Essays on Financial Markets and the Macroeconomy

Jürg Fausch

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Abstract

Asset pricing implications of a DSGE model with recursive preferences and nominal rigidities. I study jointly macroeconomic dynamics and asset prices implied by a production economy featuring nominal price rigidities and Epstein-Zin (1989) preferences. Using a reasonable calibration, the macroeconomic DSGE model is consistent with a number of stylized facts observed in financial markets like the equity premium, a negative real term spread, a positive nominal term spread and the predictability of stock returns, without compromising the model's ability to fit key macroeconomic variables. The interest rate smoothing in the monetary policy rule helps generate a low risk-free rate volatility which has been difficult to achieve for standard real business cycle models where monetary policy is neutral. In an application, I show that the model provides a framework for analyzing monetary policy interventions and the associated effects on asset prices and the real economy.

Macroeconomic news and the stock market: Evidence from the eurozone. This paper is an empirical study of excess return behavior in the stock market in the euro area around days when important macroeconomic news about inflation, unemployment or interest rates are scheduled for announcement. I identify state dependence such that equity risk premia on announcement days are significantly higher when the interests rates are in the vicinity of the zero lower bound. Moreover, I provide evidence that for the whole sample period, the average excess returns in the eurozone are only higher on days when FOMC announcements are scheduled for release. However, this result vanishes in a low interest rate regime. Finally, I document that the European stock market does not command a premium for scheduled announcements by the European Central Bank (ECB).

The impact of ECB monetary policy surprises on the German stock market. We examine the impact of ECB monetary policy surprises on German excess stock returns and the possible reasons for such a response. First, we conduct an event study to asses the impact of conventional and unconventional monetary policy on stock returns. Second, within the VAR framework of Campbell and Ammer (1993), we decompose excess stock returns into news regarding expected excess returns, future dividends and future real interest rates. We measure conventional monetary policy shocks using futures markets data. Our main findings are that the overall variation in German excess stock returns mainly reflects revisions in expectations about dividends and that the stock market response to monetary policy shocks is dependent on the prevailing interest rate regime. In periods of negative real interest rates, a surprise monetary tightening leads to a decrease in excess stock returns. The channels behind this response are news about higher expected excess returns and lower future dividends.

Keywords: *Asset pricing, business cycles, DSGE model, macroeconomic risk, monetary policy shocks, recursive preferences, stock market, VAR model, variance decomposition.*

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"If you always do what you always did, you will always get what you always got"

Albert Einstein

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Stockholm, April 2017

Jürg Fausch

Contents

Chapter 1

Introduction

This thesis consists of three self-contained essays dealing with questions at the intersection of asset pricing and macroeconomics.

The first essay **Asset pricing implications of a DSGE model with recursive preferences and nominal rigidities** studies jointly macroeconomic dynamics and asset prices implied by a production economy featuring nominal price rigidities and Epstein and Zin (1989) preferences. Using a reasonable calibration, the macroeconomic DSGE model is consistent with a number of stylized facts observed in financial markets like the equity premium, a negative real term spread, a positive nominal term spread and the predictability of stock returns, without compromising the model's ability to fit key macroeconomic variables. The interest rate smoothing in the monetary policy rule helps generate a low risk-free rate volatility which has been difficult to achieve for standard real business cycle models where monetary policy is neutral. In an application, I show that the model provides a framework for analyzing monetary policy interventions and the associated effects on asset prices and the real economy.

The second essay **Macroeconomic news and the stock market: Evidence from the eurozone** is an empirical study of excess return behavior in the stock market in the euro area around days when important macroeconomic news about inflation, unemployment or interest rates are scheduled for announcement. I identify state dependence such that equity risk premia on announcement days are significantly higher when the interests rates are in the vicinity of the zero lower bound. Moreover, I provide evidence that for the whole sample period, the average excess returns in the eurozone are only higher on days when FOMC announcements are scheduled for release. However, this result vanishes in a low interest rate regime. Finally, I document that the European stock market does not command a premium for scheduled announcements by the European Central Bank (ECB).

The third essay **The impact of ECB monetary policy surprises on the German stock market** is joint work with Markus Sigonius and examines the impact of ECB monetary policy surprises on German excess stock returns and the possible reasons for such a response. First, we conduct an event study to asses the impact of conventional and unconventional monetary policy on stock returns. Second, within the VAR framework of Campbell and Ammer (1993), we decompose excess stock returns into news regarding expected excess returns, future dividends and future real interest rates. We measure conventional monetary policy shocks using futures markets data. Our main findings are that the overall variation in German excess stock returns mainly reflects revisions in expectations about dividends and that the stock market response to monetary policy shocks is dependent on the prevailing interest rate regime. In periods of negative real interest rates, a surprise monetary tightening leads to a decrease in excess stock returns. The channels behind this response are news about higher expected excess returns and lower future dividends.

Chapter 2

Asset pricing implications of a DSGE model with recursive preferences and nominal rigidities[1]

2.1 Introduction

During the recent financial crisis policy makers have gained an increasing interest in a comprehensive understanding of the interaction between the real economy and financial markets. The standard framework used in monetary policy analysis by central banks are New Keynesian DSGE models, which have been successful in quantitatively explaining key business cycle features. However, these models ignore asset prices (e.g. see the discussion in Christiano, Eichenbaum, and Evans, 2005) or have difficulties in reproducing salient features of financial markets (e.g. Rudebusch and Swanson, 2008).

As emphasized by Cochrane (2008a), asset prices and the macroeconomy are closely linked and a failure of macroeconomic models to explain asset pricing facts implies serious flaws in the model. Moreover, risk factors in standard

[1]I am especially grateful for advice and encouragement from my advisors John Hassler and Roine Vestman. I also thank Tobias Broer, Claudio Campanale, Per Krusell and Irina Zviadadze for helpful comments.

finance models (e.g. Fama and French, 1993), which ignore the real economy, can be seen as proxies for risks associated with business cycle fluctuations as shown by Hahn and Lee (2006). This empirical evidence highlights the importance of macroeconomic models used in business cycle or monetary policy analysis being consistent with asset market data. In this paper, I build a production-based general equilibrium framework which relies on two crucial features: first, I assume that investors have recursive preferences as in Epstein and Zin (1989) and Weil (1989); second I implement nominal price rigidities and a monetary policy rule.

A key mechanism to generate sizable risk premia is the households' aversion to volatile consumption over time, which is reflected in the level of the elasticity of intertemporal substitution (EIS). A low EIS implies that the households require a higher compensation for deviations from a smooth consumption path associated with proyclical asset payoffs. In a standard expected utility framework with an infinitely lived representative household, the EIS is the inverse of the coefficient of relative risk aversion (RRA). Assuming a low enough EIS to produce substantial equity risk premia in the stock market requires a high and implausible level of RRA (e.g. Mehra and Prescott, 1985). However, using recursive preferences allows the separation between EIS and RRA such that the household is allowed to have a low EIS without imposing a high level of RRA. The decoupling of EIS from RRA leads to a preference for the timing of resolution of uncertainty. Compared to real business cycle models introducing nominal price rigidities, a monetary policy rule makes monetary policy non-neutral and allows the model to match inflation and price nominal assets. Most importantly, the model provides a framework to assess the joint effect of monetary policy interventions, such as policy rate changes, more aggressive inflation targeting, or structural breaks on asset prices and the real economy.

Previous research on asset pricing implications of macroeconomic models in production economies has mostly focused on one single asset class at at time. One strand of the literature studies the term structure of interest rates. For example Hördahl, Tristani, and Vestin (2008) and Rudebusch and Swanson (2008) analyze DSGE models with habit preferences while Rudebusch and Swanson (2012) and Van Binsbergen, Fernández-Villaverde, Koijen, and Rubio-Ramirez (2012) use recursive Epstein-Zin preferences.

Another strand of the literature uses general equilibrium-production based asset pricing models to study the equity premium. Most authors use a real business cycle model with habit preferences (e.g. Jermann, 1998; Boldrin, Christiano, and Fisher, 2001) or endogenous long-run risks (e.g. Kaltenbrunner and Lochstoer, 2010; Croce, 2014). Wei (2009) and De Paoli, Scott, and Weeken (2010) analyze the asset pricing implications of production-based asset pricing models with habit preferences and nominal price rigidities. While Wei (2009) exclusively focuses on the equity premium, De Paoli, Scott, and Weeken (2010) make a more comprehensive analysis by studying the equity premium and the behavior of the real and nominal term structure. The latter authors solve their model numerically using second-order perturbation methods which imply time invariant risk premia. In order for risk premia to vary with the state of the economy, a higher order approximation or a global nonlinear solution method like projection is required. It is well known in the literature (e.g. Rudebusch and Swanson, 2008) that first-order approximations (log-linearization) of asset pricing models around the nonstochastic steady state eliminate higher order terms that are important when analyzing financial variables like risk premia or asset returns. However, second-order approximations or the log-linear log-normal approach used in Jermann (1998) reintroduce second- order terms but imply constant risk premia. In this paper, I therefore compute a third-order approximate solution to the model around the nonstochastic steady state.

The two papers most closely related to the present one are Kung (2015) and Swanson (2016). Kung (2015) embeds an endogenous growth framework into a standard New Keynesian DSGE model and shows that the model quantitatively explains the nominal term structure and the failure of the expectations hypothesis. Swanson (2016) presents a general equilibrium production-based asset pricing model with recursive Epstein-Zin preferences and nominal price rigidities and studies the asset pricing implications of the model for the equity premium, the real and nominal term structure as well as risk premia on defaultable bonds. Compared to the model developed in this paper, Swanson (2016) abstracts from investment dynamics by assuming a fixed capital stock, so that labor is the only variable input to production. His model produces a sizable equity premium, a downward sloping real yield curve, an upward sloping nominal yield curve and a credit spread close to the data. However, these results rely on a high coefficient of relative risk aversion of 90 and a permanent technology shock. In contrast, the production-based framework studied in this paper generates a sizable equity risk premium, a negative real term spread and a positive nominal term spread with a calibration that is more consistent with a standard calibration used in dynamic macroeconomic models, without relying on high risk aversion or permanent technology shocks. In terms of monetary policy shocks, an extension of the model with elastic labor supply predicts a stock market response which is approximately consistent with empirical estimates. In sum, the present paper can be considered as complementary to the work of Kung (2015) and Swanson (2016).

The paper is organized as follows. Section 2.2 outlines the benchmark DGSE model with Epstein-Zin preferences and nominal price rigidities. Section 2.3 discusses the calibration and explores the quantitative results of the model. Section 2.4 applies an extension of the model to study the response of the stock market to monetary policy shocks. Section 2.5 concludes the paper.

2.2 A general equilibrium macroeconomic model

2.2.1 Households

The economy is inhabited by a large number of identical, infinitely lived households. The representative household has recursive preferences over uncertain consumption streams C_t as in Epstein and Zin (1989) and Weil (1989)

$$V_t = \left[(1-\beta)\, u(C_t, N_t)^{\frac{1-\gamma}{\theta}} + \beta \left(E_t V_{t+1}^{1-\gamma} \right)^{\frac{1}{\theta}} \right]^{\frac{\theta}{1-\gamma}}, \tag{2.1}$$

where $u(C_t, N_t) \equiv C_t$ is the within-period utility function and $\left(E_t V_{t+1}^{1-\gamma} \right)^{\frac{1}{\theta}}$ is the certainty equivalent function of random future lifetime utility. The parameters in these preferences are the subjective discount factor $\beta \in (0, 1)$, the coefficient of relative risk aversion $\gamma \geq 0$ and

$$\theta \equiv \frac{1-\gamma}{1-\frac{1}{\psi}}$$

where $\psi \geq 0$ is the EIS. Epstein and Zin (1989) preferences allow the EIS to be decoupled from RRA which implies nonindifference towards the temporal resolution of consumption uncertainty (e.g. Epstein, Farhi, and Strzalecki, 2014). The representative household has a preference for early resolution of uncertainty if $\gamma > \psi^{-1}$, and has a preference for late resolution of uncertainty if $\gamma < \psi^{-1}$ or is indifferent to the resolution of uncertainty if $\gamma = \psi^{-1}$. The latter case is consistent with a CRRA utility specification where the inverse of the EIS and risk aversion coincide, which implies $\theta = 1$.

The intertemporal budget constraint of the household is

$$C_t + a'_{t+1}V_t^a + \frac{B_{t+1}}{R_{t+1}P_t} = w_tN_t + a'_t\left(V_t^a + d_t\right) + \frac{B_t}{P_t} \tag{2.2}$$

where a_t is a vector of financial assets, V_t^a and d_t are vectors of real asset prices and real dividend income received from the intermediate firms, B_{t+1} is the quantity of one-period nominal bonds, R_{t+1} is the gross one-period nominal interest rate set at time t by the central bank, w_t is the real wage, N_t is the households' labor supply and P_t is the price level. The asset portfolio a contains shares of the representative intermediate goods firm and real bonds. In addition, since leisure is not an argument of the household's utility function, it is assumed that labor supply is inelastic such that the household always supplies one unit of labor ($N_t \equiv 1$) to the intermediate firm sector.

The optimizing behavior of households implies the following Euler equation

$$1 = E_t\left[M_{t+1}\frac{P_t}{P_{t+1}}R_{t+1}\right] \tag{2.3}$$

where

$$M_{t+1} = \beta\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1}\left(\frac{V_{t+1}^{1-\gamma}}{E_tV_{t+1}^{1-\gamma}}\right)^{1-\frac{1}{\theta}} \tag{2.4}$$

is the real stochastic discount factor in this economy.[2]

2.2.2 Firms

Production consists of two sectors. There is a continuum of monopolistic competitive intermediate good firms and a representative perfectly competitive final good firm.

Final goods sector

The final output Y_t is produced from a continuum of differentiated intermediate goods $Y_{j,t}$ using the following Dixit-Stiglitz aggregator

[2] The derivation of the Euler equation and the associated stochastic discount factor are outlined in appendix 2.A.

$$Y_t = \left(\int_0^1 Y_{j,t}^{\frac{\eta-1}{\eta}} \, dj \right)^{\frac{\eta}{\eta-1}}, \tag{2.5}$$

where η is the elasticity of substitution between intermediate goods. A profit maximizing behavior of the final good firm yields the standard demand function for the intermediate good $Y_{j,t}$

$$Y_{j,t} = Y_t \left(\frac{P_{j,t}}{P_t} \right)^{-\eta}, \tag{2.6}$$

where $P_{j,t}$ is the nominal price of the intermediate good j and P_t the aggregate price index

$$P_t = \left(\int_0^1 P_{j,t}^{1-\eta} dj \right)^{\frac{1}{1-\eta}}.$$

Intermediate goods sector

Intermediate goods are produced by a continuum $j \in [0, 1]$ of monopolistically competitive firms using the following Cobb-Douglas production technology

$$Y_{j,t} = e^{Z_t} K_{j,t}^{\alpha} N_{j,t}^{1-\alpha}, \quad \alpha \in (0, 1) \tag{2.7}$$

where $K_{j,t}$ is the capital stock and $N_{j,t}$ is labor used by firm j to produce output $Y_{j,t}$. Z_t is the productivity level common to all intermediate good producers and follows an AR(1)-process

$$Z_t = \rho_Z Z_{t-1} + e^{\sigma} \varepsilon_t^Z,$$

where $\varepsilon_t^Z \sim N(0, 1)$ is uncorrelated and iid and $|\rho_Z| < 1$.

Capital accumulation for each firm j is subject to capital adjustment costs and follows the following law of motion

$$K_{j,t+1} = \left[(1-\delta) K_{j,t} + \phi \left(\frac{I_{j,t}}{K_{j,t}} \right) K_{j,t} \right] \tag{2.8}$$

where $I_{j,t}$ is investment in the capital stock and $\phi(\cdot)$ is an increasing, twice-continuously differentiable function that allows for convex capital adjustment costs. The specific functional form is as in Jermann (1998) and Kaltenbrunner and Lochstoer (2010)

$$\phi\left(\frac{I_t}{K_t}\right) \equiv \frac{a_1}{1-\zeta}\left(\frac{I_t}{K_t}\right)^{1-\zeta} + a_2,$$

where ζ is the elasticity of the investment rate with respect to Tobin's q and δ denotes the depreciation rate. If ζ is high, the capital adjustment costs are high. In other words, a large value of ζ implies that positive deviations from I_t to δK_t have a decreasing, less than proportional effect on the capital stock while a negative deviation implies an increasing and more proportional effect. This provides an incentive for the firm to limit the investment variability and avoid large deviations of I_t from δK_t. The parameters a_1 and a_2 are set such that there are no capital adjustment costs in the deterministic steady state.

As in De Paoli, Scott, and Weeken (2010) and Andreasen (2012), I assume that each intermediate firm faces quadratic price adjustment costs à la Rotemberg (1982)

$$G(P_{j,t}, P_{j,t-1}) \equiv \frac{\chi^P}{2}\left(\frac{P_{j,t}}{\bar{\pi}P_{j,t-1}} - 1\right)^2 Y_t$$

where $\bar{\pi}$ is the gross steady-state inflation rate and χ^P measures the costs of adjusting prices.

Each period, profits are paid out to households as dividends

$$d_{j,t} = Y_{j,t} - w_t N_{j,t} - I_{j,t} \tag{2.9}$$

where $d_{j,t}$ and w_t are the real dividend and the real wage, respectively.

Intermediate good firm j determines $N_{j,t}$, $I_{j,t}$, $K_{j,t+1}$ and $P_{j,t}$ by maximizing the net present value of future profits using the household's stochastic discount factor

$$\max_{\{K_{t+s+1}(j), I_{t+s}(j), P_{t+s}(j)\}_{s=0}^{\infty}} E_t \sum_{s=0}^{\infty} M_{t,t+s}\left[d_{j,t+s} - \frac{\chi^P}{2}\left(\frac{P_{j,t+s}}{\bar{\pi}P_{j,t+s-1}} - 1\right)^2 Y_{t+s}\right] \tag{2.10}$$

subject to (2.6). (2.7), (2.8) and (2.9).

The corresponding first-order conditions are:

$$w_t = \Gamma_t \left(1-\alpha\right) \frac{Y_{j,t}}{N_{j,t}}, \tag{2.11}$$

$$q_t = \frac{1}{\phi' \left(I_{j,t}/K_{j,t}\right)}, \tag{2.12}$$

$$q_t = E_t M_{t+1} \left[\alpha \Gamma_{t+1} \frac{Y_{j,t+1}}{K_{j,t+1}} - \frac{I_{j,t+1}}{K_{j,t+1}} + q_{t+1} \left(1 - \delta + \varphi \left(\frac{I_{j,t+1}}{K_{j,t+1}}\right)\right)\right] \tag{2.13}$$

$$\begin{aligned} 0 = {} & (1-\eta) \left(\frac{P_{j,t}}{P_t}\right)^{-\eta} \frac{Y_t}{P_t} - \chi^P \left(\frac{P_{j,t}}{\bar{\pi} P_{j,t-1}} - 1\right) \frac{Y_t}{\bar{\pi} P_{j,t-1}} \\ & + E_t \left[M_{t+1} \chi^P \left(\frac{P_{j,t+1}}{\bar{\pi} P_{j,t}} - 1\right) \frac{P_{j,t+1} Y_{t+1}}{\bar{\pi} P_{j,t}^2} \right] + \eta \Gamma_t \left(\frac{P_{j,t}}{P_t}\right)^{-\eta-1} \frac{Y_t}{P_t} \end{aligned} \tag{2.14}$$

where Γ_t and q_t are Lagrange multipliers of the intermediate goods price (marginal costs) and capital (Tobin's q), respectively.

