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Macroeconomic announcements, volatility and interrelationships: An examination of the UK bond and stock markets

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MACROECONOMIC ANNOUNCEMENTS, VOLATILITY AND INTERRELATIONSHIPS:

AN EXAMINATION OF THE UK BOND AND STOCK MARKETS

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Declaration

This thesis contains no material that has been accepted for the award of any other degree or diploma in any tertiary institution. To the best of my knowledge and belief, this thesis contains no material previously published or written by any other person, except when due reference is made in the text.

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Bradley Jones

Abstract

This study covers considerable ground and touches on a range of issues in a rigorous investigation of the intraday and end-of-day behaviour of UK stock index and interest rate futures contracts. Firstly, the paper uses 5-minute data in an initial examination of the response of the Short Sterling, Long Gilt and FTSE100 to the release of macroeconomic announcements (assisted with the application of GMM). Secondly, in an analysis of intraday patterns in returns and volatility a GARCH(1,1) framework is employed, so that further inferences are made robust to time-varying variance. Finally, the paper draws upon some of the latest innovations in time series econometric modeling in an attempt to identify the extent to which the Short Sterling, Long Gilt and FTSE100 exhibit co-movement.

The study finds evidence suggesting investors and portfolio managers distinguish between the information content of different items of news. The results also suggest some consistency of response to news in the interest rate and stock markets. The GARCH estimation shows variance to be highly dependent on past variance and volatility. Although the three variables appear to be bound by two cointegrating relationships, the tests for lead/lag relationships and relative degrees of exogeneity produced mixed results. In sum, the results should prove intuitively appealing.

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1. Introduction

What information might be expected to move stock and bond markets? To what extent do financial assets respond in a similar manner to the arrival of new information? Furthermore, how are the markets on which such assets trade inter-linked? To address these questions, this study covers considerable ground in conducting a rigorous investigation into the pricing behaviour of UK equity index and interest rate futures contracts .

The paper uses high frequency data and precise release times in an initial examination of the response of the Short Sterling, Long Gilt and FTSE100 to the release of scheduled macroeconomic announcements. The regression results are enhanced with the application of the Generalized Method of Moments estimator, which ensures robustness to returns autocorrelation and heteroscedastistic errors. In an analysis of intraday patterns in returns and volatility, a GARCH (1,1) framework is employed, so that further inferences are made robust to time-varying variance. Finally, the paper draws upon some of the latest innovations in time series econometric modeling in an attempt to identify the extent to which the Short Sterling, Long Gilt and FTSE100 exhibit co-movement.

Unlike many of its predecessors, it is the contention of this paper to seek a generalized explanation of interest rate and share market behaviour. The current exercise should be of interest as it helps describe how markets in the UK respond to information reflecting underlying economic conditions, in the process shedding light on whether traders react rationally to new information - in other words are their responses in accordance with widely accepted views about how the economy operates. Additionally, the identification of possible relationships within and across markets may be of more general interest since the UK is considered a relatively large, open and liquid market, situated in the financial 'epicenter' of Europe.

The remainder of the paper is structured as follows. Section two contains an appraisal of how the literature on this topic has evolved. Section three contains a brief theoretical review pertaining to fundamental valuation models of stock and bond prices. Section four presents the features of the data, with a description of the statistical methodology provided in section five. In section six implications of the results are discussed. Finally, conclusions are drawn and suggestions for future research given in section seven. In brief, the results suggest traders distinguish between the information content of different news items. The analysis also suggests some consistency of response to news in the interest rate and stock markets. Whilst the three variables appear to be bound by two cointegrating relationships, the tests for lead/lag relationships and relative degrees of exogeneity produced mixed results.

2. Literature Overview and Motivation

A sizable literature has now evolved examining announcement effects, particularly in the bond and foreign exchange markets, the majority of which has emanated from the US. In stark contrast, the opposite can be said for research attempting to identify relative degrees of exogeneity across interest rate and equity markets. The following review serves two purposes: it pulls together many of the different strands of literature on this topic, and this gives rise to the motivation and underlying aims for the ensuing analysis of UK bond and stock market behaviour.

Theory says that movements in financial asset prices should reflect new information about fundamental asset values. In the case of the bond market, such theory has been confirmed repeatedly(1). Various US studies over the years have documented a significant bond market reaction arising from macroeconomic announcements, including money supply, PPI, CPI and unemployment (see table 1). The US literature also provides evidence of a 'flavour of the month' aspect to bond market behaviour, in which different announcements are regarded with varying degrees of importance depending on the prevailing state of the economy. For example, Prag (1994) documented that the effect of the unemployment announcement on bonds to be largely dependent on the existing unemployment rate.

(1) One notable exception is Elmendorf, Hirshfeld and Weil (1996), who found it difficult to relate the largest movement in UK bond prices to news arrival from 1900 to 1920.

Table 1. Frequency of Announcements Found to be Significant

US Bond Market Studies	
Announcement	*Significant*
Money Supply	15
PPI	8
CPI	6
Unemployment	6
Ind Production	4
Retail Sales	3

Source: Fleming and Remolona (1997).

Whilst not nearly as extensive as in the US, UK research has also broadly established similar patterns. The studies of Haldane and Read (1999), Dale (1993) and Goodhart and Smith (1985), have found money supply to be influential, and the same can be said for Retail Price Index releases (see Joyce and Read, 1998, and Goodhart and Smith, 1985). In an examination of very high frequency price and volume adjustments around news release times, ap Gwilym, Buckle, Clare and Thomas (1998) found RPI, unemployment, PPI and retail sales to be important for the Short Sterling. Similar results were reported by ap Gwilym and Thomas (1998) for the Long Gilt, who also highlighted the importance of US unemployment, PPI and GDP releases.

Analyzing 15-minute bond data, Becker, Finnerty and Kopecky (1995), documented a wider range of US announcements to be significant compared to domestic news. The US releases comprised PPI, CPI, non-farm payrolls, unemployment, retail sales, leading indicators and merchandise trade, while the significant domestic announcements were retail sales, visible trade, current account and Public Sector Borrowing Requirement (PSBR). More recently, the daily responses of short and long-term interest rates were also established by Brooke, Danton and Moessner (1999), to be sensitive to a range of domestic and US announcements, including average earnings and RPI in the UK, and CPI, non-farm payrolls, retail sales, industrial production and GDP in the US [(2)].

In contrast to the rich bond market announcement effect literature, there has not been a great deal of research modeling stock market responsiveness to macroeconomic news releases (this is discussed further in section three). In the US, Cutler, Poterba and Summers (1989) found that in most cases the information cited by the press as causing market movements were in fact quite unimportant. This reinforced the earlier studies of Schwert (1981), Pearce and Roley (1985) and Hardouvelis (1987), who all concluded there was little evidence to suggest that the stock market responds to macroeconomic news other than monetary information.

[(2)] Studies not distinguishing the relative importance of UK and US announcements include Clare and Courteney (1999), Buckle, ap Gwilym, Thomas and Woodhams (1998), Elmendorf et al (1994) and Becker et al (1993).

UK research distinguishing the relative importance of macroeconomic announcements has invariably established RPI releases to be important (see Joyce and Read, 2000, and Goodhart and Smith, 1985), and ap Gwilym et al (1998) found the FTSE100 to be sensitive to RPI, PPI and PSBR data. Of US news, Becker et al (1995) stated PPI, merchandise trade, non-farm payrolls and CPI were influential, with visible trade, current account and PSBR the most significant of the UK releases[3].

The overview of the existing research presented above leads us to highlight the primary aims for this study. It is envisaged the following analysis will extend the literature on UK bond and stock market behaviour in the following ways:

- By revisiting whether there is any consistency in the way the Short Sterling, Long Gilt and FTSE100 react to particular news releases. Discussed later, theory suggests short term rates should be more responsive to indicators of current activity compared to longer term rates, and that these indicators are ambiguous for stocks. The only study this author is aware of that conducts a simultaneous examination of these three financial instruments was by Goodhart and Smith (1985), however their analysis centered on daily data and just four domestic announcements. This paper examines tick data in five minute intervals and a wider array of news releases.

[3] Refer to Buckle et al (1998), and Clare and Courteney (1999) for studies which do not distinguish the relative importance of UK and US announcements for the UK equity market.

The application of high frequency data in announcement studies is absolutely critical in isolating the news release effect, as ap Gwilym et al (1998) reported that the adjustment process in both the UK interest rate and stock index futures markets is completed within three minutes;

- Only one known study (Buckle et al, 1998) has attempted to model UK interest rate and stock market returns and volatility in a framework capable of capturing the effect of changing variance in a time series. However this paper excluded the Long Gilt and also failed to distinguish between the relative importance of economic announcements. This study addresses both apparent shortcomings;
- To the authors knowledge, no paper has attempted to identify whether the three instruments that are the focus in this study are related not just in their reaction to news, but actually exhibit a degree of co-movement. The analysis that follows employs some of the more recent innovations in time series analysis, including cointegration, long run structural modeling, vector error-correction modeling, variance decomposition, impulse response analysis and persistence profiling, in order to ascertain possible causal transmission patterns and relative degrees of exogeneity amongst the Short Sterling, Long Guilt and FTSE100.

Prior to application of these procedures, it is first necessary to briefly revise the theoretical foundations on which an analysis of this type is based.

3. Theoretical Underpinnings

What drives stock and bond prices? Furthermore, how is it that the movements in one asset could be reflected by movements in another? In this section we turn to time-honoured theory to address such questions and justify any possible relationships the study may reveal.

3.1.1 Equity Index Futures

A stock index can be regarded as the price of an investment asset that pays dividends. According to Hull (1997: 62), the investment asset is the portfolio of stocks underlying the index, and the dividends paid by the investment asset are the dividends that would be received by the holder of this portfolio. To a reasonable approximation, the stocks underlying the index can be assumed to provide a continuous dividend yield. Therefore if time is T, q is the dividend yield rate, r is the discount factor, and S is the underlying spot value, the equity index futures price F can be given as

$$F = Se^{(r - q)T}$$

As an example, consider a three month futures contract on the FTSE100. Suppose that the stocks underlying the index provide a dividend yield of 3 percent per annum, that the current value of the index is 400, and that the continuously compounded riskless rate is 8 percent per annum. The futures price F is given by

$$F = 400e^{(0.08 - 0.03)0.25} = 405.03$$

3.1.2 Short Sterling Futures

In the Short Sterling futures contract, the underlying asset is the three-month Short Sterling, which is known as a discount instrument. It pays no coupon, and the investor receives the face value at maturity. Prior to the maturity of the futures contract, the underlying asset will have a maturity longer than 90 days. For example, if the futures contract matures in 160 days, the underlying asset is a 250-day Short Sterling. To provide a general analysis we can follow Hull (1997), and suppose that the futures contract matures in T_1 years and the Short Sterling underlying the futures contract matures in T_2 years, where the difference between T_1 and T_2 is three months. We can also assume that R_1 and R_2 are the continuously compounded interest rates for riskless investments maturing at times T_1 and T_2 respectively. If we suppose that the Short Sterling underlying the futures contract has a face value of 100, its present value V, can be given by

$$V = 100e^{-R_2T_2}$$

Since no income is paid on the instrument, the futures price, F, is $e^{R_1T_1}$ times V, or

$$F = 100e^{-R_2T_2}e^{R_1T_1} = 100\, e^{R_1T_1 - R_2T_2}$$

which reduces to $$F = 100e^{-R_F(T_2 - T_1)}$$

where R_F is the forward rate for the time period between T_1 and T_2. The expression shows that the futures price of a Short Sterling is the price it will have if the 90 day interest rate on the delivery date proves to be equal to the current forward rate.

