to sample selection concerns.

aggregation considerations while Bernanke and Kuttner (2005) note that such time aggregation leads to a decrease in the size of the surprise measure. On the other hand, Piazzesi and Swanson (2008) argue that risk premiums in the monthly surprises are significant whereas Bernanke and Kuttner (2005) report an endogenous response to economic news. In particular, Piazzesi and Swanson (2008) find that employment growth, bond yield spreads and corporate bond spreads significantly predict excess returns on federal funds futures, while Bernanke and Kuttner (2005) report that industrial production, inflation, retail sales and employment news affect the monthly surprise measures in (2.6). In general, monthly shocks must be treated carefully in estimation, and we will return to this issue in section 2.3.

2.3 Methodology

2.3.1 The effects of federal funds rate surprises on volatility: daily analysis

In order to investigate the effect of monetary policy moves on implied volatility, we employ event study regressions similar to those used in Bernanke and Kuttner (2005) and Chulia-Soler, Marten and van Dijk (2007). Denote by V_t the implied volatility level on the day of the target change (day d of month t)[14] observed via either the VXO or VIX

[14] For the sake of notational simplicity, we abstract again from indexing by days throughout the methodology section and use instead the subscript t to denote time. For the daily event study type analysis, it is implied that this refers to a specific day in month t (when the target change occurs) while such a distinction is not needed for the monthly analysis. We also abstract from assigning a superscript to refer to a specific daily surprise measure, since the models are always estimated for the three daily surprise and expected measures detailed in section 2.2.

index. We start by measuring the response of implied volatility to actual rate changes, as well as to the expected and surprise elements of each rate change using the regressions[15]:

$$V_t = \alpha + \phi V_{t-1} + \beta \Delta i_t + \varepsilon_t \tag{2.8}$$

$$V_t = \alpha + \phi V_{t-1} + \beta^e \Delta i_t^e + \beta^u \Delta i_t^u + \varepsilon_t \tag{2.9}$$

where V_{t-1} denotes the level of the VIX (VXO) on the previous day. The first lag of implied volatility is included to account for the high-persistence observed in implied volatility.

The preceding literature examining the effects of a multitude of variables on implied volatility, such as Chulia-Soler, Martens and van Dijk (2007) Ederington and Lee (1996), Donders and Vorst (1996), Ahoniemi (2006), Fleming, Ostidek and Whaley (1995), Chen and Clements (2007), Nikkinen and Sahlstrom (2004), Whaley (1993) and Kearny and Lombra (2004), uses the change in implied volatility rather than the level of volatility itself. Although the authors of these papers argue that implied volatility, as measured by the VIX or VXO indices, is stationary, they call for differencing the implied volatility series to avoid statistical inference problems caused by the high persistence observed in the process. Fleming, Ostidek and Whaley (1995) further argue that an investor would be interested in the returns on volatility trading, rather than the volatility level itself, so that first differencing is warranted. Chulia-Soler, Martens and van Dijk (2007) argue that considering the change in volatility allows the researcher "to control for

[15] Throughout this chapter, α , β and ε_t are used as generic symbols to denote, respectively, a regression intercept, regression slopes and regression error terms and do not imply identical values or equality.

variation in the level of volatility over time for reasons other than the FOMC announcements".

Although our tests indicate the stationarity of the levels of both series, we also report the results from the regression using the first difference of the implied volatility process for two reasons. First, our results using first differences are more comparable with the literature. Second, these results can serve as a useful robustness check. Denote as $\Delta V_t = V_t - V_{t-1}$ the change in the level of implied volatility between the day of the target rate change and the previous day. Our regressions in first differences are specified as:

$$\Delta V_t = \alpha + \beta \Delta i_t + \varepsilon_t \tag{2.10}$$

$$\Delta V_t = \alpha + \beta^e \Delta i_t^e + \beta^u \Delta i_t^u + \varepsilon_t \tag{2.11}$$

The first regression investigates the effect of actual federal funds rate change on the change in implied volatility, while the second regression assesses the impact of the expected and unexpected component of a rate change on the change in implied volatility. Given that our tests indicate the absence of a unit root from the volatility processes, opting for using the changes in implied volatility induces over-differencing and leads to the residuals of the estimated regression to display moving average autocorrelation. Thus, it is important to exercise care when conducting inference based on the unadjusted standard errors of the estimates of such regressions, and we use Newey and West (1987) heretoskedasticity and autocorrelation consistent (HAC) standard errors for inference.

2.3.2 Controlling for macroeconomic releases

A basic assumption for estimating classical regression equations as in (2.8), (2.9), (2.10), and (2.11) is that the error term is uncorrelated with the regressors. Namely, if T observations on k regressors for each of the specified regressions is denoted X, a $(T \times k)$ matrix, the vector of unknown errors ε_t is denoted ε, a $(T \times 1)$ vector, then an orthogonatility condition $E(X^T \varepsilon) = 0$ is required for unbiasdness of the equation coefficients. In fact, the error term may include any variables that affect volatility but are omitted from the specified regressions.

Previous research suggests that volatility responds to a multitude of economic and financial variables. On the financial side, the level of the market (the financial leverage of the constituent firms of the S&P 100 or S&P500), or debt to equity ratios can influence implied volatility. On the economic side, Schwert (1989a) finds that inflation, money growth and industrial production can impact volatility, while Lombra and Kearney (2004) find that the unexpected component of employment significantly affects volatility. Especially relevant for our purposes is the possible same-day effect of economic variables on implied volatility, since the Fed would supposedly not take into account financial variables when formulating its policy. The Fed responds to signs of a weakening economy, such as lower employment, slower GDP growth or a drop in industrial production by cutting interest rates. A heating economy, as signalled by increasing inflation, prompts the Fed to increase the federal funds rate. In the extreme case where monetary policy and implied volatility jointly respond to economic news, endogeneity can arise.

Inferring, a priori, the sign any possible bias is obscured by the complex interaction of financial and economic variables and the absence of a clear theoretical model relating volatility to macroeconomic variables. For instance, a common analysis is that strong employment numbers would signal a growing economy and push the FOMC to increase the federal funds rate (the converse also being true), while higher inflation would push the FOMC towards a rate hike. Thus, federal funds rate changes are positively correlated with economic variables. However, the effect of economic variables (such as employment and inflation) on the stock market is still a subject of debate. Some researchers argue that stronger employment constitutes good news to investors about the state of the economy and hence pushes stock market returns up and volatility down. Others, such as Boyd, Hu and Jagannathan, (2005) argue that bad economic news can be good for stock returns, and would therefore imply lower volatility. These conflicting analyses obscure the direction of any correlation between volatility and economic variables.

To disentangle the periods in which both volatility and the federal funds rate could possibly jointly respond to economic news, a swift examination of the FOMC's actions through the years is warranted. The Fed adopted a policy of changing the Federal funds rate by multiples of 25 basis points starting in 1989 in a bid for a more transparent monetary policy. However, the 1989-1994 period saw many target rate changes, undertaken at intermeetings of the FOMC, which arguably were prompted by news about the state of the economy. In the post-1994 period, changes in monetary policy were mostly made during regularly scheduled FOMC meetings. Since these meetings are predetermined, a joint response of implied volatility and monetary policy to economic news is remote in the post-1994 period. For the pre-1994 period, the change in monetary policy could coincide with major economic announcements of interest to the Fed's mandate of maintaining a healthy economy with stable prices. Thus, this subsample is more prone to endogeneity problems than the post-1994 sample.

Since our data are daily, major announcements which are likely to elicit a joint implied volatility and monetary policy response include the employment report and the consumer price index (CPI) report published by the Bureau of Labour Statistics (BLS). These reports constitute major indicators of the state of the economy (specifically real activity and inflation), can arguably generate a reaction by the monetary authorities and stock markets, and refer to data series that are widely used in the finance literature [ex. Fama (1990), Boudoukh and Richardson (1993), Chen, Roll and Ross (1986)]. Following Bernanke and Kuttner (2005) and using the dates of report releases from the BLS[16] we construct the dummy variables: D^{EMP} taking the value one if the rate change coincides with the employment report release and zero otherwise, and D^{CPI} taking the value one if the rate change coincides with the CPI report release and zero otherwise. To control for endogeneity, and given that the reaction of volatility to the surprise element in each rate move is of interest, we modify regressions (2.9) and (2.11) by including interaction terms involving the surprise element and each of the dummy variables. The goal of modifying

[16] The dates of the release of the CPI and employment reports by the BLS for the period 1957-2008 are archived at their webpage: http://www.bls.gov/bls/archived_sched.htm.

the models is to assess the impact of major macroeconomic releases on the estimated coefficients; in particular, to investigate whether macroeconomic news releases dampen or increase the magnitude and significance of the surprise element of monetary policy rate moves. To that effect, we estimate the following regression models:

$$V_t = \alpha + \phi V_{t-1} + \beta^e \Delta i_t^e + \beta^u \Delta i_t^u + \beta_1 D^{EMP} \Delta i_t^u + \beta_2 D^{CPI} \Delta i_t^u + \varepsilon_t \tag{2.12}$$

$$\Delta V_t = \alpha + \beta^e \Delta i_t^e + \beta^u \Delta i_t^u + \beta_1 D^{EMP} \Delta i_t^u + \beta_2 D^{CPI} \Delta i_t^u + \varepsilon_t \tag{2.13}$$

Regression (2.12) estimates the impact of adding dummy variables associated with the release of the employment and CPI reports on the sign and significance of the coefficients β^e and β^u, relating monetary policy to implied volatility; while (2.13) does so for the change in implied volatility.

2.3.3 Simultaneity

Another econometric issue to consider in our regressions is the possibility of simultaneity: in cases of market turbulence, volatility could respond to monetary policy simultaneously with the Fed intervening in the market. Rigobon and Sack (2004) and D'Amico and Farka (2006) argue that monetary policy is found to respond to changes in asset prices. Bernanke and Gertler (1999) point out that it is "neither necessary nor desirable for monetary policy to respond to changes in asset prices", while Fuhrer and

Toottell (2008) estimate Taylor rules and argue that the Fed does not respond to the stock market. Furthermore, the mandate of the Fed itself restricts it to focus on inflationary pressures and economic growth and not on stabilizing financial markets. The few episodes in which the Fed did respond to financial crises[17] were short lived and unsystematic. This makes the possibility of simultaneity remote.

2.3.4 Measurement Errors

Even on FOMC meeting days, daily changes in federal funds futures prices can be driven by a multitude of other economic and financial news arriving to futures traders. This, in turn, can lead to our surprise measures being affected by such news and would imply that our day-to-day surprise measures contain measurement errors. In the event that such a measurement error exists, we argue that the coefficient associated with the surprise measure in (2.9) and (2.11) is biased downwards, and thus attenuates the response of volatility to surprises. We adapt Poole, Rasche and Thornton's (2002) exposition to our purposes. Durbin's (1954) classic "errors in variables" model would be:

$$\Delta i_t^u = \Delta i_t^{u*} + \delta_t \tag{2.14}$$

[17] Such as the Fed's intervention following Long Term Capital Management (LTCM) insolvency described in Greespan's (2007) autobiography. Greenpsan (2007) points out that the Fed did not actually respond to the crises but intervened by encouraging the sinking fund's debtors to supply more liquidity. The recent meltdown in the subprime mortgage market that induced the Fed to intervene is another case in point.

Where Δi_t^u is the observed (computed) surprise, Δi_t^{u*} is the true surprise and δ_t is a measurement error that is uncorrelated with Δi_t^{u*}. Thus for regression equation in (2.11)[18] we have:

$$\Delta V_t = \alpha + \beta^e (\Delta i_t - \Delta i_t^u) + \beta^u \Delta i_t^{u*} + \varepsilon_t \tag{2.15}$$

That is, (2.11), in which the true value, Δi_t^{u*}, replaces Δi_t^u which is in principle measured or observed. Then,

$$\Delta V_t = \alpha + \beta^e \Delta i_t + (\beta^u - \beta^e)\Delta i_t^u - \beta^u \delta_t + \varepsilon_t \tag{2.16}$$

or,

$$\Delta V_t = \alpha + \beta^e \Delta i_t + (\beta^u - \beta^e)\Delta i_t^{u*} + \eta_t \tag{2.17}$$

where $\eta_t = \varepsilon_t - \beta^u \delta_t$. Assuming that δ_t is independent of Δi_t^{u*} for all t, then $\mathrm{cov}(\Delta i_t^u \delta_t) = \mathrm{cov}\{(\Delta i_t^{u*} + \delta_t)(\delta_t)\} = \sigma_\delta^2$ where σ_δ^2 is the variance of the measurement error. Since $-\beta^u$ is presumed to be negative, there is a downward bias in the numerical value of $(\beta^u - \beta^e)$. Thus, in the event of measurement error in the surprises, the response of

[18] A similar argument can be applied to (2.10).

implied volatility to the funds rate surprises in our daily regressions is, subject to the assumptions, not enlarged.

Notwithstanding our argument above, the possibility of a measurement error in the daily surprise measures is remote. Gurkaynak, Sack and Swanson (2005) list intradaily (computed over intervals of thirty minutes surrounding FOMC meetings) and daily measures of the current month surprise (2.1) in their appendix. The interested reader is referred to their appendix and discussion where the authors report that the intradaily and daily surprise measures correspond very closely, except for a few occasions. If computing (2.1) with intradaily data over a very short time window corresponds to minimizing, if not eliminating, the possibility of any news reaching the futures markets (within the thirty minutes used for computation) then the daily surprise measures in (2.1) can be thought of as being free of any measurement error due to news reaching the fed funds futures market.

2.3.5 The effect of timing surprises on volatility

The response of volatility to current federal funds rate surprises may be due wholly to uncertainty regarding the timing of an FOMC meeting, or it may not. To estimate the effect of uncertainty due to the timing of FOMC meetings, we modify equations (2.9) and (2.11) to include the timing surprise as:

$$V_t = \alpha + \phi V_{t-1} + \beta^e \Delta i_t^{e,0} + \beta^u \Delta i_t^{u,0} + \beta^{time} time_t + \varepsilon_t \qquad (2.18)$$

$$\Delta V_t = \alpha + \beta^e \Delta i_t^{e,0} + \beta^u \Delta i_t^{u,0} + \beta^{time} time_t + \varepsilon_t \tag{2.19}$$

The principal interest in estimating regressions (2.18) and (2.19) is the effect on the significance, direction and magnitude of β^u and the value assumed by β^{time}. Our interest centers on the current month surprise, not the other surprise measures, due to its confounding with the timing shock. By controlling for the timing surprise as in (2.18) and (2.19), the effect of any timing surprises is removed from β^u. Bernanke and Kuttner (2005, p.1238) point out that, with the regression specification in (2.18) and (2.19), the coefficient associated with the current month surprise can "be interpreted as the impact of a funds rate surprise that changes expectations by the same amount". Hence, specifying equations (2.18) and (2.19) allows the importance of timing surprises on volatility to be assessed.

2.3.6 The effect of federal funds rate surprises on volatility: monthly analysis

A regular time series regression, in line with (2.9) and (2.11), is employed to measure the effect of federal funds rate surprises on implied volatility. Using our monthly dataset, the first step consists of estimating the two equations:

$$V_t = \alpha + \phi V_{t-1} + \beta^e \overline{\Delta i}_t^e + \beta^u \overline{\Delta i}_t^u + \varepsilon_t \tag{2.20}$$

$$\Delta V_t = \alpha + \beta^e \overline{\Delta i}_t^e + \beta^u \overline{\Delta i}_t^u + \varepsilon_t \tag{2.21}$$

Given the findings of Piazzesi and Swanson (2008) and our discussion of surprise measures in section 2.2.7, we augment the regressions in (2.20) and (2.21) with macroeconomic variables that have been found to affect the monthly surprise measure in (2.6). Specifically, Piazzesi and Swanson (2008) and Bernanke and Kuttner (2005) indicate that nonfarm payroll employment significantly affects our monthly surprise measure, so we augment all our monthly regressions with the period-to-period (month to month) growth rate in nonfarm employment, Δemp_t. Based on previous findings by Bernanke and Kuttner (2005), we also include as macroeconomic factors in the monthly regressions: the growth rate of industrial production Δip_t and the consumer price inflation (inf). Monthly regressions which are augmented with macroeconomic factors are:

$$V_t = \alpha + \phi V_{t-1} + \beta^e \overline{\Delta i}_t^e + \beta^u \overline{\Delta i}_t^u + \beta_1 \Delta emp_t + \beta_2 \Delta ip_t + \beta_3 \inf_t + \varepsilon_t \tag{2.22}$$

$$\Delta V_t = \alpha + \beta^e \overline{\Delta i}_t^e + \beta^u \overline{\Delta i}_t^u + \beta_1 \Delta emp_t + \beta_2 \Delta ip_t + \beta_3 \inf_t + \varepsilon_t \tag{2.23}$$

On the financial side, based on the findings of Swanson and Piazzesi (2008) and the return predictability literature [for example, Anatolyev and Gospodinov (2009)], we include the three month T-bill rate, tb_t, and the default spread dfs_t defined as the difference between BAA and AAA rated corporate bond yields. Monthly regressions augmented with financial factors can then be written:

$$V_t = \alpha + \phi V_{t-1} + \beta^e \overline{\Delta i}_t^e + \beta^u \overline{\Delta i}_t^u + \beta_1 \Delta emp_t + \beta_2 tb_t + \beta_3 dfs_t + \varepsilon_t \tag{2.24}$$

$$\Delta V_t = \alpha + \beta^e \overline{\Delta i}_t^e + \beta^u \overline{\Delta i}_t^u + \beta_1 \Delta emp_t + \beta_2 tb_t + \beta_3 dfs_t + \varepsilon_t \tag{2.25}$$

We include the three month T-bill rate and the default spread as measures that are closely associated with the discount rate at which stocks are valued and as a broad indicator of financial leverage, respectively. Other variables, such as debt to equity ratios or other short-term interest rates can be used instead as the default spread is known to be more of a business-cycle related variable. Given that we assign an explanatory role for the T-bill rate and the default spread's effect of volatility, correct inference regarding the coefficients is of importance. Anatolyev and Gospodinov (2009) note that with highly persistent regressors, as are the three month T-Bill rate and the default spread, the limiting distribution for the regression coefficients in (2.24) and (2.25) is nonstandard. However, the authors note that when the correlation with the dependent variable is not large, the standard normal critical values can be used as a reasonable approximation. We find that the T-bill rate is nearly uncorrelated with implied volatility, while the default spread has a more significant correlation with the level of volatility (both interest rate variables have a correlation coefficient close to zero with changes in volatility) so that standard normal critical values can used as an acceptable approximation for the T-bill rate and for the default spread (except when the level of volatility is used).

The goal of modifying the monthly regressions is twofold: first, we control for any risk premiums and any correlation of the monthly surprises with macroeconomic factors. Secondly, it would be interesting, in itself, to investigate the response of volatility to such factors, since this has not been previously examined.

2.3.7 Nonlinearities in responses

Nonlinearities and asymmetries in the response of volatility to monetary policy may be investigated along two major lines. First, consider the possibility of a different reaction of implied volatility to a positive, as opposed to a negative, federal funds rate change by the Fed. Since previous research points to a considerable increase in volatility following target rate decreases, we examine whether implied volatility displays similar behaviour using nonlinear models of the form:

$$V_t = \phi V_{t-1} + (\alpha_0 + \beta_1 \Delta i_t) D(\Delta i_t > 0) + (\alpha_0 + \beta_2 \Delta i_t) D(\Delta i_t < 0) + \varepsilon_t \tag{2.26}$$

$$\Delta V_t = (\alpha_0 + \beta_1 \Delta i_t) D(\Delta i_t > 0) + (\alpha_0 + \beta_2 \Delta i_t) D(\Delta i_t < 0) + \varepsilon_t \tag{2.27}$$

Where D(.) is a indicator variable taking the value one when the condition in parentheses is satisfied, otherwise zero. The two models above help in illustrating whether volatility, or the change in volatility, displays different behaviour following a rate cut or a rate increase.

Secondly, consider the response of implied volatility to the sign of surprises. Previous research has found volatility to increase following negative news (positive federal funds rate surprises). In fact, a positive surprise element, implying that the Fed increased the interest rate more than the market expected, could lead to an increase in volatility. This is due to the fact that a positive surprise is equivalent to bad news for stocks which are now valued at a higher-than-expected discount rate (the converse is also true: negative federal funds rate surprises are good news for stocks). To examine the possible effects of negative and positive surprises on implied volatility, we estimate the following nonlinear models:

$$V_t = \phi V_{t-1} + (\alpha_0 + \beta_1 \Delta i_t^u) D(\Delta i_t^u > 0) + (\alpha_0 + \beta_2 \Delta i_t^u) D(\Delta i_t^u < 0) + \varepsilon_t \tag{2.28}$$

$$\Delta V_t = (\alpha_0 + \beta_1 \Delta i_t^u) D(\Delta i_t^u > 0) + (\alpha_0 + \beta_2 \Delta i_t^u) D(\Delta i_t^u < 0) + \varepsilon_t \tag{2.29}$$

The models (2.28) and (2.29) capture whether implied volatility, or its first difference, responds differently to positive and negative surprises.

By adding and subtracting $(\alpha_0 + \beta_1 \Delta i_t) D(\Delta i_t < 0)$ from (2.26) and (2.27) we get the equivalent representations:

$$V_t = \alpha_0 + \beta_1 \Delta i_t + \phi V_{t-1} + [(\beta_2 - \beta_1) \Delta i_t] D(\Delta i_t < 0) + \varepsilon_t \tag{2.30}$$

$$\Delta V_t = \alpha_0 + \beta_1 \Delta i_t + [(\beta_2 - \beta_1) \Delta i_t] D(\Delta i_t < 0) + \varepsilon_t \tag{2.31}$$

While adding and subtracting $(\alpha_0 + \beta_1 \Delta i_t^u) D(\Delta i_t^u < 0)$ from (2.28) and (2.29) gives:

$$V_t = \alpha_0 + \beta_1 \Delta i_t^u + \phi V_{t-1} + [(\beta_2 - \beta_1)\Delta i_t^u] D(\Delta i_t^u < 0) + \varepsilon_t \tag{2.32}$$

$$\Delta V_t = \alpha_0 + \beta_1 \Delta i_t^u + [(\beta_2 - \beta_1)\Delta i_t^u] D(\Delta i_t^u < 0) + \varepsilon_t \tag{2.33}$$

Rewriting the models in this manner affords a simple test of the possible presence of two regimes in the reaction of volatility to the sign of rate changes and surprises. Such a test involves rejecting the null of equality of the slope parameters, $H_0 : \beta_2 - \beta_1 = 0$ in (2.30) and (2.31) in the positive and negative rate change regimes with regular t-tests.

2.4 Results

2.4.1 The daily effect of funds rate surprises on implied volatility

Our estimates from the daily regressions, reported in tables 2.2, 2.3, 2.4 and 2.5 reveal interesting relationships. The actual change in the federal funds rate change has a statistically insignificant, positive effect on implied volatility. In fact, our results indicate that a unit percentage change in the actual federal funds rate causes volatility (or the change in volatility) to increase by 0.19 to 0.62 percentage points depending on the specification (level or difference of volatility) used. This effect is insignificant at any conventional level.

More elaborate and appealing results are obtained when the rate move is broken up into surprise and expected components as in (2.9) and (2.11). For all of our federal fund surprise measures defined in (2.1), (2.2) and (2.3) and for both the VIX and VXO indices, the expected component of a rate change has a negative and statistically insignificant (except for one specification) effect on implied volatility, while the surprise rate change has a large, positive and statistically significant effect on implied volatility. This general result holds true whether the level or change in volatility is used at the estimation stage.

Evidently, a unit percentage increase in a current surprise leads to a 2.76 (2.81) percentage points increase in the level of the VIX (VXO) index, while a percentage point expected rate increase elicits a 0.7 (0.76) percentage point decrease in the level of the VIX (VXO) index. When the change in volatility is considered, a percentage point increase in the surprise component increases the VIX (VXO) by 2.91(2.92) percentage points, while a percentage point increase in the expected component causes a drop of 0.13 (0.20) percentage points. While the surprise coefficient is significant at standard levels (1%, 5% and 10% levels), the expected component is insignificantly different from zero.

When the one-month ahead surprise in (2.2) is used, a percentage point increase in the surprise leads to a 4.22 (4.09) percentage points increase in the level of the VIX (VXO) while a percentage point increase in the expected component causes the VIX (VXO) to drop by 0.93 percentage points. Again, the coefficient associated with the surprise is statistically significant, while the coefficient associated with the expected component is not. While the coefficients associated with the expected component are of the same magnitude as the actual rate change, the coefficients on the surprise measures are much larger. Hamilton (2008a, 2008b) notes that the one-month ahead surprise contains more information than the current month surprise and this could explain the larger coefficient vis-a-vis the current month surprise component.

Similar results are obtained with the two-month ahead surprise in (2.3) computed for the post 1994 period: a percentage point increase in the two-month ahead surprise causes the level of the VIX (VXO) to increase significantly (at the 5% and 1% level) by 2.42 (2.59) percentage points and by 2.79 (2.96) percentage points when the change in volatility is used. The only difference is the sign of the expected component, which becomes positive but remains far from significant. The lower magnitude of the coefficient associated with the two-month ahead surprise is expected, since the computation of this surprise minimizes the timing surprise confounded in the current and one-month ahead surprises.