2.2.3 Monetary policy

The central bank follows a simple monetary policy rule (e.g. Clarida, Gali, and Gertler, 2000) and adjusts the nominal interest rate R_{t+1} in response to the deviation of inflation π_t from its target value, which is assumed to be steady-state inflation $\bar{\pi}$, and the lagged interest rate:

$$R_{t+1} = \left(R_t\right)^{\theta^R} \left(\frac{\bar{\pi}}{\beta}\right)^{\left(1-\theta^R\right)} \left(\frac{\pi_t}{\bar{\pi}}\right)^{\theta^\pi} e^{\varepsilon_t^R} \tag{2.15}$$

where $\varepsilon_t^R \sim N(0,1)$ is a monetary policy shock, $\theta^R \in [0,1)$ governs the degree of interest rate smoothing and θ^π how sensitive the central bank is to the deviation of inflation from its target. For a determinate equilibrium, $\theta^\pi > 1$ is required.

2.2.4 Equilibrium

In general equilibrium, all markets in the economy are cleared simultaneously. This implies that nominal and real bonds are in zero net supply and the supply of stocks is normalized to one. Since all firms make identical decisions in the intermediate sector, simple aggregation yields $I_{j,t} = I_t$, $K_{j,t} = K_t$ and $N_{j,t} = N_t$. Therefore, aggregate real dividends are

$$d_t = Y_t - w_t N_t - I_t = \alpha Y_t - I_t \tag{2.16}$$

and since the model abstracts from government expenditures, the economy's aggregate resource constraint in equilibrium simplifies to

$$Y_t = C_t + I_t \tag{2.17}$$

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2.2.5 Asset pricing

The price of any asset in the model can be recursively determined using the standard stochastic discount factor approach. The basic equilibrium asset pricing condition is $p_t^i = E_t\left[M_{t+1}X_{t+1}^i\right]$, where X_{t+1}^i is the payoff from asset i in $t+1$, p_t^i its price and M_{t+1} the stochastic discount factor.

Term structure

The price of a default-free n-period nominal zero coupon bond that pays one dollar at maturity satisfies:

$$p_t^{(n)\$} = E_t\left[M_{t+1}^{\$}p_{t+1}^{(n-1)\$}\right] \tag{2.18}$$

where $M_{t+1}^{\$} = \frac{M_{t+1}}{\pi_{t+1}}$ can be considered as a nominal stochastic discount factor, $p_t^{(0)\$} \equiv 1$ and the nominal one period return is $R_{t+1} = \frac{1}{p_t^{(1)\$}}$.

The term spread is measured by the slope of the yield curve which is the difference between the yield to maturity of long-term and short-term bonds. In US data, the term spread is the difference between 10-year and 3-month Treasury notes. Since the model in the present paper is calibrated at a quarterly frequency, the long-term bond has a maturity of $n = 40$ quarters and the short-term bond of $n = 1$ quarter. To simplify the computational burden associated with the introduction of a 10-year bond, I follow Rudebusch and Swanson (2008) and assume default-free bonds that pay a geometrically declining coupon in each period in perpetuity. Hence, the nominal price of the long-term bond in period t satisfies

$$p_t^{(n)\$} = 1 + \delta_c E_t\left[M_{t+1}^{\$}p_{t+1}^{(n)\$}\right] \tag{2.19}$$

where δ_c is the rate of decay of the bond coupon and $M^{\$}_{t+1}$ is the nominal stochastic discount factor. The decay factor δ_c controls the maturity of the bond. Higher values of δ_c imply an increasing maturity while $\delta_c \to 0$ means that the bond behaves like a short-term asset.

The continuously compounded yield to maturity is given by

$$i_t^{(n)\$} = \log\left(\frac{\delta_c p_t^{(n)\$}}{p_t^{(n)\$} - 1}\right) \tag{2.20}$$

Similarly, the price of an n-period real bond is then

$$p_t^{(n)} = E_t\left[M_{t+1} p_{t+1}^{(n-1)}\right] \tag{2.21}$$

and the continuously compounded yield to maturity corresponds to

$$i_t^{(n)} = \log\left(\frac{\delta_c p_t^{(n)}}{p_t^{(n)} - 1}\right) \tag{2.22}$$

Equity

Equity is defined as a claim to intermediate firm sector dividends (profits). In equilibrium, the following condition must hold

$$\upsilon_{s,t} = E_t\left[M_{t+1}\left(\upsilon_{s,t+1} + d_{t+1}\right)\right] \tag{2.23}$$

where $\upsilon_{s,t}$ denotes the real stock price in period t and d_{t+1} denotes the expected dividend in period $t+1$. The gross one period return on equity R^E_{t+1} is defined as

$$R^E_{t+1} \equiv \frac{\upsilon_{s,t+1} + d_{t+1}}{\upsilon_{s,t}} \tag{2.24}$$

and the equity premium can be written as

$$EP_{t+1} \equiv E_t R^E_{t+1} - R_{f,t} \tag{2.25}$$

where $R_{f,t}$ denotes the gross one period real risk-free rate computed as $R_{f,t} = \frac{1}{p_t^{(1)}}$ where $p_t^{(1)}$ denotes the price of a one period real bond.

2.3 Quantitative results

This section discusses the quantitative implications of the model. The model is solved by a third-order perturbation around the non-stochastic steady state to obtain policy functions. This higher order approximation allows us to account for time variation in risk premia (as in, e.g. Rudebusch and Swanson, 2008, 2012). Dynamic macroeconomic models are typically solved using a first-order approximation (log-linearization), but this solution method reduces all risk premia in the model to zero. A second-order approximation to the model generates nonzero but constant risk premia. For risk premia to vary with the state of the economy, the model must be solved at least to a third order around the deterministic steady state. Further details about the solution method can be found in appendix 2.B.

I follow Hirshleifer, Li, and Yu (2015) and simulate the model for 400 000 quarters of artificial data using normally distributed shocks to estimate first and second unconditional moments for a variety of macroeconomic and financial variables. Caldara, Fernandez-Villaverde, Rubio-Ramirez, and Yao (2012) show that in DSGE models with recursive preferences, higher order perturbation methods are competitive in terms of accuracy with Chebyshev polynomials (projection method) and value function iteration while being much faster to run. For comparison purposes, I solve the one-sector real business cycle model studied in Jermann (1998) once within the standard expected utility framework using internal habit formation and once using Epstein-Zin preferences. The version with Epstein-Zin preferences is similar to the model analyzed in Kaltenbrunner and Lochstoer (2010).

2.3.1 Calibration

Table 2.1 reports the quarterly calibration.

Most parameter values are standard in the real business cycle literature. First, I describe the preference parameters of the representative household. The subjective discount factor β is set to 0.9942 to be consistent with the level of the risk free rate. For the coefficient of relative risk aversion (RRA) γ, I choose a value of 7 which is within the range of values suggested by Mehra and Prescott (1985) where $\gamma \in (0, 10)$, and considerably lower as compared to the values used in comparable studies. For example, Campanale, Castro, and

Table 2.1: Quarterly model calibration

Parameter	Description	Model
A. Preferences		
β	Subjective discount factor	0.9942
ψ	Intertemporal elasticity of substitution	0.0355
γ	Risk aversion	7
B. Production		
α	Capital share	0.36
δ	Depreciation	0.025
ζ	Capital adjustment costs	2.00
χ^P	Price adjustment costs	260
η	Price elasticity of demand	6
C. Productivity		
ρ	Persistence of technology shock	0.95
σ	Volatility of technology shock	1.00%
D. Monetary Policy		
θ^R	Interest rate smoothing parameter	0.75
θ^π	Taylor parameter on inflation	1.5
$\bar{\pi}$	Steady state inflation	1.009

This table reports the parameter used in the quarterly calibration of the model.

Clementi (2010) use a value of 55 and Rudebusch and Swanson (2012) a value of 75. Following Kaltenbrunner and Lochstoer (2010) and Chen (2016), the elasticity of intertemporal substitution $\psi = 0.0355$ has been chosen to roughly match the level of the risk free rate, the risk premium and the Sharpe ratio (market price of risk). The chosen EIS is similar to that in Gomes, Kogan, and Yogo (2009). Moreover, this value of the EIS is also consistent with empirical estimates and falls in the confidence interval of econometric studies (e.g. Hall, 1988; Yogo, 2004). However, since the calibration of RRA and EIS implies $\psi^{-1} > \gamma$, households in the economy have a preference for a later resolution of uncertainty. Kaltenbrunner and Lochstoer (2010) show that such a calibration and the associated consumption smoothing motive of the household generate a highly persistent variation in expected consumption growth when technology shocks are transitory and identically and independently distributed. The endogenous time variation in expected consumption growth is similar to the exogenous process specified in the long-run risk model of Bansal and Yaron (2004) and helps me generate a high Sharpe ratio of stock returns, even when the volatility of consumption growth and the RRA are low.

The production side of the economy consists of a final and an intermediate goods sector. The price elasticity of demand η in the final goods sector is calibrated to a value of 6, which corresponds to a markup of 20%. For the monopolistically competitive intermediate goods firm, the capital share α is set to 0.36 as in Boldrin, Christiano, and Fisher (2001), and the depreciation rate of capital δ to 0.025 as in Jermann (1998). Following Andreasen (2012), I calibrate the price adjustment cost parameter χ^P to 260 which is equivalent to a standard Calvo-coefficient of 0.75, implying that firms change their prices once a year. For the capital adjustment cost parameter ζ, different values have been used in the literature since empirical studies do not offer a precise guidance for calibrating this parameter. Jermann (1998) uses a value of 4.35 and Hirshleifer, Li, and Yu (2015) choose a value of 1.50. For my benchmark model, I choose an intermediate value of 2.00 which improves the model's ability to jointly match macroeconomic and financial variables. Next, I calibrate the parameters for the stationary technology process. Following Jermann (1998) and Campanale, Castro, and Clementi (2010), the autocorrelation coefficient ρ_z is set to 0.95, while the parameter σ is chosen such that the model generates an output growth volatility of about 1%.

For the calibration of the parameters in the monetary policy rule, I follow De Paoli, Scott, and Weeken (2010). The parameter governing interest rate sensitivity to deviations in inflation θ_π is set to 1.5. The parameter that governs the degree of interest rate smoothing is set to 0.75. These values are consistent with the range of estimates from the literature (e.g. Clarida, Gali, and Gertler, 2000). Steady-state inflation $\bar{\pi}$ is calibrated to match the average level of inflation in the data. As in Rudebusch and Swanson (2008), I set the rate of decay of the coupon on the consol δ_c to 0.9848, which implies a maturity of 10 years in the model.

The calibration for the two comparison models is the same as in Jermann (1998) except that the persistence parameter of the technology shock is set to 0.95 as in the benchmark model. Furthermore, the coefficient of relative risk aversion in the specification with Epstein-Zin preferences is set to 15 while the elasticity of intertemporal substitution is calibrated to 0.05 as in Kaltenbrunner and Lochstoer (2010). This calibration allows the model to match a market price of risk that is roughly consistent with the data.

2.3.2 Asset Pricing implications and business cycle properties

The main results for all model specifications are found in table 2.2.

Table 2.2: Macroeconomic and asset pricing moments

	Data	Jer98	Jer98EZ	Benchmark
A. Macroeconomic moments				
$\sigma(\Delta c_t)/\sigma(\Delta y_t)$	0.52	0.40	0.66	0.38
$\sigma(\Delta i_t)/\sigma(\Delta y_t)$	2.93	3.12	1.91	2.46
$\sigma(\Delta c)$	1.42	0.40	0.68	0.38
$E(\pi)$	3.74	-	-	3.29
$\sigma(\pi)$	1.64	-	-	0.78
AC1(Δc)	0.37	0.66	-0.02	-0.01
AC1(π)	0.73	-	-	0.70
$corr(\pi, \Delta c)$	-0.56	-	-	-0.75
B. Asset pricing moments				
$E(r_{f,t})$	0.86	0.41	0.76	0.58
$\sigma(r_{f,t})$	0.97	15.87	6.23	3.61
$E(R_{E,t} - r_{f,t})$	6.33	7.50	4.32	5.02
$\sigma(R_{E,t})$	19.42	23.44	16.95	22.78
Sharpe Ratio	0.33	0.32	0.25	0.22
Real Term Spread (TS)	-2.00	3.28	0.37	-0.36
Nominal Term Spread (TSN)	1.25	-	-	3.45

This table presents the means and standard deviations for key macroeconomic and asset pricing variables. The model is calibrated at a quarterly frequency. The second and third column report the result for the Jermann (1998) model with habits and Epstein-Zin preferences, respectively. The fourth column presents the results for the benchmark model. The moments in the data column are taken from Bansal and Yaron (2004) and Kaltenbrunner and Lochstoer (2010), and are both based on a data sample from 1929-1998. Inflation data are taken from Kung (2015).

All models reproduce roughly key macroeconomic variables consistent with the data. Consumption growth is smoother, while investment growth is more volatile than output growth. Figure 2.1 plots impulse response functions for selected variables to a positive one-standard-deviation technology shock for the above benchmark calibration.

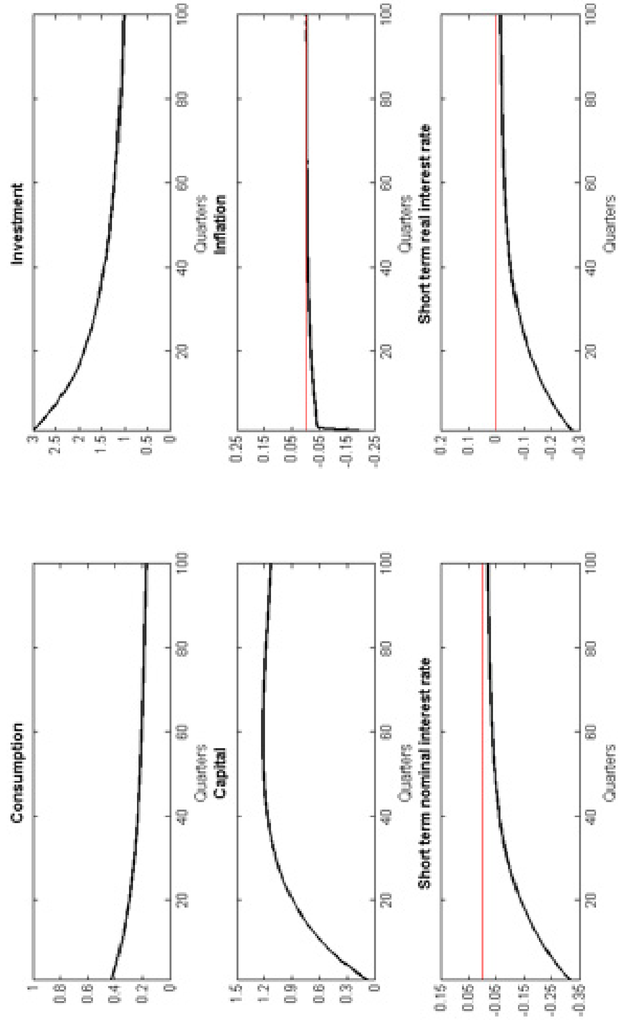

Figure 2.1: Impulse response functions for consumption, investment, capital, inflation, short-term nominal interest rates and short-term real interest rates to a positive one-standard-deviation technology shock, plotted as a percentage deviation from its steady-state value.

In response to a positive technology shock, consumption and investment take a jump upwards. An increase in technology increases output and makes households feel wealthier which in turn, leads to a rise in consumption demand. Since households want to take advantage of the higher productivity in the economy, they increase investments in the capital stock which allows them to smooth consumption over time. The magnitude of these dynamics depends on the households' preferences and in particular on the EIS. A low EIS means that the households dislike changes in consumption over time which has an important asset pricing implication for reasons explained below. In general, the initial response on consumption will be larger, the higher the EIS, the higher the capital adjustment costs and the higher the persistence of the shock.

The higher level of technology reduces the firms' marginal costs of production and causes inflation to fall. This decrease in inflation makes the monetary authority reduce the nominal interest rate according to its policy rule. The nominal interest rate declines by about 32 basis points as concerns impact as a response to the technology shock. A fall of inflation less than the nominal interest rate after the shock implies a decrease in the real interest rate.

From an asset pricing perspective, all models generate a sizable equity premium as well as a stock return volatility similar to the empirical estimates reported in the literature (e.g. Bansal and Yaron, 2004).

Jermann (1998) reproduces these facts by implementing internal habits together with relatively high convex adjustment costs in capital. Internal habits increase the curvature of the household's utility function which implies higher risk aversion and a lower EIS. As a consequence, the combination of these features allows the model to attain a sizable equity premium and a high stock return volatility while roughly matching the relative standard deviation of consumption growth to the standard deviation of output growth. The intuition for this is as follows. A low EIS implies a strong aversion to fluctuations in consumption and thus, a strong consumption smoothing motive. However, capital adjustment costs and inelastic labor supply prevent the household from doing so, which means that households require a larger compensation for holding risky assets.

With Epstein-Zin preferences, the same results can be achieved by calibrating the EIS to a relatively low level. As previously mentioned, the EIS is the parameter that controls the household's sensitivity to deviations from a smooth consumption path. A lower ψ implies a higher sensitivity. Intuitively, this means that the lower is the EIS, the higher is the compensation that the household requires for deviations from a smooth consumption stream associated with procyclical asset payoffs. The only way of transferring consumption intertemporally in this economy is through financial markets by buying stocks or riskfree bonds. A low EIS makes the household demand a higher compensation for holding risky stocks, which implies a larger equity premium. Moreover, a decomposition of the Sharpe ratio shows that the maximum amount of risk $\gamma\sigma(\Delta c)$ (short-run risk) that would arise under a standard power utility framework ($\gamma = \psi^{-1}$) is only 0.0276 (13%). The residual of 0.1861 (87%) is due to recursive preferences and the associated endogenous time variation in expected consumption growth caused by the household's consumption smoothing motive (long-run risk). However, a direct consequence of the household's reluctance to substitute current consumption with future consumption is the very low volatility and autocorrelation in consumption growth generated by the model. To generate a higher volatility and autocorrelation in consumption growth, the model requires a higher EIS. The cost of doing this is that the equity premium is much lower and at odds with the data.

Turning to the bond market, the volatility of the risk-free rate is substantially lower in the benchmark model. This is noteworthy, given the low EIS used in the model calibration. A low coefficient of intertemporal substitution implies a strong precautionary savings motive because the household is concerned to achieve a smooth consumption stream. Since changing the capital stock is costly and labor supply is inelastic, risk-free bonds provide the most efficient instrument to substitute consumption intertemporally. This consumption smoothing behavior and the fact that bonds are in zero-net supply cause the large risk-free rate variability generally observed in standard macroeconomic DSGE models. The reason why there is less variation in the risk-free rate in the benchmark model is that interest rate smoothing in the monetary policy rule ensures that the low EIS does not translate into an excessively high volatility of short-term interest rates.

The real yield curve, defined as the average of the ten-year minus one-quarter real bond spread, is downward sloping which is consistent with the empirical evidence reported in Barr and Campbell (1997) and Evans (1998) for long sample UK data. Bansal, Kiku, and Yaron (2012) extend the dataset of Evans (1998) and confirm his finding of a negatively sloped real curve for a more recent sample of 1996.07 - 2008.12. Therefore, they argue that a downward-sloping real yield curve is the appropriate target for models. However, for a short sample of US data, Piazzesi and Schneider (2007) and Beeler and Campbell (2012) report an upward sloping real yield curve.