3.1.3 Long Gilt Futures

A Long Gilt futures contract is a contract on a security providing the holder with a known income stream. According to Hull (1997: 120), the futures price, F, is related to the spot price, S, by

$$F = (S - I)e^{rT}$$

where I is the present value of the coupons during the life of the futures contract, T is the time until the futures contract matures, and r is the risk-free rate applicable to a time period of length T.

As this study examines the behaviour of interest rate and equity index futures contracts, the above analysis hopefully makes some inroads in demonstrating to the reader how such contracts are priced. One feature of this analysis can be seen in the way the futures contracts are heavily influenced by the value placed on the underlying asset. As such it is absolutely necessary that we review the fundamental valuation methods of stocks and bonds.

3.1.4 Stock Prices

In an analysis of macroeconomic variables and the stock market, Chen, Roll and Ross (1986) show that stock prices can be written as expected discounted dividends, where c is the dividend stream and k is the required return

$$P = E(c)/(1 + k)$$

If capital gains and non-constant growth are included, the intrinsic value of a stock is the present value of expected future dividends plus the present value of the expected future stock price at the end of the investment horizon h (terminal value)

$$P_o = E(c_1)/(1 + k) \quad + \quad E(c_2)/(1 + k)^2 \quad + \; . \; . \; . \; + \quad E(c_h)+P_h/(1 + k)^h$$

It follows trivially that stock prices are a negative function of the required return and a positive function of expected cash flows.

3.1.5 Bond and Bill Prices

Because coupon paying bonds and stocks are similar in a number of respects, many of the analytical ideas, theories and formulae derived for the stock market can be applied to the bond market. For instance, shareholders may expect to receive a stream of future dividends and may make a capital gain over any given holding period. Coupon paying bonds provide a stream of income (coupon payments) which are usually fixed in nominal terms for all future periods when the bond is purchased. Most bonds, unlike stocks, are redeemable at a fixed date in the future for a known price (face value). The present value of a bond maturing in year h is

$$PV = C_1/(1 + k) \quad + \quad C_2/(1 + k)^2 \quad + \; . \; . \; . \; + \quad (C_h + FV)/(1 + k)^h$$

Note that if the coupon rate is below the market rate of interest , bonds are valued below face value. Over the period of this study bonds traded above face value.

A bill has no coupon payments but its redemption price is fixed and known at the time of issue. Bills are always issued at a discount to the face value so that a positive return is earned over its life, hence they are often referred to as pure discount or zero coupon bonds.

3.2 Information Effects

What information is likely to be relevant for interest rates and stocks? Hardouvelis (1988) argued that newly arriving information can affect interest rates through two channels - either through revisions to expectations about the setting of monetary policy, which he found to dominate movements in short-term interest rates, or through revisions to expectations about inflation, which might dominate long-term interest rates. Given announcements about monetary policy are now explicit in a number of countries including the UK, such announcements might also be important. In addition it was also suggested by Edison (1997) that information about economic activity as well as inflation is likely to be important because it can affect interest rates either directly, by influencing inflationary expectations, or indirectly by encouraging expectations that such news might prompt monetary authorities to adjust interest rates.

Short-term rates are likely to be heavily influenced by expectations about the near-term setting of monetary policy. According to Campbell and Lewis (1998: 8), these expectations might be revised in response to news about monetary policy itself or, more often, about other economic announcements that might influence the policy setting. Alternatively, bonds may be expected to reflect longer-term influences. While temporary changes in monetary policy might be expected to have a smaller effect on bonds rather than short-term rates, markets never can be certain how temporary a change in monetary policy will turn out to be. In practice therefore, short-term and long-term factors are difficult to disentangle. For this reason, as discussed later, the range of news items tested for bonds is the same as for short-term rates. In the spirit of Fleming and Remolona (1997), the current study identifies in advance the news items that might be expected to affect both interest rates, rather than adopting the approach of identifying individual large movements in these instruments and then looking for causal items of news.

As alluded to in section two, the apparently weak informational effects found in the stock market are not entirely surprising. The theoretical effects of macroeconomic announcements are often ambiguous for stocks, but not for bonds. The reason for this discrepancy lies in analysis of the fundamental valuation models for stock and bond prices: whilst stock prices depend on both cashflows and the discount rate, bond prices - for which cashflows are fixed in nominal terms -

depend *only* on the discount rate. For example, an upward revision of expected real activity raises the discount rate for both stocks and bonds, which would reduce prices. However at the same time, the revision raises expected cashflows for stocks, an outcome that would push prices higher. As Fleming and Remolona (1997) suggest, the net effect on bond prices of such an announcement is clearly negative, but the issue for stocks is whether credit conditions are sufficiently tight to depress the economy and subsequently earnings. The net effect on stock prices from announcements depends on whether the cashflow or discount effect dominates.

Now that we have briefly revisited fundamental stock, bond and bill valuation theory, our attention can turn to identifying possible interrelationships common to the three variables in this study: short-term and long-term interest rate and equity index futures.

3.3.1 Inter-market Linkages: The Term Structure

How are interest rate instruments of differing terms-to-maturity related? This concept has been investigated extensively, mainly because an understanding of these dynamic relationships constitutes an indispensable tool for predicting interest rates and possibly because the shape of the yield curve may presage fluctuations in economic activity. Whilst it is not the contention of this paper to undertake an explicit test of the term structure[4], it would be remiss not to acknowledge the implications of one of the term structure theories as it relates to the current analysis.

According to the expectations theory of the term structure, forward interest rates are determined by expectations of the future path of short-term interest rates. In other words, longer maturity interest rates embody expectations of future short rates at all dates up to the maturity of the loan. To the extent that this theory holds, the front (nearest-maturity) Short Sterling futures contract indicates the market's expectation for the level of three-month interest rates at the maturity of the contract. Similarly, the Long Gilt futures contract should reflect average interest rate expectations over the life of the Gilt (ie. ten years). The bond market arbitrage condition states a two year bond should yield the same return as reinvesting the proceeds from a one year bond at the one year rate expected next year. Generalizing this approximation to an n-period bond, it follows that

$$i_{nt} = 1/n\left(i_{1t} + i^{e}_{1t+1} + \ldots + i^{e}_{1t+n-1}\right)$$

So whilst this study does not conduct a formal examination of the expectations theory, changes in the prices of these assets do provide an indication of how the different ends of the yield curve co-move and respond to news releases.

[4]This study uses prices not yields, only one short-term and one long-term instrument, and makes no adjustment for coupon payments. As such, the following cointegration analysis does not lend unambiguous support for term structure theories.

3.3.2 Cross-market Linkages

To date theory is yet to provide us with a model capable of simultaneously capturing the dynamics between bill, bond and stock prices. Instead there are a number of intriguing relationships alluding to the possibility of co-movement.

As demonstrated earlier, it has long been recognized that the value of a share is the present value of the discounted stream of future cash flows. This discount valuation model can also be adjusted to account for the company's free cash flows which are paid out to shareholders. These models require a discount rate, often defined as the shareholders rate of return. This consists of a risk-free rate plus a risk premium (consisting of idiosyncratic and market risk)[5]. Assuming the risk premium is constant, then plummeting bond prices (rising yields) increase the 'severity' of the discounting effect and thus reduce the present value of share prices. Naturally, the opposite holds for rising bond prices. The above discussion highlights that whichever valuation model is used, the yield implied by bill and bond prices is a critical determinant[6]. However this theoretical relationship has not always been supported empirically, as there have been many instances where stocks have risen in a climate of falling bond prices (rising yields), particularly over the short-term.

[5] See Dimson, Marsh and Staunton (2000) for a recent discussion of the role of the equity premium in stock pricing.
[6] For empirical evidence illustrating the strong inverse relationship between earnings growth and bill/bond rates, see Deutsche Banc Alex Brown Topical Study #44 (1999).

One of the reasons this situation may have arisen could be due to the increase in earnings expectations having more than offset the increased discount effect. Another possibility for the temporary breakdown in the theoretical model can be found in the contemporary 'weight-of-money' hypothesis, which states that as global wealth continues to rise, US pension funds (and the like) for example have to put this wealth 'somewhere', and based on the incredibly bullish equity markets in recent times, this somewhere appears to have been in stocks, even during times of falling bond prices (rising yields)(7). Whilst history shows us valuation matters, it also shows us it doesn't matter *all* the time.

This leads us to another concept suggestive of co-movement between bond and stock prices: the 'flight-to-quality' hypothesis frequently proposed by portfolio managers. The flight-to-quality shift from stocks to bonds usually occurs during either economic contractions, or volatile periods in equity markets. The impressive rally in bond prices triggered by the 1987 share market crash is a clear example. Conversely, during times of prosperity and stability, investors often prefer to seek the higher returns offered by stocks. The myriad of relationships discussed above highlight the importance of a formal test for co-movement between our variables.

(7)Naturally, if the equity risk premium falls sufficiently, rising bond yields don't necessarily imply stock prices should fall.

4. Data

The data source for this study is the LIFFE 'Euro-out' tick data made available on CD-ROM, which contains information on nearest second, delivery month, price, transaction code and traded volume. The data is potentially more informative than that used in many previous US studies (eg. Eckman, 1992) because time and sales data from the Chicago Mercantile Exchange and Chicago Board of Trade has only contained bid and ask quotes if the bid quote exceeds or if the ask quote is below the previously recorded transaction price. Also, trades on the CME and CBOT were only recorded if they involved a change in price from the last trade (see Buckle et al, 1998, for further discussion). Such a detailed dataset arguably permits a relatively richer analysis of trading behaviour.

In line with similar UK and US research, the observation window for the study is just under one trading year (243 days), from 1 December 1998, to 18 November, 1999. For the examination of announcement effects, observations are measured in intervals of 5 minutes throughout the day in order to sufficiently capture spikes in volatility. As mentioned earlier, this is an important feature given ap Gwilym and Thomas (1998) found the adjustment process in equity index and interest rate futures markets to be complete within three minutes. End-of-day observations are employed for the cointegration analysis(8).

(8) This was due to an imbalance in matrices arising from different market opening times, and software limitations.

As mentioned earlier, the analysis reported is based on the Short Sterling, Long Gilt and FTSE100 equity index futures contracts. The use of futures contracts can be justified on a number of grounds: these markets have often been established to lead spot markets, possibly due to relatively lower transaction costs and margin requirements; they are transparent and highly liquid; they are very close substitutes for the underlying spot instruments; and it is possible to follow just a single heavily traded contract(9). As the study uses a year of intraday and end of day data, there can be several contracts of differing maturity traded simultaneously. Prices of nearest-maturity (front) contracts are taken, with a switch to the next-maturity contract when its trading volume exceeds that of the front-month. For the Short Sterling and Long Gilt, this rollover typically occurs around 20 days before expiration of the front-month contract, whilst in the case of the FTSE100, rollover occurs on the last trading day of the front-month contract. The three instruments all trade on the London International Financial Futures and Options Exchange (LIFFE), and as the markets are deep and liquid, they provide reliable readings for analysis. The floor trading hours are 08:05-18:00GMT for the Short Sterling contract and 08:00-18:00GMT for the Long Gilt and FTSE100 contracts(10).