How are these coefficients to be interpreted? The negative coefficient on the expected rate move is indicative of a drop in volatility when market participants suitably anticipate the federal funds rate change. In short, we interpret the drop in volatility as reflecting confirmation of the anticipation of the federal funds change. Investors do not need to rebalance their portfolios in the light of the arrival of new information, and this drives trading, and thereby volatility, lower. The surprise element in a monetary policy move acts in an opposing direction. The positive, large and statistically significant coefficient of the surprise rate move suggests that volatility is increased due to the surprise component of an actual rate change. This result possesses an intuitive interpretation: investors adjust the allocation of their portfolios in light of the "news" contained in the monetary policy move, thereby increasing trading and, in turn, implied volatility.

Furthermore, our results are in line with the efficient markets hypothesis and the literature. According to the efficient market hypothesis, a forward looking, efficient stock market should only respond to the arrival of new information, and we find that volatility only responds to federal funds rate surprises but not to the actual or expected component of a rate move. In comparison to the literature, our results are in line with Chulia-Soler, Martens and van Dijk (2007) who examine the effect of the actual federal funds rate change, as well as the expected and surprise elements of actual rate moves on realized volatility computed from high-frequency financial return on the S&P 100 index. The authors use a similar methodology as the one employed in this paper for gauging surprises and expectations and uncover a positive, large and highly significant response of realized volatility to surprise movements (18.57) and a negative but insignificant effect of the expected rate move on realized volatility (-0.14).

We also advance a simple interpretation of our results that is line with Bernanke and Kuttner (2005). These authors find that a federal funds rate surprise decreases stock market returns, while the expected component increases stock market returns. Our coefficients have exactly the opposite sign than those of Bernanke and Kuttner (2005). The authors argue that a funds rate surprise decreases stock returns due to either an increase in the discount rate used to value cash flows (dividends) from stocks, an increase in the equity premium or a decrease in expected future dividends. Regardless of the exact channel through which surprises affect stock returns, we posit that due to the well known negative correlation between returns and volatility, any negative news for stocks (as are federal funds rate surprises) will translate into higher volatility. Thus, we conjecture that federal funds rate surprises affect volatility through the return channel. To understand the plausibility of this scenario, we refer to the observation dating back to Black (1976) which stipulates that negative returns spur an increase in volatility. This observation, known in the literature as the "leverage effect", and incorporated into popular volatility models such as the GJR-GARCH of Glosten, Jagannathan and Runkle (1993), holds for implied volatility measured by the VIX (VXO) as shown in Fleming, Ostdiek and Whaley (1995). In all, federal funds rate surprises drive stock returns down, which in turn, drive volatility up. Testing such a hypothesis calls for the use of a richer dynamic model where returns and volatility are modeled jointly. This is beyond the scope of the current chapter.

2.4.2 Controlling for macroeconomic releases

In section 2.3 of this chapter, we argued that it is conceivable for implied volatility and monetary policy to respond jointly to economic news. This can cause endogeneity if such a joint response is extreme. To account for this problem and assess its impact on the magnitude and significance of the surprise component, interaction dummy variables were included in our regressions in (2.12) and (2.13).

The results, reported in tables 2.10 and 2.11, indicate that the coefficients associated with the surprise element increase in magnitude once the dummy variables associated with the release of the employment and CPI reports are included in the model. With the dummy variables included in the model, a percentage point increase in the current surprise component increases the level of the VIX (VXO) index by 3.37 (3.60) percentage points as compared to the 2.76 (2.81) percentage points increase when the dummy variables are excluded. When the one-month ahead surprise measure is employed, a 1% surprise increases causes the level of the VIX (VXO) to increase by 4.95% (4.96%) as compared to 4.22% (4.09%) when the dummy variables are excluded.

A similar increase in the coefficient associated with the surprise is observed when the change in volatility is used. Although the surprise term interacted with the CPI release dummy is never significant, we find that the interaction term involving the employment release dummy is mostly significant. Interestingly, all the interaction term coefficients in (2.12) and (2.13) display a negative sign. For our purposes, it is interesting to note that macroeconomic releases exert an opposite effect on implied volatility, which is sometimes of equal magnitude to the surprise element (with the employment report interaction term). This, in turn, implies that the coefficient on the surprise element is weighted downwards on days where monetary policy actions correspond with macroeconomic releases. The significance of our previous results, indicating that the surprise element increases implied volatility, is equally maintained. In sum, it appears that the effect of the surprise element on implied volatility is robust to the inclusion of major macroeconomic announcements which might jointly move the stock market and monetary authorities.

2.4.3 Timing surprises and volatility

Is the positive effect of funds rate surprises on volatility due solely to market uncertainty about the timing of an FOMC action? Our results indicate that the response of volatility cannot be traced, in its entirety, to timing surprises. In fact, estimation of (2.18) and

(2.19) indicates that when the timing surprise is included alongside the current surprise in the model, the coefficient associated with the current surprise drops, but remains significant, in most specifications.

When the level of the VIX (VXO) is used, regressions (2.18) and (2.19) reveal the current surprise coefficient is significant 2.24 (2.74), while the coefficient associated with the timing surprise is an insignificant 1.38 (0.17). This suggest that uncertainty with regard to timing of an FOMC action increases volatility, but that the bulk of the response of volatility is due to the component of the current surprise that reflecting a more permanent change in the expected level of the funds rate.

2.4.4 The monthly effect of funds rate surprises on implied volatility

The results from the monthly regressions, including only the monthly federal funds rate surprise and expected components, are in line with the daily results. The monthly federal funds rate surprise increases the level (and first difference) of volatility, while the monthly expected component drives volatility down. In fact, we find that the level of the VIX (VXO) is increased by 3.27 (3.59) percentage points due to a one percentage point increase in the monthly surprise, while it is decreased by 2.91 (0.62) percentage points due a one percentage point increase in the monthly expected component.

When the change in the VIX (VXO) is considered, the results are similar: an increase of 4.09 (4.48) percentage points due to a percentage point increase in the surprise, while percentage point change in the expected component increases volatility in this specification by 1.05 (1.92) percentage points. Although the sign of the expected component is unexpected when changes in volatility are used, the monthly expected component's effect on volatility is not significant in any of the specifications. The monthly surprise significantly affects volatility in almost all specifications (except when the level of the VIX is used) and displays similarity, in terms of the magnitude of the response, with the daily coefficient.

At first, such results point to a relatively long-lived response of volatility to federal funds rate surprises. However, it was argued earlier that there should be some control for variables that affect the monthly surprise measures in (2.6). In fact, estimation of the monthly regressions augmented by macroeconomic variables, as in (2.22) and (2.23), yields a decrease in the significance of the response of volatility to surprises. Although the magnitude of the surprise coefficient does not change much, its significance decreases in all specifications and with both the VIX and VXO. The CPI inflation and the growth rate of industrial production significantly increase volatility, while the employment growth insignificantly decreases volatility. These findings imply a weaker response of volatility to federal funds rate surprises, and might be indicative of presence of macroeconomic risk premiums in volatility itself.

When the monthly regressions in (2.24) and (2.25) are estimated, the federal funds rate surprise coefficient retains its significance, while the coefficients of other interest rate variables are not significant. We interpret this result in two ways: first, it reflects the importance of separating expectations from surprises in financial markets, and secondly, such a result might be due to the fact that the federal funds rate plays a more central role than other short term interest rates in affecting the discount rate at which stocks are valued.

2.4.5 Nonlinearity and asymmetry in the response of volatility to target rate changes and surprises

The results obtained from investigating nonlinearities and symmetries in the response of implied volatility to the sign of the rate change and the sign of the surprises in (2.26), (2.27), (2.28) and (2.29) are reported in the tables 2.18, 2.19, 2.20 and 2.21. We can infer from the results that implied volatility exhibits weak evidence of nonlinearity with respect to both the sign of rate moves and the sign of the surprise element in each rate move. Implied volatility is increased following a rate cuts as evidenced by the positive coefficient associated with the negative rate change β_2 in (2.26) and (2.27). In effect, when the level of the VIX (VXO) is used, we find that rate cuts increase volatility significantly (at the 10% level) by 2.99% (2.35%), while the effect of rate increases on volatility is mixed. We find that the difference in the slope coefficient is only significant when the change in the VXO is used. Overall, evidence of a clear response of implied volatility to the sign of the rate change is elusive.

When the sign of the surprise element is used to separate the two regimes in (2.28) and (2.29), we find that positive and negative surprises increase implied volatility. We also find that, perhaps counter intuitively, the slope coefficient associated with the negative surprise term (good news for stocks) is greater and slightly more significant than the slope coefficient associated with the positive surprise term (bad news for stocks) when the level of volatility is used. Such a result is similar to Chulia-Soler, Martens and Van Dijk (2007) who find that positive news (negative funds rate surprises) increase realized volatility more than negative news (positive fund rate surprises). However, this result is weakened when the change in volatility is used. When the changes in VIX (VXO) are used, we observe that positive and negative surprises increase volatility by similar amounts and that the difference between the slope coefficients of the two regimes, governed by the sign of the surprise element, is insignificant. Again, this is suggestive of a weak response of volatility to positive versus negative news as in Chulia-Soler, Martens and Van Dijk (2007)

2.4.6 Robustness checks

We investigate the robustness of our main conclusions regarding the impact of federal funds surprises on implied volatility along two lines. First, in a bid for more transparency, the Fed changed its conduct of monetary policy in 1994 and began announcing the occurrence of policy actions. Starting in 1997, the FOMC statement included an explicit numeric value for the federal funds rate target in FOMC statements. As Lange, Sack and Whietsell (2003) and Swanson (2005) report, this change in the conduct of monetary policy is reflected by a decrease in the surprise component of monetary policy after 1994.

In order to check the robustness of our conclusions pertaining to the effect of federal funds surprises on implied volatility, we re-estimate regressions (2.9) and (2.11) with a subsample spanning the 1994 to 2007. Reassuringly, our main results (available upon request) remain robust to the change in sample. Namely, for the post 1994 period, a percentage point current month surprise increase induces a 6.08 (6.19) percentage points increase in the level of VIX (VXO) and a 6.56% (6.63%) increase in the change of VIX

(VXO). The actual rate move and the expected rate move are still insignificant in this subsample. Another robustness check undertaken relates to the specification of the regression equations in (9) and (11). Instead of using the level and change in volatility, we use the log level and changes in log levels. Even though the coefficients associated with the surprise term decrease in magnitude, we find that the significance and sign of the estimated relationship is preserved (while the actual rate move and expected rate changes are insignificant).

In all, the positive reaction of implied volatility to monetary policy appears robust to different specifications and to changes in the sample.

2.5 Concluding remarks

In this chapter, the response of implied volatility, as measured by CBOE's VIX and VXO indices, to federal funds rate surprises is investigated. Our analysis takes into account market expectations, measured using federal funds futures contracts, in order to disentangle the expected and surprise elements of each federal funds rate move by the Fed. Implied volatility responds positively and significantly across various specifications to federal funds rate surprises, while it does not respond to the expected component or actual target rate move. The results are not altered once we account for macroeconomic news releases or when timing uncertainty is introduced.

A monthly analysis relating volatility to federal funds rate surprises is also carried out. Monthly surprises significantly increase volatility, while the expected component's effect on volatility is insignificant. The monthly results are weakened by the inclusion of additional macroeconomic variables. Interestingly, we find that macroeconomic variables, such as industrial production growth and inflation, significantly affect volatility. Nonlinearities and asymmetries in the response of implied volatility to the sign of the target move and surprises are also investigated. We find that volatility is increased following rate cuts and following negative surprises. However, the evidence for such nonlinearities and asymmetries is weak.

From a policy-making perspective, the results show that the effect of the Fed on volatility is significant. A large literature in macroeconomics examines the effect of monetary policy transparency on the economy. A similar examination, relating to the stock market, might be of interest. For instance, even though the FOMC became more transparent following 1994, this did not translate into lower stock market volatility. If lower volatility is a desirable outcome, then central banks should become better at communicating not only their current policy stand, but more importantly, their expected future course of action.

From a trading perspective, our results indicate that profit seeking investors possessing a certain understanding of likely FOMC actions might be able to generate profits by taking correct options positions. In fact, it would be interesting to evaluate the ability, even with a directional forecast of volatility coming from, for example, a probit model, to generate profits on FOMC announcement days.

Lastly, other aspects of the interaction of Fed actions and volatility are interesting to explore. For instance, the transmission mechanism through which the Fed affects volatility is an interesting area to explore. Other information embedded in the term structure of federal funds futures contracts can also be used to determine the effect of different surprises on stock market returns and volatilities.

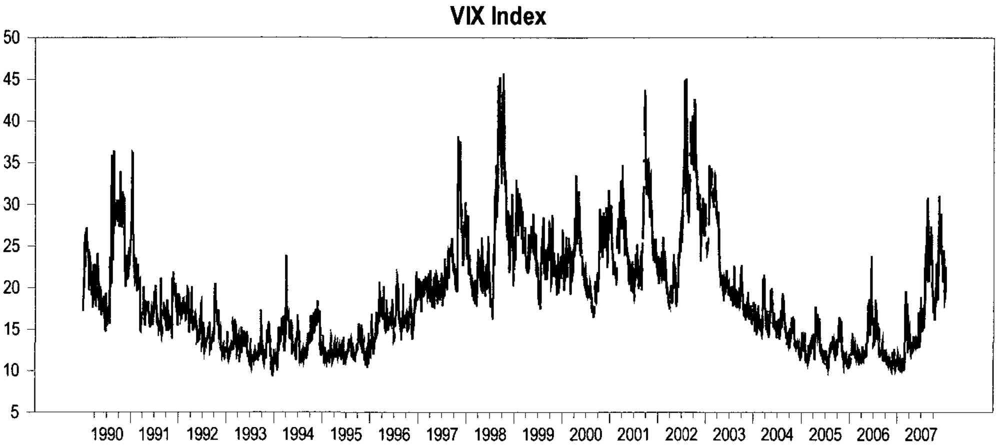

Figure 2.1: Daily time series of the VIX index for the 1990 to 2007 period

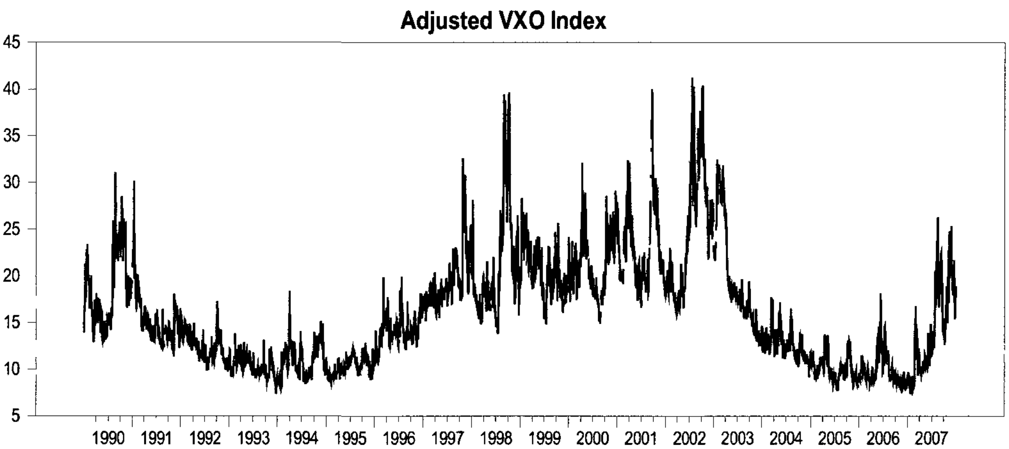

Figure 2.2: Daily time series of the adjusted VXO index for the period 1990 to 2007

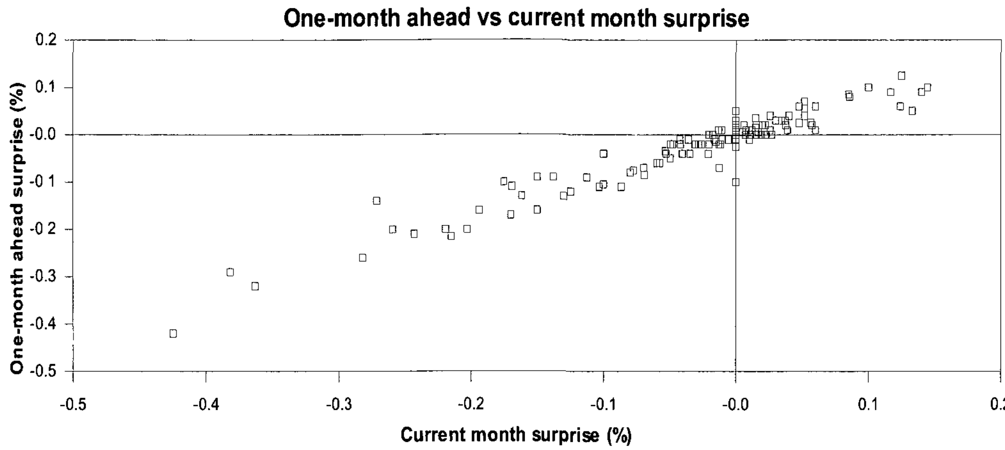

Figure 2.3: Scatter plot illustrating the relationship between the current month surprises computed in (2.1) and the one-month-ahead surprises computed in (2.2)

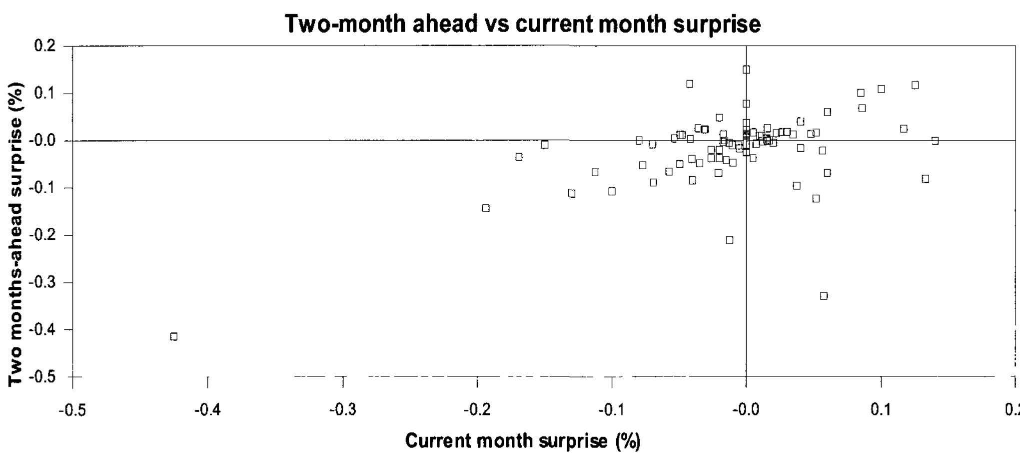

Figure 2.4: Scatter plot illustrating the relationship between the current month surprises computed in (2.1) and the two-month-ahead surprises computed in (2.3)

Table 2.1: Summary statistics and unit root tests for the VIX and adjusted VXO indices

	VIX	Adjusted VXO
Implied volatility indices-Summary statistics (levels)		
Period	02/01/1990 to 31/12/2007	02/01/1990 to 31/12/2007
Number of Observations	4535	4535
Mean	0.18	0.19
Standard Deviation	0.06	0.07
Skewness	0.97	0.98
Excess Kurtosis	0.79	0.83
First-Order Autocorrelation	0.95	0.95
Second-Order Autocorrelation	0.93	0.93
Third-Oder Autocorrelation	0.91	0.92
Fourth-Order Autocorrelation	0.90	0.91
Fifth-Order Autocorrelation	0.90	0.90
Augmented Dickey-Fuller Test	-5.34	-5.17
ADF-GLS test	-4.12	-3.87

Notes: The optimal lag length for the Augmented Dickey-Fuller (ADF) tests reported in the last row is chosen using the Akaike (AIC) and Bayesian (BIC) information criteria. The last row of the table reports the ADF test with GLS demeaning as proposed by Elliot, Rothenberg and Stock (1996). The optimal lag length for the ADF-GLS test is chosen using the Ng and Perron (2001) criterion. All unit root tests reject the null of a unit root at the 1% level

Table 2.2: Regression results from models (2.8) and (2.9) with current month surprise (levels VXO)

	Implied volatility: VXO	Implied volatility: VXO
Intercept	0.35	0.51*
	(0.25)	(0.28)
First lag of VXO	0.96***	0.95***
	(0.01)	(0.01)
Actual rate change	0.19	.
	(0.44)	.
Expected change	.	-0.76
	.	(0.54)
Current month surprise	.	2.81**
	.	(1.29)

Notes: Newey-West heteroskedasticity and autocorrelation consistent (HAC, 2 lags) standard errors in parentheses. All variables are in percent. * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.3: Regression results from models (2.8) and (2.9) with current month surprise (levels VIX)

	Implied volatility: VIX	Implied volatility: VIX
Intercept	0.33	0.49
	(0.32)	(0.36)
First lag of VIX	0.96***	0.95***
	(0.01)	(0.01)
Actual rate change	0.21	.
	(0.50)	.
Expected change	.	-0.70
	.	(0.62)
Current month surprise	.	2.76**
	.	(1.33)

Notes: Newey-West heteroskedasticity and autocorrelation consistent (HAC, 2 lags) standard errors in parentheses. All variables are in percent. * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.4: Regression results from models (2.10) and (2.11)
with current month surprise (Changes in VXO)

	Change in VXO	Change in VXO
Intercept	-0.26***	-0.20***
	(0.08)	(0.07)
Actual rate change	0.57	•
	(0.43)	•
Expected change	•	-0.20
	•	(0.43)
Current month surprise	•	2.92*
	•	(1.55)

Notes: Newey-West HAC (6 lags) standard errors in parentheses. * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.5: Regression results from models (2.10) and(2.11)
with current month surprise (Changes in VIX)

	Change in VIX	Change in VIX
Intercept	-0.42***	-0.36***
	(0.08)	(0.08)
Actual rate change	0.62	•
	(0.45)	•
Expected change	•	-0.13
	•	(0.45)
Current month surprise	•	2.91*
	•	(1.50)

Notes: The current month surprise is used. Newey-West HAC (6 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.6: Regression results from model (2.9)
with one-month-ahead surprise (levels)

	Implied volatility: VIX	Implied volatility: VXO
Intercept	0.55	0.56
	(0.36)	(0.28)
Lag of volatility	0.95***	0.95***
	(0.01)	(0.01)
Expected change	-0.93	-0.93**
	(0.56)	(0.47)
One-month surprise	4.22***	4.09***
	(1.52)	(1.39)

Notes: The one-month ahead surprise is used. Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.7: Regression results from model (2.11)
with one-month ahead surprise (changes)

	Change in VIX	Change in VXO
Intercept	-0.34***	-0.18**
	(0.08)	(0.07)
Expected change	-0.32	-0.35
	(0.41)	(0.38)
One-month surprise	4.26***	4.11***
	(1.61)	(1.59)

Notes: The one-month ahead surprise is used. Newey-West HAC (6 lags) standard errors in parentheses. * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.8: Regression results from model (2.9)
with two-month-ahead surprise (levels)

	Implied volatility: VIX	Implied volatility: VXO
Intercept	0.02	0.04
	(0.34)	(0.27)
Lag of volatility	0.97***	0.97***
	(0.01)	(0.01)
Expected change	0.43	0.13
	(0.74)	(0.65)
Two-month surprise	2.42**	2.59***
	(0.99)	(0.85)

Notes: The two-month ahead surprise is used. Newey-West HAC (2 lags) standard errors in parentheses. * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.9: Regression results from model (2.11)
with two-month-ahead surprise (changes)

	Change in VIX	Change in VXO
Intercept	-0.47***	-0.30**
	(0.10)	(0.09)
Expected change	0.75	0.39
	(0.74)	(0.64)
Two-month surprise	2.79**	2.96***
	(1.16)	(0.97)

Notes: Newey-West HAC (6 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.10: Regression results from model (2.12) with macroeconomic interaction terms (levels)

	Implied volatility: VIX	Implied volatility: VXO
Intercept	0.46	0.46*
	(0.36)	(0.28)
Lag of volatility	0.95***	0.95***
	(0.01)	(0.01)
Expected change	-0.69	-0.75
	(0.60)	(0.53)
Current month surprise	3.37**	3.60**
	(1.51)	(1.44)
Surprise × employment	-2.62	-3.63
	(2.66)	(2.23)
Surprise × CPI	-0.80	0.02
	(3.61)	(2.28)

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.11: Regression results from model (2.13) with macroeconomic interaction terms.