The average ten-year minus one-quarter nominal bond spread is about 1% to 1.5% in the data and implies a positively sloped nominal yield curve (e.g. Gürkaynak, Sack, and Wright, 2007). The macro model generates a term spread of 3.45% which is substantially larger compared to the value in the data. This result is driven by the negative correlation between inflation and consumption growth and the associated high inflation risk premium. Increases in inflation are bad news for consumption growth. In a recession when consumption is low and the household wants to have more resources to consume (small consumption growth), high inflation reduces nominal bond returns. A positive and sizable term premium is required for the household to hold long-term nominal bonds. This mechanism is also confirmed from impulse response analysis. As shown in figure 2.2, the fall in inflation associated with a positive technology shock causes bond prices to rise since long-term nominal bonds are now considered less risky to hold.

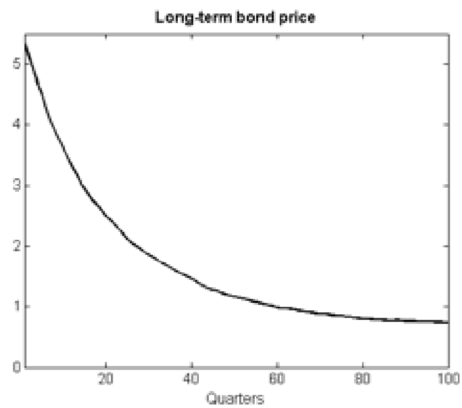

Figure 2.2: Impulse response function for long-term bond price to a positive one-standard-deviation technology shock, plotted as a percentage deviation from its steady-state value.

2.3.3 Stock return predictability

A large literature in empirical asset pricing documents the predictability of stock returns by the price-dividend ratio (e.g. Campbell and Shiller, 1988b; Fama and French, 1988; Cochrane, 2008b), which implies that the expected stock returns vary over time. This time variation has a direct business cycle implication since the expected returns are high in economic downturns or recessions when people are less willing to hold risky assets. In order to verify if the structural macroeconomic model is consistent with this empirical evidence, I run simple forecasting regressions where I regress the 1-, 3-, and 5-year excess cumulative stock market return (risk premia) onto the lagged dividend-price ratio.

$$R^e_{t \to t+k} = \alpha + \beta \frac{d_t}{p_t} + \varepsilon_{t+k}$$

The regression model is estimated 1250 times, each with a length of 80 years, consistent with the sample period in the data. Figure 2.3 shows the distribution of slope coefficients, t-statistics and R^2s.

The results of the predictability regressions reported in table 2.3 are based on median values and show that the model is able to generate slope coefficients and sizable R^2s that are consistent with the data. The slope coefficients are positive and increasing with the horizon. This implies that when stock prices are high as compared to dividends, the expected future excess returns are low. In sum, the model is able to produce a significant stock return predictability.

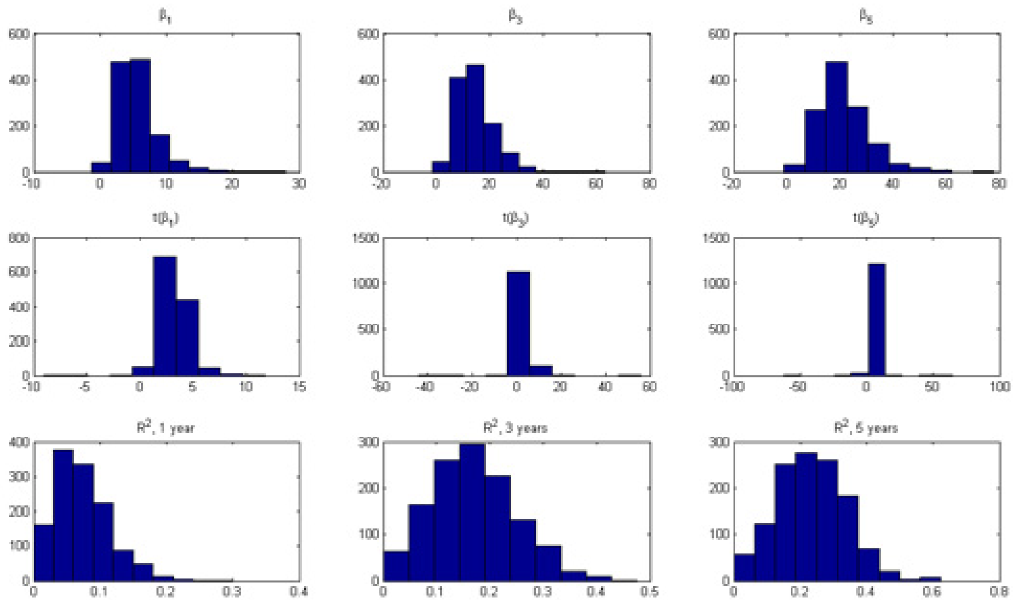

Figure 2.3: Historgram of coefficients, t-statistics and R^2s of predictability regressions

Table 2.3: Stock return predictability

	Data			Model		
	Horizon (in years)					
	1	3	5	1	3	5
β	4.0	7.9	20.6	5.10	13.19	19.73
$t(\beta)$	2.7	3.0	2.6	3.15	3.56	3.84
R^2	0.08	0.20	0.22	0.07	0.16	0.23

This table reports stock market excess return forecasts for a horizon of 1-, 2-, and 5-years for the following regression specification $R^e_{t\to t+k} = \alpha + \beta \frac{d_t}{p_t} + \varepsilon_{t+k}$. The results in the data are from Cochrane (2008a) and are based on a data sample from 1927-2005. The regression model is estimated 1250 times each with a length of 80 years which is consistent with the sample period in the data. The reported coefficients, t-statistics and R^2s are median values. Standard errors use GMM (Hansen-Hodrick) to correct for heteroskedasticity and serial correlation.

2.4 Extension to endogenous labor supply and application to stock prices and monetary policy shocks

The empirical literature reports evidence that an unanticipated monetary policy shock has a significant impact on stock prices. A 100-basis points increase in the nominal interest rate is associated with a decrease in stock market valuation ranging from -2.20% to -9.00% for US data (e.g. Bernanke and Kuttner, 2005) and -1.20% to -9.40% for European data (e.g. Kholodilin, Montagnoli, Napolitano, and Siliverstovs, 2009).[3]

In this context, I apply the benchmark model developed in section 2.2 to study the reaction of the stock market to a monetary policy shock. To be consistent with the existing literature (e.g. Challe and Giannitsarou, 2014), I extend the model by allowing elastic labor supply which further enables me to analyze the labor market impact of a monetary policy shock. The household's within-period utility function is defined as

$$u(C_t, N_t) \equiv C_t^{v} \left(1 - N_t\right)^{1-v}$$

where $v \in (0,1)$ is a parameter which controls the relative weight of consump-

[3]See Challe and Giannitsarou (2014) for a survey of empirical studies.

tion and leisure.[4] The parameter υ is calibrated in such a way that in steady state, the household devotes $N_{ss} = 1/3$ of her time endowment to work. Given the functional form of u, it follows from the household's optimization problem that the equilibrium wage satisfies

$$w_t = \frac{1-\upsilon}{\upsilon}\frac{C_t}{N_t} \tag{2.26}$$

and the stochastic discount factor in this economy is

$$M_{t+1} = \beta \left(\frac{C_{t+1}}{C_t}\right)^{\frac{\upsilon(1-\gamma)}{\theta}-1} \left(\frac{1-N_{t+1}}{1-N_t}\right)^{\frac{(1-\gamma)(1-\upsilon)}{\theta}} \left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}}\right)^{1-\frac{1}{\theta}}. \tag{2.27}$$

The optimization problem of the representative firm is the same as in the benchmark model, which means that all other equilibrium conditions remain unchanged. I solve the model with a third-order approximation around the nonstochastic steady state.

A 100-basis points contractionary monetary policy shock to the monetary policy rule outlined in equation (2.15) implies a 2.12% decrease in stock prices in the model. This model outcome is approximately consistent with the empirical estimates, even if the model implied stock market multiplier is at the lower bound of the plausible values reported for US data. For comparison, the New Keynesian macro model developed in Challe and Giannitsarou (2014) generates a stock market multiplier of -3.0691.
As shown in figure 2.2, the monetary policy shocks reduce output, consumption, investment and inflation as implied by economic theory. However, the dynamic response of these macroeconomic variables is not as hump-shaped as in empirical impulse-responses (e.g. Christiano, Eichenbaum, and Evans, 2005), meaning that they return to pre-shock levels too quickly.

Counterfactually to empirical estimates, a contractionary monetary policy shock implies an increase in dividend payments. Intuitively, a rise in the interest rate decreases aggregate demand due to intertemporal substitution of consumption. This substitution effect has a direct impact on firm sales which results in a fall in profits. The firm responds to the policy shock by producing less, which affects the labor market adversely and puts downward pressure

[4] A parametrization with inelastic labor supply as in the benchmark model would imply a value of $\upsilon = 1$.

on the equilibrium real wage and thereby reduces the firms' marginal costs. In consequence, this indirect general equilibrium effect contributes to an increase in profits. As explained in Challe and Giannitsarou (2014), the general equilibrium effect dominates the direct effect if the wages are fully flexible.

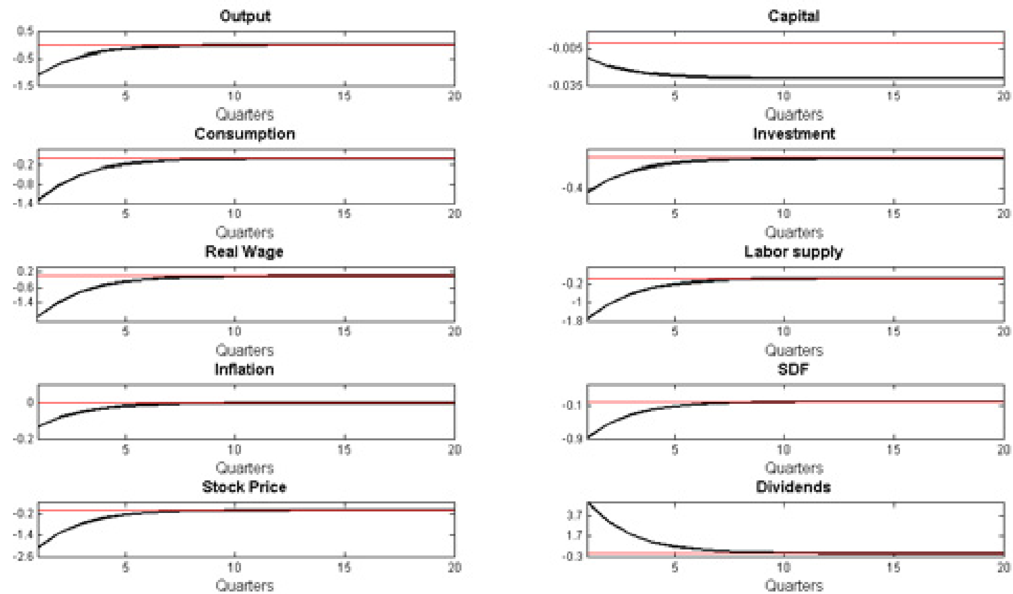

Figure 2.4: Impulse response function to a positive one-standard-deviation monetary policy shock, plotted as the percentage deviation from its steady-state value.

2.5 Conclusion

This paper studies the asset pricing implications of a New Keynesian DSGE model for major asset classes. Key features of the model are recursive preferences, nominal price rigidities and a monetary policy rule. The model is able to reproduce stylized facts observed in financial markets such as a sizable equity premium, a positively sloped nominal term structure, a negative real term spread and the predictability of stock returns without compromising the model's ability to fit key macroeconomic variables. Compared to production based asset pricing models within the real business cycle framework, the present model is able to generate a relatively low risk-free rate volatility. In terms of monetary policy shocks, the model generates a stock market multiplier which is consistent with the empirical estimates.

2.A Household's optimization problem

In this appendix, I derive the stochastic discount factor by solving the representative household's optimization problem. The elements of the asset vector a are $a = \begin{bmatrix} s_t & h_t \end{bmatrix}$ and the elements of the price vector V_t^a are $V_t^a = \begin{bmatrix} \upsilon_{s,t} & \upsilon_{b,t} \end{bmatrix}$, which implies that the household's budget constraint in equation (2.2) can be stated as

$$C_t + \upsilon_{s,t}(s_{t+1} - s_t) + \upsilon_{b,t} h_{t+1} = w_t + h_t + s_t d_t$$

For simplicity, I abstract from nominal bonds which have no impact on the formal derivation. The household then solves

$$\begin{aligned}
&\max_{C_t, s_{t+1}, h_{t+1}} V_t = \left[(1-\beta)\left(C_t\right)^{\frac{1-\gamma}{\theta}} + \beta \left(E_t V_{t+1}^{1-\gamma} \right)^{\frac{1}{\theta}} \right]^{\frac{\theta}{1-\gamma}} \\
&s.t. \\
&C_t + \upsilon_{s,t}(s_{t+1} - s_t) + \upsilon_{b,t} h_{t+1} = w_t + h_t + s_t d_t \qquad \forall t
\end{aligned}$$

which can be stated as a Lagrangian

$$\begin{aligned}
L \;=\; & \left[(1-\beta)\left(C_t\right)^{\frac{1-\gamma}{\theta}} + \beta \left(E_t V_{t+1}^{1-\gamma} \right)^{\frac{1}{\theta}} \right]^{\frac{\theta}{1-\gamma}} + \\
& \lambda_t \left(w_t + h_t + s_t d_t - C_t - \upsilon_{s,t}(s_{t+1} - s_t) - \upsilon_{b,t} h_{t+1} \right)
\end{aligned}$$

where λ_t denotes the Lagrange multiplier.

The first-order conditions with respect to C_t and s_{t+1} for an interior optimum (i.e. $C_t > 0$) are

$$\frac{\partial L_t}{\partial C_t} = \left[(1-\beta)\left(C_t\right)^{\frac{1-\gamma}{\theta}} + \beta \left(E_t V_{t+1}^{1-\gamma} \right)^{\frac{1}{\theta}} \right]^{\frac{\theta}{1-\gamma}-1} (1-\beta)\left(C_t\right)^{\frac{1-\gamma}{\theta}-1} - \lambda_t = 0$$

and

$$\frac{\partial L_t}{\partial s_{t+1}} = \frac{\theta}{1-\gamma} \left[\ldots \right]^{\frac{\theta}{1-\gamma}-1} \frac{\beta}{\theta} \left(E_t V_{t+1}^{1-\gamma} \right)^{\frac{1}{\theta}-1} (1-\gamma) \left[E_t V_{t+1}^{-\gamma} \frac{\partial V_{t+1}}{\partial s_{t+1}} \right] - \lambda_t \upsilon_{s,t} = 0$$

Using the fact that

$$\begin{aligned}
\left[(1-\beta)\left(C_t\right)^{\frac{1-\gamma}{\theta}}+\beta\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}}\right]^{\frac{\theta}{1-\gamma}-1} &= \left[(1-\beta)\left(C_t\right)^{\frac{1-\gamma}{\theta}}+\beta\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}}\right]^{\frac{\theta}{1-\gamma}\left(1-\frac{1-\gamma}{\theta}\right)} \\
&= V_t^{1-\frac{1-\gamma}{\theta}}
\end{aligned}$$

the two first-order conditions can be rewritten more compactly as

$$\frac{\partial L_t}{\partial C_t} = (1-\beta) V_t^{1-\frac{1-\gamma}{\theta}}\left(C_t\right)^{\frac{1-\gamma}{\theta}-1}-\lambda_t=0 \tag{2.28}$$

$$\frac{\partial L_t}{\partial s_{t+1}} = \beta V_t^{1-\frac{1-\gamma}{\theta}}\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}-1}\left[E_t V_{t+1}^{-\gamma} \frac{\partial V_{t+1}}{\partial s_{t+1}}\right]-\lambda_t \upsilon_{s,t}=0. \tag{2.29}$$

Now note that

$$\frac{\partial V_{t+1}}{\partial C_{t+1}}=(1-\beta) V_{t+1}^{1-\frac{1-\gamma}{\theta}}\left(C_{t+1}\right)^{\frac{1-\gamma}{\theta}-1} \text{ and } \frac{\partial C_{t+1}}{\partial s_{t+1}}=\left(\upsilon_{s,t+1}+d_{t+1}\right)$$

which implies that the term $\frac{\partial V_{t+1}}{\partial s_{t+1}}$ in equation (2.29) can be rewritten as

$$\frac{\partial V_{t+1}}{\partial s_{t+1}}=\frac{\partial V_{t+1}}{\partial C_{t+1}} \frac{\partial C_{t+1}}{\partial s_{t+1}}=(1-\beta) V_{t+1}^{1-\frac{1-\gamma}{\theta}}\left(C_{t+1}\right)^{\frac{1-\gamma}{\theta}-1}\left(\upsilon_{s,t+1}+d_{t+1}\right)$$

Combining equation (2.28) and (2.29) yields

$$\upsilon_{s,t}=\beta \frac{V_t^{1-\frac{1-\gamma}{\theta}}\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}-1}}{(1-\beta) V_t^{1-\frac{1-\gamma}{\theta}}\left(C_t\right)^{\frac{1-\gamma}{\theta}-1}} E_t\left[V_{t+1}^{-\gamma}(1-\beta) V_{t+1}^{1-\frac{1-\gamma}{\theta}}\left(C_{t+1}\right)^{\frac{1-\gamma}{\theta}-1}\left(\upsilon_{s,t+1}+d_{t+1}\right)\right]$$

which can be simplified to

$$\begin{aligned}
1 &= \beta\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}-1} E_t\left[V_{t+1}^{1-\frac{1-\gamma}{\theta}-\gamma}\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1} \frac{\left(\upsilon_{s,t+1}+d_{t+1}\right)}{\upsilon_{s,t}}\right] \\
&= \beta\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}-1} E_t\left[V_{t+1}^{\frac{\theta-1-\gamma-\theta\gamma}{\theta}}\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1} \frac{\left(\upsilon_{s,t+1}+d_{t+1}\right)}{\upsilon_{s,t}}\right]
\end{aligned}$$

Note that since $\frac{\theta-1-\gamma-\theta\gamma}{\theta}=\frac{\theta(1-\gamma)-(1-\gamma)}{\theta}=\frac{(1-\gamma)(\theta-1)}{\theta}=(1-\gamma)\left(1-\frac{1}{\theta}\right)$ it follows that

$$1=\beta\left(E_tV_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}-1}E_t\left[V_{t+1}^{(1-\gamma)\left(1-\frac{1}{\theta}\right)}\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1}\frac{(\upsilon_{s,t+1}+d_{t+1})}{\upsilon_{s,t}}\right]$$

Rewriting yields

$$1=\beta E_t\left[\left(\frac{V_{t+1}^{1-\gamma}}{\left(E_tV_{t+1}^{1-\gamma}\right)}\right)^{1-\frac{1}{\theta}}\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1}\frac{(\upsilon_{s,t+1}+d_{t+1})}{\upsilon_{s,t}}\right]$$

Defining $R_{s,t+1}\equiv\frac{(\upsilon_{s,t+1}+d_{t+1})}{\upsilon_{s,t}}$, it follows that

$$1=\beta E_t\left[\left(\frac{V_{t+1}^{1-\gamma}}{\left(E_tV_{t+1}^{1-\gamma}\right)}\right)^{1-\frac{1}{\theta}}\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1}R_{s,t+1}\right]$$

where

$$M_{t,t+1}=\beta\left(\frac{C_{t+1}}{C_t}\right)^{\frac{1-\gamma}{\theta}-1}\left(\frac{V_{t+1}^{1-\gamma}}{\left(E_tV_{t+1}^{1-\gamma}\right)}\right)^{1-\frac{1}{\theta}}$$

is the stochastic discount factor (SDF) to price reals assets in the economy. More compactly, the Euler equation can be written as

$$1=E_t\left[M_{t+1}R_{t+1}\right]$$

or

$$1=E_t\left[M_{t+1}\frac{P_t}{P_{t+1}}R_{t+1}\right]$$

to price nominal assets.