(9) Brooke, Danton and Moessner (1999), suggest swap rates may provide an alternative measure to gilts of the market's longer-term interest rate expectations, and are attracting increasing attention given the current level of gilt supply and the impact of the Minimum Funding Requirement on gilt market liquidity.

(10) This provides us with 28 917 five-minute observations for the Short Sterling and 29 160 five-minute observations for the Long Gilt and FTSE100.

Whilst many news items affect financial markets each month, the remainder of this article confines its analysis to regular items of news, which are released on pre-determined dates known to market participants. The news releases which are the focus of this analysis reflect the authors *a priori* view of which announcements are most likely to move the futures markets, and also broadly correspond with US data which is released at 08:30EST, as documented by Ederington and Lee (1993).

Since no single measure of activity is adequately comprehensive or timely, effects are expected from the following releases:

- *retail sales*, (monthly) which provides relatively timely information on one of the largest components of domestic demand;
- *industrial production*, (monthly) and of particular relevance for the stockmarket;
- *unemployment*, as the most timely indicator of the state of the economy each month, and of intrinsic interest for policy;
- *national accounts*, which contains both GDP and Balance of Payments data. Whilst it is the least timely of the news items (released quarterly), it is the most comprehensive measure of economic activity. Also, whilst the Balance of Payments (or more specifically the current account) is not an explicit objective of monetary policy, the belief that it is may linger in some circles;
- *Public Sector Borrowing Requirement,* important for the bond risk premium.

In order to assess the impact of releases related to inflation, and more importantly, the formation of inflationary expectations, the following items are considered:

- the monthly announcement of the official target measure of inflation - *retail prices less interest payments*;
- the monthly *PPI* figure, as this is sometimes thought to impact on the RPI.

Finally, the last two announcements considered to be influential are:

- *changes in UK monetary policy* settings, of which there are seven. These changes consist of five reductions (totaling a fall of 1.75 percentage points), and two increases of a quarter of one percent each;
- *changes in the US federal funds target rate*, consisting of three increases of a quarter of one percent. This release was considered relevant given the widely acknowledged global influence of the US economy, and the fact it is often viewed as something of a barometer for world interest rates.

All macroeconomic announcements are released at precisely 09:30 GMT by the Office for National Statistics. The Bank of England announced all changes in monetary policy at 12 noon (GMT). The Federal Open Market Committee made its base rate change announcements at 14:15 EST (19:15 GMT). During the sample there were 86 items of news on 73 separate days, leaving 170 days when no announcements were made.

5. Methodology

In the initial test of announcement effects in the UK stock and bond markets, a simple OLS regression model is employed. In line with Buckle et al (1998), returns and volatility around announcement times are also modeled in a series of GARCH models, in order to capture the effect of changing volatility in a time series. In a final examination of UK bond and stock market behaviour, a range of recent innovations in time series econometric modeling are applied in an attempt to ascertain the extent to which the three instruments which are the focus of this study exhibit cointegrating relationships.

5.1 Information Content of Announcements

In establishing the relative 'importance' of different items of news, this paper may shed light on whether in fact the UK market systematically differentiates between the information content of announcements. It is anticipated that say the PPI announcement would move the Short Sterling by more than say industrial production, since for a variety of reasons, the information contained in the PPI is of inherently more value to interest rate traders than that contained in industrial production.

In order to capture the impact of individual announcements on the Short Sterling, Long Gilt and FTSE100 index, this paper follows the methodology employed in a landmark US study by Ederington and Lee (1993), by defining the dependent variable in the regressions as the absolute value of the difference between the actual return R_{jt} for the five-minute interval j on announcement day t, and the mean return $\overline{R}_j$ for interval j over all 243 trading days. In summary, the regression format is

$$\left|R_{jt} - \overline{R_j}\right| = a_{oj} + \sum a_{kj} D_{kt} + e_{jt} \tag{1}$$

A series of dummy variables D_{kt} are defined where $D_{kt} = 1$ if announcement k is made on day t and $D_{kt} = 0$ otherwise. Interval j = 09:30 - 09:35 GMT for the six monthly announcements, as well as for the quarterly national accounts release; j = 12:00 - 12:05 GMT for UK base rate changes; and j = 08:00 - 08:05 GMT for the Long Gilt and FTSE100 markets, and 08:05 - 08:10 GMT for the Short Sterling market the morning preceding a change in the US federal funds target rate. This release is made at 14:15 EST - when the London markets are closed - so the first five-minute interval of trade the following morning is used to capture any UK response [11].

[11] The reported effect of US interest rate changes should be treated with a degree of trepidation, as it may be biased by the 'open effect', where volatility is often abnormally high in the first trading session of every day.

As noted in Schwert (1989) and Schwert and Sequin (1990), if log returns are normally distributed with constant mean but time-varying variance, then $E\left|R_{jt} - \overline{R_j}\right| = (2/\pi)^{0.5}\sigma_{jt}$ where σ_{jt} is the standard deviation of returns in interval j on day t. Consequently, $(\pi/2)^{0.5}a_{oj} = 1.2533\ a_{oj}$ provides an estimate of the standard deviation of returns in interval j on nonannouncement days. Whether a particular surprise is good or bad news, a_{kj} should be positive if announcement k impacts the market. The estimated standard deviation of returns in interval j on days when k (and only k) is announced is given by $1.2533\ (a_{oj} + a_{kj})$. If an announcement is ignored by the market, a_{kj} should be approximately zero.

In order to obtain meaningful estimates a_{kj} of the impact of an announcement, it is necessary that the release occur at a consistent time j and not always coincide with another announcement. Very infrequently, two announcements were made simultaneously, and in such cases the dummy variable D_{kt} was set to equal 0.5, as Ederington and Lee (1993) suggest.

As the coefficient a_{kj} provides an estimate of the impact of announcement k on absolute return deviation, it is possible to distinguish the relative 'importance' of each item of news. Campbell and Lewis (1998) point out that in this model, one factor that determines the degree of market reaction to an announcement is how hard it is to forecast the relevant variable. Indicators with large forecasting errors, that is those which tend to be associated with large surprises, might also tend to have large announcement effects, reflected in a large estimated coefficient a_{kj} in the above specification. However it is possible that an indicator may still be important even if it has a very small coefficient. For example, if PPI could be forecast perfectly, there might be no impact of the announcement on the Short Sterling or Long Gilt, yet inflation is of course an important consideration in the operation of monetary policy. Campbell and Lewis (1998) also suggest that if the coefficient for a particular variable were larger than that for another variable, while the forecasting errors of both were the same, then this would be *prima facie* evidence that the former variable contained more new information for the market than the latter.

As recommended by Buckle et al (1998), the estimation of this regression is enhanced with the application of Hansen's (1982) Generalized Method of Moments to ensure robustness to returns autocorrelation and heteroscedastistic errors, both of which are common in high frequency financial market data. Since the system is just identified, the GMM estimates are identical to those from OLS but the standard errors differ. GMM is roundly acknowledged as a robust estimator in that unlike maximum likelihood estimation, it does not require information regarding the exact distribution of the disturbances.

The efficient markets hypothesis implies that only unanticipated news should influence markets, since asset prices should already reflect prevailing market expectations about the economic outlook. As a result, past studies such as Fleming and Remolona (1997) and Campbell and Lewis (1998) have used a survey forecast to separate the announced figure into anticipated and surprise components. The change in interest rates is then regressed on the surprise and the expected change. However Ederington and Lee (1993) document that whilst this procedure captures the impact of announcements on the level of rates, it does not delineate the effect on market volatility, nor does it allow a comparison of the relative importance of various announcements. In addition, most of these announcements contain not one but several statistics which could be informative, and the success of the procedure requires that the forecast be unbiased and accurately reflect market expectations[12].

Now for a final word on the choice between these two methods: whilst the EMH suggests accounting for the unanticipated aspect of an announcement should improve announcement effect estimates, Fleming and Remolona (1997: 34) concede "nonetheless, intraday studies relying on such surprises do not identify more significant news items than do studies relying on announcement dummy variables".

[12] Exhaustive attempts were made initially to obtain relevant forecast data, as it was envisaged the application of both procedures would have made the results more robust. However such data is made available only to large institutional subscribers, and hence this latter procedure was beyond the scope of the paper.

5.2 Time-Varying Variance

The evolution of capital asset pricing models in the 1960's, which related the risk of an asset to its returns, focused attention on the need to model the variance (or volatility) of financial time series. Since then, empirical research has highlighted the tendency for the variance of speculative price series to change through time (ie. be heteroscedastistic), and for squared values of these series to exhibit autocorrelation in the form of volatility clustering. The class of ARCH models, pioneered by Engle (1982), and generalized by Bollerslev (1986), decompose the variance of a series into an unconditional and a time dependent conditional component, which allows them to capture time-varying variance and periods of relative tranquillity and volatility (Bollerslev, Chou and Kroner, 1992).

Building upon the analysis of Buckle et al (1998), this study examines intraday returns and volatility for the Short Sterling, Long Gilt and FTSE100 in a GARCH (1,1) framework. In this model, the conditional variance h is a linear function of past squared errors, e's, past conditional variance, and relevant exogenous variables. The specification applied is given by

$$R_{jt} = \alpha + \beta_1 R_{jt-1} + \beta_2 R_{jt-2} + \sum_{k=1}^{K} \gamma_k D_{kjt} + e_{jt} \qquad (2)$$

where D_{kjt}= 1 if announcement k is made in interval j on announcement day t, and set to 0 otherwise, $e_{jt} \sim N(0,h_{jt})$, and the GARCH (1,1) model defines the conditional variance of the five-minute return R for interval j on day t to be of the form

$$h_{jt} = \delta + \lambda e_{jt-1}^{2} + \omega h_{jt-1} + \sum_{k=1}^{K} \rho_k D_{kjt} \tag{3}$$

As recommended by Buckle et al (1998), the Berndt-Hall-Hall-Hausman (1974) algorithm is used to calculate the maximum likelihood estimates, based on its relative degree of stringency and because it requires stabilized parameter values.

5.3 Interrelationships

Like all other models that utilize time series data, it is important to recognize that unless the analytical tools used account for the dynamics of the relationship within a temporal causal framework, the complexity of the interrelationships involved may not be fully captured. Hence, Masih and Masih (1999) suggest there is a requirement for employing the latest advances in dynamic time series modeling within a temporal causal framework that allows for the co-existence of both short and long-run forces that drive the often ignored deviating and cyclical influences interactive with these variables.

5.3.1 Unit Root Tests

Testing for the presence or (absence) of unit roots in a time series is a critical first procedure when conducting tests for cointegration. At issue is whether shocks have a permanent effect on the Short Sterling, Long Gilt and FTSE100 (in which case they are non-stationary in that they contain a unit root), or only a temporary effect which eventually dies out, implying the variables do not contain a unit root, and as such are considered stationary.

One of the assumptions underlying OLS regression is that the variables are stationary in that the moments of their distribution are constant over time. However, there is overwhelming evidence to suggest that economic and financial time series data typically follow a non-stationary process. As such, it is appropriate to test the data in this study for the existence of a unit root(s).