	Change in VIX	Change in VXO
Intercept	-0.37***	-0.20**
	(0.09)	(0.08)
Expected change	-0.15	-0.23
	(0.53)	(0.47)
Current month surprise	3.65**	3.85**
	(1.65)	(1.60)
Surprise × employment	-3.30	-4.32*
	(2.95)	(2.44)
Surprise × CPI	-0.44	-0.01
	(3.85)	(2.56)

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.12: Regression results from model (2.18)
with the timing surprise (levels)

	Implied volatility: VIX	Implied volatility: VXO
Intercept	0.47	0.51*
	(0.33)	(0.27)
Lag of volatility	0.95***	0.95***
	(0.01)	(0.01)
Expected change	-0.59	-0.74
	(0.59)	(0.53)
Current month surprise	2.24*	2.74**
	(1.31)	(1.28)
Timing surprise	1.38	0.17
	(2.52)	(2.13)

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.13: Regression results from model (2.19)
with timing surprise (changes)

	Change in VIX	Change in VXO
Intercept	-0.37***	-0.20***
	(0.08)	(0.07)
Expected change	-0.005	-0.16
	(0.46)	(0.43)
Current month surprise	2.22	2.74*
	(1.44)	(1.60)
Timing surprise	1.82	0.46
	(2.94)	(2.58)

Notes: Newey-West HAC (6 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.14: Regression results from model (2.20)
monthly surprise (levels)

	Implied volatility: VIX	Implied volatility: VXO
Intercept	3.40***	2.47***
	(0.68)	(0.60)
Lag of volatility	0.82***	0.85***
	(0.03)	(0.04)
Expected change	-2.91	-0.62
	(1.91)	(1.28)
Monthly surprise	3.27	3.59**
	(2.47)	(1.62)

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.15: Regression results from model (2.21) with monthly surprise (changes)

	Change in VIX	Change in VXO
Intercept	0.11	0.14
	(0.18)	(0.15)
Expected change	1.05	1.92
	(1.84)	(1.26)
Monthly surprise	4.09**	4.48***
	(2.06)	(1.71)

Notes: Newey-West HAC (6 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level

Table 2.16: Regression results from model (2.22) with monthly surprise and macroeconomic variables (levels)

	Implied volatility: VIX	Implied volatility: VXO
Intercept	3.10***	2.03***
	(0.77)	(0.66)
Lag of volatility	0.83***	0.86***
	(0.03)	(0.03)
Expected change	-2.53	-0.54
	(2.08)	(1.36)
Monthly surprise	3.02	3.31*
	(2.58)	(1.76)
Employment growth	-3.18	-2.50
	(1.98)	(1.70)
Inflation rate	1.24*	1.67***
	(0.63)	(0.55)
Industrial prod. growth	1.25*	1.06*
	(0.65)	(0.59)

Notes: Newey-West HAC (2 lags) standard errors in parentheses. * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.17: Regression results from model (2.23)
with monthly surprise and macroeconomic variables (changes)

	Change in VIX	Change in VXO
Intercept	-0.33	-0.41
	(0.34)	(0.31)
Expected change	0.56	1.23
	(2.13)	(1.42)
Monthly surprise	3.34	3.73*
	(2.38)	(1.93)
Employment growth	-1.98	-1.40
	(1.91)	(1.60)
Inflation rate	1.55**	1.92***
	(0.58)	(0.53)
Industrial prod. growth	1.29**	1.09**
	(0.59)	(0.59)

Notes: Newey-West HAC (6 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level.

Table 2.18: Results from the nonlinear model (2.26)

	Implied Volatility: VIX	Implied Volatility: VXO
First lag of volatility	0.97***	0.98***
	(0.005)	(0.005)
Intercept (Positive rate change)	0.02	0.09
	(0.61)	(0.52)
Slope (Positive rate change)	0.006	-0.51
	(1.88)	(1.61)
Intercept (Negative rate change)	0.89	0.54
	(0.60)	(0.51)
Slope (Negative rate change)	2.99*	2.35
	(1.71)	(1.46)
P-value (for equality of slopes)	0.39	0.09

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level. P-values in last column refer to the null of equality of slope parameters in the 2 regimes.

Table 2.19: Results from the nonlinear model (2.27)

	Change in VIX	Change in VXO
Intercept (Positive rate change)	-0.32	-0.08
	(0.44)	(0.39)
Slope (Positive rate change)	-0.02	-0.54
	(0.84)	(0.79)
Intercept (Negative rate change)	0.48	0.33
	(0.67)	(0.54)
Slope (Negative rate change)	3.35	2.62
	(2.14)	(1.89)
P-value (for equality of slopes)	0.38	0.09

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level. P-values in last column refer to the null of equality of slope parameters in the 2 regimes.

Table 2.20: Results from the nonlinear model (2.28)

	Implied Volatility: VIX	Implied Volatility: VXO
Lag of volatility	0.97***	0.98***
	(0.008)	(0.008)
Intercept (Positive surprise)	0.18	0.11
	(0.33)	(0.28)
Slope (Positive surprise)	1.40	1.10
	(4.55)	(3.90)
Intercept (Negative surprise)	0.07	-0.11
	(0.26)	(0.22)
Slope (Negative surprise)	2.47	2.00
	(1.52)	(1.29)
P-value (for equality of slopes)	0.96	0.81

Notes: Newey-West HAC (2 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level. P-values in last column refer to the null of equality of slope parameters in the 2 regimes.

Table 2.21: Results from the nonlinear model (2.29)

	Change in VIX	Change in VXO
Intercept (Positive surprise)	-0.32	-0.08
	(0.44)	(0.39)
Slope (Positive surprise)	-0.02	-0.54
	(0.84)	(0.79)
Intercept (Negative surprise)	0.48	0.33
	(0.67)	(0.54)
Slope (Negative surprise)	3.35	2.62
	(2.14)	(1.89)
P-value (for equality of slopes)	0.38	0.09

Notes: Newey-West HAC (6 lags) standard errors in parentheses. . * denotes statistical significance at the 10%, ** at the 5% and *** at the 1% level. P-values in last column refer to the null of equality of slope parameters in the 2 regimes.

Chapter 3 Stock market volatility and monetary policy shocks

What drives stock market volatility? Do monetary policy actions affect the stock market? Economists, policy makers, analysts and the press endeavour to answer such questions. The literature investigating the link between macroeconomic conditions and stock returns, as well as monetary policy actions and stock returns, is evolving rapidly and reaching important results. Campbell's (2008) edited volume stands witness to the recent and growing interest academics and policy makers alike are attributing to the interaction of asset prices and monetary policy. Numerous researchers, such as Bernanke and Kuttner (2005), D'Amico and Farka (2002), Goto and Valkanov (2002), Thorbecke (1997), Rigobon and Sack (2004) and Patelis (1997) among others, relate monetary policy shocks to decreases in stock market returns.

Nonetheless, the literature has not made swift progress in identifying the factors which drive stock market volatility. While idiosyncratic stock volatility can be related to firm specific factors (or fundamentals) such as earnings announcements, dividend announcements or changes in firms' governance, analyzing the determinants of aggregate stock market volatility has been a more challenging endeavour. Shiller (1981) and Leroy and Porter (1981) argue, in what has been subsequently referred to as the "excess volatility" literature, that changes in dividends (being the cash flows from holding stocks) alone do not suffice in explaining stock market volatility. Campbell, Lettau, Malkiel and Xu (2001) analyze market, industry and idiosyncratic firm volatility. The authors note that while idiosyncratic stock volatility exhibits an upward trend, such a clear trend remains absent in market or industry level volatility. The authors proceed to a factor based decomposition of stock returns where the factors consist of the overall market return, the industry specific return and an idiosyncratic firm related term and advocate that variation in such factors can cause variation in returns. Notably, Campbell et al. (2001) also provide evidence that stock market volatility is a good predictor of output (GDP) growth and that volatility is countercyclical.

3.1 Literature Review

Understanding the factors that drive aggregate stock market volatility is of central importance in financial economics. In fact, volatility is a measure of portfolio risk that investors closely monitor as well as a central component of many derivative pricing models. Investors often demand the implementation of new trading procedures, such as circuit breakers or even policy interventions, in an aim to curb aggregate stock market volatility. Increases in stock market volatility have been linked by academics, investors and policy makers, as summarized in Shiller (1988), to numerous sources ranging from the rapid development of futures markets to the absence of circuit breakers and trading halts to behavioural interpretations. Schwert (1990) also reviews the debate about the causes of stock market volatility and distinguishes between factors that possibly drive long-term volatility such as financial leverage, operating leverage and the state of the economy and the factors that potentially cause a short-term movement in volatility, such as trading volume and trading in options and futures. Another theme that generated research is the effect of changes in margin requirements on stock market volatility as in Schwert (1989b), Hardouvelis (1988) and more recently Hardouvelis and Theodossiou

(2002). This strand of investigation has not yielded clear conclusions as to the efficacy of margin requirements in decreasing stock market volatility. Mishkin (1988) provides a central banking perspective to understanding and dealing with volatility in the stock markets.

One of the earliest studies of the macroeconomic determinants of aggregate stock market volatility is Schwert (1989a). Schwert (1989a) studies the association between the volatility of bond returns, inflation rates, money growth, industrial production growth (among other macroeconomic variables) and stock market volatility as well as how the level of macroeconomic volatility affects stock market volatility. Other research, such as Hamilton and Lin (1996) and Schwert (1989c), centers on the relationship between the business cycle and stock market volatility. Both studies report an increase in stock market volatility during recessions and attribute a large role to the state of business cycle in explaining increases in volatility. Another line of research focused on the interaction of monetary policy and stock market volatility. The advent of conditional heteroskedasticity (ARCH and GARCH) models following the seminal contributions of Engle (1982) and Bollorslev (1986) assisted researchers in investigating the relationship between stock market volatility and monetary policy as well as the effect of macroeconomic variables on conditional volatility. Engle, Ghysels and Sohn (2008) employ GARCH models with data sampled at different frequencies (MIDAS) to explore the relationship between stock market volatility and the level of economic activity, the volatility of macroeconomic variables and financial leverage. The authors detect causality running from macroeconomic variables to stock market volatility. Corradi, Distaso and Mele (2009) investigate the macroeconomic determinants of stock market volatility and volatility risk premiums while Fornari and Mele (2008) argue that financial volatility is a good predictor of economic activity.

Lastrapes (1989) uses ARCH models to investigate the effect of monetary policy on the volatility of five currencies. The author argues that changes in monetary policy regimes play an important role in determining the volatility of exchange rates. Lobo (2000) investigates the effect of monetary policy actions on stock market volatility by embedding changes in the federal funds rate and the discount rate into the mean and variance equations of generalized autoregressive conditional heteroskedascticity (GARCH) models. The author finds that stock market volatility, as well as returns, reacts to changes in the stance of monetary policy and that such a reaction is mostly pronounced when the federal funds target rate change coincides with a change in the discount rate. In a related analysis, Bomfim (2003) studies the reaction of stock market volatility to monetary policy announcements by including dummy variables into GARCH models. The author reports an increase in stock market volatility due to monetary policy events. Similar evidence as to a significant response of volatility to monetary policy surprises was provided by Chulia-Soler, Martens and van Dijk (2007) who find that surprises computed from federal funds futures data increase intra-daily realized volatility. Carr and Wu (2006) also report an increase in the Chicago Board of Exchange's (CBOE) volatility index (VIX) around Federal Open Market Committee (FOMC) meeting dates. Furthermore, Whitelaw (1994) presents evidence that the commercial paper-Treasury yield spread helps in predicting conditional stock market volatility. Indeed, the author states that "the significance of the commercial paper-Treasury spread introduces the possibility that monetary policy may also play an important role in determining return volatility" while Glosten, Jagannathan and Runkle (1993) provide evidence that the risk-free rate positively and significantly affects conditional stock volatility. In addition, Flannery and Protopapadakis (2002) provide an extensive GARCH based analysis of the macroeconomic factors and announcements that affect stock returns and volatility. Markedly, the authors maintain that among all seventeen variables they consider "only the money supply affects both the level and volatility of equity returns". All of these findings are indicative of an important role of monetary policy in determining stock market (and other assets') volatility.

In this chapter, we attempt to contribute to the literature investigating the economic sources that drive stock market volatility. Our main interest revolves around studying the effect of monetary policy shocks on stock market volatility. An extensive literature regarding the identification of monetary policy shocks from vector autoregressive (VARs) models developed in the last decade. Christiano, Eichenbaum and Evans (1996a, 1996b), Bernanke and Mihov (1998), Bernanke and Gertler (1995), Bernanke and Blinder (1992), Boivin and Giannoni (2002) are some examples drawn from this sizeable literature. Christiano, Eichenbaum and Evans (2000) provide a useful review of the literature and its accomplishments. In the first part of this paper, we build upon this literature and incorporate a monthly measure of aggregate stock market volatility computed from squared daily returns into monetary VARs. Previous efforts in the literature, such as Goto and Valkanov (2002), Thorbecke (1997) and D'Amico and Farka (2002), concentrate on augmenting such monetary VARs with a measure of aggregate stock returns in a bid to discern the effect of monetary policy shocks on returns. Our analysis is similar in spirit but centres around documenting the effect of monetary policy shocks on stock market volatility. In the second stage of our analysis, we attempt to make progress in answering a more challenging question: through which channels do Federal Reserve actions affect stock market volatility? To that end, we employ data on federal funds futures to identify monetary policy shocks directly from financial markets. Predicting Federal Reserve actions and measuring monetary policy shocks using federal funds futures data has been a subject of intense research[19]. Rudebusch (1998), Evans and Kuttner (1998) and Bernanke and Kuttner (2005) discuss methods for obtaining monthly monetary policy shocks from futures data. Once identified, monthly monetary policy shocks obtained from federal funds futures data are introduced as an exogenous variable into a VAR model that includes financial variables. We also introduce federal funds futures directly into VAR models, as advocated by Christiano, Eichenbaum and Evans (2000), to assess any differences in the impulse responses between the benchmark VAR models and the models that make use of financial market expectations. The response of the different financial variables to monetary policy shocks is then examined in order to study the channel through which monetary policy actions transmit to volatility. In total, five different VAR specifications are employed in this chapter. Since the models estimated include a large number of variables, we use the five different VAR specifications to avoid degrees of freedom problems in the estimated VAR models. This also allows us to introduce several stock market variables into the analysis and to disentangle the response of each of these variables to a monetary policy shock.

[19] Krueger and Kuttner (1996), Hamilton (2008a, 2008b), Gurkaynak, Sack and Swanson (2007), among others, provide evidence for the efficiency of the federal funds futures market and propose methods for computing daily monetary policy shocks from futures data.

The last exercise undertaken in this paper employs daily data - as opposed to the monthly data volatility series computed from squared returns used in VARs – and involves both returns and conditional volatility. Following Hamilton (2008b), we compute daily monetary policy shocks from federal funds futures data. Namely, we define slope, level and curvature surprises from near-dated federal funds futures contracts and assess their impact on stock market returns and volatility using different GARCH models.

At the time of writing, this chapter is the first to report an attempt to estimate the quantitative effect of monetary policy shocks on volatility and seek to establish the channels though which Fed actions transmit to volatility. This is also the first contribution to adapt Hamilton's (2008b)[20] methodology to study the effect of daily monetary policy shocks (defined in terms of the implied term structure of federal funds futures contracts) on stock market returns and volatility. The rest of the chapter is organized as follows: section 3.2 discusses the data used; section 3.3 introduces the econometric methodology and elaborates on our assumptions, section 3.4 presents the results and section 3.5 offers some concluding remarks.

[20] Hamilton's (2008c) research concerns the effect of daily monetary policy shocks on the housing market, but does not touch upon stock market returns or volatility. The literature has previously used only the spot-month (current month) federal funds futures contracts to examine the effect of monetary policy shocks on returns. Hamilton (2008c) and others argue that near dated federal funds futures contracts contain significantly more information and advocate their use.

3.2 Data

3.2.1 Macroeconomic and interest rate data

As is widely accepted in the monetary VAR literature, a measure of aggregate economic activity, inflation, commodity prices and a monetary policy instrument series are included as a minimum in any VAR estimation that attempts to identify monetary policy shocks. We obtain monthly data on the industrial production index, the consumer price index, the federal funds rate, the three month Treasury Bill (T-Bill) rate and the yield on Moody's rated AAA and BAA corporate bonds for the period 1983:1 to 2007:12 from the Federal Reserve of St. Louis Economic Database (FRED). Furthermore, we obtain a daily time series for the federal funds target rate from FRED for the 1988 to 2007 period. The Commodity Research Bureau's (CRB) spot commodity price index is obtained from CRB's website for the same period. Augmented Dickey-Fuller (ADF) unit root tests are performed on the series and the null of a unit root cannot be rejected at conventional levels. It is well known that ADF tests exhibit low power when the alternative is near unit root behaviour. Therefore, we follow Elliot, Rothenberg and Stock's (1996) efficient procedure to test for unit roots in the macroeconomic series of interest. Again, the results show that the null of a unit root cannot be rejected at conventional significance levels. The monetary VAR literature has differed on whether such non-stationary variables should be included in log-levels or in differenced (stationary) form. We employ the period to period growth rates in all the variables except the federal funds rate (which is taken to be the monthly average of daily figures) and the default premium. Thus, the variables used in the VAR are: industrial production growth (IPG), inflation (INF), commodity price inflation (DPCOM), the federal funds rate (FF) and the default premium (DEFP) defined as the difference between BAA and AAA rate corporate bond yields. Unit root tests, summary statistics and times series of our macroeconomic variables used at the estimation stage are reported in table 3.1.

We make the following remarks about our variables and sample choice. First, we opt to use industrial production as a measure of aggregate economic activity due to the lack of any output series which are usually sampled only quarterly. Although different, and arguably more suitable, monthly measures of economic activity [employment as in Christiano, Eichenbaum and Evans (1996a, 1996b); unemployment rate as in Bernanke and Gertler (1992)] exist; the finance literature has widely used industrial production to measure economic activity. In fact, note that Fama (1981,1990), Geske and Roll (1983) and Bodoukh et al. (1994) argue in favour of using industrial production growth as a "proxy for theoretical dividend growth" [Goto and Valkanov (2002)]. Thorbecke (1997) also employs industrial production growth to investigate the effect of monetary policy shocks on returns in VAR system. Moreover, the use of industrial production growth has not been unfamiliar to the macroeconomics literature such as Sims (1980) or Eichenbaum and Singleton (1986). Second, we use the federal funds rate as the monetary policy instrument.

Alternative monetary policy instruments have been proposed in the literature. Christiano, Eichenbaum and Evans (1996a, 1996b) opt for using non-borrowed reserves, Strongin (1995) maintains that the best indicator for monetary policy can be obtained by extracting the part of non-borrowed reserve that is orthogonal to total reserves, while

Cosimano and Sheehan (1994) argue for using borrowed reserves. This has led researchers to incorporate more than a one policy instrument in their VARs. First, Bernanke and Mihov (1998) analyze a VAR model that nests the different choices for monetary policy instruments and find that the federal funds rate can be used as a monetary policy instrument for the post-1982 period. Second, many authors [Bernanke and Gertler (1995), Bernanke and Blinder (1992), Boivin and Giannoni (2002) among many others] use the federal funds rate as the monetary policy instrument. Third, the commodity price inflation rate is included as a proxy for future inflation. Researchers typically include this variable to limit the extent of the "price puzzle"[21]. Fourth, our sample starts in 1983:1 as this date marks a federal funds rate targeting operating procedure by the Fed [Thornton (2006) and Bernanke and Mihov (1998)] and is line with other studies in the finance literature such as Goto and Valkanov (2002). Finally, the default premium, being a forward-looking variable, is included to allow the VAR to span a larger information set as noted in Goto and Valkanov (2002).

3.2.2 Stock index returns, aggregate stock market volatility and financial variables

Daily and monthly closing price data on the Standard and Poor's S&P500 index is obtained from Yahoo! Finance. The S&P500 is chosen for being a broad index that was

[21] The price puzzle is the counterintuitive finding (reported in many studies) that indicates an increase in the price level following a contractionary monetary policy shock.

analyzed in many studies like Schwert (1989a). Let P_{it} denote the closing index price on day i of month t. Continuously compounded daily returns are computed as[22]:

$$r_{it} = \ln(P_{it}) - \ln(P_{(i-1)t}) \tag{3.1}$$

We proceed to compute a monthly measure of aggregate stock market volatility by summing daily demeaned squared returns as in Schwert (1989a), Campbell et al. (2001) or Ludvigson and Ng (2007) :

$$\hat{\sigma}_t^2 = \sum_{i=1}^{D} \tilde{r}_{it}^2 \tag{3.2}$$

where $\tilde{r}_{it}$ denote the returns obtained by subtracting the average return over the sample, and i =1,2...,D denotes the number of days in month t. In our subsequent VAR estimation, we refer to the square root of (3.2) as VOL. As noted in Andersen et al. (2003) and reiterated in Ludvigson and Ng (2007), such an estimator for volatility possesses good properties in that it is an unbiased estimate of actual volatility.

We also compute the monthly excess returns on the S&P500 index by subtracting the three months T-Bill rate (considered as a proxy for the risk-free rate) from the monthly returns on the S&P500. The resulting series is referred to as ER in our VAR estimation. Finally, monthly dividend yield data on the S&P500 index is obtained from

[22] Monthly returns on the S&P 500 are computed from monthly closing price data in a similar manner but with monthly changes instead.

Datastream. The time series of our financial variables are reported in Figure 3.1 (on p.113 at the end of this chapter).

3.2.3 Federal funds futures contracts and market-based monetary policy shocks

Federal funds futures, officially known as 30-day interest rate futures, are interest rate futures that settle on the average of the month's overnight funds rate. These futures contracts trade on the Chicago Board of Trade (CBOT) and contracts ranging from the current (spot) month to several months ahead exist. Krueger and Kuttner (1996), Gurkaynak, Sack and Swanson (2007) and Hamilton (2009) provide evidence for the efficiency of federal funds futures market[23]. Given the evidence in favour of the efficiency of federal funds futures, several authors have proposed extracting monthly and daily measures of monetary policy shocks from these contracts.

A daily dataset of federal funds futures closing prices was obtained from the Commodity Research Bureau (CRB)[24]. Our dataset spans the October 1988 to December 2007 period and contains futures contracts of different maturities. Following Bernanke and Kuttner (2005), we compute a monthly monetary policy shocks series from futures prices. Let $f_{i,t}^{1}$ denote the one-month ahead implied rate for day i of month t, and

[23] We also test for unbiasdness and efficiency of federal funds futures in predicting the federal funds funds rate. Regression results indicate that unbiasdness and efficiency cannot be rejected for the one-month-ahead contract used in this paper.

[24] I would like to thank my supervisor for providing me with this data.

$\xi_{i,t}$ denote the federal funds target rate for day i in month t. A monthly monetary policy shocks series is computed from market data as in Bernanke and Kuttner (2005):

$$MP_t = \frac{1}{D}\sum_{i=1}^{D}\xi_{i,t} - f_{D,t-1}^{1} \tag{3.3}$$

where $f_{D,t-1}^{1}$ denotes the one month ahead futures rate on the last (Dth) day of month t-1.

We also compute daily monetary policy shocks series from federal funds futures rates. Gurkaynak (2005) introduced the possibility of defining several different monetary policy shocks from near dated futures contracts. Similarly to the term structure of interest rate literature[25], the author suggests defining timing, level and slope surprises from current month, two-month and four-month ahead futures rates. The author argues that the shocks defined in this manner are more informative about future interest rate expectations than current month surprises. Hamilton (2008a, 2008b) further developed this approach by providing evidence that daily changes in current, one-month and two-month ahead futures contracts can be used for computing several monetary policy shocks. Let $f_{i,t}^{0}$ the implied rate on day d of month t from the current month contract, $f_{i,t}^{1}$ the implied rate on day d of month t from the one month contract and $f_{i,t}^{2}$ the implied rate on day i of month t from the two month contract. Then, we follow Hamilton (2008b) and define daily

[25] In the literature on the term structure of interest rates, researchers resorted to a principal component analysis of the factors that affect the yield curve. See for example Litterman and Scheinkman (1991). The first three factors found to affect the yield curve have been labelled the slope, level and curvature.

monetary policy shocks (slope, level and curvature) in terms of the implied term structure of federal funds futures contracts as:

$$s_{it} = f_{it}^{1} - f_{(i-1)t}^{1} \tag{3.4}$$

$$l_{it} = (f_{it}^{2} - f_{(i-1)t}^{2}) - (f_{it}^{1} - f_{(i-1)t}^{1}) \tag{3.5}$$

$$c_{it} = (f_{it}^{2} - f_{(i-1)t}^{2}) - 2(f_{it}^{1} - f_{(i-1)t}^{1}) + (f_{it}^{0} - f_{(i-1)t}^{0}) \tag{3.6}$$

These computations are undertaken for all days of the month except the first. On the first day of the month, the rate from the nearest futures contract on the last day of the month is used instead of $f_{(i-1)t}^{2}$, $f_{(i-1)t}^{1}$, and $f_{(i-1)t}^{0}$. Hamilton (2008b) notes that a coefficient associated with the slope surprise measures the impact of a one basis point increase per month in the federal funds rate for the next two months. Furthermore, the author interprets a coefficient associated with the level surprise as measuring the effect of a joint one basis point increase in market expectations of the federal funds rate from the current month out to two months. The coefficient on the curvature shock, again according to Hamilton (2008b), measures the effect of an increase in funds rate that is expected to be faster between next month and the following relative to the increase between the current month and the next.

3.3 Methodology

3.3.1 (Structural) Vector Autoregressions (SVARs)

In this section, we introduce our econometric methodology and clarify the assumptions underlying our analysis. A good treatment of vector autoregressive models is given in Hamilton (1994), Lutkepohl (2006) or Boivin and Giannoni (2002) and we follow a similar exposition here. Denote by Y_t an $(n \times 1)$ vector of macroeconomic and financial time series of interest. A pth-order vector autoregressive [VAR(p)] model relates each series in Y_t to p of its own lags (and possibility contemporaneous) as well as p lags of all other variables in the system:

$$Y_t = a + A_1 Y_{t-1} + A_2 Y_{t-2} + \cdots + A_p Y_{t-p} + u_t \tag{3.7}$$

Or alternatively,

$$A(L) Y_t = a + u_t \tag{3.8}$$

Where a denotes an $(n \times 1)$ vector of constants, $A_1 \cdots A_p$ are $(n \times n)$ matrices of coefficients, L denotes the lag operator, and u_t an $(n \times 1)$ vector of mean zero residuals with an $(n \times n)$, positive definite variance-covariance matrix $E(u_t u_t^T) = \Sigma_u$.