2.B Perturbation

In this exposition, I follow Schmitt-Grohé and Uribe (2004) and van Binsbergen (2008). The set of equilibrium conditions that solve the DSGE model can be written as

$$E_t f\left(y_{t+1}, y_t, x_{t+1}, x_t\right) = 0 \tag{2.30}$$

where x_t is an $n_x \times 1$ vector of state variables and y_t is an $n_y \times 1$ vector of control variables. The vector x_t of state variables can be partitioned into endogenous state variables and exogenous state variables. The total number of variables and equations in the model is $n = n_x + n_y$. The function f maps $\mathbb{R}^{n_y} \times \mathbb{R}^{n_y} \times \mathbb{R}^{n_x} \times \mathbb{R}^{n_x}$ into $\mathbb{R}^n$.

As shown in Schmitt-Grohé and Uribe (2004), the solution of the model takes the form:

$$y_t = g\left(x_t, \sigma\right) \tag{2.31}$$

$$x_{t+1} = h\left(x_t, \sigma\right) + \sigma\varepsilon_{t+1} \tag{2.32}$$

where the function g maps $\mathbb{R}^{n_x}$ into $\mathbb{R}^{n_y}$ and the function h maps $\mathbb{R}^{n_x}$ into $\mathbb{R}^{n_x}$. The scalar $\sigma \geq 0$ is the perturbation parameter and ε_{t+1} is an $n_x \times 1$ vector of shocks. The main idea of perturbation is to interpret the solution to the model as a function of the state vector x_t and of the perturbation parameter σ which scales the amount of uncertainty in the economy. In the deterministic steady state of the model, σ equals 0 and in the stochastic version of the model, σ equals 1. Typically, the functions g and h are unknown and a perturbation method finds a local approximation around the non-stochastic steady state. More specifically, a local approximation is an approximation that is valid in the neighborhood of a certain point $(\bar{x}, \bar{\sigma})$. For ease of exposition and to simplify the notation, I only use one control variable, one endogenous state variable x_1 and one exogenous state variable x_2. A Taylor series approximation of the functions g and h around the point $(x, \sigma) = (\bar{x}, \bar{\sigma})$ yields

$$
\begin{aligned}
g(x_1,x_2,\sigma) = g(\bar{x}_1,\bar{x}_2,\bar{\sigma}) &+ \sum_{i=1}^{2} g_{x_i}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_i-\bar{x}_i) + g_{\sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma}) \\
&+ \sum_{i=1}^{2} g_{x_i x_i}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_i-\bar{x}_i)^2 + g_{x_1 x_2}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_1-\bar{x}_1)\,(x_2-\bar{x}_2) \\
&+ \sum_{i=1}^{2} g_{x_i \sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_i-\bar{x}_i)\,(\sigma-\bar{\sigma}) + g_{\sigma\sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma})^2 + \ldots
\end{aligned}
\tag{2.33}
$$

and

$$
\begin{aligned}
h(x_1,x_2,\sigma) = h(\bar{x}_1,\bar{x}_2,\bar{\sigma}) &+ \sum_{i=1}^{2} h_{x_i}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_i-\bar{x}_i) + h_{\sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma}) \\
&+ \sum_{i=1}^{2} h_{x_i x_i}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_i-\bar{x}_i)^2 + h_{x_1 x_2}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_1-\bar{x}_1)\,(x_2-\bar{x}_2) \\
&+ \sum_{i=1}^{2} h_{x_i \sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(x_i-\bar{x}_i)\,(\sigma-\bar{\sigma}) + h_{\sigma\sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma})^2 + \ldots
\end{aligned}
\tag{2.34}
$$

The influence of uncertainty in the model on the control and state variables is measured by the terms $g_{\sigma\sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma})^2$ and $h_{\sigma\sigma}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma})^2$. Moreover, this second-order approximation will generate constant risk premia. To obtain risk premia that vary linearly with the state variables, terms of the form

$$
g_{\sigma\sigma x_i}(\bar{x}_1,\bar{x}_2,\bar{\sigma})\,(\sigma-\bar{\sigma})^2\,(x_i-\bar{x}_i) \tag{2.35}
$$

are required. To obtain these terms, a third-order approximation of the policy function is needed. To identify these n^{th} order derivatives of the function g and h evaluated at the point $(x,\sigma)=(\bar{x},\bar{\sigma})$, substitute the proposed solution in equations (2.31) and (2.32) into (2.30) to obtain

$$
F\,(x,\sigma) = E_t f\left(g\left(h\,(x,\sigma)+\sigma\varepsilon',\sigma\right), g\,(x,\sigma)\,, h\,(x,\sigma)+\sigma\varepsilon', x\right) = 0 \tag{2.36}
$$

where I drop time subscripts and use a prime to indicate variables dated in period $t+1$. Due to the fact that $F(x,\sigma)$ must be equal to zero for any possible values of x and σ, it must be the case that the derivatives of any order of F must also be equal to zero:

$$F_{x^k\sigma^j}(x,\sigma) = 0 \quad \forall x, \sigma, j, k \tag{2.37}$$

where $F_{x^k\sigma^j}(x,\sigma)$ denotes the derivative of F with respect to x taken k times and with respect to σ taken j times. Solving this exactly identified system of equations gives the values for each of the n^{th} order derivatives.

Chapter 3

Macroeconomic news and the stock market: Evidence from the eurozone[1]

3.1 Introduction

A fundamental question in economics is the interaction between financial markets and the macroeconomy. In a recent study using a long sample (1958-2009) of US stock market data, Savor and Wilson (2013) show that excess returns are significantly higher on days when important macroeconomic news about inflation, unemployment, or interest rates is scheduled for announcement. Their main explanation for these higher risk premia is that investors are compensated for higher macroeconomic risks around announcement days. Intuitively, the stock market tends to perform poorly when news about the state of the economy creates uncertainty, which requires risk-averse investors to demand a higher expected excess return on risky assets. Previous research mainly focused on the sensitivity of realized returns to announcement surprises. Two important contributions in that literature using US data are Boyd, Hu, and Jagannathan (2005) and Bernanke and Kuttner (2005). Boyd, Hu, and Jagannathan (2005) study the sensitivity of stock returns to unemployment surprises and find a positive stock market response to news of rising unemployment sur-

[1]Thanks to Roine Vestman for his advice and to Rickard Sandberg and Irina Zviadadze for their comments.

prises during economic expansions and a negative response during economic contractions. Bernanke and Kuttner (2005) use an event study approach to investigate the impact of Federal Open Market Committee (FOMC) interest rate announcement surprises on stock market returns and document a sizable and significant response to unexpected policy rate decisions.

In more recent work, Haitsma, Unalmis, and de Haan (2016) examine how stock markets respond to the policies of the European Central Bank (ECB) and find that especially unconventional monetary policy surprises affect the European stock market.

Compared to this body of research, the focus of this paper is, as in Savor and Wilson (2013), to study the effect of prescheduled announcements on expected returns. In terms of standard factor models, this means that I study factor risk premia as opposed to factor betas. Next to the work of Savor and Wilson (2013), there are a few other papers that focus on the effect of macroeconomic announcement days on financial markets. Jones, Lamont, and Lumsdaine (1998) document an increase in volatility and higher excess returns on announcement days for long-term US Treasury bonds. In more recent work, Lucca and Moench (2015) find that since 1994, over 80% of the equity premium are earned in the 24 hours prior to scheduled FOMC announcements. Cieslak, Morse, and Vissing-Jorgensen (2015) document that since 1994, the equity premium in the US and worldwide is earned entirely in even weeks (0,2,4 and 6) in the FOMC cycle (the FOMC meets about every 6 weeks) starting from the last FOMC meeting.

In this paper, I extend the results in Savor and Wilson (2013) and investigate whether there exists an equivalent announcement day premium for the eurozone. For that purpose, I identify the difference between expected returns on announcement days and expected returns on other trading days for the aggregate stock market.The sample period (3 January 2000 to 30 January 2015) includes a substantial variation in economic conditions and the crisis period which led to historically low interest rates in most advanced economies. A major implication of this development is that the sample contains a long period where the policy rate in the eurozone is in the vicinity of the zero lower bound. This offers the opportunity to test for structural breaks and to analyze if the stock market behaves differently on announcement days in an environment characterized by higher uncertainty and a monetary policy regime with interest rates close to zero.

Previewing the results, I identify state dependence such that equity risk premia for the Euros Stoxx index on announcement days are significantly higher when interests rates are in the vicinity of the zero lower bound. Moreover, I provide evidence that for the whole sample period, average excess returns in the eurozone are only higher on days when FOMC announcements are scheduled for release. However, this result vanishes in a low interest rate regime. Finally, I document that the European stock market does not command a premium for scheduled announcements by the ECB.

The remainder of the paper is organized as follows. In section 3.2, I describe the data set. Section 3.3 presents and discusses the empirical results. Section 3.4 concludes the paper.

3.2 Data and sample

The dates of prescheduled monthly macroeconomic news announcements are obtained from Eurostat and the ECB from January 2000 to January 2015. Although not as long as in Savor and Wilson (2013), the sample period includes many eminent financial events such as the bust of the dot-com bubble in 2000, the global financial crisis of 2008-2009, and the more recent eurozone crisis. In the choice of announcement types, I follow Savor and Wilson (2013). Over the sample period, there are 178 prescheduled CPI announcements[2], 182 unemployment announcements and 203 scheduled ECB interest rate announcements. 12 announcement days in my sample had more than 1 announcement. Overall, the sample consists of 551 announcement days and 3318 non-announcement days.

I obtain daily prices for the EURO STOXX stock market index[3] from the STOXX website. From these daily price data, I compute continuous stock returns, $r_t = ln(P_t) - ln(P_{t-1})$. To calculate excess returns, $r_t^e \equiv r_t - r_f$, I infer a daily risk-free rate, r_f, from the 1 month EURIBOR (obtained from Datastream) assuming it to be constant over the month.

[2]Consumer price inflation in the euro area is measured by the Harmonised Index of Consumer Prices (HICP).

[3]The index represents large, mid and small capitalization companies of 12 eurozone countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain.

3.3 Stock returns around scheduled macroeconomic announcements

3.3.1 Baseline analysis

In this section, I verify how scheduled macroeconomic announcements affect asset prices in the eurozone. For that purpose, I run a simple time series regression of daily stock market excess returns (r_t^e) on a constant and a dummy variable $Announcement$, which equals 1 on a trading day where macroeconomic news is scheduled for release and 0 otherwise

$$r_t^e = \alpha + \beta Announcement_t + \varepsilon_t. \tag{3.1}$$

The regression coefficients are estimated using ordinary least squares (OLS) and t-statistics are computed using Newey and West (1987) robust standard errors with 5 lags. The coefficient α in this regression model measures the unconditional mean excess return earned on all non-announcement days; the coefficient β is equal to the difference between the mean excess return on all announcement days and that earned on all non-announcement days.
The results for the full sample are reported in table 3.1 and indicate that stock market returns on trading days where macroeconomic announcements are scheduled for release are not significantly different from non-announcement days. This finding is in contrast to empirical evidence reported in Savor and Wilson (2013) for US data where announcement days are fundamentally riskier than other days and associated with higher excess stock returns.

Table 3.1: Regression analysis for daily stock market excess returns

Sample period	Obs	T	α	β	$Adj.R^2$
Jan-00 to Jan-15	511	3869	-29.21	2.21	0.00
(full sample)			$(-11.98)^{***}$	(0.33)	

The table reports results from a time series regression of daily excess stock market returns on a constant and an announcement-day dummy variable. $Announcement$ (β) is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. Obs indicate the number of scheduled macroeconomic announcements and T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

3.3.2 Structural change

Based on the previous analysis, it seems that scheduled macroeconomic announcements do not affect risk premia in the eurozone. However, the sample period under investigation shows a substantial variation in economic conditions, characterized by three recessions (2.5 business cycles) and a sovereign debt crisis which forced the European Central Bank (ECB) to implement several unconventional policy measures to support liquidity and prevent a credit crunch in the euro area[4].

This change in the macroeconomic environment raises the question if the full sample findings can be explained by structural instability in the stock market response to scheduled macroeconomic announcements due to the impact of the global financial crises and the European sovereign debt crises. To formally verify this hypothesis, I test the regression specified in equation (3.1) for a structural break in the relationship between announcement days and excess stock returns. For that purpose, I apply the Quandt-Andrews (Andrews, 1993) and the Bai-Perron (Bai and Perron, 1998, 2003) test, which both test for structural breaks with an unknown break point. However, while the Quandt-Andrews method only tests for a single structural break, the approach of Bai and Perron is constructed to test the time series for multiple breaks.

Table 3.2 shows the results using the test methodology suggested by Bai and Perron (Bai and Perron, 1998, 2003). The testing strategy is to verify, by implementation of a $supF$ test, if at least a structural break exists. The number of breaks can then be determined based upon a sequential examination of the $\sup F(l+1|\, l)$ statistics constructed using the break dates estimates obtained from a global minimization of the sum of squared residuals. Bai and Perron (2003) conclude that this approach leads to the best results and is recommended for empirical applications. I follow the standard procedure in the Bai-Perron

[4]During the crisis, the ECB mainly implemented three programs consisting of non-standard monetary policy measures. First, the *Enhanced Credit Support* (ECS) with the aim of restoring liquidity in the banking sector. Second, the *Securities Market Program* (SMP) which involved purchases of euro area government bonds to ensure liquidity in secondary markets. Third, the *Outright Monetary Transactions* (OMT) program which consists of buying government bonds and lowering bond yields with the aim of reducing borrowing costs. Falagiarda and Reitz (2015) provide a more detailed overview of these different programs.

Table 3.2: Bai and Perron endogenous break point test

$supF$ test		$\sup F(l+1\|\,l)$		Break date
$supF(1)$	76.26^{***}	$l=0$	76.26^{***}	March 10, 2009
$supF(2)$	40.97^{***}	$l=1$	9.67	
$supF(3)$	30.76^{***}			
$supF(4)$	32.40^{***}			
$supF(5)$	25.03^{***}			
UD_{max}	76.26^{***}			
WD_{max}	76.26^{***}			

This table reports the results of a structural break point test using the test procedure suggested by Bai and Perron (Bai and Perron, 1998, 2003). The test considers the null hypothesis of no structural change versus the alternative of unknown structural breaks in the regression model specified in equation (3.1). The null hypothesis of no structural break is rejected at the 1% significance level. For conducting the test, the maximum number of breaks is set to be 5 and the trimmage percentage is chosen to be 15% of the sample size. The test statistics employ heteroskedasticity and autocorrelation consistent covariances (Newey and West, 1987). *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

method and assume at most five breaks in the testing regime, i.e. $supF(m)$, $m = 1, 2, ..., 5$. The results in table 3.2 show that all $supF$ statistics, the UD_{max} and the WD_{max} statistic are highly significant at the 1% level confirming the existence of at least one significant structural break in the time series. To define the number of structural breaks, I implement a sequential test. The results indicate that $\sup F(1|\,0) = 76.26$ is highly significant while $\sup F(2|\,1) = 9.67$ rejects the existence of a second structural break. Based on these tests, I can conclude that there is only one significant structural break in the period January 2000 to January 2015. The date of the break is located to March 10, 2009.

I use the Quandt-Andrews test to corroborate the results of the Bai and Perron structural break point test. The results of the test and the associated break date are reported in table 3.3.

The Quandt-Andrews test considers the null hypothesis of no structural change versus the alternative of one unknown structural break in the regression model specified in equation (3.1). The test identifies a significant structural break at March 10, 2009 and confirms the result obtained by the Bai-Perron test. The question is now how to explain this finding? One potential explanation is linked to the monetary policy regime. The structural break occurs

Table 3.3: Quandt-Andrews endogenous break point test

	Test statistic	p-value	Break date
QA Max LR F-statistic	29.71	0.00	March 10, 2009

This table reports the Quandt-Andrews (Andrews, 1993) Maximum Likelihood Ratio F-statistic along with the estimated structural break date. The test considers the null hypothesis of no structural change versus the alternative of one unknown structural break in the regression model specified in equation (3.1). The null hypothesis of no structural break is rejected at the 1% significance level. The p-value is calculated using Hansen's method (Hansen, 1997). A trimmage percentage of 15% has been used to conduct the tests.

around the time when interest rates approach the vicinity of the zero lower bound in the eurozone. This can also be seen graphically from figure 3.1 which shows the time series of the euro overnight rate (EONIA)[5].

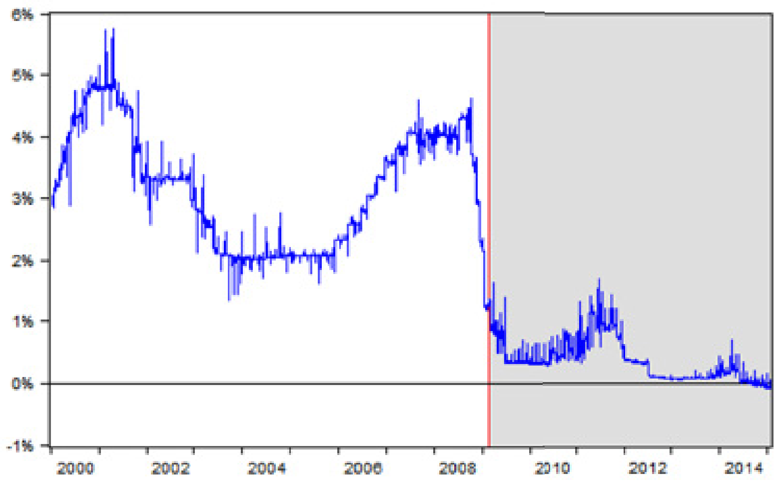

Figure 3.1: Euro overnight rate. The vertical line marks the structural break identified by the Quandt-Andrews and Bai-Perron structural break point tests. The gray shaded area indicates a regime where interest rates approach the vicinity of the zero lower bound.

The graph illustrates a sharp reduction in interest rates implemented by the ECB to stimulate the economy in the aftermath of the collapse of Lehman

[5]EONIA is the market interest rate that is most closely linked to the policy rate and constitutes the first step in the monetary policy transmission mechanism.

Brothers on September 15, 2008. Moreover, the vertical line marks the structural break identified by the Quandt-Andrews (Andrews, 1993) and Bai-Perron (Bai and Perron, 1998, 2003) structural break point tests and separates the time series into two monetary policy regimes. This analysis suggests that the period with low interest rates in the vicinity of the zero lower bound is different from the previous one.