The most commonly used strategy is to draw on the Augmented Dickey-Fuller (ADF) procedure. This test consists of regressing the first difference of a series against a constant, the series lagged one period, the differenced series at *n* lag lengths and a time trend, that is

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \sum_{i=1}^{n} \beta_i \Delta y_{t-i} + \gamma_1 t + \varepsilon_t \qquad (4)$$

If the coefficient α_1 is significantly different from zero, then the hypothesis that y is non-stationary is rejected. A potential dilemma in the application of this test is determining the lag length to be specified, and also whether to include a constant and/or time trend. McKenzie and Brooks (1999: 14) state that the superiority of one version of the test over another cannot be established *a priori*, and as such, this paper calculates a range of different versions of the test statistic.

A second strategy in testing for the presence of a unit root is provided by the Phillips-Perron (PP) test, in which the appropriate lag length is usually set via the Newey-West procedure (see Hamilton, 1994: 506-16). This test regresses the differenced dependent variable on a constant, lagged dependent variable and a time trend, where

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \gamma_1 t + \varepsilon_t \quad (5)$$

The primary difference between the ADF and PP tests is the method by which each test controls for higher order autocorrelation. Whilst the ADF test corrects autocorrelation through the addition of lagged differenced terms of the dependent variable, the PP procedure , via application of Newey-West, adjusts for autocorrelation and heteroscedasticity.

5.3.2 Cointegration and Granger (Temporal) Causality

The almost universal application of the two-step Engle Granger (1987) procedure prior to the early 1990's is a testimony to its contribution in testing for cointegrating relationships and Granger-causality. In situations where variables are cointegrated - that is they exhibit long run equilibrium relationships if they share common trends - then Granger-causality must exist in at least one direction (see Granger (1969) and Sims (1972) tests). According to Masih and Masih (1999), this also dismisses the possibility of the relationship being 'spurious'. The direction of Granger causality is detected through implementation of the vector error correction model, which in turn is derived from the long run cointegrating vectors.

The Engle-Granger (1987) approach has been widely applied in tests for cointegration, primarily as it has the advantage of being straightforward to implement, relying on a single OLS estimation. However, the more recent Johansen - Juselius procedure possesses a number of advantages over the Engle Granger approach.

These advantages can be summarized as follows (according to Masih and Masih, 1999) : (i) the JJ procedure does not, a priori, assume the existence of at most one cointegrating vector, instead testing for a number of cointegrating relationships; (ii) unlike the Engle-Granger procedure which is sensitive to the choice of the dependent variable, the JJ method assumes all variables to be endogenous; (iii) when extracting the residual from the cointegrating vector, the JJ procedure avoids the arbitrary choice of the dependent variable used in the Engle-Granger approach, and is unresponsive to the variable being normalized; (iv) the JJ procedure is established on a unified framework for estimating and testing cointegrating relations within the VECM formulation; and (v) JJ provide the appropriate statistics and the point distributions to test hypothesis for the number of cointegrating vectors and tests of restrictions upon the coefficients of the vectors.

It is demonstrated in Johansen (1991) that the procedure involves the identification of rank of the *m* by *m* matrix Π in the specification given by

$$\Delta X_t = \delta + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + \varepsilon_t \qquad (6)$$

where X_t is a column vector of the *m* variables, Γ and Π represent coefficient matrices, Δ is a difference operator, *k* denotes the lag length, and δ is a constant. If Π has zero rank, no stationary linear combination can be identified. In other words, the variables in X_t are non-cointegrated. If the rank *r* of Π is greater than zero, however, there will exist *r* possible stationary linear combinations and Π may be decomposed into two matrices α and β, (each $m \times r$) such that $\Pi = \alpha\beta'$. In this representation, β contains the coefficients of the *r* distinct cointegrating vectors that render $\beta' X_t$ stationary, even though X_t is itself non-stationary, and α contains the speed-of-adjustment coefficients for the equation (Masih and Masih, 1999).

5.3.3 Long-Run Structural Modeling

Once the number of cointegrating vectors has been determined, long-run structural modeling endeavours to estimate theoretically meaningful cointegrating relations by imposing identifying and over-identifying restrictions based on theory.

5.3.4 Vector Error-Correction Modeling

Once a number of variables exhibit cointegration, Engle and Granger (1987) found there will always be a corresponding error-correction representation implying changes in the dependent variable are a function of the level of disequilibrium in the cointegrating relationship, captured by the error correction term, as well as changes in other explanatory variables. If we assume variables which trend together in finding a long run stable equilibrium, which are $I(0)$ after applying a first order differencing filter, and the dynamic behaviour of the vector X_t, we may express the unrestricted reduced form of a VAR model as

$$\Delta X_t = \sum_{i=1}^{n} A_i \Delta X_{t-1} + \varepsilon_t \tag{7}$$

where X_t is an $n \times 1$ vector of variables, the A's are estimable parameters, Δ is a difference operator, ε_t is a vector of impulses which represent the unanticipated movements in X_t and $E(\varepsilon_t, \varepsilon_t') = \Omega$ which is diagonal. Provided that the variables in X_t are also cointegrated of order *r*, we may impose the following constraint upon the unrestricted VAR to enable a VECM formulation as follows

$$\Delta X_t = \sum_{i=1}^{n} A_i \Delta X_{t-i} + \sum_{i=1}^{r} \xi_i \Theta_{t-1} + \upsilon_t \tag{8}$$

where Θ contains the *r* individual error-correction terms derived from the *r* long-run cointegrating vectors using the JJ maximum likelihood procedure. Given that there are *r* cointegrating vectors, equation (6) may be reformulated assuming (*n*-1) common trends.

Masih and Masih (1999) demonstrate that a consequence of relationships described by equation (6) when expanded out is that either $\Delta x_{1t},\ldots,\Delta x_{n,t}$, or a combination of any of them must be caused by Θ_{t-1} which is itself a function of $[x_{1t-1},\ldots,x_{nt-1}]$. Intuitively, if $[x_{1t-1},\ldots,x_{nt-1}]$ shares a common trend, then the current change in x_{1t} (say the dependent variable) is partly the result of x_{1t} moving into alignment with the trend value of $x_{2t},\ldots,x_{nt}$ (say the independent variables). Through the ECT, the ECM opens up an additional channel for Granger-causality to emerge which is completely ignored by the standard Granger and Sims tests. The Granger -

causality can be exposed either through the statistical significance of: (i) the lagged ECTs by separate t-test; (ii) a joint test applied to the significance of the sum of the lags of each explanatory variables (A's) in turn, by a joint F or Wald χ^2 test; or (iii) a joint test of all the set of terms just described in (i) and (ii) by a joint F or Wald χ^2 test, ie. taking each of the parenthesised terms separately: the (A's, ξ's). The non-significance of both the t and F or Wald χ^2 tests in the VECM indicates econometric exogeneity of the dependent variable (Masih and Masih, 1999).

As well as indicating the direction of causality amongst variables, the VECM allows short run and long run forms of Granger-causality to be distinguished. When variables are cointegrated, in the short-term, deviations from this long-run equilibrium will feed back on the changes in the dependent variable in order to force the movement towards the long-run equilibrium. If the dependent variable is driven by this long-run equilibrium error, then it is responding to this feedback. If not, it is responding only to short-term shocks to the stochastic environment (Hodgson, Masih and Masih, forthcoming). The F-tests of the 'differenced' explanatory variables give us an indication of the 'short-term' causal effects, whereas the 'long run' causal relationship is implied through the significance or otherwise of the 't' test(s) of the lagged error-correction term(s) which contains the long term information since it is derived from the long run cointegrating relationship(s). The coefficient of the lagged error-correction term, however, is a short-term adjustment coefficient and represents the proportion by which the long-run imbalance in the dependent variable is being corrected in each period. The non-significance of any of the 'lagged error-correction terms' affects the implied long-run relationship and may be a violation of theory. The non-significance of any of the 'differenced' variables, which reflect only short-run relationships, does not however involve such violations since theory typically has little to say about short-term relationships (see Thomas, 1993).

5.3.5 Variance Decompositions (VDCs)

'Out-of-sample' causality tests decompose the variance of the forecast error of variables into proportions attributable to shocks in each variable in the system, including its own (Masih and Masih, 1999). VDC's also provide a literal breakdown of the changes in the value of the variable in a given period arising from changes in the same variable in addition to the changes in other variables in previous periods. Section six demonstrates a variable optimally forecast from its own lagged values will have its forecast error variable accounted for by its own disturbances (Sims, 1982).

5.3.6 Impulse Response Functions

The information contained in the VDCs can be equivalently represented by graphs of the impulse response functions (IRFs). IRFs graphically map out the dynamic response path of a variable arising from a one-period standard deviation shock to another variable (Masih and Masih, 2001).

5.3.7 Persistence Profiles

The persistence profiles estimate the speed with which markets return to equilibrium owing to a system wide shock on the cointegrating relations. Both the persistence profiles and the IRFs map out the dynamic response path of the long-run relations. The distinguishing feature between the two is persistence profiles trace out the effects of a *system wide* shock, while the IRFs trace out the effects of a *variable specific* shock on the long-run relations.

6. Application and Estimation of Results

6.1.1 Announcement Effects Regression: Short Sterling

We can use the regression results from equation (1) to test the hypothesis that short term interest rate expectations should be more responsive to indicators of prevailing economic conditions, while gilt movements should be more responsive to factors that influence long-term inflation expectations and the economy's equilibrium real rate of interest.

Table 2. Regression Tests for the Significance of Announcement Effects

	Short Sterling	
	Coefficient	*(t-statistic)*
Intercept	25.11	*(53.86)*
Ret Sales	**120.16**	***(2.21)***
PSBR	0.46	*(0.02)*
RPI	**72.50**	***(2.27)***
PPI	**26.21**	***(2.10)***
Ind Prod	2.39	*(0.20)*
Unempl	18.11	*(1.05)*
Nat Stats	18.81	*(1.06)*
UK MPC	**272.31**	***(2.28)***
US MPC	**80.92**	***(3.11)***

Results for the estimation of equation (1) with GMM.
All coefficients are multiplied by 10^6.

The results in Tables 2 and 3 unambiguously highlight the volatile response of short-term interest rates to indicators of the current state of the economy. Table 2 shows that no less than five of the nine announcements examined over the course of the study are statistically significant. In descending order of statistical significance, they comprise changes in the US federal funds target rate, UK monetary policy changes, retail price index, retail sales and producer price index news releases. In an explicit measure of volatility, Table 3 reports that on days when the Bank of England announced a shift in monetary settings, the estimated standard deviation of 12:00-12:05 Short Sterling returns is 1.2533 (272.31 + 25.11) (10^{-6}) = 372.76 (10^{-6}) versus 1.2533 (25.11) (10^{-6}) = 31.47 (10^{-6}) on non-release days.

In other words, the Short Sterling's announcement day standard deviation of 12:00-12:05 returns is almost 12 times higher than on days of no announcements. Ranking the other significant announcements in terms of their regression coefficients (in effect measuring how many times volatility is higher than on non-announcement days), retail sales, US interest rate changes, RPI and PPI announcements are all found to be associated with relatively large surges in volatility. Of the statistically insignificant announcements, only the unemployment release appears to have any noticeable impact on volatility, relative to that experienced on days when no announcements are made (1.72 times higher). Perhaps with the exception of the National Accounts release (which includes GDP and Balance of Payments data), these results should prove intuitively plausible given the theoretical analysis in sections 2, 3 and 4.