Since there is no economic structure underlying (3.7), the vector of residuals resulting from estimating (3.7) is typically referred to as "reduced form" disturbances. Since no obvious economic interpretation can be given to these shocks, the use of more elaborate structural models has been undertaken in the literature. A structural vector autoregression (SVAR) can be written as:

$$B_0 Y_t = b + B_1 Y_{t-1} + B_2 Y_{t-2} + \cdots + B_p Y_{t-p} + \varepsilon_t \tag{3.9}$$

Where again $B_0 \cdots B_p$ are $(n \times n)$ matrices (B_0 is assumed to be non-singular and is sometimes referred to as the contemporaneous impact matrix, normalized to have ones on the diagonal), and ε_t are the structural innovations (shocks) that the researcher seeks to identify.

Furthermore, suppose that the variance-covariance matrix of the structural shocks is $E(\varepsilon_t \varepsilon_t^T) = \Sigma_\varepsilon$, where Σ_ε is an $(n \times n)$ variance-covariance matrix of the structural innovations that is assumed to be diagonal (commonly normalized to be the identity matrix). Note that (3.7) and (3.9) are related by: $a = B_0^{-1} b$; $A_j = B_0^{-1} B_j$ for j=1,...,p and $u_t = B_0^{-1} \varepsilon_t$. Thus, the reduced form residuals and the structural shocks are related by $\varepsilon_t = B_0 u_t$, and it is clear that the structural shocks are a linear combination of the reduced form residuals. It follows that, the variance-covariance matrices of the structural and reduced form shocks are related by $\Sigma_u = B_0^{-1} \Sigma_\varepsilon (B_0^{-1})^T$.

Imposing identifying assumption on the contemporaneous impact matrix has been common in the literature [Sims (1986), Bernanke (1986) among others]. In the context of identifying monetary policy shocks, the most common (though by no means the only) identification strategy imposes a recursive (or Wold recursive) assumption on a model such as (3.9). Imposing a recursive ordering assumption would imply that the contemporaneous impact matrix B_0 is lower triangular (commonly normalized to have ones on the main diagonal) and provides enough restrictions to recover the structural shocks from the reduced form disturbances. We note that this assumption is equivalent to a Choleski orthogonolization of the variance-covariance of the residuals from (3.8). Such an assumption implies that variables that appear earlier in the ordering of the vector Y_t affect the other variables in the system contemporaneously, while the variables coming later in the ordering affect the variables ranked prior to them only with a lag. Several studies have used such an identification scheme to recover monetary policy shocks [Bernanke and Blinder (1992), Bernanke and Gertler (1995), Christiano, Eichenbaum and Evans (1996a,1996b), Boivin and Giannoni (2002), Goto and Valkanov (2002) among others].

We use the following two orderings for the variables in our VARs:

$$Y_t = [IPG \quad INF \quad DPCOM \quad FF \quad DEFP \quad VOL] \tag{M1}$$

$$Y_t = [IPG \quad INF \quad DPCOM \quad FF \quad DEFP \quad ER \quad VOL] \tag{M2}$$

We refer to the first of these models as (M1) and to the second model as (M2). The orthogonalized shocks from the federal funds rate equation, ε_t^{ff}, are identified as the monetary policy shocks. Let y_t denote one of the time series in the VAR. The impulse responses to a monetary policy shock, $\frac{\partial y_{t+s}}{\partial \varepsilon_t^{ff}}$, are then computed from our model along with their associated Monte Carlo (with 2500 replications) confidence bands.

Our ordering scheme implies the following assumptions: The federal fund rate responds contemporaneously to all the macroeconomic variables while macroeconomic variables respond to a change in the federal funds rate only with a time lag. Our identification strategy allows stock market volatility to respond contemporaneously to all the variables in the system, while the monetary authorities do not respond contemporaneously to stock market volatility. Previous research investigating the response of the Fed to the stock market yields no consensus. On the one hand, Rigobon and Sack (2004) and D'Amico and Farka (2006) note a positive response of the Fed to the stock market. These authors, among others, emphasize the importance of financial variables in the monetary transmission mechanism and argue that such forward-looking financial variables can be viewed as early indicators of future output growth and inflation by the monetary authority. Consequently, the Fed might respond to financial variables inasmuch as they signal future inflation or deflation. On the other hand, Bernanke and Gertler (1999, 2001) report using simulations and estimation results that central banks do not respond to asset price movements over and above their response to inflation and output growth. Indeed, the authors state that central banks should monitor and respond to asset price fluctuations to the extent that they reveal any new information regarding the path of future inflation and output growth[26]. Fuhrer and Toottell (2008) estimate forward looking Taylor rules using "Greenbook" forecasts[27] and find no evidence that the Fed responds to asset prices beyond its response to inflation and output gaps. Our identifying assumptions preclude any response by the Fed to the stock market, and the results obtained should be viewed in light of this assumption.

3.3.2 Incorporating information from financial markets into VARs

Rudebusch (1998) criticized the recursive VAR methodology outlined in section 3.3.1 on several grounds. Among the criticisms expressed by the author we note the following: (i) VARs incorporate small information sets, (ii) monetary policy shocks derived from recursive VARs, as in section 3.3.1, display weak correlation with monetary policy shocks derived from federal funds futures and (iii) unanticipated changes in the federal funds rate may be indicative of an endogenous response by the Fed to the economy.

Christiano, Eichenbaum and Evans (2000) propose incorporating federal funds futures data into VAR models directly. Given that federal funds futures incorporate market participants' expectations of Fed actions, their inclusion allows for a larger information set to be spanned by the VAR. Christiano, Eichenbaum and Evans (2000) observe that the central interest of VAR analyses lies in the dynamic (impulse) responses generated by these models and that "policy shock measures can display a low correlation,

[26] There also exists a large body of literature that investigates the role of asset markets in the transmission of monetary policy. Mishkin (2007) provides an account of this literature.

[27] Greenbook forecasts are forecasts prepared by Federal Reserve's Board of Governors Staff before FOMC meetings.

while not changing inference about the economic effects of monetary policy shocks". Therefore, the authors suggest replacing the federal funds rate in models (M1) and (M2) by the difference between the federal funds rate and the lag of the one-month-ahead federal funds futures rate. The goal of replacing the federal funds rate in the VARs is to inspect whether the dynamic responses of the variables are affected by the inclusion of federal funds futures directly into the model.

Let $FM_t = FF_t - f_{t-1}^1$ be the difference between the federal funds rate in month t and the one-month-ahead futures rate in month (t-1). Then, models (M1) and (M2) are replaced by the following VAR ordering:

$$Y_t = [IPG \quad INF \quad DPCOM \quad FM \quad DEFP \quad VOL] \qquad \text{(M3)}$$

$$Y_t = [IPG \quad INF \quad DPCOM \quad FM \quad DEFP \quad ER \quad VOL] \qquad \text{(M4)}$$

The monetary policy shocks in (M3) and (M4) are the orthogonalized disturbances from the FM equation. As in Christiano, Eichenbaum and Evans (2000) we use the FM monetary policy shocks to investigate changes in the magnitude and sign of the response of volatility, returns, inflation and industrial production. Since federal funds futures data are available only starting 1988, models (M3) and (M4) have the additional virtue of precluding the October 1987 stock market crash from the sample.

3.3.3 Financial vector autoregressions

The second step of our analysis involves investigating the more challenging question regarding the channels through which monetary policy shocks affect stock market volatility. To this end, we refer back to the contribution of Campbell and Shiller (1988) in which they provide a decomposition of unexpected returns in terms of news regarding dividends and expected future returns. This contribution was further extended by Patelis (1997) and Bernanke and Kuttner (2005) who provide a decomposition (based on a log-linearization) of excess returns explicitly demonstrating the role of monetary policy in determining excess returns. Denote the excess return at time t by er_t. Then, the decomposition due to Campbell and Shiller (1988), and extended by Patelis (1997) and Bernanke and Kuttner (2005) can be expressed as[28]:

$$er_t - E_t(er_{t+1}) = (E_{t+1} - E_t)\left\{\sum_{j=0}^{\infty}\rho^j \Delta d_{t+1+j} - \sum_{j=1}^{\infty}\rho^j rr_{t+1+j} - \sum_{j=1}^{\infty}\rho^j er_{t+1+j}\right\} \tag{3.10}$$

where $E_t(.)$ denotes the conditional expectation operator (given the information set at t), d_t denotes the log dividend at time t, rr_t denotes the real interest rate at time t and ρ is a constant (generally set close to one) which equals the ratio of ex-dividends to the cum-dividends. As convincingly argued in Patelis (1997) and Bernanke and Kuttner

[28] We abstract from indexing by both the day i and the month t for notational simplicity and include only a subscript t to denote time. For a derivation of this decomposition, and the assumptions underlying it, the reader is referred to Campbell and Shiller (1989), Patelis (1997) or Bernanke and Kuttner (2005) or for a textbook demonstration to Campbell, Lo and Mackinlay (1997) or Cochrane (2005).

(2005), the decomposition in (3.10) illustrates the various channels through which monetary policy can affect returns. Contractionary monetary policy can exert an effect by decreasing expected future dividends, increasing the expected excess returns or thorough an increase in the future expected real interest rate used to discount dividends (the cash flows from holding stocks).

Campbell and Ammer (1993) were the first to propose the use of VARs containing financial variables, such as dividend yields, real interest rates and bond spreads, to analyze the response of returns to different shocks in line with what (3.10) illustrates [Patelis (1997) follows a similar analysis]. Black (1976) first documented a negative association between returns and volatility which has been subsequently referred to as the "leverage effect", while other researchers [Bernanke and Kuttner (2005), Goto and Valkanov (2002)] point to a decrease in returns to a monetary policy shock. We postulate that a monetary policy shock increases stock market volatility because it decreases returns. In order to investigate this hypothesis, we employ a VAR that only incorporates financial variables in the spirit of Campbell and Ammer (1993) or Patelis (1997). However, we adopt Bernanke and Kuttner's (2005) methodology by embedding the market based monetary policy shocks, derived from federal funds futures contracts as defined in (3.3), as an exogenous variable in our VAR. The aim of this model is to investigate the dynamic reponses of the different financial variables to a monetary policy shock. Let Z_t denote an $(n \times 1)$ vector of financial variables. Our VAR, following Bernanke and Kuttner (2005), is defined as:

$$Z_t = AZ_{t-1} + \Phi MP_t + \omega_t \tag{3.11}$$

In this setting, the $(n \times 1)$ vector Φ summarizes the contemporaneous response of Z_t to an exogenous monetary policy shock identified from futures data. As described in Bernanke and Kuttner (2005), writing the VAR in this manner allows for decomposing the reaction of the vector Z_t into components relating to monetary policy shocks and components that relate to variables other than policy. Such an approach has also been used to evaluate the effect of monetary policy shocks in the macroeconomics literature [see Favero and Bagliano (1999) or Faust, Swanson and Wright (2004)]

We include in the vector Z_t the following variables: the monthly excess returns on the S&P500 defined as the difference between the monthly continuously compounded returns and the three-month's treasury bill rate, the real interest rate defined as the difference between the three month's bill rate and the consumer price inflation, the dividend yield on the S&P500, the change in the three-month's T-Bill rate and monthly volatility as defined in (3.2). Due to the availability of federal funds futures data and the construction of the shocks in (3.3), the estimation period is confined to the 1988:10 to 2007:12. By investigating the impulse responses of financial variables to an exogenous monetary policy shock as in (3.11), the researcher can gain insight into the possible channels through monetary policy affects volatility. For instance, a monetary policy shock can drive excess returns down, and thus volatility up. due to the negative return-volatility correlation. One additional benefit of the model in (3.11) is that it allows us to check the robustness of our results obtained from the VAR model in sections 3.3.1 and 3.3.2. In fact, by confining the sample to the 1988:10 to 2007:12 period, we exclude the October

1987 stock market crash from our sample. This is a useful assessment of the effect the stock market crash on our results.

3.3.4 Daily analysis with a conditional heteroskedasticity model

The VAR analyses proposed in sections 3.3.1, 3.3.2 and 3.3.3 made use of monthly data. We proceed with a higher frequency, daily analysis that uses conditional heteroskedasticity models to investigate the effect of daily monetary policy shocks on both stock market returns and volatility. In contrast to the previous VAR analyses, stock market volatility is now treated as a latent variable and we incorporate into our model the daily monetary policy shocks (3.4), (3.5) and (3.6) derived from the implied term structure of federal funds futures contracts. In order to investigate the effect of slope, level and curvature surprises on S&P500 return and volatility, we modify Nelson's (1991) EGARCH (1,1) [i.e. Exponential GARCH] by including the monetary policy surprises into the mean and the variance equations. Specifically, we have:

$$r_{tt} = \omega + \phi r_{(t-1)t} + \beta_s s_{(t-1)t} + \beta_l l_{(t-1)t} + \beta_c c_{(t-1)t} + \sqrt{h_{tt}}\varepsilon_{tt}$$

$$\varepsilon_{tt} \sim iidN(0,1)$$

$$\log(h_{tt}) = \varpi + \alpha_1 \frac{\left|\varepsilon_{(t-1)t}\right|}{\sqrt{h_{(t-1)t}}} + \gamma_1 \log(h_{(t-1)t}) + \lambda_s s_{(t-1)t} + \lambda_l l_{(t-1)t} + \lambda_c c_{t(-1)t} \tag{3.12}$$

where h_{it} denotes the daily conditional variance series. The goal of incorporating the three daily monetary policy shocks into the EGARCH (1,1) model is to check both the significance, direction and magnitude of the shocks on daily returns and volatility We opt for using the EGARCH model to circumvent the need for constraining the parameters (associated with the three shock series) in order to ensure a positive volatility series. Our daily data set covers the 01/01/1994 to 31/12/2007 period for a total of 3,504 observations. We elect to use post-1994 futures data as in Kuttner (2008) who argues that, in a commentary on Hamilton (2008a), interpreting daily changes in the futures price prior to 1994 as reflecting monetary policy shocks (or news regarding monetary policy) is complicated by uncertainty as to whether the Fed changed its target funds rate or not.

3.4 Results

3.4.1 Comparing VAR and market-based monetary policy shocks

We begin our results section by a swift comparison of the two monetary policy shock series used in the monthly analyses of this paper. Figure 3.4 (p. 115) displays the monthly time series of the federal funds rate, the federal funds target rate and the one-month-ahead federal funds futures rate. Figure 3.5 (p. 115) displays the time series of the orthogonalized VAR shocks corresponding to the VARs federal funds rate equation as well as the monthly time series of the monetary policy shocks obtained from futures data as in (3.3). The time series of the two shock series exhibit significant co-movement, in the sense that both series tend to agree on qualifying monetary policy as expansionary or contractionary.

Despite such co-movement, the correlation coefficient between the two series is 0.37. In light of a similarly small correlation coefficient, Rudebusch (1998) argues against using VARs to characterize monetary policy shocks. The author maintains that since federal funds futures markets are forward-looking and efficiently incorporate all market participants' expectations about Fed actions, monetary policy shocks obtained from these contracts is superior to VAR based monetary policy shocks (since VARs necessarily incorporate only a relatively small information set due to degrees of freedom considerations associated with larger models). However, Sims (1997), Evans and Kuttner (1998) and Christiano, Eichenbaum and Evans (2000) address Rudebsuch's (1998) criticisms. Specifically, Evans and Kuttner (1998) argue that the correlation metric might not be the best benchmark to judge the ability of VARs to characterize monetary policy. Christiano, Eichembaum and Evans (2000) argue that monetary policy shocks can have low correlation all the while yielding similar impulse responses for the variables of interest. By using both shock measures in this paper, we ensure that our results are not subject to the identifying assumptions we impose on the VAR.

Another interesting feature that emerges from casual inspection of the two shock series is the significant increase in the market's ability to infer Fed actions starting around 2003. Figure 3.5 (p. 115) shows a decrease in the variability of monetary policy shocks derived from futures data when compared to those obtained from VARs. This observation has previously been noted in several studies in the literature such as Swanson (2006), Hamilton (2008b) or Swiston (2007). In fact, Hamilton (2008b) notes the decrease in futures forecast errors (that is when futures are used to predict the federal funds rate) during the 2003 to 2007 period while Swiston (2007) relates this improved accuracy to better communication of monetary goals and objectives by the Federal Open Market Committee (FOMC).

3.4.2 Results from recursive VARs

In this section, we discuss the results obtained from estimating the VAR models (3.9) of section 3.3.1. Our main interest is to document the response of aggregate stock market volatility and aggregate stock market returns to a monetary policy shock obtained from orderings (M1) and (M2) in section 3.3.1. The VAR models are estimated with two lags as determined by the Akaike information criterion (AIC). Figure 3.6 (p. 116) displays the response of industrial production growth (IPG), inflation (INF) and stock market volatility (VOL) to a one standard deviation contractionary monetary policy shock (or an increase in the federal funds rate) using the first model (M1) while Figure 3.7 (p. 117) presents the same set of responses using model two (M2) (i.e. the VAR augmented with returns).

Both models show an increase in stock market volatility of around 1.5% following a monetary policy shock. The dynamics of the response of volatility to a monetary policy shock are such that volatility reaches a peak in one month and starts slowly to revert back to its original level. We note that the increase in volatility is persistent and dies out only after ten periods. The VAR model that embeds returns (M2) illustrates interesting joint dynamics of excess returns and aggregate stock market volatility to a monetary policy shock. Our results indicate that a contractionary monetary policy shock decreases excess stock returns by 2% while simultaneously increasing stock market volatility by 1.5%. We note that the negative response of excess returns to a monetary policy shock remains for around five months and that this effect is the largest contemporaneously. Similar results regarding the effect of a monetary policy shock on excess returns was found by Goto and Valkanov (2002). The effect of the policy shock on excess returns dies out in around five months. Again, the dynamic response [obtained from model (M2)] of volatility indicates an increase in volatility that reaches its peak in one month and slowly reverts back to its initial level.

We attribute the following interpretation to the effect of a monetary policy shock: due to the tight correlation between the federal funds rate and several short-term interest rates, a monetary policy shock drives interest rates up and hence increases the discount rate at which future cash flows (future dividends) are valued. This, in turn, decreases contemporaneous excess returns. As widely recognized since Black's (1976) contribution, negative return "news" spurs an increase in volatility. Whether the effect of monetary policy shocks on interest rates is the major determinant of the increase in volatility (and decrease in returns) can be better investigated using the model (3.11) and we will turn to this question in a later subsection of this chapter (see section 3.3.4).

Further examination of the impulse responses reported in figures 3.6 and 3.7 reveals an initial increase in industrial production growth. Industrial production growth starts to fall with a lag of around three months. We find that inflation increases following a contractionary monetary policy shock even though commodity price inflation is included in our models. This finding has been known in the literature as the "price puzzle" and is commonly found in VAR analyses [such as Goto and Valkanov (2002) for monthly data, Ludvigson, Steindel and Lettau (2002) for quarterly data]. In this context, we note that several studies in finance reported a negative correlation between excess returns and inflation [Fama and Schwert (1977)] and other studies have attempted to provide an interpretation for this robust empirical feature [Fama (1981), Marshall (1992), Boudoukh, Richardson and Whitelaw (1994), Geske and Roll (1993), Goto and Valkanov (2002) and Kaul (1987)]. Given the negative return-inflation correlation, our VAR results can be interpreted in the following manner. Following a contractionary monetary policy shock, stock market participants expect a decrease in aggregate economic activity (i.e. industrial production) leading to an immediate decrease in excess stock returns. Due to the decrease in excess returns, volatility increases given the negative excess returns-volatility correlation. Insofar as this correlation is accurate, an initial increase in inflation following a contractionary policy shock will lead to a decrease in stock returns and an increase in volatility.

Several possible different channels contribute to the simultaneous increase in volatility and decrease in stock returns: the aggregate economic activity channel, the inflation channel and the interest rate (discount rate) channel. Granger causality tests reported in tables 3.2 and 3.3 fail to detect causality running from the macroeconomic variables to the financial variables. Such a finding is not surprising in view of similar results previously reported in the literature. Indeed, Ghysels, Engle and Sohn (2008) indicate the need to refine the volatility measure used in (3.2) by decomposition into long- and short-run volatility components, in order to detect Granger causality running from a macroeconomic series to stock market volatility.

3.4.3 Results from incorporating financial market information into VARs

Figures 3.8 and 3.9 show the effect of a one standard deviation FM monetary policy shock, as defined in section 3.3.2, for models (M3) and (M4) respectively. The VARs is estimated with 2 lags as selected by the AIC over the period 1988:12 to 2007:12. The response of stock market volatility and returns are similar, quantitatively and qualitatively, to those found with models (M1) and (M2). The results confirm that the dynamic response of the variables of interest is altered only slightly once federal funds futures are included directly in the VAR: the response of volatility peaks in one month and volatility displays a persistent response to monetary policy shocks.

One notable difference occurs regarding the response of excess returns to an FM monetary policy shock. When models (M3) and (M4) are used, the initial negative response of excess returns is significantly larger, standing at an around a 9% initial decrease as compared with a decrease of 2% from models (M1) and (M2). With models (M3) and (M4), the decrease in excess returns persists for only one month following which excess returns become slightly positive in two months. This response of excess returns differs from the relatively more persistent (yet smaller) decline is excess returns suggested by models (M1) and (M2). The response of economic variables to an FM monetary policy shock resembles their counterparts from (M1) and (M2). Industrial production growth declines with a lag of two months while the "price puzzle" (as indicated by the initial increase in inflation) remains[29]. Granger causality tests (not

[29] One way to eliminate the price puzzle would be to include inflation expectations into the VAR. Inflation expectations could be measured using the University of Michigan surveys or as the difference between the rates on Treasury Inflation Protected Securities (TIPS) and appropriate bonds. A more recent and

reported) fail to detect any causality running from the macroeconomic variables to stock market variables in the dynamic setting of the models examined.

3.4.4 Results from financial VAR

The impulse responses of the different financial variables of model (3.11) to a percentage point federal funds rate surprise are reported in Figure 3.10 (p. 120-121). We note that a federal funds rate surprise computed from federal funds futures data generates a similar response of returns and volatility as with the VAR models (M1) and (M2) discussed above. Namely, we find that a federal funds rate surprise (or a monetary policy shock) decreases returns and increases volatility contemporaneously. However, we note two differences. The magnitude of the contemporaneous effect of the shock on excess returns when the monetary shock is derived from futures data is -9.6% and is considerably larger than the response of excess returns to a monetary policy shock derived from models (M1) and (M2) The effect of a federal funds rate surprise on excess returns lasts for two periods following which returns are close to zero or slightly positive seven periods ahead.

The response of volatility also displays similar patterns as compared to previous results in that volatility peaks in one period at 0.8%. However, this response dies out much more quickly (in 3 periods) than the response of volatility displayed in models (M1) and (M2). In turn, this implies a shorter time for investors to exploit the increase in volatility than with models (M1) and (M2). The dividend yield shows a persistent

successful approach is the Factor Augmented Vector Autoregressions (FAVAR) of Bernanke, Boivin and Eliasz (2005).

increase that does not die out twenty periods ahead, while the real interest rate and the change in the bill rate display a relatively shorter lived increase.

We note the similarity of the results obtained here to those of Bernanke and Kuttner (2005) who use the Center for Research in Security Prices (CRSP) value weighted index and report an initial decrease of 11.6% in stock returns due to a percentage point federal funds rate surprise. The impulse responses obtained from model (3.11) and discussed in this section can be interpreted as relating the initial decrease in excess stock returns and increase in stock market volatility to an increase in the real interest rate (an interest rate effect). Once the effect of the increase in the real interest rate dies out, the dynamic behaviour of excess returns and volatility is dominated by the increase in the dividend yield causing excess returns to increase and volatility to decrease. We note that the results obtained with model (3.11) use data for period November 1988 to December 2007 and thus the stock market crash of October 1987 is excluded from the sample. Thus, the result that volatility is increased and excess stock returns are decreased contemporaneously due to a monetary policy shock does not hinge upon monetary policy shock identification or the specific sample period used for estimation.

Finally, we note the conformability of the dynamic responses of excess returns obtained from models (M4) and (3.11). Both models show an initial decrease in returns, lasting for around two periods, following which excess returns become slightly positive before eventually returning to their original level. In this sense, we view the results from the VAR models that incorporate federal funds futures data directly as intermediary between VAR models that do not include federal funds futures data [models (M1) and (M2)] and those that incorporate them as exogenous variables [model (3.11)].

3.4.5 Results from conditional heteroskedasticity model

The results from maximum likelihood estimation of the EGARCH (1,1) of section 3.3.4 are discussed in this section. Table 3.4 reports the results of our estimation which assumes conditional normality of the residuals. The results indicate that the three daily monetary policy shocks derived from the term structure of federal funds futures contracts decrease returns while they increase volatility. This is consistent with the hypothesis that volatility is increased due to a decrease in returns and is also consistent with the results obtained from monthly VAR models. In more detail, we find that daily monetary policy shocks do not significantly affect returns; as the three shock series do not enter the mean equation of the EGARCH model significantly. Two of the shocks significantly drive volatility upwards. Namely, we find that the slope and level surprises increase volatility while the curvature surprises do not. In turn, this implies that when market expectations of interest rate are increased or that when the market becomes aware of an impending increase in federal funds rate, volatility is significantly increased. In contrast, the speed at which market participants expect changes in interest rates to occur, measured by the curvature surprise, increase volatility but this effect is insignificant.