Table 3.4: Regression analysis for daily stock market excess for two different monetary policy regimes

Sample period	Obs	T	α	β	$Adj.R^2$
Jan-00-03 to Mar-09-09	302	2354	-42.57	-7.70	0.00
(high interest rate)			$(-13.95)^{***}$	(-0.85)	
Mar-09-11 to Jan-15-30	209	1514	-9.10	19.66	0.17
(low interest rate)			$(-2.76)^{***}$	$(1.94)^{**}$	

The table reports results from a time series regression of daily excess stock market returns on a constant and an announcement-day dummy variable. *Announcement* (β) is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. Obs indicate the number of scheduled macroeconomic announcements and T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

To analyze if risk premia on announcement days depend on the monetary policy regime, I run the regression in equation (3.1) once for the period before the break (January 3, 2000 to March 9, 2009) and once for the period after the break (March 11, 2009 to January, 30 2015). The estimation results are reported in table 3.4 and show that the risk premia are significantly higher on announcement days during the low interest rate regime. More specifically, in such an environment the risk premium is 19.66 bp higher on days where macroeconomic news is scheduled for release. Furthermore, the difference in β coefficients for the two sample periods is about 27 bp with a t-statistic of 2.04. Intuitively, announcement days are trading days when investors learn about the state of the economy. These days are thus considered to be riskier in a low interest regime, which makes investors require a higher premium to compensate for a higher exposure to state variable risk.

3.3.3 US macroeconomic news announcements

Previous research (e.g. King, Sentana, and Wadhwani, 1994; Funke and Matsuda, 2006) finds some evidence that the economic and financial development in European economies is closely linked to those in the United States. This raises the question if prescheduled US macroeconomic news announcements have a significant impact on risk premia in Europe. In order to answer this question, I modify the regression in equation (3.1) by adding an indicator variable *USannouncement* that equals 1 for trading days with a prescheduled macroeconomic news announcement (employment, inflation, FOMC interest rates) for the US and zero otherwise

$$r_t^e = \alpha + \beta_1 Announcement_t + \beta_2 USannouncement_t + \varepsilon_t. \tag{3.2}$$

Dates for prescheduled macroeconomic news announcements for the US are obtained from the Bureau of Labor Statistics and the Federal Reserve. As in Savor and Wilson (2013), announcement days are trading days when the producer price index (PPI), employment figures or Federal Open Market Committee (FOMC) decisions are scheduled for release. The sample consists of 176 PPI announcements, 178 employment announcements and 132 FOMC interest rate announcements.

Table 3.5 shows that controlling for US announcement days has no impact on the results obtained in the baseline estimation in tables 3.1 and 3.4. More specifically, the coefficient β_2 for US macroeconomic news announcements is highly insignificant which confirms the result that risk premia in the eurozone are only significantly different from non-announcement days in a period of low interest rates.

Previous research has shown that central bank policies and decisions have a large impact on financial markets. Among others, Bernanke and Kuttner (2005) for US data and Wang and Zhu (2013) for international data show that stock markets react strongly to Federal Open Market Committee (FOMC) interest rate decisions. Moreover, Savor and Wilson (2013) and Lucca and Moench (2015) provide empirical evidence that on days where the FOMC announces its policy, US stock returns have on average been more than thirty times larger than on non-announcement days. Based on this evidence, I analyze whether FOMC announcement days impact risk premia in the eurozone. In order to assess the average excess returns earned on FOMC announcement

Table 3.5: Regression analysis for daily stock market excess for two different monetary policy regimes controlling for US announcements

Sample period	Obs	T	α	β_1	β_2	$Adj.R^2$
Jan-00 to Jan-15	511	3869	-29.61	2.23	3.12	0.00
(full sample)			$(-11.81)^{***}$	(0.33)	(0.43)	
Jan-00-03 to Mar-09-09	302	2354	-43.85	-7.31	10.00	0.00
(high interest rate)			$(-14.16)^{***}$	(-0.84)	(1.06)	
Mar-09-11 to Jan-15-30	209	1514	-8.27	19.67	-6.68	0.13
(low interest rate)			$(-2.46)^{**}$	$(1.95)^{**}$	(-0.62)	

The table reports results from a time series regression of daily excess stock market returns on a constant and two announcement-day dummy variables. $Announcement$ (β_1) is a dummy variable equaling 1 if day t is an announcement day in the eurozone, and 0 otherwise. $USannouncement$ (β_2) is a dummy variable equaling 1 if day t is an announcement day in the USA, and 0 otherwise. Obs indicate the number of scheduled macroeconomic announcements and T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

days, I proceed by running another simple dummy variable regression model

$$r_t^e = \alpha + \beta_1 Announcement_t + \beta_2 FOMC_t + \varepsilon_t \tag{3.3}$$

where $FOMC$ is a dummy variable, which is equal to one on scheduled FOMC announcement dates and zero on all other days and all other variables are the same as in the baseline regression in equation (3.1).

The results reported in table 3.6 show for the full sample and the high interest rate period that the average excess return is significantly higher on FOMC announcement days. For the full sample period, the coefficient β_2 indicates that on FOMC announcement days the excess return is 25.8 bp higher while for the high interest rate period, the FOMC announcement effect is even stronger with a coefficient β_2 of 42.9 bp. However, European scheduled macroeconomic announcements do not affect risk premia in these sample periods. For both samples, the coefficient β_1 is only slightly positive (1.98 bp) or even negative (-8.12 bp) and no longer statistically significant. Analyzing the low interest rate regime shows that the average excess return on FOMC announcement days is virtually zero and no longer significant. Importantly, the coefficient β_1 is now significant at the 5% level. This means that average ex-

Table 3.6: Regression analysis for daily stock market excess for two different monetary policy regimes controlling for FOMC announcements

Sample period	Obs	T	α	β_1	β_2	$Adj.R^2$
Jan-00 to Jan-15	511	3869	-30.06	1.98	25.79	0.04
(full sample)			$(-12.20)^{***}$	(0.29)	$(1.82)^{*}$	
Jan-00-03 to Mar-09-09	302	2354	-43.97	-8.12	42.88	0.18
(high interest rate)			$(-14.28)^{***}$	(-0.93)	$(2.15)^{**}$	
Mar-09-11 to Jan-15-30	209	1514	-9.13	19.66	0.80	0.10
(low interest rate)			$(-2.79)^{***}$	$(1.95)^{**}$	(-0.04)	

The table reports results from a time series regression of daily excess stock market returns on a constant and two announcement-day dummy variables. $Announcement$ (β_1) is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. $FOMC$ (β_2) is a dummy variable equaling 1 if day t is a day where a FOMC announcement is scheduled for release, and 0 otherwise. Obs indicate the number of scheduled macroeconomic announcements and T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

cess returns in the low interest rate regime are higher on days when European macroeconomic news is scheduled for release.

Based on the previous analysis, it can be concluded that investors demand a risk premium on the European stock market to bear risks associated with FOMC decisions and that this behavior depends on the state of the economy and the associated interest rate regime. To further investigate why average risk premia are higher on days related to European macroeconomic announcements in the low interest rate period, I distinguish between scheduled monetary policy announcements (ECB announcements) and macroeconomic announcements not related to monetary policy (non-policy announcements, i.e. CPI and unemployment news). The regression I estimate is

$$r_t^e = \alpha + \beta_1 ECB_t + \beta_2 NonECB_t + \beta_3 FOMC_t + \varepsilon_t \tag{3.4}$$

where ECB is a dummy variable that takes a value of 1 on a trading day where interest rate decisions by the ECB are scheduled for announcement and zero otherwise. The dummy variable $NonECB$ denotes all other announcement days not related to monetary policy decisions in the eurozone and takes a value of one on trading days where CPI and unemployment announcements are scheduled for release and zero otherwise.

Table 3.7: Regression analysis for daily stock market excess for two different monetary policy regimes controlling for FOMC announcements and distinction between monetary policy related announcements and other scheduled macroeconomic announcements.

Sample period	Obs	T	α	β_1	β_2	β_3	$Adj.R^2$
Jan-00 to Jan-15	511	3869	-30.22	-8.03	9.63	25.02	0.06
(full sample)			$(-12.25)^{***}$	(-0.66)	(1.22)	$(1.76)^{*}$	
Jan-00-03 to Mar-09-09	302	2354	-44.52	-15.56	3.14	41.80	0.16
(high interest rate)			$(-14.54)^{***}$	(-1.04)	(0.30)	$(2.10)^{**}$	
Mar-09-11 to Jan-15-30	209	1514	-8.83	12.81	19.82	0.49	0.00
(low interest rate)			$(-2.69)^{***}$	(0.63)	$(1.64)^{*}$	(0.03)	

The table reports results from a time series regression of daily excess stock market returns on a constant and three announcement-day dummy variables. ECB (β_1) is a dummy variable equaling 1 if day t is a day where a ECB announcement is scheduled for release, and 0 otherwise. $NonECB$ (β_2) is a dummy variable equaling 1 if on announcement day t CPI or unemployment news is scheduled for release and 0 otherwise. $FOMC$ (β_3) is a dummy variable equaling 1 if day t is a day where a FOMC announcement is scheduled for release, and 0 otherwise. Obs indicate the number of scheduled macroeconomic announcements and T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

The results in table 3.7 mainly confirm the results obtained in table 3.6. For the full sample and the high interest rate sample, average excess returns are higher on FOMC announcement days. However, there is no such effect on ECB announcement days. For the full sample and the high interest rate, sample stock returns on the ECB announcement day are even lower. In the low interest rate regime, the risk premium is positive and economically significant (12.81 bp) but the associated t-statistic is far from significant levels. Based on these results, it seems that market participants in Europe only demand a higher risk premium for risks associated with FOMC decisions but not for monetary policy decision by the ECB.

Moreover, in the low interest rate regime, macroeconomic news announcements not related to monetary policy command a positive but only borderline significant risk premium while the FOMC announcement premium has vanished. One obvious interpretation of this finding is that when interest rates in the eurozone and the US are close to the zero lower bound, announcements related to fundamental macroeconomic data such as unemployment and inflation are of greater importance for investors in financial markets.

To summarize, the results in this section show that FOMC announcements are a main driver for equity risk premia in the eurozone and holding stocks on such announcement days must be perceived to be risky. However, this effect cannot be observed in a macroeconomic environment with low interest rates. Announcement days related to CPI and unemployment news do, on average, earn higher risk premia compared to other trading days. In this context, a puzzling finding is that ECB announcements are not associated with higher excess stock returns. It seems that Federal Reserve pronouncements have a substantial impact on European investors' risk aversion and are an important driver for equity risk premia in Europe.

3.3.4 Macroeconomic news and the business cycle

Several papers examine whether the stock market's reaction to macroeconomic news depends on the state of the economy. Among others, Basistha and Kurov (2008) and Jansen and Tsai (2010) document significant cyclical variation in the impact of monetary policy on stock prices with a stronger effect when monetary expansions coincide with a contracting economy, bear stock markets and tight credit market conditions[6]. Boyd, Hu, and Jagannathan (2005) show that the stock market reacts differently to unemployment news depending on whether the economy is in a contraction or expansion. In this section, I extend this work by verifying if average stock market returns on announcement days exhibit a similar state dependence and vary over the business cycle. For that purpose, I use business cycle turning points data for the Euro area which are available on the OECD website to determine contraction and expansion periods. I follow the methodology used by the National Bureau of Economic Research in the US which defines a contraction phase in the business cycle from peak to trough and an expansion phase from trough to the next peak. A cycle is defined from peak (trough) to the next peak (trough). From this definition, it follows that the sample period consists of 2.5 business cycles. To test the hypothesis that risk premia on days where macroeconomic news is announced are state dependent, I interact all dependent variables in the regression model specified in equation (3.3) with a business cycle dummy which indicates the state of the economy[7] The dummy variable $D_t^{\text{contr.}}$ equals 1 if on day t the economy is in a contraction phase and 0 otherwise.

$$r_t^e = \alpha + (\beta_1 Announcement_t + \beta_2 FOMC_t)\, D_t^{contr.} + \varepsilon_t \tag{3.5}$$

Analogously, the dummy variable $D_t^{\text{expan.}}$ equals 1 if on day t, the economy is in an expansion phase and 0 otherwise.

$$r_t^e = \alpha + (\beta_1 Announcement_t + \beta_2 FOMC_t)\, D_t^{\exp an.} + \varepsilon_t \tag{3.6}$$

I estimate the regression model in equations (3.5) and (3.6) jointly using the Generalized Method of Moments (GMM) estimator. Standard errors and the associated t-statistics are estimated using a heteroskedasticity and autocorrelation consistent (HAC) covariance matrix (Newey and West, 1987) with 5 lags. Table 3.8 shows the results for these regressions. In terms of statistical significance, only FOMC announcements days in a recession offer higher equity risk premia with a β_2 coefficient of 49.41 bp and an associated t-statistic of

[6]In these studies, the news is measured as monetary policy surprises and is derived from federal funds futures data as in Bernanke and Kuttner (2005).

[7]I use the model specification in equation (3.3) since the previous analysis showed that stock returns on FOMC announcement days are significantly higher as compared to other trading days.

1.89. This stronger effect during economic downturns can be understood as an increase in uncertainty about the state of the economy and investors in Europe demand an even higher risk premium for risks associated with FOMC decisions in recessions.

Table 3.8: Regression analysis for daily stock market excess returns controlling for FOMC announcements under state dependence

	Contraction	Expansion
α	-28.91	-29.98
	$[-11.96]^{***}$	$[-11.66]^{***}$
β_1	-12.37	9.61
	[-0.87]	[1.34]
β_2	49.41	10.13
	$[1.89]^{*}$	[0.67]
N	3869	3869
Adj. $R^2(\%)$	0.15	0.04

This table reports results from a time series regression of daily excess stock market returns on a constant and two announcement-day dummy variables. *Announcement* (β_1) is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. *FOMC* (β_2) is a dummy variable equaling 1 if day t is a day where a FOMC announcement is scheduled for release, and 0 otherwise. Column 1 corresponds to the contraction and column 2 to the expansion phase. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in square brackets. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

Moreover, I use a Wald test of coefficient equality to test the joint hypothesis $\beta_1^{contr.} = \beta_1^{\exp an.}$ and $\beta_2^{contr.} = \beta_2^{\exp an.}$. The test gives a χ^2-statistic of 3.39, corresponding to a p-value of 0.1834 which indicates that the coefficient estimates are not different in a contracting and expanding economy from a statistical point of view. However, the difference in coefficients matters economically, since in a recessionary economy, the average excess return on FOMC announcement days is approximately 39 bp higher as compared to FOMC announcement days in an expanding economy.

3.3.5 Heterogeneity among European economies

So far, this paper has focused on one common stock market for the eurozone by using EURO STOXX data. However, the recent financial crises and the onset of a sovereign debt crisis in some euro area countries caused a heterogeneous development in fiscal positions and macroeconomic fundamentals within the eurozone. In this context, financial markets evaluate the country risk of these economies differently which can be observed by looking at the pattern of sovereign credit default swaps (CDS) spreads. Countries such as Greece, Ireland, Italy, Portugal and Spain were particularly badly hit by the economic downturn in the aftermath of the financial crises while economies like Germany were mainly left unscathed.

In this section, I extend the empirical analysis and use equity data for different euro area countries in order to verify the hypothesis that the stock markets of these economies react differently to scheduled macroeconomic news announcements. For that purpose, I estimate the regression model in equation (3.3) where I use excess stock returns for France (CAC 40), Germany (DAX), Greece (Athex 20), Ireland (ISEQ Overall), Italy (FTSE MIB), Portugal (PSI 20) and Spain (IBEX 35) as a dependent variable. As before, I infer the daily risk free rate from the 1 month EURIBOR. Data are obtained from Datastream and Bloomberg.

The results in table 3.9 provide evidence that there is substantial heterogeneity in the stock market's reaction to scheduled macroeconomic announcements. Compared to the aggregate European stock market, the average excess returns on FOMC announcements are only significantly higher for Germany. As shown in the second row of table 3.9, the β_2 coefficient on the FOMC announcement day dummy is 27.22. This indicates that the excess return of the DAX on FOMC announcement days on average exceeded the excess return on non-announcement days by 27.22 bp.

For Greece and Ireland, the coefficient β_1 for scheduled macroeconomic announcements related to the eurozone is positive and highly significant meaning that risk premia on announcement days are higher as compared to trading days with no announcements. For all other countries in the sample, no such effect can be observed. Next I estimate the regression model in equation (3.4) which makes a distinction between scheduled monetary policy announcements (ECB announcements) and macroeconomic announcements not related to monetary policy (non ECB announcements, i.e. CPI and unemployment news).

Similar to the results in table 3.9, the average excess return on FOMC an-

Table 3.9: Regression analysis for daily stock market excess for 7 European economies controlling for FOMC announcements.

	T	α	β_1	β_2	$Adj.R^2$
France	3869	-29.17	-0.31	21.66	0.01
		$(-12.14)^{***}$	(-0.05)	(1.43)	
Germany	3869	-27.37	-0.37	27.22	0.05
		$(-10.17)^{***}$	(-0.06)	$(1.98)^{**}$	
Greece	3869	-35.94	17.63	4.19	0.06
		$(-10.38)^{***}$	$(2.31)^{**}$	(0.21)	
Ireland	3868	-29.41	17.56	-10.49	0.15
		$(-11.39)^{***}$	$(2.85)^{***}$	(-0.49)	
Italy	3869	-30.20	-1.44	19.55	0.00
		$(-11.33)^{***}$	(-0.21)	(1.29)	
Portugal	3869	-29.90	1.77	-5.76	0.00
		$(-12.59)^{***}$	(0.34)	(-0.45)	
Spain	3869	-29.41	4.29	20.51	0.00
		$(-11.53)^{***}$	(0.61)	(1.33)	

The table reports results from a time series regression of daily excess stock market returns on a constant and two announcement-day dummy variables for 7 eurozone economies. $Announcement$ (β_1) is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. $FOMC$ (β_2) is a dummy variable equaling 1 if day t is a day where a FOMC announcement is scheduled for release, and 0 otherwise. T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

nouncement days is only significantly higher for Germany. However, the non-existence of a ECB announcement premium is more remarkable. For all regressions in table 3.10, none of the β_1 coefficients are statistically significant. This corroborates the finding for the aggregate European stock market that ECB announcements are not associated with higher excess stock returns.

Table 3.10: Regression analysis for daily stock market excess controlling for FOMC announcements and distinction between monetary policy related announcements and other scheduled macroeconomic announcements for 7 European economies.