Table 3. Estimated Standard Deviation of Returns around Announcements

	Short Sterling	
	Est std deviation on announcement days	Times higher vs. non-release days
Ret Sales	**182.07**	**5.79**
PSBR	32.05	1.02
RPI	**122.33**	**3.89**
PPI	**64.32**	**2.04**
Ind Prod	34.47	1.10
Unempl	54.17	1.72
Nat Stats	43.92	1.39
UK MPC	**372.76**	**11.84**
US MPC	**132.89**	**4.22**

All coefficients multiplied by 10^6. Est std dev on non-announcement days is 31.47 (10^6).

6.1.2 Long Gilt

Whilst short-term rates appear to very sensitive to indicators of current economic activity, the results in Table 4 suggest that the same cannot be said for longer-term rates. Only announcements for domestic changes in interest rates and retail sales are statistically significant here. Table 5 shows that on days of domestic rate changes, the estimated standard deviation of 12:00 - 12:05 Long Gilt futures returns is 1.2533 (455.21 + 262.06) (10^{-6}) = 898.96 (10^{-6}), compared with only 1.2533 (262.06) (10^{-6}) = 328.44 (10^{-6}) for non-announcement days. Therefore, the Long Gilt's standard deviation of 12:00 - 12:05 returns is almost three times higher in comparison to days where no news is released.

Table 4. Regression Tests for the Significance of Announcement Effects

	Long Gilt	
	Coefficient	*(t-statistic)*
Intercept	262.06	*(73.27)*
Ret Sales	**379.05**	***(2.76)***
PSBR	75.61	*(0.86)*
RPI	151.32	*(1.56)*
PPI	40.12	*(0.76)*
Ind Prod	39.53	*(0.70)*
Unempl	25.91	*(0.67)*
Nat Stats	17.62	*(0.35)*
UK MPC	**455.21**	***(2.99)***
US MPC	249.33	*(1.34)*

Results for the estimation of equation (1) with GMM.
All coefficients are multiplied by 10^6.

Table 5. Estimated Standard Deviation of Returns around Announcements

	Long Gilt	
	Est std deviation on announcement days	Times higher vs. non-release days
Ret Sales	**803.51**	**2.45**
PSBR	423.21	1.29
RPI	518.09	1.58
PPI	378.72	1.15
Ind Prod	377.98	1.15
Unempl	360.91	1.09
Nat Stats	350.52	1.07
UK MPC	**898.96**	**2.74**
US MPC	640.93	1.95

All coefficients multiplied by 10^6. Est std dev on non-announcement days is 328.44(10^6).

One should note however that this is still considerably less volatile than the reaction of the Short Sterling to the same item of news. Theory may provide us with a couple of possible explanations. Firstly, perhaps a move in the cash rate of 0.25 or 0.5 percent by the Bank of England is not considered by gilt traders to have a substantial impact on the economy's long-run performance. Referring back to the expectations theory of the term structure may provide an alternative explanation. This theory states that the long rate is a geometric weighted average of the current and expected future short rates. Given that the Long Gilt should reflect average interest rate expectations over a decade, and assuming only one or two of the many interest rates incorporated in this model has altered slightly (ie. the current rate), then it is not entirely surprising the Long Gilt remains statistically significant whilst not reacting as severely to base rate changes compared to the Short Sterling.

The other statistically significant release, retail sales, was also characterized by a spike in volatility about 2.5 times higher than that on non-announcement days. Of the statistically insignificant news items, the overnight change in US interest rates saw a jump in volatility about double that experienced on days of no news.

Perhaps conspicuous by absence is the lack of reported impact of inflation announcements on the Long Gilt. Both the RPI and PPI announcements are statistically insignificant and there is little ‘above-average’ surge in volatility immediately preceding their release. As alluded to earlier, an explanation for this result may lie in the structure of the regression model employed (see equation (1)). As the efficient market hypothesis states only unanticipated news should move asset prices, then if both the RPI and PPI can be accurately forecast, then it is conceivable these announcements would have only a minimal impact on prices. The statistical insignificance of both news items does not necessarily imply a diversion from well-established economic theory, but instead may highlight a shortcoming in methodology.

6.1.3 FTSE100 Equity Index

The results presented in Table 6 provide reasonable support for the notion that the theoretical effects of macroeconomic announcements are often ambiguous for stocks, but not for bonds. In these results we find PPI, industrial production and domestic interest rate changes to be statistically significant. However the same regression also shows that the unemployment, PSBR, retail sales and RPI releases are clearly ignored by the stock market (as evidenced by their negative coefficients).

Table 6. Regression Tests for the Significance of Announcement Effects

	FTSE 100	
	Coefficient	*(t-statistic)*
Intercept	822.23	*(52.65)*
Ret Sales	-89.81	*(-0.69)*
PSBR	-112.22	*(-0.79)*
RPI	-33.16	*(-0.23)*
PPI	**1004.13**	***(3.12)***
Ind Prod	**905.32**	***(2.31)***
Unempl	-192.11	*(-1.32)*
Nat Stats	502.12	*(1.21)*
UK MPC	**5031.36**	***(2.27)***
US MPC	439.31	*(1.86)*

Results for the estimation of equation (1) with GMM.
All coefficients are multiplied by 10^6.

The importance of domestic monetary settings is further reinforced by this analysis. On days of UK base rate changes, the estimated standard deviation of 12:00-12:05 FTSE100 returns is 1.2533 (5031.36 + 822.03) (10^{-6}) = 7336.30 (10^{-6}), which is over seven times higher than the standard deviation experienced on non-announcement days. The PPI and industrial production releases are both associated with volatility spikes over twice that experienced on days of no news.

The significance of monetary policy and industrial production announcements are supported on theoretical grounds given both variables form integral components of the share valuation models presented in section 3. However the reported importance of producer prices is not as clear cut. Perhaps a healthy PPI figure could lead investors to believe the economy is expanding at a rate which will continue to boost earnings. Conversely, a robust PPI figure may also result in traders react swiftly in anticipation of a tightening in monetary settings, which could depress the economy, earnings, and subsequently stock prices.

Table 7. Estimated Standard Deviation of Returns around Announcements

	FTSE 100	
	Est std deviation on announcement days	Times higher vs. non-release days
Ret Sales	n.a.	n.a.
PSBR	n.a.	n.a.
RPI	n.a.	n.a.
PPI	**2288.98**	**2.22**
Ind Prod	**2165.14**	**2.10**
Unempl	n.a.	n.a.
Nat Stats	1659.81	1.61
UK MPC	**7336.30**	**7.12**
US MPC	1581.09	1.53

All coefficients multiplied by 10^6. Est std dev on non-announcement days is 1030.5 (10^6).
Results where n.a. is reported imply extreme insignificance via negative coefficients.

6.1.4 Overview of Announcement Effects

So what can we conclude about the responsiveness of financial markets to news releases in the UK? The most intuitively appealing finding is that changes in domestic monetary policy settings are absolutely critical to all three markets which form the focus of this study. In the case of the Short Sterling and Long Gilt, it is the most statistically significant announcement, and for all three instruments, its coefficient is clearly the largest, and hence this release invokes the largest spikes in volatility of all news items considered.

The results show that the Short Sterling and Long Gilt both respond significantly to retail sales, while the PPI announcement is also important for both the Short Sterling and FTSE100. This brings us to an interesting anomaly - why is it that the retail price index is significant only for the Short Sterling market? One reason may be attributable to the notion that producer prices are the leading variable of the two inflation announcements - that is any pick up in inflationary pressures is first reflected in the PPI. Another possible explanation is that because the PPI release was made public earlier in the month relative to the RPI figure in every instance over the course of the study, then traders perceived the majority of price-related information was already known to the market.

Unemployment, PSBR, and the quarterly National Accounts announcements were insignificant for all three markets. As mentioned earlier, the National Accounts release is comprised of GDP and Balance of Payments data. In the case of GDP, there is a three-monthly information cycle in place. In the month immediately preceding the end of a quarter, the Office for National Statistics publishes a release entitled, "GDP: Preliminary Estimates", which is followed by the "UK Output, Income and Expenditure" publication the next month. In the third month following the end of a quarter, the ONS publishes "Quarterly National Accounts". The reason the final National Accounts release appears fairly insignificant for all three markets may be due to the fact relevant information is gradually 'leaked' to the market over a long period of time.

6.2 GARCH Estimation

In a further test of the significance of intraday patterns in returns and volatility, Tables 8, 9 and 10 present the results from the GARCH model defined in section 4. This estimation framework allows the error term variance to be time-varying, so that any inferences drawn are likely to be more robust. Before progressing, it is appropriate at this stage to highlight an important caveat: the following analysis examines intraday pricing behavior - it is not intended to supersede the explicit tests for announcement effects presented above. The specification of the two models are quite distinct as they are testing different concepts.

For the Short Sterling contract, there appears to be strong negative autocorrelation in the two lagged returns, indicating that consecutive five-minute returns tend to have opposite signs, which as noted by Buckle et al (1998), is suggestive of a bid-ask bounce. The results of the mean equation (equation (2)) show significantly large positive returns following changes in domestic monetary settings, and to a lesser extent following changes in US interest rates, as well as announcements for unemployment and the quarterly National Accounts. Statistically significant negative returns are shown to be linked to the PSBR and RPI releases.

Table 8. Results for GARCH Estimation

	Short Sterling			
	Equation (2)		*Equation (3)*	
α	0.16	(0.59)	n.a.	
β_1	-196132	(-30.16)	n.a.	
β_2	-65140	(-10.57)	n.a.	
δ	n.a.		0.0009	(11.89)
λ	n.a.		150799	(11.13)
ω	n.a.		602910	(42.63)
Ret Sales	-2.82	(-0.84)	-0.00009	(-0.77)
PSBR	-54.11	(-4.88)	0.00005	(0.03)
RPI	-22.41	(-2.45)	-0.0002	(-0.28)
PPI	1.65	(0.14)	-0.00004	(-0.02)
Ind Prod	3.63	(0.31)	-0.00004	(-0.03)
Unempl	24.74	(2.29)	-0.00007	(-0.04)
Nat Stats	30.81	(1.96)	-0.00005	(-0.02)
UK MPC	228	(55.35)	0.0002	(1.48)
US MPC	37.12	(4.44)	-0.00003	(-0.04)

All coefficients are multiplied by 10^6. T-statistics in parentheses. Returns are calculated as the natural logarithm of the last transactions price in the current five-minute interval minus the logarithm of the last transactions price in the preceding interval.

The results for the variance equation show the coefficients on both the ARCH and GARCH terms to be positive and significant, hence it appears variance of the error term is conditional on information contained in the previous five-minute periods volatility (ARCH term), and the previous five-minute periods variance (GARCH term). The sum of the coefficients is comfortably below unity, indicating the errors are covariance stationary and that shocks to the system are not overly persistent. This implies forecasts of the conditional variance converge to the steady state at a moderate pace.

In the case of the Long Gilt, there is also evidence suggestive of a bid-ask bounce, as the negative autocorrelation in the lagged returns indicates consecutive five-minute returns have opposite signs. The results for the mean equation show statistically significant positive returns following unemployment releases and changes in domestic interest rates, whilst significant negative returns follow RPI announcements. In the variance equation, again error term variance is conditional on the volatility and variance in the previous five-minute period, and shocks to the system are not overly persistent. In this model, volatility is found to be statistically significant following the release of the release of quarterly National Accounts, unemployment, retail sales and PPI data.