We note that the largest increase in volatility is due to the slope surprise followed by level innovations while curvature surprises have the smallest impact on volatility. We interpret this result as reflecting the fact that slope surprises capture a certain federal target rate increase. Level surprises relate only to market expectations of federal funds rate increases which might not materialize.

3.4.6 Robustness of the results

The robustness of the results obtained from the different models estimated in this chapter is investigated. First, relating to models (M1) and (M2) of section 3.3.1, we change the ordering of the variables and place the federal funds rate as first in our ordering. Christiano, Eichenbaum and Evans (1996b) refer to this ordering as the "monetary policy first" ordering and we note that this ordering implies different assumptions about the reaction of monetary policy to macroeconomic variables. The implication of placing the federal funds rate first in the ordering of the VAR is that monetary policy can affect the other macroeconomic variables in the VAR contemporaneously. With this new ordering, we find that the response of volatility and excess returns remains similar to our earlier findings.

Another change relating to models (M1) and (M2) involves removing the default premium (DEFP) from the estimated models. Again, we find that our results regarding the responses of excess returns and volatility remain largely similar. In terms of econometric testing, we perform stability tests for our VAR, a Lagrange Multiplier test for autocorrelation of the residuals as well as a test of joint multivariate normality of the residuals. Models (M1) and (M2) are stable and display no residual autocorrelation but the null of joint multivariate residual normality is rejected.

With regard to the financial VAR, Bernanke and Kuttner (2005) indicate that monthly futures based monetary policy shocks in (3.3) exhibit some response to economic news prior to 1994 (while this is not the case for post-1994 data). Therefore, we re-estimate model (3.11) using post-1994 data and find that our impulse responses are qualitatively similar to those reported in the paper with small changes quantitatively.

Finally, we re-estimate the EGARCH (1,1) model under different distributional assumptions. One of the stylized facts in financial economics is that daily returns display excess kurtosis. Estimation of the EGARCH (1,1) is undertaken under different distributional assumptions [Generalized Error Distribution (GED) and student t distribution] to account for this property of daily returns. When the GED distribution is used, only the level surprise significantly affects volatility (the other surprises having a positive but insignificant effect), while the results remain closer to the benchmark case (normal distribution) when the t distribution is used.

3.5 Concluding remarks

As indicated in section 3.1, the principal aim of this chapter has been to estimate, within the framework of well-specified dynamic models, the quantitative effect of monetary policy shocks on stock market volatility. The dynamic models specified incorporate, by design, financial variables as a means of establishing the channels through which Fed actions are transmitted to stock market volatility.

It has been established that stock market volatility is consistently negatively associated with excess stock market returns arising from a monetary policy shock. It has also been established that three channels contribute to a decrease in returns and an increase in volatility. These are the interest rate channel, the economic activity channel and the dividend channel. The dynamics of the inter-relations between these channels contained in the results of this chapter indicate: (i) a relatively short-lived interest rate effect followed by (ii) the effect of dividends which then impinge on the level of returns and volatility.

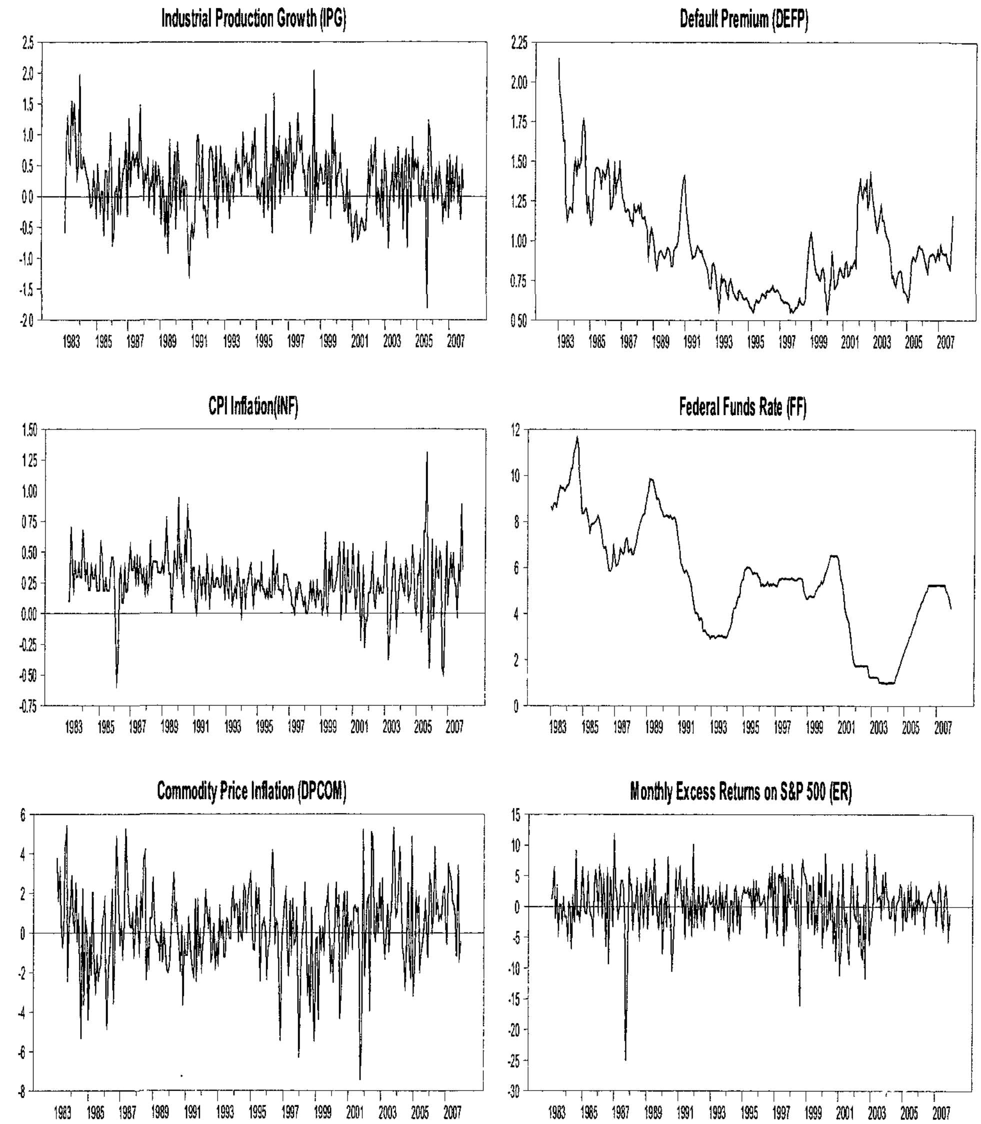

Figure 3.1: The time series of the variables used in the VAR estimation.

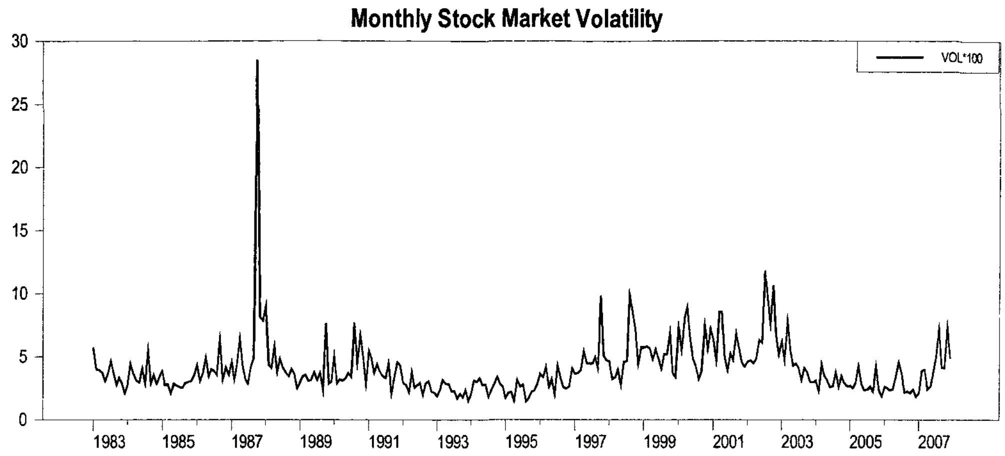

Figure 3.2: Time Series of monthly stock market volatility computed from daily returns as in (3.2).

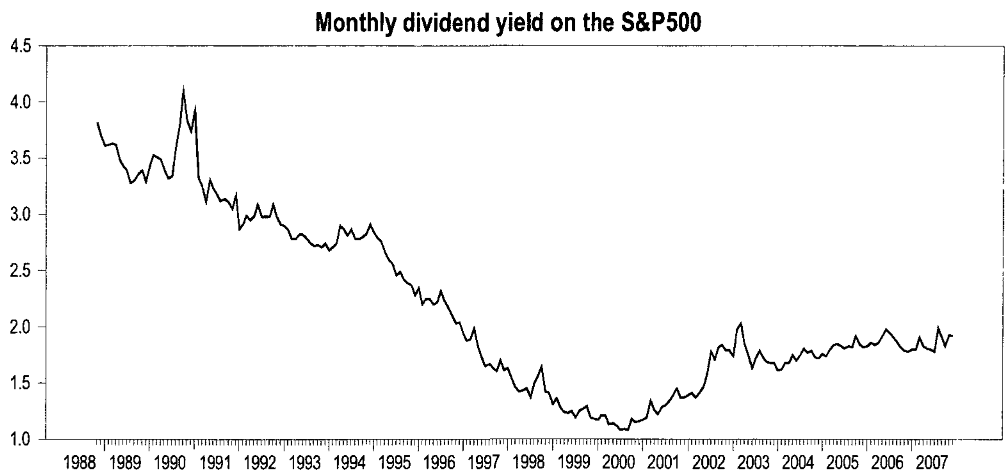

Figure 3.3: Time series of the monthly dividend yield on the S&P500.

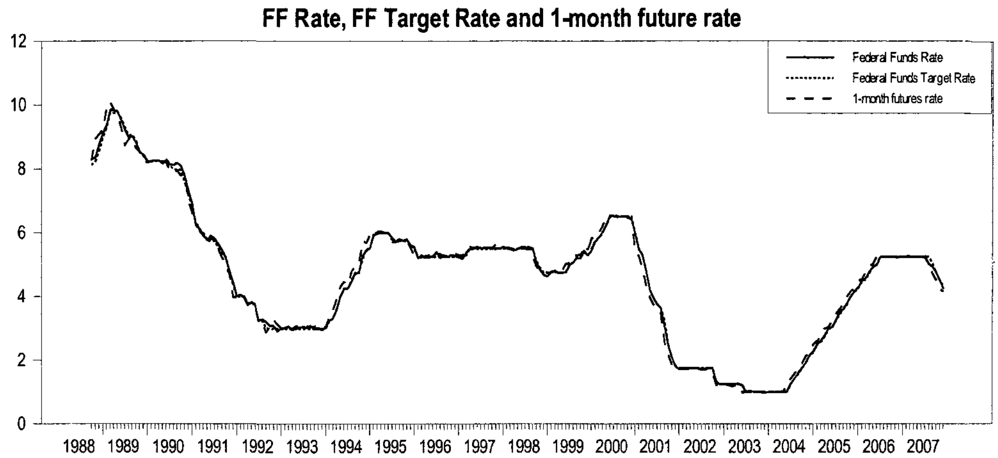

Figure 3.4: Monthly time series of the federal funds rate, the federal funds target rate and the one-month-ahead futures rate

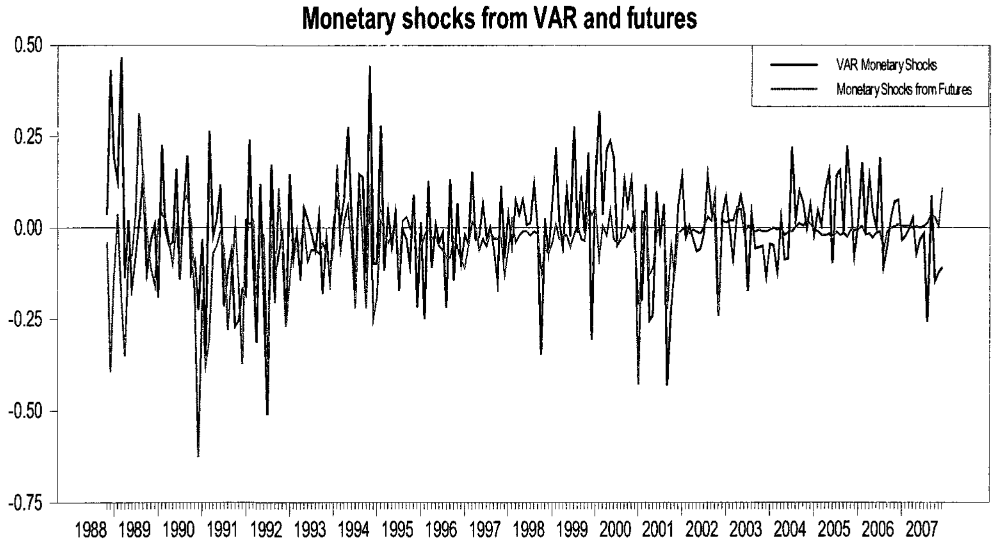

Figure 3.5: Monetary policy shocks obtained from the VAR model and from futures data for the period 1988:11 to 2007:12

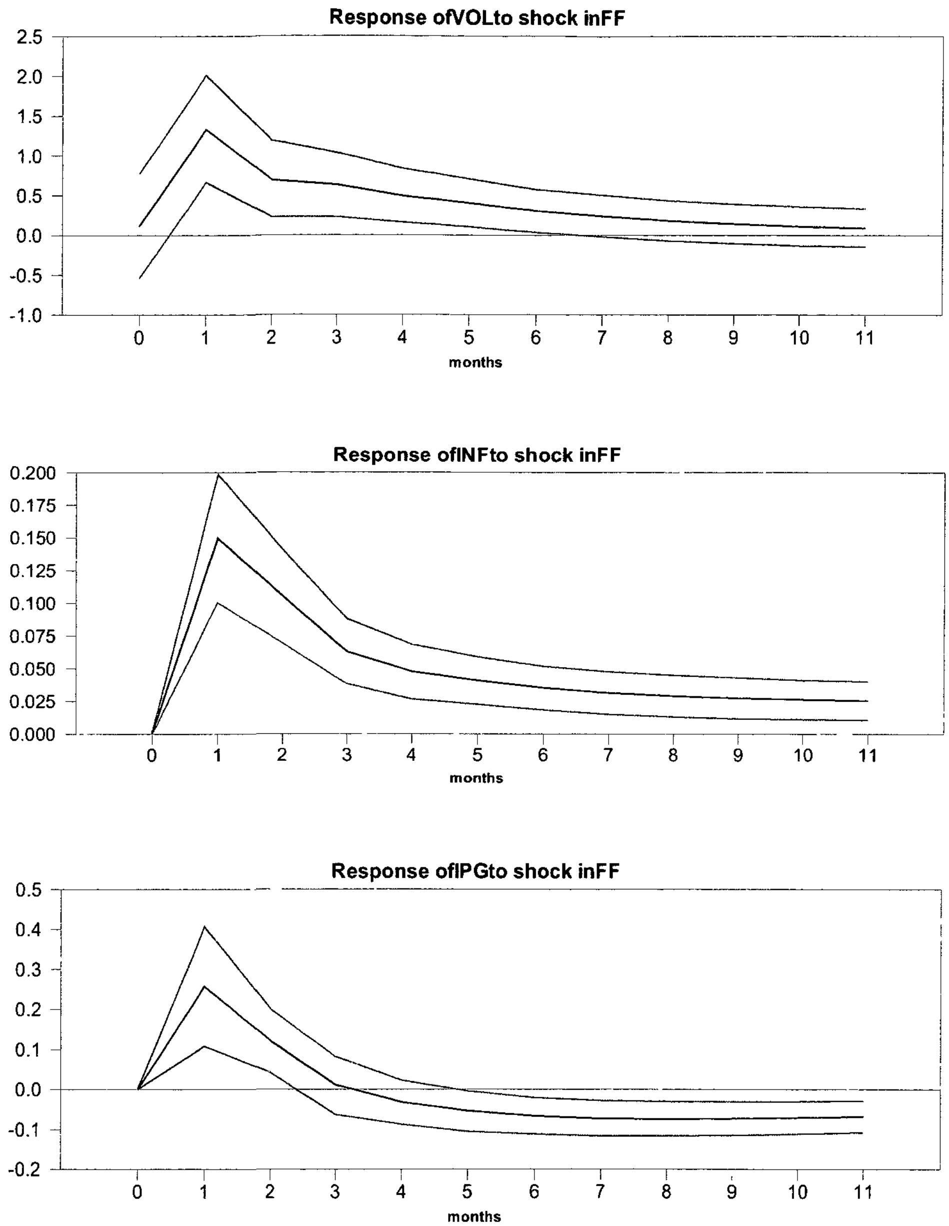

Figure 3.6: Response of aggregate stock market volatility, inflation and industrial production growth to a one standard deviation monetary policy shock from VAR model (M1) with 68% confidence bands constructed by Monte Carlo simulation.

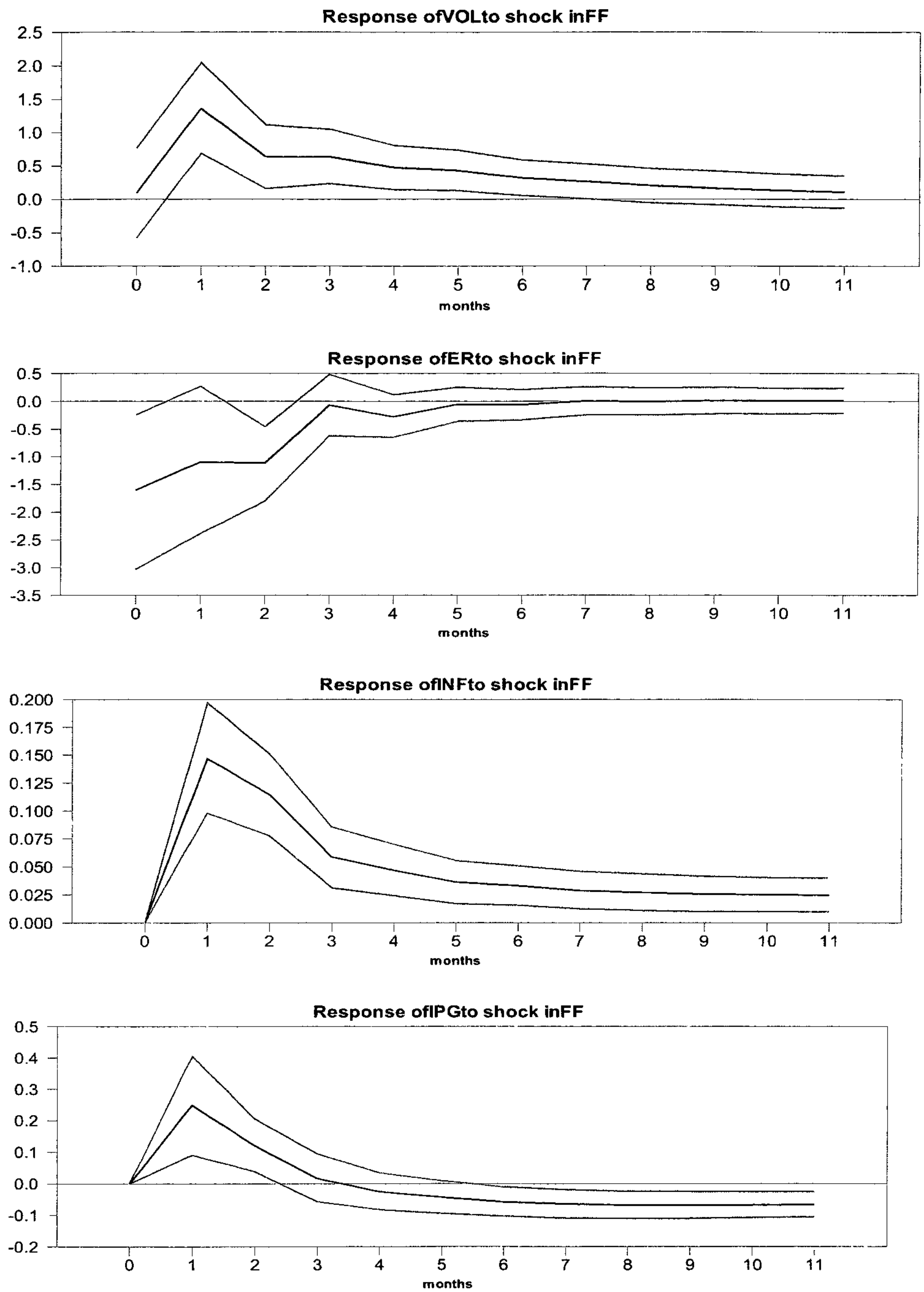

Figure 3.7: Response of aggregate stock market volatility, excess returns on the S&P500, inflation and industrial production growth to a one standard deviation monetary policy shock from VAR model (M2) with 68% confidence bands constructed by Monte Carlo simulation.

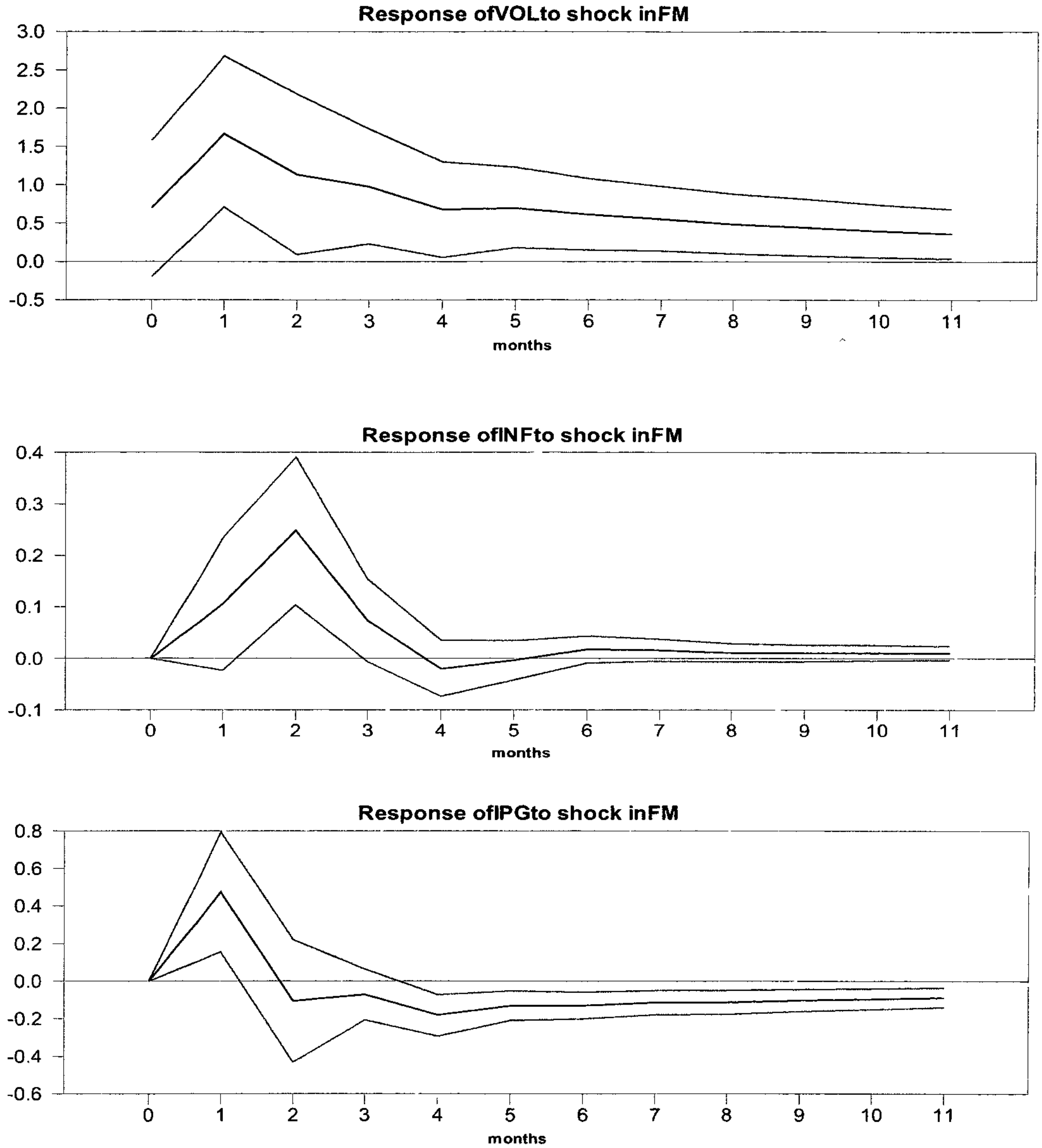

Figure 3.8: Response of aggregate stock market volatility, inflation and industrial production growth to a one standard deviation FM monetary policy shock from VAR model (M3) with 68% confidence bands constructed by Monte Carlo simulation.