	T	α	β_1	β_2	β_3	$Adj.R^2$
France	3869	-29.31	-11.83	8.11	20.79	0.05
		$(-12.21)^{***}$	(-1.00)	(1.04)	(1.37)	
Germany	3869	-27.48	-14.10	8.96	26.23	0.09
		$(-10.22)^{***}$	(-1.16)	(1.13)	$(1.90)*$	
Greece	3869	-35.97	16.90	17.90	4.20	0.04
		$(-10.36)^{***}$	(1.45)	$(2.02)^{**}$	(0.22)	
Ireland	3868	-29.47	3.75	26.08	-11.40	0.22
		$(-11.37)^{***}$	(0.37)	$(3.37)^{***}$	(-0.54)	
Italy	3869	-30.36	-14.35	7.98	18.58	0.04
		$(-11.41)^{***}$	(-1.18)	(1.03)	(1.22)	
Portugal	3869	-30.06	0.26	4.36	-5.96	0.00
		$(-12.58)^{***}$	(0.03)	(0.74)	(-0.46)	
Portugal	3869	-29.58	-6.39	12.27	19.70	0.05
		$(-11.56)^{***}$	(-0.52)	(1.57)	(1.28)	

The table reports results from a time series regression of daily excess stock market returns on a constant and three announcement-day dummy variables for 7 eurozone economies. ECB (β_1) is a dummy variable equaling 1 if day t is a day where a ECB announcement is scheduled for release, and 0 otherwise. $NonECB$ (β_2) is a dummy variable equaling 1 if on announcement day t CPI or unemployment news is scheduled for release and 0 otherwise. $FOMC$ (β_3) is a dummy variable equaling 1 if day t is a day where a FOMC announcement is scheduled for release, and 0 otherwise. T denotes the number of trading days in the sample. t-statistics are calculated using Newey and West (1987) robust standard errors (with 5 lags) and are given in parenthesis. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

3.4 Conclusion

In this paper, I study the excess return behavior in the euro area around days when important macroeconomic news about inflation, unemployment or interest rates is scheduled for announcement. Average excess returns on announcement days are significantly higher when interest rates are in the vicinity of the zero lower bound. Moreover, I document that FOMC announcement days have a significant impact on risk premia in the eurozone while ECB announcement days do not affect excess stock returns. However, announcement days related to CPI and unemployment news do, on average, earn higher risk premia compared to non-announcement days in a low interest rate regime. The finding that average excess returns on scheduled ECB announcement days are not higher for the aggregate European stock market is in strong contrast to the empirical evidence for US data where average excess stock returns are higher on FOMC announcement days. Since the ECB is an independent central bank that pursues an active monetary policy, it is puzzling that there is no ECB announcement day premium and this requires further research.

Chapter 4

The impact of ECB monetary policy surprises on the German stock market[1]

4.1 Introduction

In recent years, monetary policy has become the main tool for stabilizing the economy. A change in monetary policy does not only have important effects on the real economy, it is also transmitted to financial markets. Therefore, policy makers in central banks as well as financial market investors have a great interest in understanding this transmission mechanism. In the last few years, a number of studies have examined the impact of monetary policy shocks, i.e. unanticipated changes in monetary policy, on asset prices such as interest rates and stock returns. For the USA, the effect of monetary policy surprises on various other interest rates is studied by, among others, Kuttner (2001) and Cochrane and Piazzesi (2002), while Rigobon and Sack (2004), Ehrmann and Fratzscher (2004), Bernanke and Kuttner (2005) and more recently Kontonikas, MacDonald, and Saggu (2013) examine how surprise changes in monetary policy affect the stock market. Moreover, Wang and Zhu (2013) analyze the reac-

[1]This is joint work with Markus Sigonius. We thank David Domeji and Roine Vestman for their advice and to Rickard Sandberg, Irina Zviadadze and two anonymous referees for helpful comments and suggestions. Thanks to Reinder Haitsma for poviding us with his data.

tion of international stock markets to unanticipated changes in the US policy rate. The main finding of these studies is that asset markets react significantly to monetary policy shocks. In terms of the stock market an unexpected increase (decrease) in the policy rate is associated with a decrease (increase) in stock prices. However, much less is known about what are the sources underlying these market reactions. For the stock market, the linearized present value identity of stock returns of Campbell and Shiller (1988a) implies that a decline in stock prices may be associated with a decrease in expected future dividends, a rise in the future expected real interest rates used to discount those dividends, or an increase in the expected excess returns (equity risk premia) associated with holding stocks. Campbell (1991) and Campbell and Ammer (1993) developed an empirical approach to decompose surprise changes in excess returns into revisions in future dividends, real interest rates or future excess returns. Bernanke and Kuttner (2005) use this methodology to examine the impact of US monetary policy surprises on current US stock market returns and its components.

In this paper, we study the impact of ECB monetary policy changes on the German stock market. Germany is the largest economy in the eurozone and therefore of particular interest for such an analysis. So far, several previous studies provide evidence that unexpected ECB interest rate cuts (hikes) increase (decrease) German stock prices (e.g. Angeloni and Ehrmann, 2003; Bohl, Siklos, and Sondermann, 2008; Hussain, 2011; Hayo and Niehof, 2011). However, none of these studies provide further insights into the underlying sources of the observed stock market movements.

In the first part of our empirical analysis, we use an extended sample and examine the impact of ECB monetary policy shocks on German aggregate stock returns using a standard event study methodology as in Bernanke and Kuttner (2005). A very recent study of Haitsma, Unalmis, and de Haan (2016) provides empirical evidence that the impact of the ECB monetary policy is different in crisis and non-crisis years. Their analysis indicates that unconventional monetary policy, measured by changes in the yield spread between German and Italian 10-year government bonds at the day of the policy announcements, affect the EURO STOXX 50 index significantly while conventional monetary policy shocks extracted from future prices (e.g. Kuttner, 2001; Bernanke and Kuttner, 2005) have no impact on the stock market during the crisis.

In the second part of this study, we shed some light on the question of what drives the stock market movements. More specifically, we adopt the approach of Campbell and Ammer (1993) and Bernanke and Kuttner (2005) to investigate the reasons behind the response of ECB monetary policy shocks to current German excess stock returns. Moreover, we study the effects of the recent financial and sovereign debt crisis by first analyzing a pre- and post crisis sample. Finally, we divide our sample into a period with low (e.g. negative) and high (e.g. positive) real interest rates to gain some insights into whether the impact of ECB monetary policy shocks on stock return dynamics is dependent on the prevailing interest rate regime. The novel finding of our analysis is that a strong and significant stock market reaction is only observed when the real interest rates are negative.

The paper is organized as follows. Section 2 briefly discusses the identification of monetary policy. Section 3 describes the event study methodology, the econometric framework associated with the variance decomposition and how we measure the impact of monetary policy surprises on news regarding current excess stock returns and their respective components. Section 4 describes the data and Section 5 provides the empirical results. Section 6 concludes the paper.

4.2 Identification of monetary policy

The study of the impact of monetary policy on stock prices raises the question of how to measure policy changes. Theories based on the efficient markets hypothesis (Fama, 1970) suggest that only unanticipated changes in monetary policy, often referred to as monetary policy shocks, should have an impact on stock prices. The anticipated component is already in the investors' information set and thus priced into the value of the stock prior to the policy announcement. Based on this argument, changes in monetary policy must be decomposed into expected and unexpected components. Failure of such a decomposition may lead to biased results due to an error in the variables problem. The most common method used in the literature to obtain the surprise element of a conventional monetary policy change is based on futures market data (e.g. Kuttner, 2001; Bernanke and Kuttner, 2005). Gürkaynak, Sack, and Swanson (2007) find that future markets measures are most accurate in terms of capturing the market expectation of monetary policy. In line with this evidence, we adopt the same approach and derive measures of monetary policy shocks based on changes in appropriate futures contracts. For US data Bernanke and Kuttner (2005) use federal fund futures to gauge the market expectations regarding the level of the policy rate. Since there exists no comparable future instrument that tracks the Euro area policy rate, we use interest rate futures contracts that are close substitutes since these contracts are likely to be strongly influenced by the market expectations of future policy rates. More specifically, Bernoth and von Hagen (2004) find that the 3-month Euribor futures rate is an unbiased and reliable predictor of Euro area policy rate changes. As in Bredin, Hyde, and O'Reilly (2010), we thus proxy surprise changes in the ECB policy rate by changes in the 3-month Euribor futures rate. To measure unexpected unconventional monetary policies, we follow Rogers, Scotti, and Wright (2014) and Haitsma, Unalmis, and de Haan (2016) and proxy the surprise component by the change in the spread between 10-year German and Italian government bond yields. The rationale for this choice is that a large part of the ECB's unconventional monetary policies was aimed at decreasing sovereign spreads in the eurozone. An increase in the spread implies a tighter monetary policy than expected.

4.3 Econometric methodology

4.3.1 Event study

In the first part of our empirical analysis, we investigate the impact of ECB monetary policy shocks on German aggregate stock returns. We use the specification of Haitsma, Unalmis, and de Haan (2016) and estimate the following regression model

$$\begin{aligned} r_t &= \alpha + \beta_1 (1 - C_t) \Delta r_t^u + \gamma_1 (1 - C_t) \Delta r_t^e + \beta_2 C_t \Delta r_t^u + \\ &\quad \gamma_2 C_t \Delta r_t^e + \varphi \Delta r_t^{u,c} + \delta \mathbf{X}_t + \varepsilon_t \end{aligned} \tag{4.1}$$

where r_t denotes the stock return on day t, α is a constant, C_t is a crisis dummy that takes a value of 1 during the crisis and a value of zero before. The start of the crisis period is defined as August 22 2007 where the ECB announced the first unconventional monetary policy.[2] Δr_t^u, Δr_t^e and $\Delta r_t^{u,c}$ denote the conventional monetary policy surprise, the expected policy rate change and the

[2]During the crisis, the ECB mainly implemented three programs consisting of non-standard monetary policy measures. First the *Enhanced Credit Support* (ECS) with the aim of restoring liquidity in the banking sector. Second, the *Securities Market Program* (SMP) which

unconventional monetary policy surprise on day t. $\mathbf{X}_t$ is a vector of controls that consists of a global stock market index excluding Europe to control for general economic developments in the rest of the world and the crisis dummy.

4.3.2 Variance decomposition

The empirical analysis for the second part of this paper is based on the log-linear representation of the rational valuation formula developed in Campbell and Shiller (1988a):

$$r_{t+1} = k + \rho p_{t+1} + (1-\rho)\, d_{t+1} - p_t \tag{4.2}$$

where r_{t+1} is the log return in period $t+1$, p_t is the log price at the end of period t, p_{t+1} is the log price at the end of period $t+1$ and d_{t+1} denotes the log dividend during period $t+1$. k is a constant and $\rho \equiv 1/\left(1+\exp\left(d-p\right)\right)$ is a weight term, both associated with the linearization. The weight term is a function of the long-run mean of the log dividend-price ratio, $d-p$, and slightly below unity. Equation (4.2) relates current stock prices to future stock prices, dividends and returns. Imposing the condition to rule out rational bubbles, $\lim_{j\to\infty} E_t \rho^j p_{t+j} = 0$, and taking expectations, equation (4.2) can be solved forward to give

$$p_t = \frac{k}{1-\rho} + E_t\left[\sum_{j=0}^{\infty} \rho^j \left((1-\rho)\, d_{t+j+1} - r_{t+j+1}\right)\right] \tag{4.3}$$

By substituting equation (4.3) into (4.2), Campbell (1991) shows that unexpected changes of current stock returns either reflect news (revision of expectations) of dividend growth (Δd_{t+1}) or future returns (r_{t+1}).

$$r_{t+1} - E_t r_{t+1} = \left(E_{t+1} - E_t\right)\sum_{j=0}^{\infty} \rho^j \Delta d_{t+j+1} - \left(E_{t+1} - E_t\right)\sum_{j=0}^{\infty} \rho^j r_{t+j+1} \tag{4.4}$$

Campbell and Ammer (1993) suggest an excess return version of the Campbell (1991) present value relation stated in equation (4.4) such that unexpected ex-

involved purchases of euro area government bonds to ensure liquidity in secondary markets. Third, the *Outright Monetary Transactions* (OMT) program which consists of buying government bonds and lowering bond yields with the aim of reducing the borrowing costs. Falagiarda and Reitz (2015) provide a more detailed overview of these different programs.

cess returns innovations are decomposed into revisions in expectations with respect to future excess returns (risk premia), future dividends and real interest rates. Define the excess stock return over a short-term interest rate as $rx_{t+1} \equiv r_{t+1} - rr_{t+1}$, where r_{t+1} is the expected return and rr_{t+1} is the real interest rate. The unexpected excess return can then be written as

$$rx_{t+1} - E_t rx_{t+1} = (E_{t+1} - E_t) \left\{ \begin{array}{l} \sum_{j=0}^{\infty} \rho^j \Delta d_{t+j+1} - \\ \sum_{j=0}^{\infty} \rho^j rr_{t+j+1} - \sum_{j=0}^{\infty} \rho^j rx_{t+j+1} \end{array} \right\} \tag{4.5}$$

In order to simplify the notation, equation (4.5) can be more compactly expressed

$$e^{rx}_{t+1} = \tilde{e}^{d}_{t+1} - \tilde{e}^{rr}_{t+1} - \tilde{e}^{rx}_{t+1} \tag{4.6}$$

where the superscript tilde ($\sim$) represents an innovation in a variable. For example

$$\tilde{e}^{d}_{t+1} = (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta d_{t+j+1} \tag{4.7}$$

Intuitively, equation (4.6) implies that a positive surprise movement in the current excess stock return (e^{rx}_{t+1}) is related to positive dividend news ($\tilde{e}^{d}_{t+1}$), lower than expected real interest rates ($\tilde{e}^{rr}_{t+1}$) or lower than expected future excess returns ($\tilde{e}^{rx}_{t+1}$) or an arbitrary combination of these components.

In order to determine what factors that drive stock market volatility (Campbell, 1991; Campbell and Ammer, 1993), one can decompose the variance of news regarding excess returns by taking the variance of both sides of equation (4.6)

$$\begin{aligned} \mathrm{var}(e^{rx}_{t+1}) \quad = \quad & \mathrm{var}(\tilde{e}^{d}_{t+1}) + \mathrm{var}(\tilde{e}^{rr}_{t+1}) + \mathrm{var}(\tilde{e}^{rx}_{t+1}) - \\ & 2\,\mathrm{cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rr}_{t+1}) - 2\,\mathrm{cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rx}_{t+1}) + 2\,\mathrm{cov}(\tilde{e}^{rr}_{t+1}, \tilde{e}^{rx}_{t+1}) \end{aligned} \tag{4.8}$$

This variance decomposition gives further insights into the relationship between risk premia news and news about fundamentals and allows us to study if the different news components can be considered as correlated or distinct shocks. To measure the relative magnitude of each of the news components, we normalize the variance and covariance terms by the variance of the unexpected excess return, $\mathrm{var}(e^{rx}_{t+1})$ such that the sum of the variances and covariances in equation (4.8) adds up to one.

Since the revisions in expectations of excess returns and their components are not directly observable, the empirical implementation of equation (4.2) requires a data generating process for future excess returns, future dividends and future real interest rates. We follow the approach outlined in Campbell (1991) and Campbell and Ammer (1993) and assume that a forecasting VAR model adequately captures the dynamics of excess stock returns. The VAR system is based on defining a vector of state variables $\mathbf{z_t}$ that helps measure or forecast excess returns and be able to compute each of the components in (4.6). For that purpose, $\mathbf{z_t}$ needs to include at least the excess stock return and the real interest rate. Following Engsted, Pedersen, and Tanggaard (2012), we also include the log dividend-price ratio as a state variable in our VAR approach in order to alleviate the concerns raised by Chen and Zhao (2009) about the choice of stock market news components obtained as residuals from the VAR. In addition to these variables, we also include variables that have been successful in predicting excess stock returns and real interest rates, the change in the nominal interest rate, the term spread, [3] or the relative bill rate defined as the level of the short-term interest rate relative to a 1-year backwards moving average of short rates. The first-order VAR can then be written as

$$\mathbf{z_{t+1}} = \mathbf{A}\mathbf{z_t} + \omega_{t+1} \tag{4.9}$$

where $\mathbf{A}$ is the matrix of VAR coefficients and ω_{t+1} is the vector of error terms. From this VAR model, the estimate of $\mathbf{z_{t+1}} - E_t\mathbf{z_{t+1}}$ is ω_{t+1}. Similarly, the VAR estimate to compute revisions in long-horizon expectations $(E_{t+1} - E_t)\,\mathbf{z}_{t+1+j}$ is $\mathbf{A}^j\omega_{t+1}$. Defining two appropriate selection vectors, $\mathbf{e1}$ and $\mathbf{e2}$, that pick out the first and second element of $\mathbf{z_t}$, VAR estimates of (4.6) can be obtained as follows:

The news of the current excess stock return can be computed directly from equation (4.9) as

[3]The term spread is defined as the spread between a long-term government bond yield and the short-term interest rate.

$$e_{t+1}^{rx} = \mathbf{e1}'\omega_{t+1} \tag{4.10}$$

because the excess stock return is the first element of $\mathbf{z_t}$. News of future expected stock market excess returns is obtained from the VAR estimates as

$$\tilde{e}_{t+1}^{rx} = \mathbf{e1}'\sum_{j=1}^{\infty}\rho^j\mathbf{A}^j\omega_{t+1} = \mathbf{e1}'\rho\mathbf{A}\left(\mathbf{I}-\rho\mathbf{A}\right)^{-1}\omega_{t+1} \tag{4.11}$$

while news of future real interest rates is computed from

$$\tilde{e}_{t+1}^{rr} = \mathbf{e2}'\sum_{j=0}^{\infty}\rho^j\mathbf{A}^j\omega_{t+1} = \mathbf{e2}'\left(\mathbf{I}-\rho\mathbf{A}\right)^{-1}\omega_{t+1} \tag{4.12}$$

Given the two components in equation (4.11) and (4.12), dividend news is computed from the identity (4.6) as the residual:

$$\tilde{e}_{t+1}^{d} = e_{t+1}^{rx} + \tilde{e}_{t+1}^{rr} + \tilde{e}_{t+1}^{rx} \tag{4.13}$$

4.3.3 Monetary policy surprise

According to the Campell-Shiller decomposition, an unanticipated cut in the policy rate may lead to an increase in stock prices, via higher expected dividends, a lower future real interest rate, a decrease in future excess returns (risk premia) or an arbitrary combination. Following Bernanke and Kuttner (2005), we include the surprise element in monetary policy as an exogenous variable in the forecasting VAR to analyze the impact of monetary policy shocks on revisions in expected excess returns, real interest rates and dividends. The extended VAR model is

$$\mathbf{z_{t+1}} = \mathbf{Az_t} + \phi\Delta i_{t+1}^{u} + \omega_{t+1}^{\perp} \tag{4.14}$$

where the $n \times 1$ coefficient vector ϕ captures the contemporaneous response of the endogenous variables in $\mathbf{z_{t+1}}$ to the unanticipated change in monetary policy Δi_{t+1}^{u}. The disturbance term $\omega_{t+1}^{\perp}$ is by construction orthogonal to the monetary policy shock. To obtain consistent estimates of both $\mathbf{A}$ and ϕ, we first estimate the parameters of the VAR in equation (4.9) and then regress the VAR's 1-step-ahead forecast errors on the monetary policy surprise. This two-step procedure allows us to estimate the VAR dynamics over a sample longer than the period in which our measure of the monetary policy shock is available which improves the estimates' precision of the underlying VAR parameters. Alternatively, we could simply have estimated the model in equation (4.14) where the monetary policy shock is included in the forecasting VAR.[4]

Next, we compute the impact of the monetary policy shock on each of the discounted sums of expected future excess returns, real interest rates and dividends. For that purpose, we substitute $\phi \Delta i^u_{t+1} + \omega^{\perp}_{t+1}$ into equation (4.10) to (4.12). Incorporating this surprise component as an exogenous variable then implies

$$\tilde{e}^{rx}_{t+1} = \mathbf{e1}'\rho\mathbf{A}\left(\mathbf{I} - \rho\mathbf{A}\right)^{-1}\left(\phi \Delta i^u_{t+1} + \omega^{\perp}_{t+1}\right). \tag{4.15}$$

Hence, the response of the present value of expected future excess returns to the monetary policy shock is just:

$$\mathbf{e1}'\rho\mathbf{A}\left(\mathbf{I} - \rho\mathbf{A}\right)^{-1}\phi \tag{4.16}$$

Similarly, the response of the present value of current and expected future real interest rates is:

$$\mathbf{e2}'\left(\mathbf{I} - \rho\mathbf{A}\right)^{-1}\phi \tag{4.17}$$

and the response of the present value of future and current dividends is:

$$\left(\mathbf{e1} + \mathbf{e2}\right)'\left(\mathbf{I} - \rho\mathbf{A}\right)^{-1}\phi. \tag{4.18}$$

4.3.4 Threshold VAR

The macroeconomic environment in recent years is characterized by a period of exceptionally low, i.e. negative, real interest rates and an unprecedented expansionary monetary policy. However, the recent period of negative real interest rates is not unique since the German economy experienced other phases of low real interest rates, for example in the late 1970s or early 1990s. In such an environment, the causes of stock market fluctuations might differ from a regime with positive real interest rate levels and investors might put more weight on news about macroeconomic risks. In order to assess these possible nonlinearities, we follow Nitschka (2014) and use a threshold VAR based on

[4] As a robustness test, we have checked for this alternative and the results of our analysis do not change qualitatively.

the distinction between negative and positive real interest rate levels. The exogenous threshold level is 0 and real interest rates below the threshold level constitute a low interest rate regime (l), and real interest rates above the threshold level constitute a high interest rate regime (h). The threshold VAR specification can then be written as

$$\mathbf{z_{t+1}} = \begin{cases} \mathbf{A}^h\mathbf{z_t} + \omega^h_{t+1} & \text{if } rr_{t+1} > 0 \\ \mathbf{A}^l\mathbf{z_t} + \omega^l_{t+1} & \text{if } rr_{t+1} < 0 \end{cases} \tag{4.19}$$

where rr denotes the monthly, ex post measured real interest rate, which serves as a proxy for future expected real interest rates in period t in line with equation (4.5). From this framework, we obtain VAR estimates for both interest rate regimes and perform the variance decomposition. Furthermore, we study the impact of monetary policy shocks on the corresponding stock return news components in both regimes. The respective VAR model corresponds to

$$\mathbf{z_{t+1}} = \begin{cases} \mathbf{A}^h\mathbf{z_t} + \phi^h\Delta i^u_{t+1} + \omega^{\perp,h}_{t+1} & \text{if } rr_{t+1} > 0 \\ \mathbf{A}^l\mathbf{z_t} + \phi^l\Delta i^u_{t+1} + \omega^{\perp,l}_{t+1} & \text{if } rr_{t+1} < 0 \end{cases} \tag{4.20}$$

To asses the impact of monetary policy surprises on each of the discounted sums of expected future excess returns, dividends and real interest rates, we once more use a two-step estimation procedure as outlined in section 4.3.3.