Table 9. Results for GARCH Estimation

	Long Gilt			
	Equation (2)		*Equation (3)*	
α	1.06	(0.28)	n.a.	
β_1	-39777	(-4.37)	n.a.	
β_2	-11086	(-2.14)	n.a.	
δ	n.a.		0.08	(146.49)
λ	n.a.		57422	(27.47)
ω	n.a.		653910	(31.91)
Ret Sales	74.81	(1.14)	0.05	(2.71)
PSBR	-72.72	(-0.92)	-0.03	(-1.08)
RPI	-150	(-1.93)	-0.03	(-1.27)
PPI	-131	(-1.37)	-0.08	(-2.08)
Ind Prod	35.41	(0.36)	-0.06	(-1.74)
Unempl	165	(2.12)	-0.11	(-4.56)
Nat Stats	46.91	(0.78)	-0.12	(-6.19)
UK MPC	245	(2.05)	0.13	(1.68)
US MPC	-298	(-0.70)	0.03	(0.14)

All coefficients are multiplied by 10^6. T-statistics in parentheses. Returns are calculated as the natural logarithm of the last transactions price in the current five-minute interval minus the logarithm of the last transactions price in the preceding interval.

Unlike the two earlier results, in the case of the FTSE100 contract, there is no evidence suggestive of a bid-ask bounce, nor are there any significant trends in returns following announcements. However the ARCH and GARCH parameters are similar to those reported for the Short Sterling and Long Gilt. Volatility appears significant around the time of retail sales data. For a discussion on relevant GARCH model selection criteria and diagnostics, refer to the appendix.

Table 10. Results for GARCH Estimation

	FTSE 100			
	Equation (2)		*Equation (3)*	
α	-141	(-4.59)	n.a.	
β_1	-22292	(-0.97)	n.a.	
β_2	35098	(1.26)	n.a.	
δ	n.a.		2.99	(92.01)
λ	n.a.		123818	(25.72)
ω	n.a.		621296	(46.18)
Ret Sales	144	(0.16)	-3.30	(-2.13)
PSBR	484	(0.59)	-1.25	(-1.07)
RPI	37.51	(0.03)	-2.86	(-1.74)
PPI	-604	(-1.03)	-2.34	(-1.32)
Ind Prod	563	(0.89)	-1.08	(-0.76)
Unempl	52.32	(0.06)	-3.01	(-1.86)
Nat Stats	870	(1.32)	-3.88	(-1.56)
UK MPC	670	(0.93)	0.48	(0.59)
US MPC	648	(0.97)	-5.38	(-0.92)

All coefficients are multiplied by 10^6. T-statistics in parentheses. Returns are calculated as the natural logarithm of the last transactions price in the current five-minute interval minus the logarithm of the last transactions price in the preceding interval.

The results presented in the GARCH (1,1) models bear a slight resemblance to those reported in tables 2-7. However it is difficult to draw direct comparisons as the models employed have an inherently different structure, and as such their focuses are different - the regression model explicitly tested for announcement effects whilst the main thrust of the GARCH models is centered on examining conditional dependencies.

6.3 Interrelationships

In a final examination of patterns in the behaviour of the Short Sterling, Long Gilt and FTSE100, the paper attempts to identify whether these variables exhibit a degree of co-movement, or are in fact exogenous to one another. It is envisaged that the application of some of the more recent innovations in time series econometric modeling will add robustness to any inferences drawn from the results.

6.3.1 Pre-requisites for Cointegration: Unit Root Tests

As discussed in section five, testing for the presence of unit roots is a critical first procedure when conducting cointegration analysis. The section that follows examines the dynamic properties of the aforementioned variables.

Table 11 presents the results of the ADF and PP tests for non-stationarity for the Short Sterling, Long Gilt and FTSE100 in original 'level' form. From this table it is possible to see that each of the six versions of the ADF test clearly fail to reject the null of a unit root for each series at the 5 percent level (ie. the test score is greater than the MacKinnon critical value in each instance). It is also obvious the PP tests tell a similar story. Each of the three PP tests generate coefficients which are greater than the critical test value (at the 5 percent level), and as such, strongly suggest the existence of a unit root in each of the series(13).

(13) As both the ADF and PP tests suffer from low power, it is not uncommon to accept the null even when it is false.

Table 11. ADF and PP Stationarity Testing of Series in Levels

	Short Sterling	Long Gilt	FTSE100
ADF (lag=5)	-0.06 *(-2.57)*	0.03 *(-2.57)*	1.06 *(-2.57)*
ADF with constant (lag=5)	-1.45 *(-3.46)*	-1.62 *(-3.46)*	-2.56 *(-3.46)*
ADF with constant and time trend (lag=5)	-2.11 *(-4.00)*	-2.10 *(-4.00)*	-2.58 *(-4.00)*
ADF (lag=10)	-0.32 *(-2.57)*	-0.20 *(-2.57)*	1.05 *(-2.57)*
ADF with constant (lag=10)	-1.04 *(-3.46)*	-1.18 *(-3.46)*	-3.18 *(-3.46)*
ADF with constant and time trend (lag=10)	-1.65 *(-4.00)*	-1.63 *(-4.00)*	-3.18 *(-4.00)*
PP	0.24 *(-2.57)*	-0.13 *(-2.57)*	0.89 *(-2.57)*
PP with constant	-1.68 *(-3.46)*	-2.14 *(-3.46)*	-2.81 *(-3.46)*
PP with constant and time trend	-2.46 *(-4.00)*	-2.78 *(-4.00)*	-3.00 *(-4.00)*

Critical t-statistics in parentheses. A value greater than the critical t-value indicates non-stationarity.

Given the overwhelming evidence supporting the presence of non-stationarity in the data, convention dictates the taking of differences until stationarity is established. As such, the log price relative was calculated (log returns = $\ln(P_t / P_{t-1})$), which approximates the continuously compounded percentage return for each price series.

Table 12 presents the ADF and PP tests applied to the first-differenced data. The results strongly suggest the series becomes stationary after first-differencing. Each of the test scores were below the critical 5 percent level, and again this result is insensitive to the lag structure or to the presence of an intercept and/or time trend.

Table 12. ADF and PP Stationarity Testing of Series after First-Differencing

	Short Sterling	Long Gilt	FTSE100
ADF (lag=5)	-5.29 *(-2.57)*	-5.15 *(-2.57)*	-8.15 *(-2.57)*
ADF with constant (lag=5)	-5.27 *(-3.46)*	-5.13 *(-3.46)*	-8.24 *(-3.46)*
ADF with constant and time trend (lag=5)	-5.39 *(-4.00)*	-5.23 *(-4.00)*	-8.25 *(-4.00)*
ADF (lag=10)	-5.03 *(-2.57)*	-4.96 *(-2.57)*	-4.44 *(-2.57)*
ADF with constant (lag=10)	-5.02 *(-3.46)*	-4.94 *(-3.46)*	-4.56 *(-3.46)*
ADF with constant and time trend (lag=10)	-5.13 *(-4.00)*	-5.02 *(-4.00)*	-4.60 *(-4.00)*
PP	-16.27 *(-2.57)*	-15.90 *(-2.57)*	-18.49 *(-2.57)*
PP with constant	-16.24 *(-3.46)*	-15.88 *(-3.46)*	-18.54 *(-3.46)*
PP with constant and time trend	-16.41 *(-4.00)*	-16.11 *(-4.00)*	-18.52 *(-4.00)*

Critical t-statistics in parentheses. A value greater than the critical t-value indicates non-stationarity.

6.3.2 Johansen - Juselius (JJ) Cointegration Tests

Having concluded the series in this analysis follow a non-stationary process (in level form), the paper now shifts its focus onto establishing whether stable long-run relationships exist between the Short Sterling, Long Gilt and FTSE100.

Prior to the application of the JJ procedure, it is necessary to first select the optimal order of the VAR. After setting the maximum order of the VAR to 6, Table 13 shows the highest SBC value suggests 2, whilst the AIC selects 5 as the optimal order. Since a reasonably short time series is being investigated, Pesaran and Pesaran (1997: 293) highlight there is a risk of over-parameterization, and as such it is recommended that the SBCs suggestion of 2 as the optimal order of the VAR is implemented.

Table 13. Criteria for Selecting the Optimal Order of the VAR

Order	AIC	SBC
6	3624.1	3530.5
5	*3630.5*	3552.5
4	3617.9	3555.5
3	3620.4	3573.6
2	3617.1	*3588.5*
1	3604.1	3585.9
0	1034.5	1034.5

AIC = Akaike Information Criterion
SBC = Schwarz Bayesian Criterion

The results based on the Johansen-Juselius multivariate tests, provided in Table 14, indicate that the three variables in this study are more likely than not bound together by two long-run equilibrium relationships (ie. r = 2). The maximum eigenvalue statistic rejects the null of no cointegration (r = 0) and the null $r \leq 1$ at the 95 percent critical value. The trace statistic also rejects the null of no cointegration, but can only accept the alternative hypothesis that there exists two cointegrating vectors at the 90 percent level. Both tests unambiguously find no support to suggest three cointegrating relations.

Table 14. Johansen & Juselius Test for Multiple Cointegrating Vectors

Ho:	H1:	*Max Eigenvalue*	95% Crit Value	90% Crit Value
$r = 0$	$r = 1$	*49.66*	21.12	19.02
$r \leq 1$	$r = 2$	*15.19*	14.88	12.98
$r \leq 2$	$r = 3$	*1.75*	8.07	6.50

Ho:	H1:	*Trace*	95% Crit Value	90% Crit Value
$r = 0$	$r \geq 1$	*66.60*	31.54	28.78
$r \leq 1$	$r \geq 2$	*16.94*	17.86	15.75
$r \leq 2$	$r \geq 3$	*1.75*	8.07	6.50

Given at least one, and more likely two, cointegrating vectors appear evident, the finding of no causality in any direction can be ruled eliminated, as can be the existence of 'spurious' correlations. Rogers and Wang (1993) state cointegration rules out the modeling of dynamic relationships through first-differenced ordinary VARs as these models do not impose cointegrating constraints.

The finding of (more likely than not) two cointegrating vectors also implies there will be two residual series and consequently two error-correction terms embedded as exogenous variables appearing in lagged-levels as part of the vector error-correction model (Masih and Masih, 1999). The direction of Granger causality may be detected via this VECM derived from the long-run cointegrating relations[14].

6.3.3 Long-run Structural Modeling

Having reasonably established there to be two cointegrating vectors linking the three variables, the JJ procedure also allows to impose on these vectors just and over-identifying restrictions based on theory. Given that *a priori* relationships between the Short Sterling, Long Gilt and FTSE100 may be considered slightly ambiguous, the study progresses using cointegrating vectors obtained under Johansen's just -identifying restrictions (see Pesaran and Pesaran, 1997, for details).

6.3.4 Vector Error-Correction Modeling

In order to ascertain lead/lag relationships and possible directions of causality the VECM may be employed, given the presence of cointegration does not itself suggest the direction of causality. This is achieved based on the earlier analysis identifying at least one and most likely two cointegrating vectors, which in turn provides us two error-correction terms for constructing models.

[14] This estimation follows Pesaran and Pesaran (1997; 295) in accepting the setting of 'unrestricted intercepts and no trends'. Other combinations were also estimated, and the results proved consistent with those reported above.