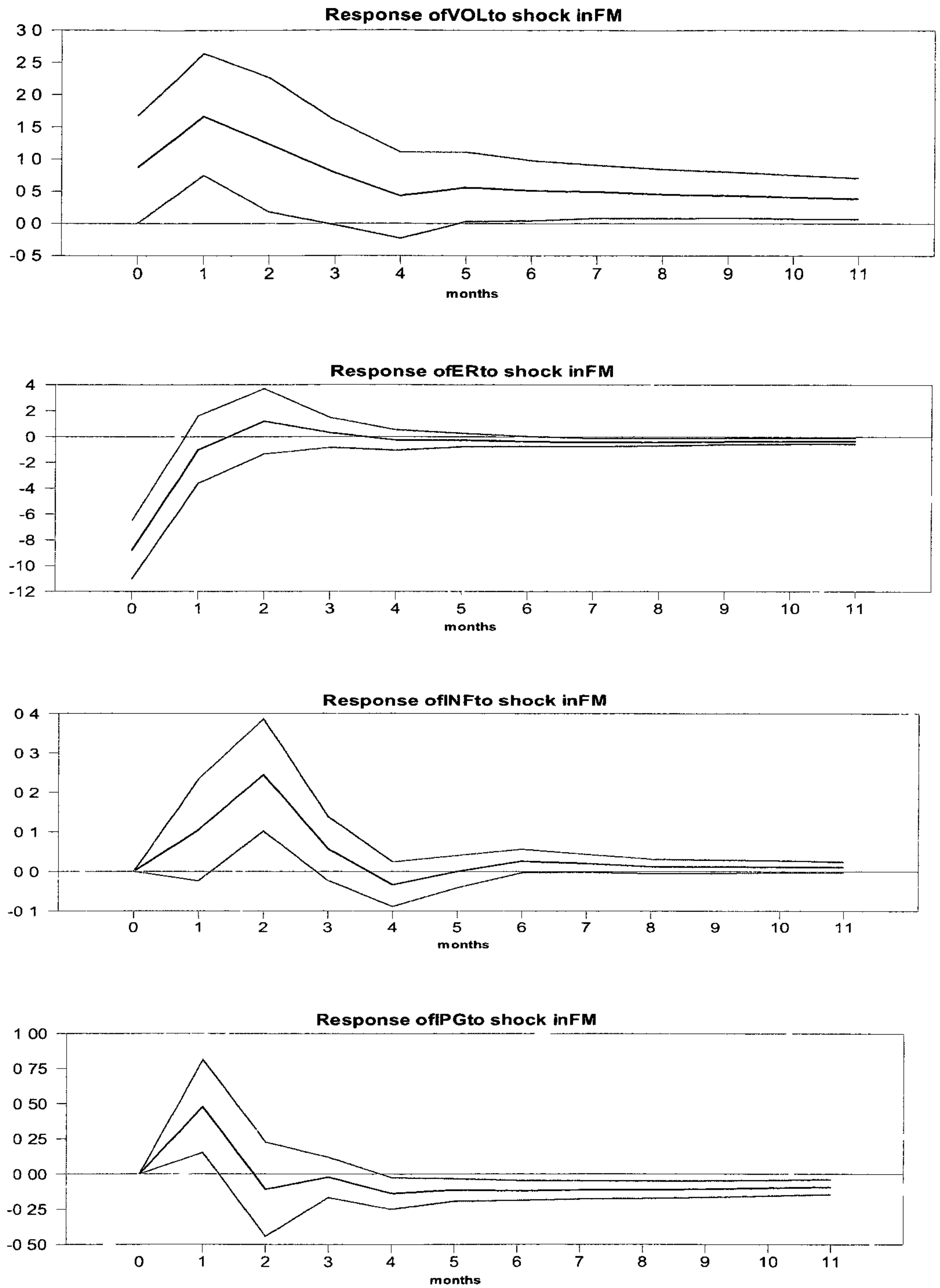

Figure 3.9: Response of stock market volatility, excess returns on the S&P500, inflation and industrial production growth to a one standard deviation FM monetary policy shock from VAR model (M4) with 68% confidence bands constructed by Monte Carlo simulation.

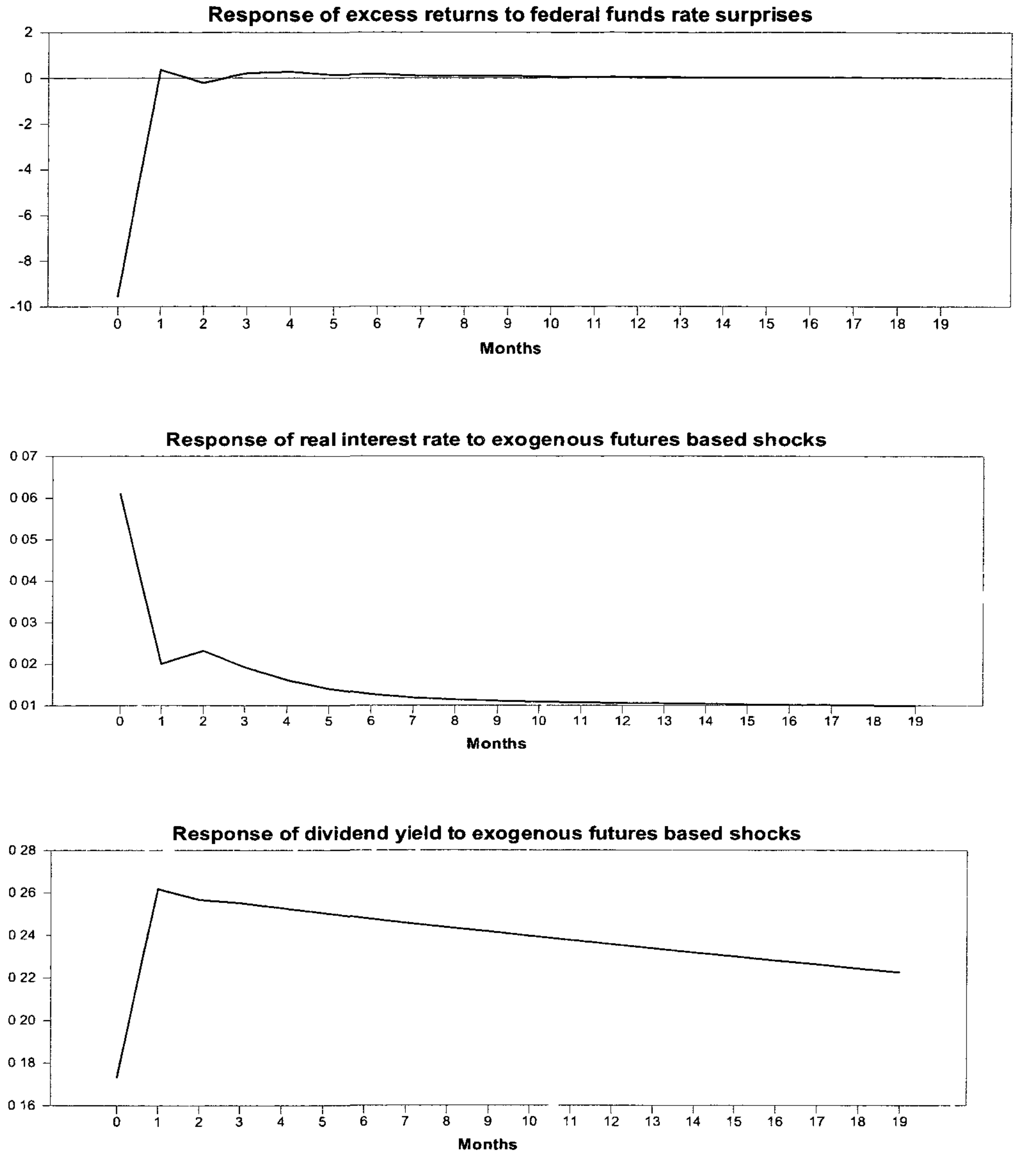

Figure 3.10a: Responses of financial variables to a one percentage point federal funds rate surprise as computed from model (3.11)

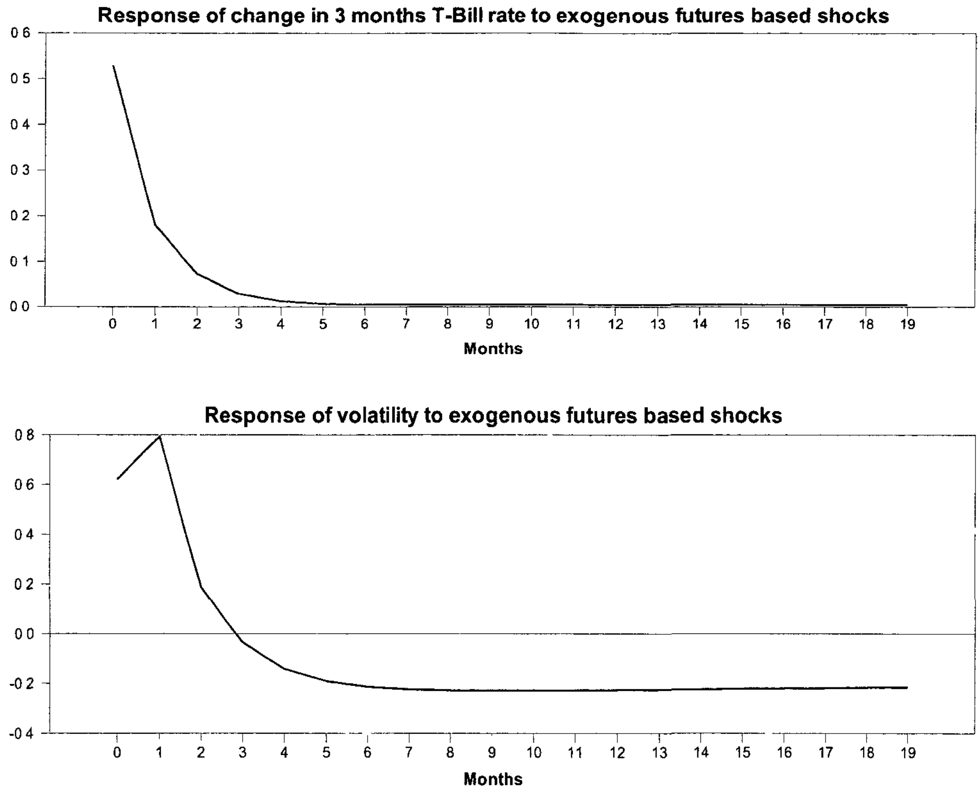

Figure 3.10b: Responses of financial variables to a one percentage point federal funds rate surprise as computed from model (3.11)

Table 3.1: Summary statistics and unit root tests

Variable	Mean	Std. Dev.	ADF	ADF-GLS
Industrial Prod. Growth (IPG)	0.24	0.52	-5.14	-3.01
Inflation (INF)	0.25	0.21	-7.21	-2.90
Commodity Price Inflation (DPCOM)	0.19	2.09	-6.68	-1.21
Federal Funds Rate (FF)	5.42	2.44	-1.83	-3.49
Default Premium (DEFP)	0.96	0.29	-3.09	-0.80
Volatility (VOL)	4.05	2.24	-4.69	-3.45
Excess Returns (ER)	0.35	4.21	-7.77	-9.91

Notes: The last two columns report, respectively, the Augmented Dickey Fuller (ADF) test and Elliot, Stock and Rothenberg (1996) ADF test with Generalized Least Squares (ADF-GLS) detrending. The lag length for the ADF test is chosen using an Akaike Information criterion (AIC) while the Ng and Perron (2001) criterion is used for choosing the lag length for the ADF-GLS test.

Table 3.2: Granger causality tests for volatility using model (M1)

Variable	F-Test	P-value
IPG	0.59	0.55
INF	0.18	0.83
DPCOM	0.42	0.65
FF	1.53	0.21
DEFP	0.43	0.64
VOL	45.61	0.00

Notes: The Granger causality test is an F-test testing the null that the variable (to the left in the table) does not significantly enter the volatility equation in the VAR.

Table 3.3: Granger causality tests for volatility using model (M2)

Variable	F-Test	P-value
IPG	1.15	0.31
INF	0.10	0.90
DPCOM	0.59	0.55
FF	1.45	0.23
DEFP	0.55	0.57
ER	3.23	0.04
VOL	32.74	0.00

Notes: The Granger causality test is an F-test testing the null that the variable (to the left in the table) does not significantly enter the volatility equation in the VAR.

Table 3.4: Results from estimation of
EGARCH (1,1) model in (3.12)

Mean equation		Variance equation	
constant (ω)	0.000570***	constant (ϖ)	-0.239165***
	(0.000138)		(0.027609)
Lag return (ϕ)	-0.012336	EGARCH term (α_1)	0.152191***
	(0.018395)		(0.011443)
Slope (β_s)	-0.000196	lagged variance (γ_1)	0.986650***
	(0.016427)		(0.002265)
Level (β_l)	-0.007701	Slope (λ_s)	1.808607***
	(0.009484)		(0.599778)
Curvature (β_c)	-0.000819	Level (λ_l)	1.059576**
	(0.012012)		(0.486368)
		Curvature (λ_c)	0.691141
			(0.735715)
Log Likelihood	11292.59		

Notes: Standard errors in parentheses (under normality assumption). * denotes significance at the 10%, ** at the 5% and *** at the 1% level.

Chapter 4 Risk premiums and predictability in a Canadian short-term interest rate futures market

Testing the expectations hypothesis is a long researched topic in financial economics. While numerous studies concentrate on testing the expectations hypothesis per se, only recently have researchers turned their attention to establishing the presence and determinants of risk premiums[30]. In fact, instances in which the expectations hypothesis was rejected were attributed by researchers to the presence of time-varying risk premiums without a clear identification of the magnitude or determinants of such risk premiums. More recently, researchers have employed several techniques, such as Kalman filtering, factor models, asset pricing models or predictive regressions to disentangle such risk premiums. This line of research led to important findings across a number of financial assets and commodities.

4.1 Literature Review

Cochrane and Piazzesi (2008) and Ludvigson and Ng (2009a, 2009b) document the presence of risk premiums in bond markets. Cochrane and Piazzesi (2008) show that bond returns are forecastable with a single factor, while Ludvigson and Ng (2009a,

[30] Risk premiums were studied using several approaches. In the context of asset pricing models, such as the Capital Asset Pricing Model, the Fama and MacBeth (1973) procedure was used to estimate risk premiums. Other researchers interpret predictability in excess returns on an asset as evidence of risk premiums. Still, other studies model the risk premium as a latent (unobservable) process. This last definition is arguably more coherent with the theoretical fact that risk premiums are unobservable. Theoretically, Cochrane (2001,p.17) explains how an asset's risk premium depends on the correlation between its returns and the stochastic discount factor (pricing kernel).

2009b) use large datasets and factor models to relate excess bond returns to macroeconomic factors. In the context of commodity (futures) markets, Gorton, Hayashi and Rouwenhorst (2008) and de Roon, Nijman, Szymanowska and van der Goorbergh (2009) are recent examples of the growing attention that the study of commodity risk premiums is eliciting. The traditional approach to testing efficiency is various futures markets involves testing for cointegration between spot and futures prices as in Antoniou and Holmes (1996) or Brooks, Rew and Ritson (2001) for stock index futures.

Testing the expectations hypothesis (and the presence of risk premiums) conceivably received the most attention in the context of foreign exchange markets. Since the contribution of Fama (1984), a sizeable literature relating to the forward rate unbiasdness hypothesis developed. This important literature concurs that forward rates do not perform well in terms of predicting future spot rates. In fact, some of the estimation results emanating from this literature and relating exchange rate returns to the forward premium were largely unexpected and have been referred to as the "forward premium anomaly/puzzle". Baillie and Bollerslev (1994, 2000), Liu and Maynard (2005), Maynard and Phillips (2001) and Sakoulis and Zivot (2002) are some papers drawn from this literature. In a recent contribution, Gospodinov (2009) argues that the presence of a risk premium coupled with econometric problems such as high persistence, low signal-to-noise ratio as well as strong endogeneity explain the anomalous estimation results obtained in the literature. Earlier studies also maintain the presence of risk premiums in foreign exchange forwards. Cheung (1993) and Wolff (1987) employ state space models and Kalman filtering (or signal extraction techniques) to estimate a time-varying latent risk premium (the signal) in foreign exchange forwards. Both studies find a time-varying risk premium in forward exchange markets. On the other hand, McCurdy and Morgan (1992) find evidence of risk premiums in foreign exchange futures markets.

The presence of risk premiums in (relatively) longer-term interest rate futures and forwards has also been documented. Hess and Kamara (2003) study risk premiums in U.S. Treasury bill futures and find evidence of a time-varying risk premium. Gospodinov (2002a) uses a grid bootstrap as well as Kalman filtering to find that forwards on U.S. bonds display time-varying risk premiums. A parallel literature investigates the presence of risk premiums and efficiency in various shorter-term interest rate futures and forwards. Numerous factors render such analyses of significant importance. First, financial theory entails that efficiency in futures markets should hold. Second, predictability of excess futures (forward) returns implies profit opportunities for investors. Third, and by virtue of such contracts being written on interest rates, efficiency (and absence of time-varying risk premiums) in interest rate futures markets allows for the futures rates' natural use as a market based measure of monetary policy expectations. This is especially true for futures contracts written on short-term interest rates due to the high correlation between the monetary policy rate set by central banks and various short-term interest rates. In fact, implied rates from several U.S. short-term interest rate futures (for example, federal funds futures and Eurodollar futures) have been widely used by central banks, academics and practitioners to measure interest rate and monetary policy expectations.

Nevertheless, the presence of risk premiums in short-term interest rate futures has implications for their adequacy as a market gauge of interest rate expectations. In the case in which the risk premium is constant, a simple risk adjustment of the rates implied by these futures is sufficient. However, time-varying risk premiums cannot be straightforwardly corrected, and this distorts the predictive ability of futures contracts.

The usefulness of shorter-term interest rate futures from an academic and a policy making perspective generated considerable research into their possibly embedded risk premiums as well as their efficiency. This strand of the literature did not always reach conformable conclusions. In the U.S. context, Gurkaynak, Sack and Swanson (2007) provide a thorough analysis of the predictive performance and risk premiums of different short-term interest rate futures. The authors also argue that, in contrast with other futures contracts they consider, federal funds futures (which are effectively written on the Federal Reserve monetary policy instrument, the federal funds rate) exhibit a small constant risk premium and are very good predictors of the federal funds rate. A similar conclusion was reached by Kruger and Kuttner (1996) who analyze the efficiency of federal funds futures as predictors of the monetary policy rate. The authors note that very few variables help in forecasting excess returns on federal funds futures; a point that is indicative of the absence of time-varying risk premiums, but that there is evidence of a constant risk premium. In a daily analysis, Hamilton (2009) also conforms to the view that changes in federal funds futures rates are largely unpredictable. Furthermore, he suggests that the time-varying risk premium found by some researchers, such as the work of Piazzesi and Swanson (2008) to be discussed shortly, could be due to the presence of outliers. Sack (2004) demonstrates how to extract monetary policy expectations from federal funds and Eurodollar futures under the assumption of a constant risk premium. In contrast to the findings of the previous papers, some authors found evidence of time-varying risk premiums in federal funds futures. Durham (2004) uses regressions within the capital asset pricing model (CAPM) and arbitrage pricing theory (APT) to conclude that a small, possibly time-varying risk premium is present. The strongest evidence of time-varying risk premiums in federal funds futures is reported in Piazzesi and Swanson (2008). These authors use predictive regressions to assess the magnitude of the risk premium and report sizeable estimates of constant and time-varying risk premiums. Piazzesi and Swanson (2008) similarly find that Eurodollar futures display time-varying risk premiums.

The efficiency of Canadian interest rate futures and forwards has also been investigated in a number of papers. Paquette and Streliski (1998) argue that the expectations hypothesis holds for forwards on Canadian three months banker acceptances (BAs). In contrast, Gravelle, Muller and Streliski (1998) use vector error correction models to the find a time-varying risk premium in forwards on three months banker acceptances. Hijazi, Lai and Yang (2001) study Canadian term structure data and find that, perhaps surprisingly, the conditional variances of U.S. macroeconomic variables drive risk premiums while Canadian macroeconomic variables do not. Gravelle and Morley (2005) use state space methods and the Kalman filter to characterize the latent risk premium in forwards on three months Canadian banker acceptances. In a paper related to this work and similar to Gurkaynak, Sack and Swanson (2007), Johnson (2003) surveys a number of Canadian short-term interest forwards and futures as predictors of monetary policy expectations. One of the instruments surveyed is the future contract on three months banker acceptances (BAX) which will be introduced and studied in detail in this chapter. Specifically, the author investigates whether the expectations hypothesis holds for such instruments and briefly discusses evidence of risk premiums in BAX contracts, without specifying the determinants or magnitude of such premiums.

This chapter provides a detailed analysis of the existence and determinants of risk premiums in futures on three month Canadian bankers' acceptances. Specifically, at the time of writing, this is the first paper to use a predictive regression framework to determine whether Canadian macroeconomic or financial factors are useful in forecasting excess returns on BAX contracts. In addition, we evaluate unbiasdness and efficiency in the BAX market. In light of the results obtained, we impose restrictions on the estimated model and undertake a forecasting exercise for spot returns. We find that the model whose coefficients are restricted in accordance with the unbiasdness hypothesis outperforms autoregressive, random walk and error correction models in terms of forecasting performance. The implications of our findings for deriving monetary policy expectations are discussed. In contrast to earlier research using BAX data, such as Johnson (2003), we provide a detailed analysis of the existence, determinants and magnitude of risk premiums in the BAX market. Johnson (2003) briefly discusses the possible presence of risk premiums. We also demonstrate the importance of our results in terms of out-of-sample forecasting performance. We also differ from earlier studies by considering risk premiums in futures on three months bankers' acceptances rather than forwards [Paquette and Stretitski (1998), Gravelle, Muller and Streliski (1998), Gravelle and Morley (2005)]. Although forward rate agreements and futures share many similarities, the fact that the former are not exchange listed (and thus standardized) introduces an important element of credit risk that might affect estimates of the risk premium. In addition, forwards are less liquid than BAX futures. We thus circumvent problems of credit or liquidity risks by studying BAX contracts.

The plan of the chapter is as follows: section 4.2 introduces the data as well as the variables used, section 4.3 discusses the methodology employed and the results while section 4.4 offers some concluding remarks.

4.2 Data

4.2.1 Bankers' acceptances (BA's) and bankers' acceptances futures (BAX)

Bankers' acceptances are one of the most widely used money market instruments in Canada. As described by the Montreal Exchange circulars, they constitute short-term commercial debt obligations issued by one of Canada's major banks on behalf of a customer. The issuing bank ensures that the principal and interest are repaid in full so that default risk in these instruments is negligible. Futures contracts written on three months bankers' acceptances are known as BAX futures. These futures contracts started trading on the Montreal Exchange[31] in April 1988 and have quarterly expirations. The contract size is one million Canadian dollars (the futures contract is written on an underlying one million dollars banker's acceptance). The Montreal Exchange lists, at all times, three years of quarterly contracts. Settlement of BAX futures occurs two business days prior to the third Wednesday of the contract month, based on the prevailing three month banker's

[31] The Montreal Exchange is Canada's derivative exchange. Equity derivatives, options on exchange traded funds, index derivatives and interest rate derivatives are traded on the Montreal Exchange.

acceptance rate. As noted by the Montreal Exchange and the academic literature, trading volume, open interest and liquidity in BAX contracts increased considerably starting 1994 as they have become a popular tool for short-term interest rate hedging. Further details can be obtained from the website of the Montreal Exchange.

As noted in Gagnon and Lypny (1995), BAX futures share the same characteristics as Eurodollar (time deposit) futures[32]. In fact, BAX contracts have been modeled after the Eurodollar contracts (in terms of expiration, maturity and settlement among other things) due to the latter's wide success in financial markets. Previous studies that make use of BAX and BA data centre on the use of these contracts for interest rate hedging and optimal hedge ratio estimation. These studies include Gagnon and Lypny (1995), Watt (1996) and Siam (2000). Tests of efficiency of the BAX market have been considered in Johnson (2003). We obtain daily data on BAX futures prices starting in April 1988 and ending in June 2008 from the website of the Commodity Research Bureau[33]. Daily data on Bankers' acceptance rates starting in 1994 is obtained from CANSIM, while weekly data on Bankers' acceptances was obtained starting 1988. The chapter will follow Piazzesi and Swanson (2006) in only making use of quarterly data to match the expiration cycle of BAX contracts. We report our estimation results for two samples. The full sample spans the second quarter of 1988 till the second quarter on 2008. We also use the

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[32] The Montreal Exchange also introduced the 30 day overnight repo rate futures (ONX) as the Canadian counterpart of the U.S. 30 day interest rate futures (federal funds futures). However, open interest and trading volume in these contracts have remained very low. The exchange also lists futures on Government of Canada bonds which were more successful.

[33] I would like to thank my supervisor for providing me with this data.

1994Q1 to 2008Q2 subsample (corresponding to the dates in which BAX contracts became more liquid) to sidestep any issues of liquidity and market depth in BAX futures.

4.2.2 Macroeconomic and interest rate data

This paper employs quarterly macroeconomic data (end of quarter observations from monthly data) on real output and the consumer price index (CPI) obtained from CANSIM. We also use several interest rate variables in our examination of risk premiums: the commercial paper rate, the three month Treasury bill rate (T-bill) and the zero coupon yields on one year and ten year bonds. While the commercial paper rates and the three months T-bill data are obtained from CANSIM, the zero coupon yields are due to Bolder, Johnson and Metzler (2004) of the Bank of Canada[34]. We define output growth and inflation as the period to period growth rates in real GDP and the CPI, respectively. The spread between the yields on zero coupon bonds (one and ten year) and the three months T-bill rate as well as the spread between the commercial paper rate and the three months T-bill rate are used at the estimation stage. Other U.S. macroeconomic variables are obtained from the Federal Reserve of St. Louis Economic Database (FRED) and entertained as possible predictors of excess and holding period returns. The U.S. variables used are: real output, GDP deflator, nonfarm payroll employment, the yield on ten year bonds, the yield on one year bonds as well as the yield on BAA rated corporate bonds.