4.3.5 Bootstrap simulation

Small sample standard errors and associated confidence intervals of the components of the variance decomposition and the impacts of monetary policy shocks are computed using a nonparametric bootstrap method with a simulation size of 10,000 runs. This means that we sample 10,000 observations with replacement from the original empirical sample distribution in order to calculate these statistics. Bootstrapped statistics are likely to be more accurate as it is well known that the delta method used in Bernanke and Kuttner (2005) understates true standard errors. Furthermore, compared to a Monte Carlo simulation which assumes a given distribution, the bootstrap procedure is based on the empirical distribution which allows for a potential non-normality of the data and thus improves the accuracy of the estimated standard errors.

4.4 Data and sample period

For the event study, our sample period runs from January 1999 to December 2014; that is beginning when the ECB became responsible for monetary policy in the euro area. Stock returns are calculated as $r_t = ln(P_t) - ln(P_{t-1})$ where P_t is the closing price of the respective stock market index. We use the Morgan Stanley Capital International (MSCI) Germany stock index for the German economy and the MSCI World Ex Europe as a global stock market index excluding Europe. The unanticipated change in conventional monetary policy is computed as

$$\Delta r_t^u = f_{s,t} - f_{s,t-1} \tag{4.21}$$

where Δr_t^u represents the monetary policy surprise at day t, and $f_{s,t} - f_{s,t-1}$ the difference between the futures spot rate at day t and the prevailing rate at the day before the announcement, $t-1$. The futures implied rate on each day t is computed as 100 minus the futures contract price. The expected change in the policy rate is then the difference between the actual rate change (Δr_t) and the policy surprise calculated above:

$$\Delta r_t^e = \Delta r_t - \Delta r_t^u. \tag{4.22}$$

To measure unexpected unconventional monetary policies, we follow Rogers, Scotti, and Wright (2014) and Haitsma, Unalmis, and de Haan (2016) and proxy the surprise component $\Delta r_t^{u,c}$ by the change in the spread between 10-year Italian and German government bond yields

$$\Delta r_t^{u,c} = \left(y_{s,t}^I - y_{s,t}^G\right) - \left(y_{s,t-1}^I - y_{s,t-1}^G\right) \tag{4.23}$$

where $y_{s,t}^I$ and $y_{s,t}^G$ are the Italian and German 10-year government bond yields at day t, respectively.

The announcement dates for unconventional monetary policy measures are provided by Haitsma, Unalmis, and de Haan (2016) for the entire sample period. The data are obtained from Datastream. For the variance decomposition, we follow Campbell and Ammer (1993) and Bernanke and Kuttner (2005) and include six variables in the state vector $\mathbf{z_t}$ of the forecasting VAR: excess return on the stock market (rx), the short-term real interest rate (rr), changes in the nominal short-term rate (dy), the spread between yields on long-term government bonds and a short-rate (term spread, s), the log dividend-price ratio

(dp), and the short-rate relative to a one-year moving average (relative bill-rate, rb). Descriptive statistics for the variables used in the VAR together with the proxy for the monetary policy shock (Δi^u) are reported in table 4.1.

Table 4.1: Summary statistics of the time series

	Mean	Std. dev.	Min.	Max.	Skew.	Kurt.	JB-test	Obs.
rx	0.14	5.86	-29.04	16.80	-0.86	5.34	159.56***	456
rr	0.18	0.33	-0.99	0.94	-0.55	3.40	25.81***	456
dy	0.00	0.03	-0.16	0.39	3.68	54.47	51351.94***	456
s	0.98	1.28	-3.52	3.17	-0.74	3.01	41.91***	456
dp	-3.86	1.20	-9.57	-2.27	-3.13	14.08	3074.59***	456
rb	-0.01	0.08	-0.23	0.40	0.36	5.62	140.83***	456
Δi^u	-0.02	0.20	-1.23	0.71	-1.59	11.89	712.18***	192

This table reports descriptive statistics for the variables employed in the VAR models, excess return on the stock market (rx), the short-term real interest rate (rr), changes in the nominal short-term rate (dy), the spread between yields on long-term government bonds and a short-rate (term spread, s), the log dividend-price ratio (dp), and the short-rate relative to a one-year moving average (relative bill-rate, rb), as well as the monetary policy shock, Δi^u. The sample period starts in January 1977 and ends in December 2014 for all variables except the monetary policy shock that starts later, in January 1999. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively. JB-test stands for the Jarque-Bera test for normality.

The data frequency is monthly. The baseline sample for the VAR estimates and the variance decomposition runs from January 1977 to December 2014. Although the VAR is estimated for a sample period running from January 1977 to December 2014, our measure for monetary policy shocks is restricted to a shorter sample running from January 1999 to December 2014.[5] This is due to the introduction of the Euro and a common monetary policy for the members (including Germany) of the European monetary union in January 1999. Stock market excess returns are calculated as the stock return minus the short-term risk-free rate. Continuous stock returns are computed from MSCI Germany stock index price data, $r_t = ln(P_t) - ln(P_{t-1})$, while the 1-month deposit rate is

[5]The VAR dynamics needed as input for this analysis are, however, estimated over a sample longer than the period in which our measure of the monetary policy shock is available which improves the estimates' precision of the underlying VAR parameters, as discussed in 4.3.3.

used as a proxy for the short-term interest rate. To calculate real interest rates (rr), we subtract the log difference in the non-seasonally adjusted consumer price index (CPI) from the nominal short-term interest rate. The change in the nominal short-rate (dy) is computed as the monthly difference in the 1-month deposit rate. As in Campbell and Ammer (1993), the term spread is calculated as the difference between 10-year government bond yields and the 3-month interbank rate. The relative bill rate is defined as the 1-month deposit rate minus its one-year lagged moving average. Monthly dividends are obtained as the difference between the return on the total return index (including dividend payments) in $t + 1$ minus the return on the price index (excluding dividends) in $t + 1$ times the stock price index in t for the MSCI Germany stock market index. Monthly dividends are annualized in order to avoid seasonal patterns. Following the standard procedure in the literature, the monthly (smoothed) log dividend-price ratio is obtained as the log of the sum of monthly dividends over the previous year minus the log of the stock price index in the current month. The time series data used in the VAR estimation are plotted in figure 4.1.

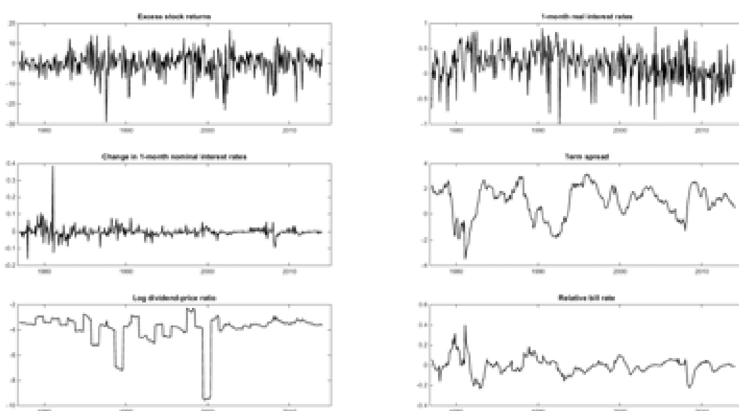

Figure 4.1: Time series of variables included in the VAR estimation.

The parameter ρ from the log-linear approximation procedure needs to be estimated. As outlined in section 4.3, we use the expression $\rho = 1/\left(1 + \exp\left(d - p\right)\right)$ where $d - p$ is the mean log dividend price ratio. Using German stock market data, the estimation result is $\rho = 0.9794$. Following Bredin, Hyde, and O'Reilly

(2010), we proxy surprises in the ECB policy rate by the 1-month change in the 3-month Euribor futures rate.

The time series data for the 10-year government bond, the 3-month interbank rate, and the CPI are obtained from FRED Economic Data of the St. Louis FED. The remaining data series are collected from Datastream. Augmented Dickey-Fuller (Dickey and Fuller, 1979) and Phillips-Perron (Phillips and Perron, 1988) unit root tests (not reported; details are available upon request) confirm stationarity of all variables used in the VAR analysis.

4.5 Empirical results

4.5.1 Event study

In table 4.2, we report the impact of conventional and unconventional monetary policy surprises on German stock returns. For the pre-crisis period, we find a significant effect of monetary policy surprises on the stock market. Quantitatively, the results imply that a 25 basis point surprise increase in the ECB policy rate is associated with a 1.64% decline in stock returns. For the crisis period, the coefficient of conventional monetary policy is not significant. However, a Wald test cannot reject the null hypothesis that the pre-crisis parameter is equal to the crisis parameter. Moreover, we also find a highly significant inverse relationship between the unconventional monetary policy surprise and the stock market. More precisely, a decrease in the yield spread between Italian and German government bonds implies an increase in stock returns. During the sample period, the average change in the yield spread on event days was a 0.06% decrease which caused an increase in stock returns of 0.31%. Finally, we find that expected policy changes have a highly negative impact on the stock market. This result contradicts the efficient markets paradigm since expected policy changes are reflected in market prices, so that expected policy changes should not influence stock returns. These results are consistent with those in Haitsma, Unalmis, and de Haan (2016) and confirm the importance of controlling for unconventional monetary policy measures applied during the recent crisis when studying the impact of ECB monetary policy on the stock market.

Table 4.2: Influence of ECB monetary policy changes on German stock returns

Constant	0.000
	(0.276)
Conventional surprise pre-crisis	-0.066**
	(-2.093)
Expected change pre-crisis	-0.146***
	(-3.348)
Conventional surprise crisis	0.076
	(0.904)
Expected change crisis	0.036
	(0.360)
Surprise unconventional	-0.052***
	(-4.007)
MSCI World Ex Europe	0.726***
	(25.062)
Crisis dummy	-0.000
	(-0.678)
$\bar{R}^2$	0.307
Wald test F-statistic	2.472

This table reports the results of the regression model estimated in equation (4.1). t-statistics are based on Newey and West (1987) standard errors and given in parenthesis. The reported measure of fit, $\bar{R}^2$, is the adjusted R^2. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively. Number of observations: 4173.

4.5.2 VAR estimates

We follow Campbell and Ammer (1993) and Bernanke and Kuttner (2005) and estimate a parsimonious, six-variable one-lag VAR system[6] that includes excess returns on stocks, the short-term real interest rate, changes in the nominal short-term rate, the term spread, the log dividend-price ratio, and the relative bill rate.

[6]Information criteria (e.g. Akaike, Schwartz, Hannan-Quinn) do not unambiguously indicate the optimal VAR lag length for our models. Due to this fact, we did also estimate VAR models with higher lag orders. However, the results of the variance decomposition are not affected as compared to the parsimonious one-lag VAR.

Table 4.3: VAR estimates

Panel A: Standard VAR estimates: January 1977 to December 2014

	rx_{t+1}	rr_{t+1}	dy_{t+1}	s_{t+1}	dp_{t+1}	rb_{t+1}
rx_t	0.086*	-0.005*	0.000	0.003	-0.003	0.000
	(1.817)	(-1.806)	(0.665)	(1.373)	(-0.696)	(0.746)
rr_t	-0.414	0.089*	-0.007	0.013	-0.101	-0.007
	(-0.480)	(1.895)	(-1.618)	(0.341)	(-1.395)	(-1.601)
dy_t	10.624	-0.061	-0.158***	0.278	-1.147	-0.107**
	(1.040)	(-0.110)	(-2.935)	(0.597)	(-1.337)	(-1.991)
s_t	0.307	-0.060***	0.003**	0.950***	-0.026	0.004***
	(1.237)	(-4.438)	(2.216)	(83.929)	(-1.248)	(3.132)
dp_t	0.015	-0.030**	-0.002	0.019*	0.908***	-0.002
	(0.063)	(-2.401)	(-1.357)	(1.795)	(46.156)	(-1.419)
rb_t	-4.341	-0.192	0.103***	-0.933***	-0.029	0.969***
	(-0.987)	(-0.807)	(4.469)	(-4.655)	(-0.078)	(41.977)
$\bar{R}^2$	0.008	0.078	0.049	0.956	0.832	0.858

Panel B: Threshold VAR estimates, threshold: $rr_{t+1} > 0$

	rx_{t+1}	rr_{t+1}	dy_{t+1}	s_{t+1}	dp_{t+1}	rb_{t+1}
rx_t	0.035	-0.002*	0.000	0.002	-0.006	0.000
	(0.516)	(-1.420)	(0.838)	(0.597)	(-1.362)	(0.8847)
rr_t	0.920	0.006	-0.005	0.021	-0.268***	-0.004
	(0.715)	(0.201)	(-0.6128)	(0.346)	(-3.086)	(-0.578)
dy_t	2.754	0.040	-0.118	-0.292	-0.513	-0.069
	(0.201)	(0.124)	(-1.480)	(-0.445)	(-0.554)	(-0.869)
s_t	0.134	-0.051***	0.003	0.954***	-0.003	0.004**
	(0.388)	(-6.193)	(1.495)	(57.88)	(-0.148)	(1.995)
dp_t	-0.106	-0.011	-0.002	0.018	0.904***	-0.002
	(-0.329)	(-1.43)	(-0.869)	(1.192)	(41.752)	(-0.904)
rb_t	2.14	-0.271*	0.099***	-0.807***	0.194	0.964***
	(0.331)	(-1.761)	(2.641)	(-2.605)	(0.444)	(25.549)
$\bar{R}^2$	-0.021	0.154	0.018	0.954	0.888	0.823

Panel C: Threshold VAR estimates, threshold: $rr_{t+1} < 0$

	rx_{t+1}	rr_{t+1}	dy_{t+1}	s_{t+1}	dp_{t+1}	rb_{t+1}
rx_t	0.097	-0.002	0.000	0.006**	0.001	-0.000

	(1.480)	(-0.563)	(-0.936)	(2.271)	(0.068)	(-0.807)
rr_t	-1.031	-0.059	-0.008	-0.011	0.084	-0.008*
	(-0.901)	(-1.081)	(-1.587)	(-0.228)	(0.684)	(-1.647)
dy_t	19.702	0.141	-0.256***	1.482**	-2.724	-0.200***
	(1.250)	(0.189)	(-3.659)	(2.183)	(-1.602)	(-2.883)
s_t	0.463	0.024	0.003*	0.936***	-0.052	0.005***
	(1.261)	(1.364)	(1.876)	(59.233)	(-1.323)	(2.821)
dp_t	0.196	-0.012	-0.001	0.014	0.915***	-0.001
	(0.582)	(-0.756)	(-0.814)	(0.993)	(25.144)	(-0.899)
rb_t	-11.565**	-0.097	0.112***	-1.113***	-0.175	0.978***
	(-2.033)	(-0.361)	(4.427)	(-4.543)	(-0.285)	(39.001)
$\bar{R}^2$	0.037	-0.004	0.099	0.955	0.756	0.908

This table reports coefficient estimates from a vector autoregression (VAR) with a lag-length of one month that includes excess return on the stock market (rx), the short-term real interest rate (rr), changes in the nominal short-term rate (dy), the spread between yields on long-term government bonds and a short-rate (term spread, s), the log dividend-price ratio (dp), and the short-rate relative to a one-year moving average (relative bill-rate, rb). Panel A gives the estimates from a standard VAR using data over the full sample period from January 1977 to December 2014. Panels B and C report estimates from a threshold VAR where the threshold variable is the level of the monthly real interest rate. Panel B shows the threshold VAR estimates when the real interest rates were positive and panel C presents the estimates of the threshold VAR when the real interest rates were negative. t-statistics are given in parentheses. The reported measure of fit, $\bar{R}^2$, is the adjusted R^2. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

Panel A of table 4.3 gives the estimates from a standard VAR using data over the full sample period from January 1977 to December 2014. The first column of panel A indicates that stock market excess returns are only predicted by their own lag. The other variables included in the VAR system do not significantly predict future excess returns. Our results are consistent with recent empirical work (e.g. Nitschka, 2014) and common when assessing a one-month ahead time variation in stock market returns. Panels B and C of table 4.3 report estimates from a threshold VAR where the threshold variable is the level of the monthly real interest rate. Panel B shows the threshold VAR estimates for states of the world when the real interest rate was positive and panel C presents the threshold VAR estimates for states of the world when the real interest rate was negative. The results in panel B are based on 234 months in which the

German real interest rate was positive. The results in panel C are obtained from the 221 months in which the monthly real interest rate was negative. The results show that in a positive interest rate regime, none of the variables used in the stock return forecasting equation are significant, while in a negative interest rate regime, stock market excess returns are predicted by the relative bill rate. The VAR estimates for the other variables are similar to the baseline sample estimation in panel A.