The function of the error-correction term (ECT) is primarily to pick up short-run fluctuations before guiding variables back to equilibrium. In situations where the ECT is significant, this term contains additional information than that implied by only lagged changes in the explanatory variables, and has a significant feedback effect on the changes in the dependent variable in order to force temporary deviations back towards long-run equilibrium (Masih and Masih, 1999).Conversely, when this ECT is insignificant, the dependent variable is responding only to short term shocks, not to deviations from long-run relations. The coefficient of this ECT describes the speed and direction of adjustment of each series back to the long-run equilibrium.

The results in table 15 indicate that in the Short Sterling equation in the VECM, the ECT is statistically insignificant (the t-ratio is below two). However in the case of both the Long Gilt and FTSE100, at least one of the error-correction terms is statistically significant. This implies that after an exogenous shock to equilibrium, the Long Gilt and FTSE100 bear the brunt of short term adjustment in order to restore equilibrium. It follows that the Short Sterling is the initial receptor of an exogenous shock , and over the long-run, is not greatly influenced by either the Long Gilt or FTSE100. The Short Sterling appears to be the leading long term variable in the system. With regard to short-term interactions, the Short Sterling also provides significant information leadership to the Long Gilt, whilst the Long

Guilt and FTSE100 play a subdued and subordinated role in information search and dissemination in the short-run.

As discussed previously, the magnitude of the error-correction term coefficient indicates the single period response to a shock. In the case of the Long Gilt, almost 46 percent of the resulting imbalance is corrected within a single day. This speed of adjustment is considerably swifter than in the FTSE100 equation, where less than 11 percent of the disequilibria is corrected within a day.

Table 15. Temporal Causality Results Based on VECM

	Short-Run Lagged Differences			*Error-Correction Terms*	
	▲*LSS*	▲*LLG*	▲*LFTSE*	**ECT1** $[\xi_{i,t-1}]$	**ECT2** $[\xi_{i,t-1}]$
Dep Variable		*F*-statistics		*t*-statistic	
▲*LSS*	-	0.79	-0.81	-0.44	-1.03
▲*LLG*	3.77	-	-1.71	-5.44	1.96
▲*LFTSE*	0.33	-1.14	-	1.72	-3.69

6.3.5 Generalized Variance Decomposition

The VECM, F- and t-tests can be interpreted as within-sample causality tests, as they indicate only the Granger exogeneity (or endogeneity) of the dependent variable within the sample period. VDCs however provide an indication of the dynamic properties of the system, and do allow us to gauge the relative strength of the Granger causal chain and the degree of exogeneity amongst the variables beyond the sample period.

The results in table 16 are based on generalized VDCs. These tests differ from orthogonalized VDCs in that they are unaffected by the ordering of the variables, and the other variables in the system are not switched off when a particular variable is shocked. This is an important attribute because in this field, it is it is quite rare for a series of variables to be 'immune' to shocks in others[15].

Table 16. Generalized Decomposition of Variance

Percentage of Forecast Variance Explained by Innovations in:

		▲ LSS	▲ LLG	▲ LFTSE
Periods	Relative Variance in:			
1	▲ LSS	68.62	31.26	0.12
2		69.22	30.67	0.11
5		70.44	29.33	0.23
10		71.54	27.79	0.67
20		72.28	25.90	1.82
60		72.35	22.98	4.67
Periods	Relative Variance in:			
1	▲ LLG	48.27	51.48	0.25
2		51.81	47.99	0.20
5		59.60	40.28	0.12
10		65.13	28.26	0.15
20		69.07	29.97	0.96
60		71.44	24.41	4.15
Periods	Relative Variance in:			
1	▲ LFTSE	0.36	0.66	98.98
2		0.49	0.79	98.72
5		0.59	1.52	97.89
10		0.50	2.62	96.88
20		0.85	4.29	94.86
60		10.53	8.52	80.95

Note: Figures in the first column refer to horizons (ie. days). All other figures rounded to two decimal places.

[15] In any case, the results for the orthogonalized VDCs were consistent with those reported in Table 16.

The outstanding feature of the decomposition analysis can be seen examining the highly exogenous nature of the FTSE100 amongst the trivariate system. It appears that after 1 trading month (20 days), approximately 95 per cent of its variance is still explained by its own shocks, and so it stands the variances of the Short Sterling and Long Gilt contribute very little. In the case of the Short Sterling, it also appears to be characterized by a degree of exogeneity, although the variance in the Long Gilt accounts for around 20-30 percent following a shock. Variance in the Long Gilt is increasingly being explained by the variance in the Short Sterling.

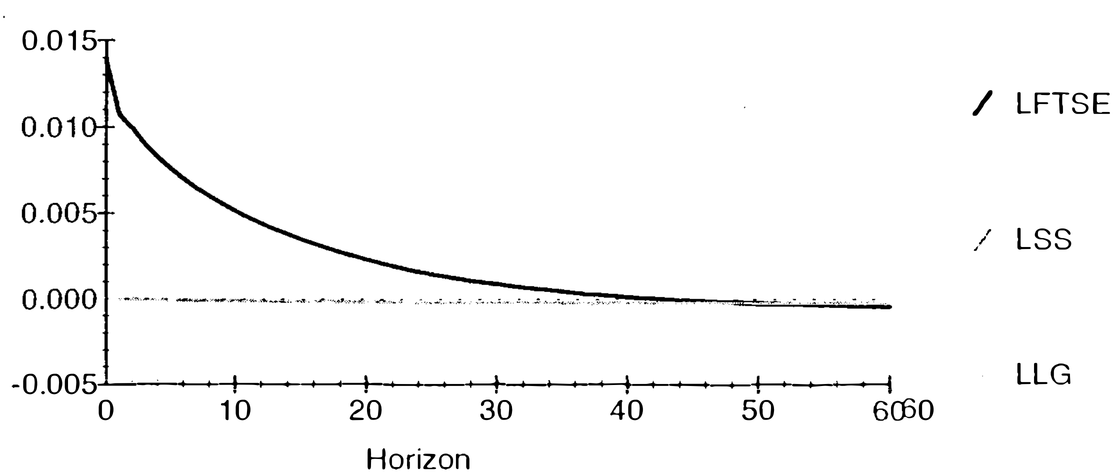
The results of the VDCs do not precisely resemble those of the error-correction model, although it must be pointed out that both procedures are actually testing different concepts over different sample periods. The analysis here suggests that the FTSE100 share index is the most exogenous variable within the trivariate system, followed by the Short Sterling. However both procedures confirm that the Short Sterling is exogenous and has an influential impact on the Long Gilt, and so it may be inferred these results lend a degree of support to the expectations theory of the term structure of interest rates (bearing in mind the limitations previously acknowledged). Whilst the FTSE100 is not particularly influenced by either short term or long term interest rates, conversely, the results also confirm neither the Short Sterling nor the Long Gilt are particularly vulnerable to movements in the FTSE100 share index.

6.3.6 Generalized Impulse Response Analysis

Another method of representing the findings of the variance decomposition analysis is through impulse response functions, which are designed to map out the dynamic response path of a variable arising from a one-period standard deviation shock to another. In essence, impulse response functions portray the extent to which the shocking of one variable has a persistent effect on the other variables in the system. Figure 1 below illustrates the generalized impulse responses of the Short Sterling and Long Gilt to shocks in the FTSE100 share index depict a negligible impact, in line with the above discussion. The GIRF also highlights all three financial instruments converge to 0 after the effect of the shock dies away, approximately two and a half months (50 days) later.

Figure 1.

Generalized Impulse Responses to one S.E. Shock in the Equation for the LFTSE100

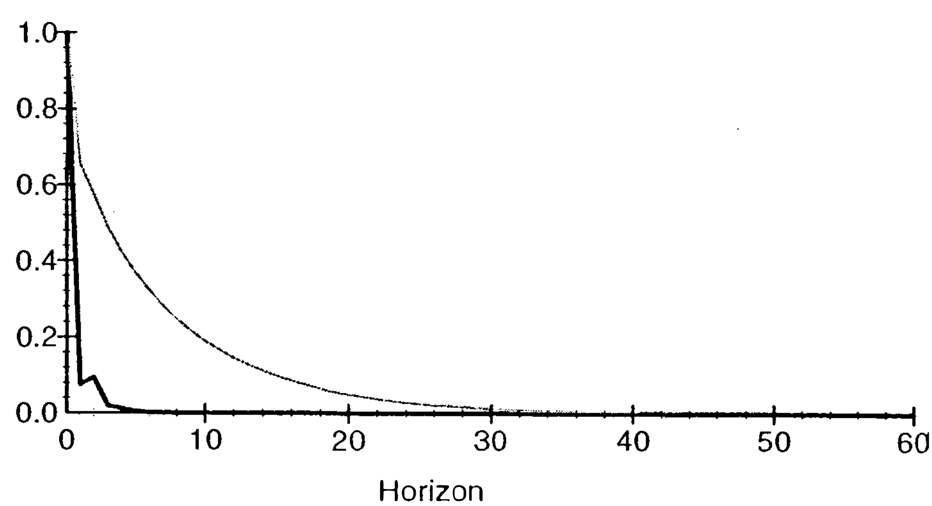


6.3.7 Persistence Profiles

Persistence Profiles map out the speed with which the economy or markets return to equilibrium owing to a system wide shock on the cointegrating relations (IRFs which trace out the effects of a variable specific shock on long-run relations). Figure 2 shows whilst both cointegrating relations have a strong tendency to converge to their respective equilibria, the speed of this adjustment does vary between vectors.

Figure 2.

Persistence Profile of the Effect of a System Wide Shock on the Cointegrating Vectors



7. Conclusions

This paper has covered considerable ground and touched on a wide range of issues. The most important conclusions and contributions of this investigation into the behaviour of the UK interest rate and stock index futures markets follow below.

A number of interesting inferences can be derived from the tests for announcement effects. Firstly, The reaction of markets to macroeconomic data suggests investors and portfolio managers distinguish between the information content of different news items, in line with previous central bank studies in the US (see Fleming and Remolona, 1997) and Australia (see Campbell and Lewis, 1998).

Secondly, the paper finds considerable support of the hypothesis that short term interest rate expectations are highly sensitive to indicators of prevailing economic conditions, as evidenced by the sharp reaction to announcements of changes in UK and US interest rates, retail prices, retail sales and producer prices. As the Long Gilt responded significantly to changes in domestic interest rate and retail sales announcements, only moderate support can be extended to the notion that longer-term interest rates respond to factors influencing long-run inflation expectations and the economy's equilibrium rate of interest.

Thirdly, the results also imply the effects of macroeconomic announcements to be somewhat ambiguous for the stock market. For instance, whilst FTSE100 traders responded significantly to changes in domestic monetary policy settings, industrial production and PPI announcements, four of the remaining items of news appeared to be completely ignored (shown by the negative coefficients).

Taken in sum, the results suggest some consistency in how the three markets respond to announcements, and reassuringly, in the vast majority of cases, these responses should prove intuitively appealing to both market analysts and economic theorists alike.

The main feature of the GARCH estimation of intraday returns and volatility is the extent to which error term variance is dependent on the information contained in the variance and volatility in the previous five-minute interval. Another feature common to all three contracts is that volatility shocks displayed no real signs of persistence, which may be considered slightly unusual in high frequency data. In another interesting aspect of the GARCH analysis, both the Short Sterling and the Long Gilt were characterized by consecutive five-minute returns of opposite signs (negative autocorrelation), which is suggestive of a bid-ask bounce. However no such pattern emerged for the FTSE100.