[34] Zero coupon yield data can be downloaded from: http://www.bank-banque-canada.ca/en/rates/yield_curve.html

4.2.3 Excess returns on BAX futures

Denote by f_t^1 and f_t^2 the rates on the first (nearest) and second (next-to-nearest) BAX futures contracts in quarter t (end of quarter). We use Piazzesi and Swanson's (2006) two definitions of excess returns. Given there is no cost for taking a futures position, an investor long in a BAX contract achieves a profit of rx_{t+1} where:

$$rx_{t+1} = f_t^1 - r_{t+1}, \tag{4.1}$$

r_{t+1} denoting the banker's acceptance rate on the day of expiration of the BAX contract. Given that daily data on bankers' acceptances are available only starting 1994Q4, excess returns defined in (4.1) are computed starting 1994Q4[35].

We also compute the returns realized by holding the BAX contract for one period (holding- period returns or hpr) for the full sample (1988Q2 to 2008Q2) as:

$$hpr_{t+1} = f_t^2 - f_{t+1}^1 \tag{4.2}$$

Given that BAX futures are marked to market daily, an investor holding the futures contract for one period can realize the profits (or losses) associated with his trading position. The nearest futures rate f_t^1 can be thought of as the spot (cash) rate in (4.2).

[35] Excess returns make use of BA rate data on the day of expiration. This requires the use of daily data for dating precision.

The time series plots of the returns defined in (4.1) and (4.2) are found in figure 4.1. Casual inspection of the graph shows a decrease in the volatility of the holding-period returns starting around 1994, the date in which trading volume and open interest in BAX markets increased significantly.

We investigate whether the returns defined in (4.1) and (4.2) exhibit any predictability which, in case is found, can be indicative of the presence of risk premiums.

4.3 Methodology and results

4.3.1 Constant risk premiums

Using returns on BAX futures defined in (4.1) and (4.2), we estimate simple regression equations including only a constant to check for evidence of constant risk premiums. Specifically, we estimate[36]:

$$rx_{t+1} = \alpha + \varepsilon_{t+1} \tag{4.3}$$

and

$$hpr_{t+1} = \alpha + \varepsilon_{t+1} \tag{4.4}$$

[36] Throughout this chapter, α, β and ε_{t+1} are used as generic symbols to denote, respectively, a regression intercept, regression slopes and regression error terms and do not imply identical values or equality.

for the two returns defined in (4.1) and (4.2). Table 4.1 reports the estimation results from (4.3) and (4.4) with Newey and West (1987) heteroskedasticity and autocorrelation consistent (HAC) standard errors. The results indicate the presence of a constant risk premium because the constant is mostly significant across the sample and subsamples. The estimates of the risk premium range from 15 to 22 basis points depending on the definition of returns used as well as the sample in question. We note that our estimates of the constant risk premium are smaller than those reported by Piazzesi and Swanson (2006) for Eurodollar futures, which is 47.2 basis points for the nearest Eurodollar futures contract.

4.3.2 Time-varying risk premiums: Business-cycle effects

Given that we have established the presence of a constant risk premium in BAX futures, we proceed to investigate whether the behaviour of excess and holding period returns changes over the business cycle. To the extent that excess (holding period) returns display different behaviours across economic contractions and expansions, this might be indicative that the risk premium found previously is time-varying.

In order to assess the behaviour of returns over the business cycle, we define a recession dummy variable, D_t^R, taking the value one when the quarterly growth rate in real output is negative for two consecutive periods. To provide enough usable observations, (when the dummy is different from zero) we restrict analysis to holding-period returns and estimate:

$$hpr_{t+1} = \alpha + \beta^R D_t^R + \varepsilon_{t+1} \tag{4.5}$$

Similarly, we define a dummy variable, D_t^C, taking the value one when quarterly output growth contracts and estimate:

$$hpr_{t+1} = \alpha + \beta^C D_t^C + \varepsilon_{t+1} \tag{4.6}$$

Estimation results from equations (4.5) and (4.6) are reported in Tables 4.2 and 4.3. Time series plots of the holding period returns and the fitted values from regressions (4.5) and (4.6) are presented in Figures 4.2 and 4.3.

The estimation results indicate that returns on BAX futures increase during (what we define as) recessions and economic contractions. In fact, holding-period returns on BAX futures increase by 0.14% during recessions, while they are 0.42% higher during contractions. We note that while the coefficient of the recession dummy is not significant, the corresponding coefficient of the contraction dummy is significant at the 5% level. Overall, our results are indicative of countercyclical return behaviour, in that investors require a higher return during economic downturns. A comparison of figures 4.2 and 4.3 also shows that the fitted values from the regression with the contraction dummy tracks returns better. In light of the evidence that returns display a different behaviour during economic downturns, we turn next to studying the predictive power of Canadian macroeconomic variables and interest rate spreads.

4.3.3 The predictive power of macroeconomic variables

In this section, we analyze the predictive ability of Canadian macroeconomic variables in forecasting excess and holding period returns. As discussed previously, researchers have related predictability in excess returns on futures contracts to time-varying risk premiums. We use a predictive regression setting to uncover signs of predictability in excess and holding period BAX futures returns. The first set of regressions uses the following macroeconomic variables: quarterly real output growth and consumer price inflation[37]. The estimated regressions take the form:

$$rx_{t+1} = \alpha + \beta_1 \Delta gdp_t + \beta_2 \Delta cpi_t + \varepsilon_{t+1} \tag{4.7}$$

and

$$hpr_{t+1} = \alpha + \beta_1 \Delta gdp_t + \beta_2 \Delta cpi_t + \varepsilon_{t+1} \tag{4.8}$$

Where gdp_t denotes the natural logarithm of real gross domestic product and cpi_t denotes the natural logarithm of the consumer price index. The results obtained from estimating (4.7) and (4.8) are reported in Table 4.4, while the fitted values from (4.8) are shown in Figure 4.4. In line with our earlier findings and with those of Piazzesi and Swanson (2008), excess and holding-period returns are countercyclical because the

[37] We also undertake the estimation with annual growth rates for real output and the CPI. The results are similar and are not reported.

estimated regressions uncover a negative relationship between returns and output growth. A percentage point drop in output growth leads to a 15 to 30 basis points increase in returns, depending on the definition used for returns as well as the sample in question. While the effect of output growth on excess and holding-period returns is strongly significant over the 1994Q4 to 2008Q2 subsample, the magnitude and significance drops 15 basis points in returns due to a percentage point decrease in output growth with holding-period returns over the full sample [1988Q2 to 2008Q2]. Inflation is never found to be a significant predictor of returns. In terms of goodness of fit, the regressions have R^2 ranging from 2% when the holding period returns are used over the full sample to a larger 13% when holding-period returns are used over the 1994Q4 to 2008Q2 sample. Overall, we view the results as merely suggestive of a role for output growth in forecasting returns. However, we note the following: the fact that goodness of fit as well as the significance of the coefficient associated with output growth both decrease over the full sample casts doubt on the predictive capacity of output growth.

Piazzesi and Swanson (2008) find that the growth in nonfarm payroll employment is a good predictor of returns. We investigate whether this holds for BAX futures returns and estimate the following regressions:

$$rx_{t+1} = \alpha + \beta_1 \Delta emp_t + \varepsilon_{t+1} \tag{4.9}$$

and

$$hpr_{t+1} = \alpha + \beta_1 \Delta emp_t + \varepsilon_{t+1} \tag{4.10}$$

In which emp_t denotes the natural logarithm of employment. Results from estimating (4.9) and (4.10) are presented in Table 4.5. Due to the availability of employment numbers, the full sample is confined to start in 1991Q2.

The results of regressions (4.9) and (4.10) are similar to our previous findings and those of Piazzesi and Swanson (2008) for Eurodollar futures. Again, returns are found to be countercyclical and a percentage point drop in employment growth leads to a 16 (32) basis point increase in returns when holding period returns are used over the full sample (1994Q4 to 2008Q2 sample). Employment growth is found not to be a significant predictor of returns over the full sample but is significant across subsamples at the 10% and 5% level. Furthermore, we note that the goodness of fit of the regressions is lower than the regressions with output growth and inflation and stands at a maximum of 7%. The insignificance of employment growth over the full sample combined with the considerably lower t-statistics obtained here in comparison to Piazzesi and Swanson (2008) also cast doubt about the significance of employment growth in predicting returns.

4.3.4 Predictive ability of macroeconomic variables in "real time"

We now subject the macroeconomic variables used in the previous section to a more stringent predictability test. According to the website of Statistics Canada, output and employment numbers are released with a delay of around two months. In contrast, the consumer price index is released with a shorter delay of two weeks. Since output and employment numbers become known to futures markets participants only with a time lag, we re-estimate the regressions including lags of the macroeconomic variables. We refer to the models with lags of macroeconomic variables as real-time regressions, since they presumably include information known to futures markets participants only at the time they decide to take a certain position in the futures contract[38]. We employ these regressions to check for the robustness of our previous results regarding the predictability of returns. The estimated regressions take the form:

$$rx_{t+1} = \alpha + \beta_1 \Delta gdp_{t-1} + \beta_2 \Delta cpi_{t-1} + \varepsilon_{t+1} \tag{4.11}$$

$$hpr_{t+1} = \alpha + \beta_1 \Delta gdp_{t-1} + \beta_2 \Delta cpi_{t-1} + \varepsilon_{t+1} \tag{4.12}$$

and

$$rx_{t+1} = \alpha + \beta_1 \Delta emp_{t-1} + \varepsilon_{t+1} \tag{4.13}$$

$$hpr_{t+1} = \alpha + \beta_1 \Delta emp_{t-1} + \varepsilon_{t+1} \tag{4.14}$$

The results from (4.11) and (4.12) are reported in Table 4.6, while the results from (4.13) and (4.14) are shown in Table 4.7. The estimation results indicate the following: With the exception of a single significant coefficient associated with inflation (with holding period returns over the period 1994Q4, 2008Q2) the macroeconomic variables

[38] The use of lags is also due to the lack of a real-time dataset in Canada

are not useful predictors of excess or holding-period returns. We also note that the significance of the coefficient associated with inflation dissipates completely and changes in sign when the full sample is used. The countercyclical behaviour of returns display is maintained when lagged output growth is used, but not when lagged employment growth is used (lagged employment growth is far from significant). In addition, the estimated models have lower goodness of fit measures than models not including the lags of macroeconomic variables. In all, we view these results as further weakening the predictability and time-varying risk premiums evidence found in the previous section.

4.3.5 The predictive power of interest rate spreads

We turn next to consider the predictive power of Canadian interest rate spreads. Interest rate data are readily available to futures markets participants on a daily basis (interest rates can be assumed to be in the investors' information sets contemporaneously). We use the following interest rate variables: the spread between the rate on 10 year zero coupon bonds and the three months Treasury bill rate (denoted sp10), the spread between the one year zero coupon yield and the three month Treasury bill rate (denoted sp1) and the spread between the commercial paper rate and the three month Treasury bill rate (denoted cpsp). We check the predictive ability of interest rate spreads by estimating the following regressions:

$$rx_{t+1} = \alpha + \beta_1 sp10_t + \beta_2 sp1_t + \beta_3 cpsp_t + \varepsilon_{t+1} \tag{4.15}$$

and

$$hpr_{t+1} = \alpha + \beta_1 sp10_t + \beta_2 sp1_t + \beta_3 cpsp_t + \varepsilon_{t+1} \tag{4.16}$$

The results from estimating (4.15) and (4.16) are reported in Table 4.8 and the fitted values of obtained from (4.16) are plotted in Figure 4.5. As evidenced by Table 4.8, none of the interest rate spreads are of assistance in predicting holding period or excess returns. In addition, while returns exhibit a countercyclical movement across the business cycle, no obvious relationship exists between returns and the interest rate spreads used in our analysis.

4.3.6 U.S. macroeconomic data

We entertain the possibility that returns on BAX futures display time-varying risk premiums which depend on U.S. macroeconomic variables. This possibility is investigated since Hijazi, Lai and Yang (2001) find that the conditional variances of U.S. variables drive risk premiums in Canadian forwards, while Gravelle and Moessner (2001) find that U.S. macroeconomic announcements have a large effect on Canadian interest rate markets. The authors of the latter study attribute the importance of U.S. variables to the fact that Canada is small open economy with sizeable trade links with the U.S. Moreover, the Montreal Exchange reports that U.S. investors account for 42% of open interest in BAX contracts, and that a popular interest arbitrage strategy consists of attempting to benefit from the spread between BAX rates and Eurodollar rates (what is known as the BED spread). We test the predictive ability of the growth rate in real U.S. gross domestic product and inflation (as measured by the change in the natural log of the GDP deflator) in our predictive regression setting:

$$rx_{t+1} = \alpha + \beta_1 \Delta gdp_t^{US} + \beta_2 \Delta def_t^{US} + \varepsilon_{t+1} \quad (4.17)$$

and

$$hpr_{t+1} = \alpha + \beta_1 \Delta gdp_t^{US} + \beta_2 \Delta def_t^{US} + \varepsilon_{t+1} \quad (4.18)$$

The results from (4.17) and (4.18) are found in Table 4.9. The results indicate that U.S. real output growth does not predict excess or holding period returns. Inflation is found not to be a useful predictor of excess returns but a significant predictor of holding-period returns. We note that when the lag of output growth or inflation is used, any predictability found in our regressions with U.S. macroeconomic variables lessens or disappears. The predictive ability of other U.S. macroeconomic variables is tested and we do not find signs of predictability using these variables (These results are not reported in the sake of brevity). Specifically, the growth rate in U.S. nonfarm payroll employment (used by Piazzesi and Swanson (2008)) is not found to be significant while interest rate spreads (the spread between the rates on the U.S. ten year bond and the Canadian T-Bill and between Moody's BAA rated corporate and Canadian T-bills) are also found not to be significant predictors of returns. We view these results as indicating that U.S.

macroeconomic and interest rate variables are weak predictors of returns on BAX contracts.

4.3.7 Forecast efficiency regressions

The evidence obtained thus far in the paper points to the presence of a constant risk premium in BAX futures while no strong evidence in favour of time-varying risk premiums has been detected. Our analysis proceeds by considering forecast efficiency regressions of the type analyzed previously in the literature [as in Gospodinov (2009), Chernenko, Schwarz and Wright (2004) or Inci and Lu (2005)]. Let $y_{t+1}^{BA} = r_{t+1} - r_t = \Delta r_{t+1}$ denote the spot returns from Bankers' acceptances and $x_t^{BA} = f_t^1 - r_t$ denote the futures basis (here, r_t denotes the BA rate sampled at the end of the quarter, and $f_t^{\prime}$ denotes again the futures rate). The following regression model will be considered:

$$y_{t+1}^{BA} = \alpha_1 + \beta_1 x_t^{BA} + \varepsilon_{t+1} \tag{4.19}$$

Also define the returns from the nearest futures contract as $y_{t+1}^{BAX} = f_{t+1}^1 - f_t^1$ and the difference between the rates on the nearest and next-to-nearest contracts as $x_t^{BAX} = f_t^2 - f_t^1$. We consider the regression:

$$y_{t+1}^{BAX} = \alpha_2 + \beta_2 x_t^{BAX} + \varepsilon_{t+1} \tag{4.20}$$

The unbiasdness hypothesis and economic theory stipulate that models (4.19) and (4.20) should satisfy the restrictions $H_0 : \alpha_1 = 0, \beta_1 = 1$ and $H_0 : \alpha_2 = 0, \beta_2 = 1$. Usual tests of efficiency (cointegration) in futures and forward markets consider regressions in levels. Gospodinov (2009) notes that (4.19) and (4.20) constitute restricted error correction formulations of such tests. Researchers view rejecting $\beta_i = 1$ as evidence of a time-varying risk premium and the forward exchange literature has usually obtained puzzling estimates of the parameters. These anomalous results have been largely attributed to the presence of time-varying risk premiums in foreign exchange markets. We report the results from (4.19) and (4.20) in Tables 4.10 and 4.11.

Our results indicate that the unbiasdness hypothesis is rejected at the 10% level in all but one instance. The parameter estimates show that the intercept is relatively large, negative and significantly different from zero over the 1994Q4 to 2008Q2 period (but not significantly different from zero over the full sample). In turn, the null that the slope parameter is equal to one is not rejected in any of our regressions. Combining both results, we view the rejection of the unbiasdness hypothesis as well as the magnitude and significance of the intercept parameter (over the 1994Q4 to 2008Q2 period) as a manifestation of the presence of a constant risk premium that is confounded in the intercept term. The fact that the slope coefficient is not significantly different from one indicates the absence of a time-varying risk premium. Both results obtained from the forecast efficiency regressions are in line with our earlier results. In fact, Table 4.1 establishes the presence of a highly statistically significant constant risk premium over the 1994Q4 to 2008Q2, while the evidence of a constant risk premium over the full sample is weaker. The inability to reject the null of unity slope parameters is also in accordance with the predictive regressions that detect weak evidence of predictability in excess and holding-period returns.

4.3.8 Out-of-sample predictability of spot (BA) returns

Forecasting returns of different assets is of foremost importance to investors. In this section, we turn next to assessing the out-of-sample forecasting performance of different models in predicting spot (BA) returns. To this end, we employ a number of models to obtain one-step ahead forecasts of spot returns. For all the models considered the in-sample period (estimation period) is 1988Q2 to 1998Q2, while the out-of-sample forecasting period is 1998Q3 to 2008Q2. Our out-of-sample forecasting exercise starts with the forecast efficiency regression in (4.19):

$$\Delta r_{t+1} = \alpha_1 + \beta_1 (f_t^1 - r_t) + \varepsilon_{t+1} \tag{4.21}$$

We estimate the forecast efficiency regression (4.21) over the 1988Q2 to 1998Q2 sample and produce a series of one-step-ahead forecasts for the period 1998Q3 to 2008Q2. We obtain the forecasts from (4.21) in two ways: (i) recursive one-step-ahead forecasts, referred to as "recursive regression forecasts", obtained by estimating the model over the 1988Q2 to 1998Q2 and producing a forecast for 1998Q3 and then re-estimating the model over the 1988Q2 to 1998Q3 and obtaining a forecast for 1998Q4 and so on; (ii) rolling forecast referred to as "rolling regression forecasts"; where a moving window that adds one observation and drops one observation is used.

Specifically, rolling one-step-ahead forecasts are obtained by estimating (4.21) over the 1988Q2 to 1998Q2 sample and forecasting one-step-ahead, and then re-estimating over the 1988Q3 to 1998Q3 period and producing a forecast for 1998Q4 and so forth. Forecasts from (4.21) are also produced by imposing a unit slope and estimating the intercept parameter over the recursive and rolling samples. The former set of forecasts is referred to as "Recursive intercept, unit slope" and the latter set is referred to as "Rolling intercept, unit slope" forecasts. The final set of forecasts is obtained by restricting (4.21) according to economic theory: conforming to the unbiasdness hypothesis, we impose the null of a zero intercept and unit slope parameter in (4.21) and forecast over the 1988Q3 to 2008Q2 period from the model. The forecasts from this model are referred to as "restricted forecasts"

The forecasting ability of these three models is benchmarked against an autoregressive model for the spot (BA) rate. Estimation and diagnostic checks such as Box-Jenkins identification, lag length selection by the Akaike and Bayesian information criteria (AIC and BIC) and checking the residuals for any remaining autocorrelation indicate that the spot rate can be modeled accurately using an autoregressive model of order one [AR(1)]. We therefore use an AR(1) for the spot rate:

$$r_{t+1} = \alpha + \phi r_t + \varepsilon_{t+1} \tag{4.22}$$

Which can be re-written in terms of the spot returns, y_{t+1}^{BA}, as:

$$\Delta r_{t+1} = \alpha + \rho r_t + \varepsilon_{t+1} \tag{4.23}$$

where $\rho = (\phi - 1)$. Recursive and rolling one-step-ahead forecasts (as described above) of the spot returns are produced for the 1999Q3 to 2008Q2 using (4.23). We refer the former as "Recursive AR(1) forecast" and to the latter as "Rolling AR(1) forecast". Estimation of (4.22) indicates that the autoregressive parameter is very close to unity (the coefficient is 0.98). In fact, Augmented Dickey Fuller tests (ADF) cannot reject the null of a unit root in the spot rate series. In light of this evidence, we impose the unit root null and produce random walk forecasts of the spot returns over the 1998:03 to 2008:02 sample. These forecasts as referred to as "random walk".

The last forecasting model employed exploits cointegration between the spot and nearest futures rates. Gospodinov (2009) explains that (4.19) is a restricted error correction representation where the short and long-run behaviour of the variables is constrained. Naturally, we consider whether an unconstrained Error Correction Model (ECM) will be able to forecast spot returns more accurately than (4.21) out-of-sample. Testing for unit roots in the spot and futures rates indicates that the null of a unit root cannot be rejected at conventional levels in both series and both series can be characterized as integrated of order one [I(1)]. Futures efficiency and economic theory requires that the two series be cointegrated. We test for cointegration between the spot and futures rate using the Engle and Granger (1987) two step procedure. Namely, we estimate a regression model relating the spot rate to the futures rate and test the residuals from this regression for a unit root. The null of a unit root in the residuals is rejected at any conventional level, and we conclude that the spot and nearest futures rate are cointegrated. Thus, we proceed with an Error Correction representation of the relationship between the spot and nearest futures rates that can be written:

$$\Delta r_{t+1} = \beta_0 + \delta.\hat{z}_t + \sum_{i=0}^{p} \beta_i \Delta r_{t-i} + \sum_{j=0}^{q} \gamma_j \Delta f_{t-i}^{1} + \varepsilon_{t+1} \tag{4.24}$$

where $\hat{z}_t$ denotes the residuals from the cointegrating regression of spot returns on the nearest futures. The lag length in (4.24) is set to 3 as chosen by the AIC. We produce out-of-sample "recursive ECM forecasts" from (4.24) by recursively estimating the cointegrating relationship and model (4.24) and forecasting one-step-ahead, and obtain "rolling ECM forecasts" by estimating the cointegrating relationship and (4.24) with a rolling window and forecasting one-step-ahead.

Figure 4.6 is a time series graph displaying spot (BA) returns and the recursive, rolling and restricted forecasts from (4.21). Figure 4.7 displays the random walk, recursive AR(1) and rolling AR(1) forecasts along with spot returns whereas Figure 4.8 displays the "Rolling ECM" and "Recursive ECM" forecasts. Visual inspection of Figures 4.6, 4.7, 4.8 illustrates the good performance of the forecast efficiency regressions in (4.21) (especially the restricted one) when compared to the random walk or AR(1) forecasts. Since this is not a formal procedure for comparing forecasts, we turn next to more a formal forecast evaluation criteria.

4.3.9 Forecast evaluation: Statistical criteria and Mincer-Zarnowitz regressions

Economic and financial forecasts can be evaluated in numerous ways. Statistical criteria have been widely used but are subject to sampling error. One alternative to statistical criteria is to evaluate forecasts on grounds of the profits generated by investors. For instance, Gospodinov (2002b) uses such an approach when evaluating forecasts for interest rates while Brooks, Rew and Ritson (2001) use profit-based evaluation criteria in the context of forecasting stock index returns. In spite of its appeal, this approach will not be pursued in this chapter.

Another alternative for evaluating forecasts is the regression based approach of Mincer and Zarnowitz (1969). Anatolyev and Gospodinov (2009) employ this approach to evaluate stock return forecasts. In this paper, we will use both statistical criteria as well as Mincer-Zarnowitz regressions to evaluate forecasts from the competing models. In terms of statistical criteria, we use two loss functions which are the Mean Absolute Error (MAE) and the Root Mean Square Error (RMSE):

$$MAE = \frac{1}{h}\sum_{t=T+1}^{T+h}\left|y_t - \hat{y}_t\right|$$

$$RMSE = \sqrt{\frac{\sum_{t=T+1}^{T+h}\left(y_t - \hat{y}_t\right)^2}{h}}$$

Where $j = T+1,\ldots,T+h$ is the forecast sample, y_t is the actual spot return and $\hat{y}_t$ is the spot return forecast obtained from one the models. As noted in Gospodinov, Gavala and Jiang (2006), the MAE and RMSE do not treat over and under-predictions differently.

We also employ the regression based framework of Mincer and Zarnowitz (1969) to test forecast unbiasdness. The approach consists of regressing actual spot returns on the forecasts (from all the models) and allows for a simple test of forecast unbiasdness. Specifically, the Mincer-Zarnowitz regression setting uses:

$$y_t = a_0 + a_1 . \hat{y}_t + error$$

for the forecast sample $j = T+1,\ldots,T+h$. Testing forecast unbiasdness entails testing the null $H_0 : a_0 = 0, a_1 = 1$. The coefficient of fit from the Mincer-Zarnowitz regressions is also examined in order to compare forecasts.

Table 4.12 reports the MAE and RMSE of the different forecasts and ranks the competing models. Both statistical evaluation criteria confirm the results conjectured by visual inspection of Figures 4.6, 4.7 and 4.8. Namely, we find that the forecast efficiency regression whose slope and intercept coefficients are restricted according to the unbiasdness hypothesis (referred to as "restricted forecasts") outperforms all the models and that the unrestricted forecast efficiency regressions (rolling and recursive) closely follow. Regardless of the method chosen (recursive, rolling or by restricting the model) or the statistical criterion chosen, forecasts from (4.21) always outperform the random walk. In contrast, the error correction model and the AR(1) produce worst forecasts than the random walk.