4.5.3 Variance decomposition

In table 4.4, we report the variance decomposition for German stock excess returns for two sample periods and the threshold VAR. Panel A gives the results for the baseline sample period from January 1977 to December 2014. Panel B gives the results excluding the recent financial and sovereign debt crisis. The sample period runs from January 1977 to December 2007. Panel C provides the corresponding results based on the threshold VAR estimates for months with a positive real interest rate. Panel D reports the results for months in which the real interest rates were negative. Both threshold VAR models are estimated for the baseline sample period from January 1977 to December 2014.

The variance decompositions show that dividend news dominates the volatility of unexpected stock return movements in the German economy, independent of the prevailing interest rate regime. This finding is different to existing US evidence where stock market volatility is mainly driven by future excess return news (e.g. Campbell and Ammer, 1993; Bernanke and Kuttner, 2005; Wang and Zhu, 2013). However, in a recent study, Nitschka (2014) shows that in periods with negative real interest rates, dividend news is mostly responsible for the variation in US excess stock returns.

Table 4.4: Variance decomposition of excess stock returns

Panel A: January 1977 - December 2014			
$\mathrm{Var}(\tilde{e}^{rx}_{t+1})$	0.123	(2.831)	[0.066 , 0.224]
$\mathrm{Var}(\tilde{e}^{rr}_{t+1})$	0.007	(7.421)	[0.005 , 0.009]
$\mathrm{Var}(\tilde{e}^{d}_{t+1})$	1.216	(7.501)	[0.942 , 1.571]
$2\,\mathrm{Cov}(\tilde{e}^{rr}_{t+1}, \tilde{e}^{rx}_{t+1})$	-0.038	(-0.197)	[-0.064 ,-0.021]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rx}_{t+1})$	-0.336	(-0.253)	[-0.516 , -0.215]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rr}_{t+1})$	0.028	(0.131)	[0.006 , 0.054]

Panel B: January 1977 - December 2007			
$\mathrm{Var}(\tilde{e}^{rx}_{t+1})$	0.067	(2.853)	[0.035 , 0.122]
$\mathrm{Var}(\tilde{e}^{rr}_{t+1})$	0.006	(6.562)	[0.005 , 0.009]
$\mathrm{Var}(\tilde{e}^{d}_{t+1})$	1.158	(6.291)	[0.859 , 1.572]
$2\,\mathrm{Cov}(\tilde{e}^{rr}_{t+1}, \tilde{e}^{rx}_{t+1})$	-0.021	(-0.149)	[-0.040 , -0.008]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rx}_{t+1})$	-0.229	(-0.331)	[-0.321 , -0.157]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rr}_{t+1})$	0.018	(0.109)	[-0.000 , 0.040]
Panel C: threshold, $rr_{t+1} > 0$			
$\mathrm{Var}(\tilde{e}^{rx}_{t+1})$	0.010	(3.669)	[0.006 , 0.017]
$\mathrm{Var}(\tilde{e}^{rr}_{t+1})$	0.003	(3.294)	[0.002 , 0.005]
$\mathrm{Var}(\tilde{e}^{d}_{t+1})$	1.127	(4.833)	[0.764 , 1.672]
$2\,\mathrm{Cov}(\tilde{e}^{rr}_{t+1}, \tilde{e}^{rx}_{t+1})$	0.000	(0.008)	[-0.002 , 0.003]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rx}_{t+1})$	-0.131	(-0.349)	[-0.170 , -0.094]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rr}_{t+1})$	-0.008	(-0.065)	[-0.022 , 0.004]
Panel D: threshold, $rr_{t+1} < 0$			
$\mathrm{Var}(\tilde{e}^{rx}_{t+1})$	0.351	(3.860)	[0.206 , 0.560]
$\mathrm{Var}(\tilde{e}^{rr}_{t+1})$	0.002	(8.347)	[0.002 , 0.003]
$\mathrm{Var}(\tilde{e}^{d}_{t+1})$	1.484	(6.696)	[1.115 , 1.972]
$2\,\mathrm{Cov}(\tilde{e}^{rr}_{t+1}, \tilde{e}^{rx}_{t+1})$	0.004	(0.073)	[-0.005 , 0.011]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rx}_{t+1})$	-0.828	(-0.416)	[-1.165 , -0.547]
$-2\,\mathrm{Cov}(\tilde{e}^{d}_{t+1}, \tilde{e}^{rr}_{t+1})$	-0.013	(-0.155)	[-0.025 , 0.000]

This table reports results from the variance decomposition of the revision in expectations about current excess returns into variances and covariances of three news components: news about future excess returns $\tilde{e}^{rx}_{t+1}$, real interest rate $\tilde{e}^{rr}_{t+1}$ and dividends $\tilde{e}^{d}_{t+1}$. These statistics are normalized by the variance of current excess return news, such that they add up to one. The numbers in parentheses and brackets are t-statistics and 95% confidence intervals, respectively, based on 10,000 bootstrap simulations. Panel A gives the results for the baseline sample period from January 1977 to December 2014. Panel B gives the results excluding the recent financial and sovereign debt crisis. The sample period runs from January 1977 to December 2007. Panel C provides the corresponding results based on the threshold VAR estimates for months with a positive real interest rate. Panel D reports the results for months in which the real interest rates were negative. Both threshold VAR models are estimated for the baseline sample period from January 1977 to December 2014.

The distinct role of dividend news can be linked to the vast literature on stock return predictability. Variations in the dividend-price ratio must reflect future dividend growth or expected returns. For US data, there is strong evidence that the variability in the dividend-price ratio can mainly be explained by expected returns. In general, this result is derived from predictive regressions which suggest that the dividend-price ratio predicts future returns and not dividend growth (e.g. Cochrane, 2008b, 2011). However, there has been some recent empirical work which shows that dividend growth can be predicted by dividend yield. Chen (2009) demonstrates that the dividend yield did predict dividend growth in the US in a period pre-WWII but that the relationship was reversed in the more recent past. Engsted and Pedersen (2010) show for Sweden and Denmark that dividend growth is strongly predictable by the dividend-price ratio, whereas returns are not predictable. In a cross-country analysis, Rangvid, Schmeling, and Schrimpf (2014) provide evidence that dividend-growth predictability by the dividend-price ratio is the rule rather than the exception in global stock markets. For Germany, news about future dividends constitutes the main driver of stock market valuation.

Another interesting finding of our analysis is that in a low real interest rate regime, which is typical for a slowly growing or recessionary economy, the variance in expected future excess returns accounts for a larger part of the variance in the current stock return as compared to the other sample periods. This is consistent with theoretical stock market models of time-varying risk-aversion like the habit formation model of Campbell and Cochrane (1999) which implies that people's risk aversion and hence, expected risk premia (future excess returns) increase during periods with lower economic growth.

4.5.4 Impact of monetary policy

Next we study the impact of ECB monetary policy shocks on German future excess returns, real interest rates and dividends. The contemporaneous response to a conventional monetary policy surprise is estimated on the January 1999 to December 2014 sample and reported in panel A of table 4.5. The ECB policy rate surprises have a marginally significant impact on the discounted sum of future excess returns, while accounting for almost all of the contemporaneous excess return response of -2.559.[7] Overall, the signs for the variance decomposition considering the shocks to monetary policy are consistent with the results in Bernanke and Kuttner (2005).

In panel B of table 4.5, the impact of surprise Euro area monetary policy changes for the high and low interest rate regime is reported. For positive real interest rates, we only observe a borderline significant stock market reaction at the 10% level. However, the results for the low interest rate regime show that unexpected changes in the ECB policy rate have a large and statistically highly significant impact on contemporaneous excess stock returns. Specifically, a 1% surprise increase in the policy rate implies a 12.394 contemporaneous decrease in excess stock returns. In terms of coefficients of the decomposition, the impact of future excess returns is -3.549 and significant at the 1% level while dividend news contributes with -8.931 to the overall impact at the 1% significance level. The impact of the real interest rates is very small and statistically insignificant. In summary, the results indicate that in a regime with negative interest rates, the impact of monetary policy surprises on the current excess stock return can be attributed to future excess return and dividend news. The economic intuition that a contractionary monetary policy implies an increase in risk premia and a decrease in stock prices can, for example, be explained by the fact that higher interest rates increase investors' risk aversion induced by a decrease in the expected levels of consumption (Campbell and Cochrane, 1999).

[7]In this study, using monthly data, the contemporaneous response excess stock returns to a monetary policy surprise are not significant. However, the studies cited in section 4.1 report a statistically significant impact of monetary policy shocks on German stock prices based on daily or intraday data.

Table 4.5: Impact of monetary policy shocks on future returns, interest rates and dividends

Panel A: January 1999 - December 2014

	e^{rx}_{t+1}	$\tilde{e}^{rx}_{t+1}$	$\tilde{e}^{rr}_{t+1}$	$\tilde{e}^{d}_{t+1}$
	-2.293	-2.273**	0.214	-0.234
	(-0.707)	(-1.968)	(1.459)	(-0.092)

Panel B: threshold VAR

	e^{rx}_{t+1}	$\tilde{e}^{rx}_{t+1}$	$\tilde{e}^{rr}_{t+1}$	$\tilde{e}^{d}_{t+1}$
$rr_{t+1} > 0$	7.511*	0.497	0.017	6.997
	(1.659)	(0.496)	(0.642)	(1.411)
$rr_{t+1} < 0$	-12.394***	-3.549***	0.086	-8.931***
	(-3.587)	(-3.856)	(0.636)	(-2.919)
Diff.	19.905***	4.045***	-0.069	15.929***
	(3.495)	(3.065)	(-0.407)	(3.056)

Panel C: Pre-crisis and crisis sample

	e^{rx}_{t+1}	$\tilde{e}^{rx}_{t+1}$	$\tilde{e}^{rr}_{t+1}$	$\tilde{e}^{d}_{t+1}$
1999-2007	-9.124	-1.071	-0.061	-7.992
	(-1.237)	(-0.466)	(-0.193)	(-1.275)
2008-2014	-2.703	-3.158**	0.307	0.148
	(-0.624)	(-2.005)	(1.537)	(0.044)
Diff.	-6.420	2.087	-0.368	-8.140
	(-0.753)	(0.752)	(-0.998)	(-1.144)

This table reports the impact of monetary policy surprises on the current excess stock return and the discounted sums of future excess stock returns, current and future real interest rates, and current and future dividends. The six-variable VAR(1) used to model excess stock return dynamics is estimated over the sample period January 1999 to December 2014 for the baseline analysis reported in panel A, the threshold VAR reported in panel B and the crisis and pre-crisis sample reported in panel C. The VAR(1) sample period for the pre-crisis period is January 1977 to December 2007. The contemporaneous response to a monetary policy surprise is estimated on the January 1999 to December 2014 subsample for the baseline analysis reported in panel A, the threshold VAR reported in panel B, while the subsample for the pre-crisis period runs from January 1999 to December 2007 and the subsample for the crisis period goes from January 2008 to December 2014. t-statistics reported in parentheses are based on 10,000 bootstrap simulations. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

In panel C, we divide our sample into a pre-crisis period running from January 1999 to December 2007 and a period covering the recent global financial and European sovereign debt crisis running from January 2008 to December 2014. The decomposition for the pre-crisis period shows that none of its components is statistically significant. The results for the crisis period are similar to those for the overall sample. However, the difference in point estimates between the two samples is statistically insignificant which implies that the stock market reaction to monetary policy shocks did not change in the aftermath of the recent crisis. Kontonikas, MacDonald, and Saggu (2013) report in an event study using daily US data that throughout the crisis period, stocks did not react positively to unexpected cuts in the federal funds rate, which were interpreted as signals of worsening future economic conditions. However, we do not observe any such effects at a monthly data frequency for the German stock market.

The interesting and novel finding of our analysis in table 4.5 is that a strong and significant stock market reaction under conventional monetary policy is only observed when the real interest rates are negative. Low interest rates are typical for an economy with higher uncertainty and volatile consumption since people want to save more (precautionary savings behavior), thereby driving down the interest rates. The empirical evidence derived in this paper shows that in such an environment, unexpected changes in monetary policy have pronounced effects on stock market excess returns. Moreover, these results indicate that the effect of monetary policy on the stock market is state dependent and matters particularly in bad economic times.

In the event study conducted in section 4.5.1, we found a significant impact of unconventional monetary policy (measured by the monthly change in the Italian-German government bond yield spread) on the German stock market. Next, we analyze what is the impact of these unconventional monetary policy shocks on German future excess returns, real interest rates and dividends. Panel A of table 4.6 reports the contemporaneous response of German excess stock returns to an unconventional monetary policy shock for the baseline sample period January 1999 to December 2014. The result implies that an unconventional monetary policy surprise that causes a decrease in the Italian-German yield spread causes an increase in stock returns. The variance decomposition indicates that the main channel behind this response can be attributed to dividend news at a 1% significance level. The impact of future excess returns and real interest rates is very small and insignificant.

Since non-standard monetary policy measures were introduced with the onset of the financial crisis in late 2007, it is reasonable to split the sample into a pre-crisis and and a crisis period. Panel B of table 4.6 shows that none of the components of the decomposition for the pre-crisis sample is statistically significant. However, as expected, during the crisis period, unconventional monetary policy shocks have a highly significant impact on German excess stock returns.

Table 4.6: Impact of unconventional monetary policy on future returns, interest rates and dividends

Panel A: January 1999 - December 2014

	e^{rx}_{t+1}	$\tilde{e}^{rx}_{t+1}$	$\tilde{e}^{rr}_{t+1}$	$\tilde{e}^{d}_{t+1}$
	-6.717***	-0.301	0.054	-6.470***
	(-2.762)	(-0.535)	(0.402)	(-2.878)

Panel B: Pre-crisis and crisis sample

	e^{rx}_{t+1}	$\tilde{e}^{rx}_{t+1}$	$\tilde{e}^{rr}_{t+1}$	$\tilde{e}^{d}_{t+1}$
1999-2007	-3.369	8.422	-4.304	-7.486
	(-0.054)	(0.571)	(-1.606)	(-0.131)
2008-2014	-6.480***	-0.319	0.098	-6.259***
	(-2.759)	(-.600)	(0.744)	(-2.899)
Diff.	3.112	8.741	-4.402	-1.227
	(0.050)	(0.592)	(-1.641)	(-0.021)

This table reports the impact of unconventional monetary policy surprises on the current excess stock return and the discounted sums of future excess stock returns, current and future real interest rates, and current and future dividends. The six-variable VAR(1) used to model excess stock return dynamics is estimated over the sample period January 1999 to December 2014 for the baseline analysis reported in panel A and the crisis and pre-crisis sample reported in panel B. The VAR(1) sample period for the pre-crisis period is January 1977 to December 2007. The contemporaneous response to an unconventional monetary policy surprise is estimated on the January 1999 to December 2014 subsample for the baseline analysis reported in panel A while the subsample for the pre-crisis period runs from January 1999 to December 2007 and the subsample for the crisis period goes from January 2008 to December 2014 (panel B). t-statistics reported in parentheses are based on 10,000 bootstrap simulations. *, **, ***, denote statistical significance at the 10, 5, and 1 % level, respectively.

As in the baseline analysis, we identify news about future dividends as the main propagation channel of unconventional monetary policy. One explanation for this observation is that investors believe that lower long-term interest rates have a positive impact on the economy by increasing the corporate profits through lower debt-service payments and stronger economic growth. This implies an increase in future dividend payments and ceteris paribus a higher stock market valuation.

4.6 Conclusion

This paper studies factors explaining time variation in excess stock returns and examines the impact of unanticipated changes in ECB monetary policy on the German stock market. An essential part of the study is the decomposition of the variation in current excess stock returns and the respective contemporaneous response to monetary policy shocks into revisions of the expectations regarding future excess returns, future dividends and future real interest rates. The main findings are that overall variations in German stock returns mainly reflect revisions in expectations about dividends and that the stock market response to monetary policy shocks is dependent on the prevailing interest rate regime. Specifically, a strong and significant stock market reaction to surprise changes in monetary policy is only observed when real interest rates are negative. We find that the effects of unanticipated changes in monetary policy on expected excess returns and future dividends account in equal proportions for all the response of stock prices in such an interest rate regime. Moreover, our analysis provides empirical evidence that unconventional monetary policy has a significant impact on stock returns. Future dividends are identified as the main propagation channel of unconventional monetary policy on the observed stock market response.

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Sammanfattning

Denna avhandling består av tre fristående uppsatser som behandlar frågor i skärningspunkten mellan tillgångsprissättning och makroekonomi.

Den första uppsatsen **Asset pricing implications of a DSGE model with recursive preferences and nominal rigidities** studerar den makroekonomiska dynamiken och tillgångspriser i en produktionsekonomi med nominella prisstelheter och Epstein och Zin (1989) preferenser. Baserat på en rimlig kalibrering är den makroekonomiska DSGE-modellen förenlig med ett antal stiliserade fakta som observeras på finansmarknaderna såsom en riskpremie för aktier, en negativ real räntekurva, en positiv nominell räntekurva och prediktiv avkastning på aktiemarknaden. Trots detta komprometteras inte modellens förmpåga att matcha dynamiken i de makroekonomiska variablerna. Utjämning av räntor över tid, baserat på en regel för penningpolitiken, bidrar till att generera en låg riskfri volatilitet vilket har varit svårt att åstadkomma för reala konjunkturcykelmodeller i vilka penningpolitiken är neutral. I en tillämpning visar jag att modellen ger ett ramverk för att analysera de penningpolitiska interventionerna och de relaterade effekterna på tillgångspriser och den reala ekonomin.

Den andra uppsatsen **Macroeconomic news and the stock market: Evidence from the eurozone** är en empirisk studie av överavkastning på aktiemarknaden i euroområdet kring de dagar när viktiga makroekonomiska nyheter om inflation, arbetslöshet eller räntor tillkännages. Jag identifierar så kallad state dependence, dvs att riskpremien är signifikant högre när räntorna ligger i närheten av noll procent (zero lower bound). Vidare finner jag belägg för att under hela undersökningsperioden är riskpremien inom eurozonen enbart högre de dagar då den amerikanska centralbankens (FOMCs) tillkännagivanden är planerade. Detta resultat försvinner emellertid i tider av låg ränta. Slutligen dokumenterar jag att den europeiska aktiemarknaden inte prissätter planerade tillkännagivanden från den Europeiska Centralbanken (ECB).

Den tredje uppsatsen **The impact of ECB monetary policy surprises on the German stock market** är ett gemensamt arbete med Markus Sigonius och studerar effekten av ECBs penningpolitiska beslut på tyska aktiemarknadens riskpremie. Först så utför vi en eventstudie för att bedöma effekten av konventionell och icke-konventionell penningpolitik på aktiemarknadens riskpremie. Sedan använder vi Campell and Ammer's VAR-modell från 1993 för att separera riskpremien i nyheter om förväntad riskpremie, förväntade utdelningar och förväntade realräntor. Vi mäter konventionella penningpolitiska chocker genom att använda data över terminsräntor. Våra främsta resultat är att variationen i tyska aktiemarknadens riskpremie främst speglar revideringar av förväntade riskpremier och att aktiemarknadens respons på penningpolitiska chocker beror på den rådande ränteregimen. I perioder med negativa realräntor leder en överraskande återhållen penningpolitik till en minskning av aktiemarknadens riskpremie. De bakomliggande kanalerna för denna respons är nyheten om högre förväntad riskpremie och lägre förväntade utdelningar.

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