In a final examination of patterns in the pricing behaviour of the Short Sterling, Long Gilt and FTSE100 futures contracts, the paper attempted to establish whether these variables exhibited a degree of co-movement. Initially it appeared the three financial instruments were in fact bound by at least one, but most likely two, cointegrating relationships. The (within sample) error-correction analysis identified the Short Sterling as the leading variable amongst the trivariate system, however the variance decompositions suggested that beyond the sample period, the FTSE was actually the most exogenous variable, followed by the Short Sterling. The ambiguity surrounding causality directions and relative degrees of exogeneity was not entirely unexpected, given theory is yet to provide us with an accepted framework capable of simultaneously capturing the dynamics between stock and bond prices. However, the most robust finding emanating from the tests for interrelationships appeared to be consistent with the expectations theory of the term structure of interest rates.

This leads us onto possible areas of future research. Any practicing macroeconomist worth their salt would overwhelm an interested onlooker with ideas behind the relationship between stock and bond prices. However, whilst cross-market correlation structures and excess return relationships have been modeled extensively, theory has relatively little to say about these dynamics, and as such, this disparity highlights an interesting opportunity for future empirical work.

Appendix

- Prior to the estimation of the GARCH (1,1) models, it was first established that ARCH effects were present in the data. This was achieved through a visual examination of volatility clustering, the significance of the Ljung-Box Q-statistic, and finally with Engle's (1982) LM test (see McKenzie and Brooks, 1999).
- No ARMA terms were fitted to the mean equation for two reasons: not only did it prove incredibly difficult to adequately capture correlation structures, but once such structures were finally accounted for, it was found this had virtually no impact on the GARCH estimations. However this is not uncommon, as some of the leading exponents of ARCH models also conclude that the application of AR/MA terms have no real impact on models estimated in continuous time (see Nelson, 1990a, 1990b, and Gannon, 1996a, 1996b), or in discrete time (McKenzie,1997).
- A number of ARCH (p), GARCH (p,q) and TARCH (p,q) specifications were found to meet accepted criteria (that is): the models converged; the t-statistics on the coefficients were significant and positive; and the ARCH and GARCH parameters summed to less than unity. In such cases, McKenzie and Brooks (1999) suggest choosing the optimal model specification by referring to the lowest AIC and SBC statistic, as well as the highest R^2 value. All fitted models adequately captured ARCH effects, and were adjusted for non-normal errors and leverage effects.

References

Becker, K., Finnerty, J. and Kopecky, K. (1993), 'Economic News and Intraday Volatility in International Bond Markets', *Financial Analysts Journal*, May/June, p.81-86.

Becker, K., Finnerty, J. and Kopecky, K. (1995), 'Domestic Macroeconomic News and Foreign Interest Rates', *Journal of International Money and Finance*, 14, (6), p.763-83.

Becker, K., Finnerty, J. and Kopecky, K. (1996), 'Macroeconomic news and the Efficiency of International Bond Futures Markets', *Journal of Futures Markets* 16,(2), p.131-45.

Berndt, E., Hall, B., Hall, R., and Hausman, J. (1974), 'Estimation and Inference in Nonlinear Structural Models', *Annals of Economic and Social Measurement*, 3, p.653-65.

Bollerslev, T. (1986), 'Generalized Autoregressive Conditional Heteroskedasticity', *Journal of Econometrics*, Vol. 31, p.307-27.

Bollerslev, T., Chou, R., and Kroner, K. (1992), "ARCH Modeling in Finance: A review of the Theory and Empirical Evidence', *Journal of Econometrics*, 52,p.5-59.

Brooke, M., Danton, G. and Moessner, R. (1999), 'News and the Sterling Markets', *Bank of England Quarterly Bulletin*, November, p.374-83.

Buckle, M., ap Gwilym, O., Thomas, S. and Woodhams, M. (1998), 'Intraday Empirical Regularities in Interest Rate and Equity Index Futures Markets, and the Effect of Macroeconomic Announcements', *Journal of Business, Finance and Accounting*, 25 (7) & (8), September/October, p.921-44.

Campbell, F. and Lewis, E. (1998), 'What Moves Yields in Australia?', Reserve Bank of Australia Research Discussion Paper no. 9808.

Clare, A. and Courtenay, R. (2000), 'Financial Market Reactions to Interest Rate Announcements and Macroeconomic Data Releases' *Bank of England Quarterly Bulletin*, Vol 38(3), p.266-73.

Chen, N., Roll, R., and Ross, S. (1986), 'Economic Forces and the Stock Market', *Journal of Business*, 59, p.383-403.

Cutler, D., Poterba, J. and Summers, L. (1989), 'What Moves Stock Prices?', *Journal of Portfolio Management*, 15, p.4-12.

Dale, S. (1993), 'The Effect of Changes in Official UK rates on Market Rates since 1987', *The Manchester School*, Vol LXI, p.76-94.

Dimson, E., Marsh, P. and Staunton, M. (2000), 'Risk and Return in the 20th and 21st Centuries', *Business Strategy Review*, Summer.

Eckman, P. (1992), 'Intraday Patterns in the S&P Index Futures Market', *Journal of Futures Markets*, 12, p.365-81.

Ederington, L. and Lee, J. (1993), 'How Markets Process Information: News Releases and Volatility', *Journal of Finance* 48, (4), p.1161-91.

Edison, H. (1997), 'The Reaction of Exchange Rates and Interest Rates to News Releases', *International Journal of Finance and Economics*, 2(2), p. 87-100.

Elmendorf, D., Hirschfeld, M. and Weil, D. (1994), 'The Effect of News on Bond Prices: Evidence from the United Kingdom, 1900-1920', *Review of Economics and Statistics* (78), p.341-4.

Engle, R. (1982), 'Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of United Kingdom Inflation', *Econometrica*, Vol. 14, p.987-1007.

Engle, R. and Granger, C. (1987), 'Co-integration and Error Correction: Representation, Estimation and Testing', *Econometrica* 55, p.391-407.

Fleming, M. and Remolona, E. (1997), 'What Moves the Bond Market', *Federal Reserve Bank of New York Economic Policy Review*, 3(4), p. 31-50.

Gannon, G. (1996a), 'Unconditional First and Conditional Second Moment Effects: Index Portfolios and Index Futures', *Research in Finance Supplement to 1996 Eds. Chen, A. ,and Chan, K. p.143-58.*

Gannon, G. (1996b), 'Conditional Second Moment Effects. Australian Evidence: Index Futures and Stock Price Processes', *Paper presented at 1996 APFA/PACAP Finance Conference and CFA Annual Meetings, July, 1996 Taipei.*

Goodhart, C. and Smith, R. (1985), 'The Impact of News on Financial Markets in the United Kingdom', *Journal of Money, Credit and Banking*, 17, 4, p.507-11.

Granger, C. (1969), 'Investigating Causal Relations by Econometric Models and Cross Spectral Methods', *Econometrica*, 37, p.424-38.

ap Gwilym, O. and Thomas, S. (1998), 'Which Macroeconomic Announcements Have Most Impact on Long Gilt Futures Market Volatility', *Derivatives Use, Trading and Regulation*, 4, (2), p.164-73.

ap Gwilym, O., Buckle, M., Clare, A., and Thomas, S. (1998), 'The Transaction-by-Transaction Adjustment of Interest Rate and Equity Index Futures Markets to Macroeconomic Announcements', *Journal of Derivatives*, Winter, p.7-17.

Haldane, A. and Read, V. (1999), 'Monetary Policy and the Yield Curve', *Bank of England Quarterly Bulletin*, Vol 34(3), p.232-4.

Hamilton, J. (1994), 'Time Series Analysis', Princeton University Press, Princeton.

Hansen, L. (1982), 'Large Sample Properties of Generalized Method of Moments Estimators', *Econometrica*, 50, p.1029-54.

Hardouvelis, G. (1987), 'Macroeconomic Information and Stock Prices', *Journal of Economics and Business*, 39, p.131-40.

Hodgson, A., Masih, A. and Masih, R. (forthcoming), 'Price Discovery Between Informationally Linked Markets During Different Trading Phases: Evidence from Australia', *Journal of Financial Research.*

Hull, J. (1997), 'Introduction to Futures and Options Markets', Prentice-Hall, New Jersey.

Johansen, S., and Juselius, K. (1990), 'Maximum Likelihood Estimation and Inference on Cointegration with Applications to Money Demand', *Oxford Bulletin of Economics and Statistics*, 52, p.169-210.

Joyce, M. and Read, V. (1999), 'Asset Price Reactions to RPI Announcements', *Bank of England Quarterly Bulletin*, No 94.

Masih, A. and Masih, R. (1999), 'Are Asian Stock Market Fluctuations Due Mainly to Intra-regional Contagion Effects?' Evidence Based on Emerging Stock Markets', *Pacific-Basin Finance Journal*, 7 (3-4), p.251-82.

Masih, A. and Masih, R. (2001), 'Long and Short Term Dynamic Causal Transmission Amongst International Stock Markets', *Journal of International Money and Finance*, Vol. 20 (4), August.

McKenzie, M., and Brooks, R. (1999), 'Research Design Issues in Time-Series Modeling of Financial Market Volatility', McGraw-Hill, Sydney.

McKenzie, M. (1997), 'ARCH Modeling of Australian Bilateral Exchange Rate Data', *Applied Financial Economics*, 7, p.147-74.

Nelson, D. (1990a), 'Stationarity and Persistence in the GARCH (1,1) Model', *Econometric Reviews*, 6, p.318-34.

Nelson, D. (1990b), 'ARCH Models as Diffusion Approximations', *Journal of Econometrics*, 45, p.7-38.

Nelson, D. (1991), 'Conditional Heteroskedasticity in Asset Returns: A New Approach', *Econometrica*, 59, p.347-70.

Pearce, D., and Roley, V. (1985), 'Stock Prices and Economic News', *Journal of Business*, 58, no.1, 49-67.

Pesaran, M., and Pesaran, B. (1997), 'Microfit 4.0', *Oxford University Press*, Oxford.

Prag, J. (1994), 'The Response of Interest Rates to Unemployment Rate Announcements: Is There a Natural Rate of Unemployment?', *Journal of Macroeconomics*, 16, no.1, p.171-84.

Rogers, J. and Wang, P. (1993), Sources of Fluctuations in Relative Prices: Evidence from High Inflation Countries', *Review of Economics and Statistics*, 75(4), p.589-605.

Schwert, G. (1981), 'The Adjustment of Stock Prices to Information about Inflation', *Journal of Finance*, 36, no.1, p.15-29.

Schwert, G. (1989), 'Why Does Stock Market Volatility Change Over Time?', *Journal of Finance*, 44, p.1115-1153.

Schwert, G., and Sequin, P. (1990), 'Heteroscedasticity in Stock Returns', *Journal of Finance*, 45, p.1129-1155.

Sims, C. (1972), 'Money, Income and Causality', *American Economic Review*, 62, p.540-52.

Thomas, R. (1993), 'Introductory Econometrics: Theory and Applications', Longman, London.

Zakoian, J. (1990), 'Threshold Heteroskedastic Models', *manuscript, CREST, INSEE, Paris.*