Tables 4.13a, 4.13b and 4.13c report the results from the Mincer-Zarnowitz regressions. The results indicate that the null of unbiasdness for the AR(1) cannot be rejected, while unbiasdness is rejected for a number of the regression and ECM forecasts. We attribute the inability of the Mincer-Zarnowitz regressions to reject the null of unbiasdness and efficiency for the AR(1) forecasts to the large standard errors associated with the slope and intercept parameters when the AR(1) forecasts are used. The only instance in which the null of unbiasdness is not rejected (at the 1% level) when model (4.21) is used occurs with the "restricted forecast". Recall that this is the forecast obtained by restricting the slope parameter in (4.21) to zero and the slope to unity. In addition, the Mincer-Zarnowitz regression fit provides a different account. All the forecasts obtained from (4.21) have impressive R^2 ranging from a minimum of 37% to a maximum, again corresponding to the "restricted forecast", of 41%. The ECM forecasts have significantly lower R^2 with a maximum of 8% while the AR(1) forecasts have a near zero R^2.

4.4 Concluding remarks

This chapter provides evidence of the presence of a constant risk premium in BAX futures. Predictive regression techniques uncover countercyclical behaviour in excess and holding-period returns on BAX contracts. The evidence in favour of predictability in returns remains weak, as a number of important macroeconomic variables and interest rate spreads are not robust predictors of returns. In light of this evidence, we argue that there exist no time-varying risk premiums in BAX futures returns. We also argue that forecast efficiency regressions relating spot (Bankers' acceptances) returns to the futures basis are in line with our predictive regressions results. In fact, while unbiasdness is sometimes rejected using the forecast efficiency regressions, we find that efficiency is not rejected (the slope parameter is not significantly different from unity). We therefore attribute the rejection of the unbiasdness hypothesis to the presence of the constant risk premium which we detect. We also undertake out-of-sample forecasting of spot (BA) returns and find that forecasts imposing unbiasdness in the relationship between spot returns and the futures basis outperform autoregressive, random walk and error correction models using two different criteria.

Our results have interesting implications from two perspectives. From a policy making perspective, the evidence suggesting the lack of a time-varying risk premium in the most liquidly traded Canadian short-term interest rate futures implies that, once the constant risk premium found is adjusted for, the Bank of Canada can use BAX futures as market based measures of interest rate expectations. From a trading perspective, our out-of-sample forecasting results are suggestive of a possible profit opportunity for investors. In fact, an interesting exercise to undertake would consist of calculating profits generated by trading based on the different models proposed in the chapter.

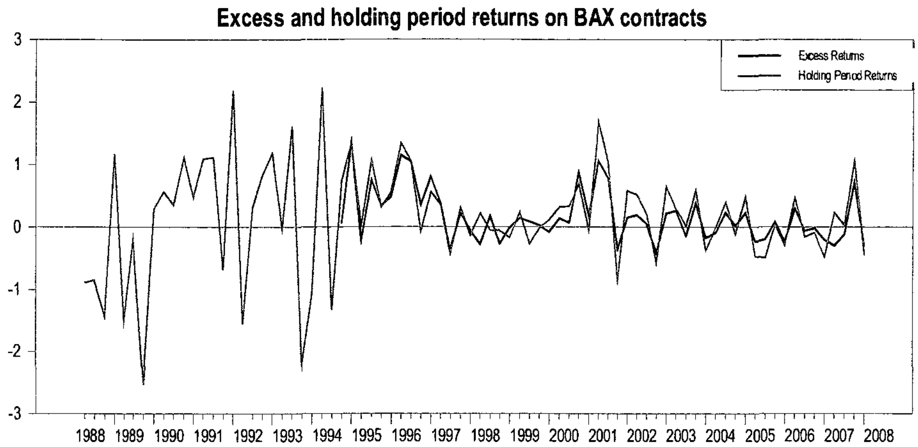

Figure 4.1: Time series of excess and holding-period returns on BAX futures (in percent)

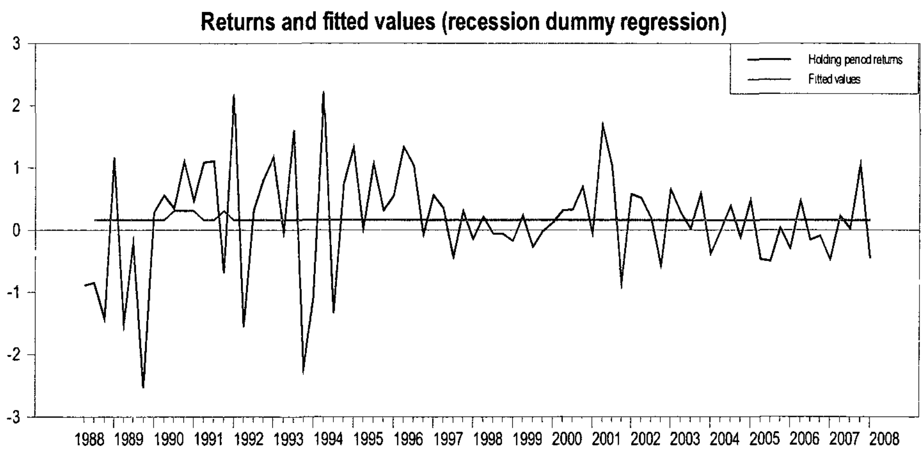

Figure 4.2: Holding-period returns and fitted values from regression with recession dummy

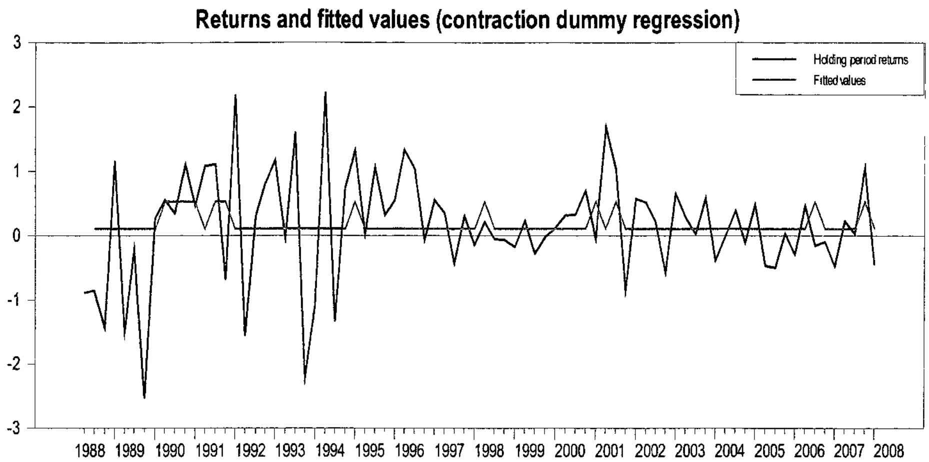

Figure 4.3: Holding-period returns and fitted values from regression with contraction dummy

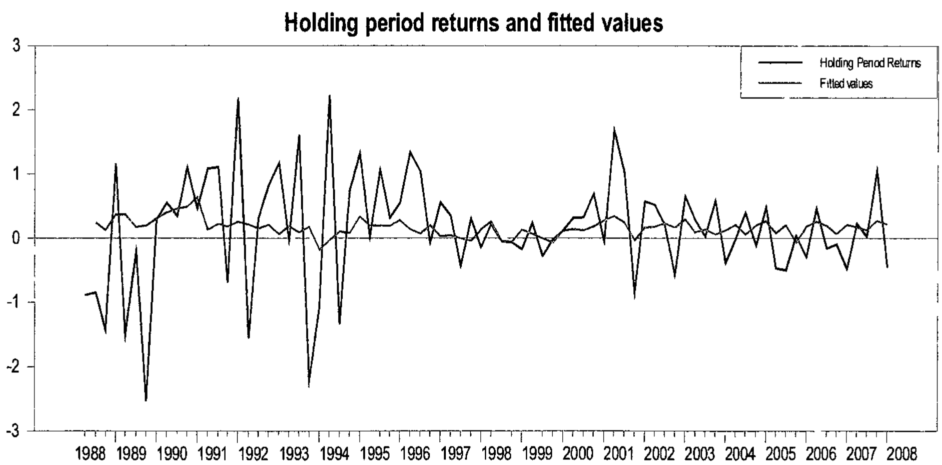

Figure 4.4: Holding-period returns and fitted values from from regression with output growth and inflation

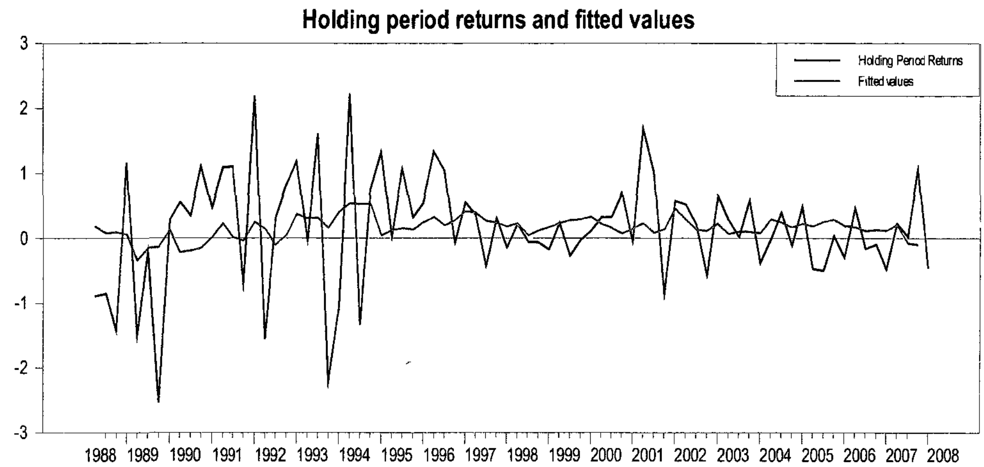

Figure 4.5: Holding-period returns and fitted values from regression with interest rate spreads

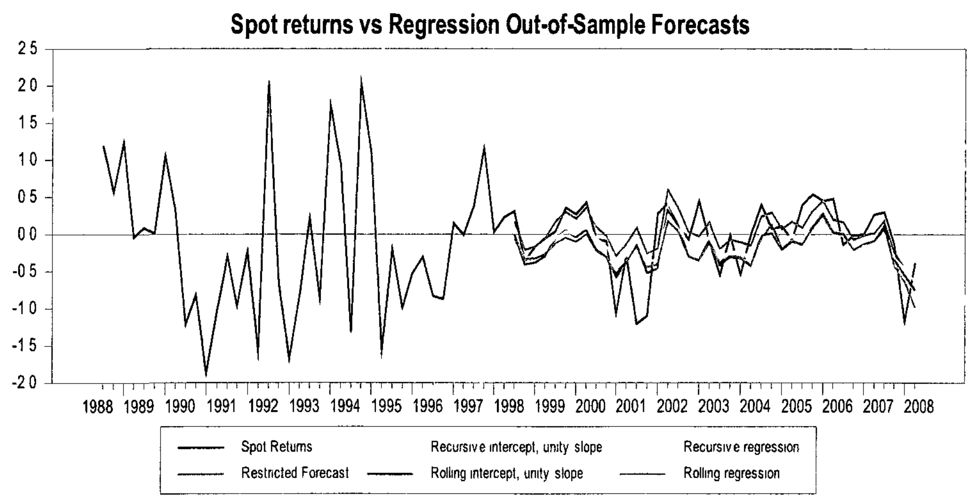

Figure 4.6: Spot (BA) returns and out-of-sample forecasts obtained from forecast efficiency regressions

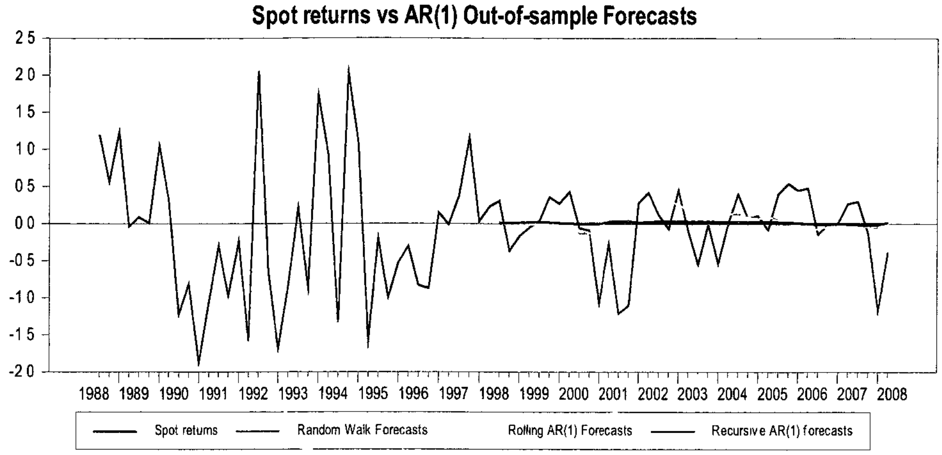

Figure 4.7 Spot (BA) returns and out-of-sample forecasts from AR(1) models

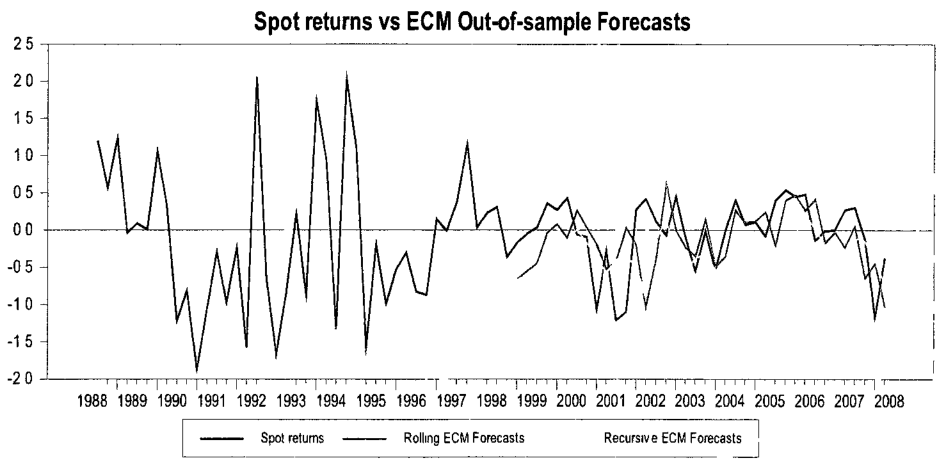

Figure 4 8 Spot (BA) returns and out-of-sample forecasts from error correction models

Table 4.1: Estimation results from regressions (4.3) and (4.4) of excess and holding-period returns on a constant

Dependent Variable	Excess returns	Holding Period Returns	Holding Period Returns
Constant	0.17***	0.22***	0.15*
	(0.06)	(0.07)	(0.08)
Sample period	94Q4-08Q2	94Q4-08Q2	88Q2-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2 lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.2: Estimation results from regression (4.5) of holding-period returns on a constant and recession dummy

Dependent Variable	Holding Period Returns
Constant	0.16*
	(0.09)
Recession Dummy	0.14
	(0.36)
R^2	0.00
Sample period	1988Q2-2008Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2 lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.3: Estimation results from regression (4.6) of holding-period returns on a constant and contraction dummy

Dependent Variable	Holding Period Returns
Constant	0.10
	(0.09)
Contraction Dummy	0.42**
	(0.17)
R^2	0.03
Sample period	1988Q2-2008Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2 lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.4: Estimation results from regressions (4.7) and (4.8) with output growth and inflation

Dependent Variable	Excess Returns	Holding Period Returns	Holding Period Returns
Constant	0.29***	0.35***	0.20*
	(0.08)	(0.09)	(0.10)
Output growth	-0.23**	-0.30***	-0.15*
	(0.09)	(0.09)	(0.08)
Inflation	0.06	0.18	0.11
	(0.08)	(0.14)	(0.10)
R^2	0.08	0.13	0.02
Sample period	94Q4-08Q2	94Q4-08Q2	88Q2-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.5: Estimation results from regressions (4.9) and (4.10) with employment growth

Dependent Variable	Excess Returns	Holding Period Returns	Holding Period Returns
Constant	0.30***	0.38***	0.29***
	(0.11)	(0.11)	(0.10)
Emp. growth	-0.26*	-0.32**	-0.16
	(0.14)	(0.14)	(0.15)
R^2	0.07	0.06	0.01
Sample period	94Q4-08Q2	94Q4-08Q2	91Q2-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level

Table 4.6: Estimation results from regressions (4.11) and (4.12) with lagged macroeconomic variables

Dependent Variable	Excess Returns	Holding Period Returns	Holding Period Returns
Constant	0.21**	0.24*	0.30**
	(0.10)	(0.12)	(0.12)
Lag output growth	-0.11	-0.13	-0.13
	(0.09)	(0.13)	(0.11)
Lag inflation	0.10	0.15**	-0.06
	(0.06)	(0.07)	(0.14)
R^2	0.03	0.04	0.01
Sample period	94Q4-08Q2	94Q4-08Q2	88Q2-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.7: Results from regressions (4.13) and (4.14) with lagged employment growth

Dependent Variable	Excess Returns	Holding Period Returns	Holding Period Returns
Constant	0.09	0.09	0.19
	(0.12)	(0.14)	(0.11)
Lag emp. growth	0.16	0.27	0.07
	(0.19)	(0.23)	(0.19)
R^2	0.02	0.04	0.00
Sample period	94Q4-08Q2	94Q4-08Q2	91Q3-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.8: Estimation results from regressions (4.15) and (4.16) with interest rate spreads

Dependent Variable	Excess Returns	Holding Period Returns	Holding Period Returns
Constant	0.04	0.11	0.21
	(0.10)	(0.14)	(0.15)
1 year -Tbill spread	-0.19	0.01	0.36
	(0.15)	(0.20)	(0.27)
10 year -Tbill spread	0.11	0.06	-0.03
	(0.06)	(0.08)	(0.10)
Comm. Paper- Tbill spread	0.11	0.17	-0.29
	(0.24)	(0.38)	(0.28)
R^2	0.05	0.02	0.03
Sample period	94Q4-08Q2	94Q4-08Q2	88Q2-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.9: Estimation results from regressions (4.17) and (4.18) with U.S. output growth and inflation

Dependent Variable	Excess Returns	Holding Period Returns	Holding Period Returns
Constant	0.41**	0.61***	0.65***
	(0.16)	(0.17)	(0.23)
U.S. Output growth	-0.12	-0.15	-0.24
	(0.11)	(0.12)	(0.17)
U.S. Inflation	-0.28	-0.49**	-0.53**
	(0.20)	(0.24)	(0.26)
R^2	0.02	0.04	0.02
Sample period	94Q4-08Q2	94Q4-08Q2	88Q2-08Q2

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.10: Results from the forecast efficiency regression in (4.19)

Dependent Variable	Spot Returns	Spot Returns
Constant (α_1)	-0.13	-0.18***
	(0.08)	(0.06)
Futures basis (β_1)	0.76	1.26
	(0.20)	(0.20)
F-test (H_0: α_1=0,β_1=1)	2.32	3.79
P-value (F-test)	0.10	0.02
t-test (H_0: β_1=1)	-1.17	1.23
P-value (t-test)	0.24	0.20
Sample period	88Q2-08Q2	94Q4-08Q2

Notes: Standard errors in parentheses. * denotes statistical significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 4.11: Results from forecast efficiency regressions in (4.20)

Dependent Variable	Nearest Futures Returns	Nearest Futures Returns
Constant	-0.11	-0.21**
	(0.09)	(0.08)
Difference in futures rates	0.53	1.08
	(0.30)	(0.30)
F-test (H_0: α_2=0,β_2=1)	2.49	3.34
P-value (F-test)	0.08	0.04
t-test (H_0: β_2=1)	-1.53	0.27
P-value (t-test)	0.12	0.78
Sample period	88Q2-08Q2	94Q4-08Q2

Notes: Standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level.

Table 4.12: Statistical forecast evaluation criteria

Forecast From:	MAE	RMSE
Recursive Regression	0.2922 (3)	0.3802 (4)
Rolling Regression	0.3230 (5)	0.3937 (5)
Recursive intercept, unit slope	0.2742 (2)	0.3610 (2)
Rolling intercept, unit slope	0.3037 (4)	0.3792 (3)
Restricted Regression	0.2490 (1)	0.3610 (1)
Recursive AR(1)	0.3315 (7)	0.4602 (7)
Rolling AR(1)	0.3577 (8)	0.5074 (8)
Random Walk	0.3312 (6)	0.4585 (6)
Recursive ECM	0.4332 (10)	0.5196 (10)
Rolling ECM	0.4048 (9)	0.5094 (9)

Notes: Number in parenthesis indicates the rank of the model/forecast in terms of the criterion considered. The in-sample period is 1988Q2 to 1998Q2 and the forecasting period runs from 1998Q3 to 2008Q2.

Table 4.13a: Results from Mincer-Zarnowitz Regressions

	Recursive Regression	Rolling Regression	Restricted Regression
Intercept (a_0)	0.171***	0.189***	-0.089
	(0.050)	(0.052)	(0.056)
slope (a_1)	1.627	1.185	1.195
	(0.236)	(0.211)	(0.171)
P-value (H_0:a_1=1)	0.007	0.382	0.253
P-value(H_0:a_0=0,a_1=1)	0.000	0.001	0.253
R^2	0.40	0.37	0.41

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level (for intercept). P-values correspond to Wald tests of null of efficiency H_0: a_1=1 and unbiasdness H_0: a_0=0,a_1=1, respectively

Table 4.13b: Results from Mincer-Zarnowtiz regressions

	Recursive Intercept, Unit Slope	Rolling Intercept, Unit Slope
Intercept (a_0)	0.101**	0.185***
	(0.049)	(0.046)
slope (a_1)	1.203	1.246
	(0.168)	(0.181)
P-value (H_0:a_1=1)	0.228	0.172
P-value (H_0:a_0=0, a_1=1)	0.057	0.000
R^2	0.41	0.41

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level (for intercept). P-values correspond to Wald tests of null of efficiency H_0: a_1=1 and unbiasdness H_0: a0=0,a_1=1, respectively.

Table 4.13c: Results from Mincer-Zarnowitz regressions

	Rolling AR(1)	Recursive AR(1)	Rolling ECM	Recursive ECM
Intercept (a_0)	-0.038	-0.040	0.008	0.104
	(0.097)	(0.114)	(0.076)	(0.096)
slope (a_1)	-0.095	-0.445	0.359	0.507
	(0.526)	(6.743)	(0.191)	(0.294)
P-value (H_0:a_1=1)	0.037	0.830	0.000	0.094
P-value (H_0:a_0=0, a_1=1)	0.049	0.805	0.003	0.006
R^2	0.00	0.00	0.08	0.07

Notes: Newey and West (1987) Heteroskedasticity and Autocorrelation (HAC, 2lags) consistent standard errors in parentheses. *denotes significance at the 10% level, **at the 5% and ***at the 1% level (for intercept). P-values correspond to Wald tests of null of efficiency H_0: a_1=1 and unbiasdness H_0: a0=0,a_1=1, respectively.

Chapter 5 Conclusion

This thesis provides an account of the effect of monetary policy surprises on stock market volatility. It also explores and exploits the use futures contracts to measure interest rate (monetary policy) expectations. The second chapter (first essay) studies, at the daily and monthly frequencies, the effect of federal funds rate surprises on implied stock market volatility. The results establish an increase in volatility due to the surprises and we attribute this increase to news arrival to stock market participants.

The third chapter (second essay) employs more elaborate (multivariate) dynamic models to examine the effect of monetary policy shocks on realized stock market volatility and excess returns. The transmission mechanism through which monetary policy shocks affect returns and volatility is studied by incorporating a number of financial variables into the dynamic models. We find that volatility is increased and stock market volatility is decreased contemporaneously due to a monetary policy shock. A daily analysis which embeds three different interest rate surprise measures within conditional heteroskedasticity models (thus treating volatility as a latent variable) similarly indicates a contemporaneous increase in volatility and a decrease in returns. Analogously to the first essay, the second essay also utilizes futures contracts to measure interest rate expectations and incorporates the information provided by futures contracts directly into the estimated models.

The fourth chapter (third essay) examines the properties of Canadian interest rate futures and discusses the importance of these contracts from a policy making and trading perspectives. In contrast to the first two essays that draw upon the literature to use U.S. interest rate futures, the third essay tests the presence, magnitude and determinants of risk premiums in Canadian interest rate futures. The presence of risk premiums has direct implications for the use of such contracts to measure interest rate expectations. We find evidence of a constant risk premium, but no evidence of a time-varying risk premium. We note that such contracts can be used as market expectations of monetary policy expectations, akin to their U.S. counterparts, once the constant risk premium is adjusted for. Furthermore, out-of-sample forecasting results reported in this essay demonstrate an important element of predictability in spot returns.

Different avenues of future research that build upon the findings in this thesis can be proposed. First, investigating volatility co-movements across several markets is a promising avenue for future research. It would be interesting to investigate whether volatility changes across several markets can be traced to specific shocks. Second, it would be interesting to investigate whether BAX futures help with identifying monetary policy shocks in VAR models. Third, an analysis of the effect of monetary policy shocks on Canadian stock market returns and volatility can be undertaken. In this context, it would also be interesting to investigate any systematic differences in the response of Canadian stock and future markets to improved communication by the Bank of Canada following the fixed announcement dates regime implemented in 2000.

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