# Have markets reacted differently to macroeconomic announcements since 1997? An empirical analysis of UK intraday trades and prices

*Oliver Burrows**
*and*
*Anne Vila Wetherilt***

August 2004

* Monetary Instruments and Markets Division, Monetary Analysis, Bank of England, Threadneedle Street, London, EC2R 8AH.
E-mail: O*liver.burrows*@bankofengland.co.uk
** Monetary Instruments and Markets Division, Monetary Analysis, Bank of England.
E-mail: *Anne.vila-wetherilt* @bankofengland.co.uk

The views expressed in this paper are those of the authors, and not necessarily those of the Bank of England. We thank Richhild Moessner, Paul Robinson and Peter Westaway and seminar participants at the Bank of England for their comments, and Lorna Pickford and Martin Owen for their assistance in preparing the data.

**Contents**

**Abstract**

In this paper, we examine the high-frequency reaction of the yield and trading volume of a short-term interest rate futures contract to the release of macroeconomic indicators, in an attempt to find evidence of increased monetary policy transparency since the Bank of England was granted independence in June 1997. We put forward two hypotheses; first, that increased transparency should have led to a larger change in price following macroeconomic news, and, second, that increased transparency should lead to prices incorporating macroeconomic news more quickly. We use simple OLS regressions to test the reaction of both the yield and the trading volume over periods of one to sixty minutes. Consistent with earlier studies, we find evidence to reject the first hypothesis, with the size of the change in yield response to macroeconomic news appearing, if anything, a little smaller post-Bank independence. By contrast, we find some evidence in favour of the second hypothesis, with price adjustment in the Short Sterling market appearing to take place more quickly post-Bank independence.

Key words:
*JEL* classification:

## 1 Introduction

This paper extends and updates earlier work on the reaction of UK financial markets to economic news. Specifically, the paper examines the reaction of short and long-term interest rate futures to a range of macroeconomic indicators. The response of market interest rates to macroeconomic announcements should be of particular interest to monetary policy makers as it may contain information about market perceptions of the policy maker's reaction function. Changes over time in the reaction of interest rates to macroeconomic announcements may be of further interest, as they may suggest that the market's understanding of monetary policy has changed. With this motivation, we look for evidence of a changed reaction of interest rates following the granting of independence to the Bank of England in 1997. We are particularly interested in finding evidence of increased transparency.

Monetary theory tells us that increased central bank transparency affects the behaviour of interest rates along the entire yield curve. This follows from the commonly-held view that interest rates embody market participants' beliefs about future policy rate decisions. Greater transparency, it is argued, provides market participants with more information, both about the central bank's policy objectives and the process of decision making. Theoretical models also show that interest rate expectations are affected by the market's knowledge of the central bank's policy objectives and by their belief that these objectives will not be altered in the short-run. In other words, the behaviour of interest rates is related to central bank credibility.

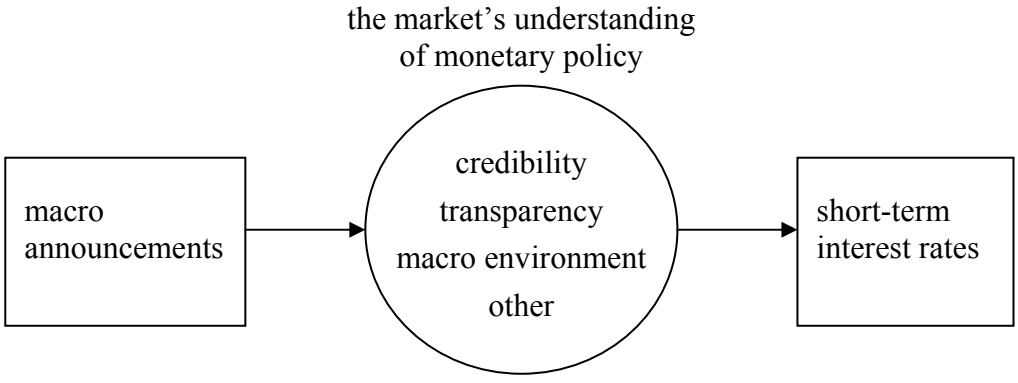
Haldane and Read (2000) models these relationships and show that if markets have a greater understanding of the central bank's objectives and are more confident that these objectives will be adhered to, then monetary policy announcements should cause less surprises, as market participants will be able to forecast them with greater accuracy. Moreover, if market participants understand the nature of the macroeconomic variables that affect the central bank's policy decisions, then they will most likely pay particular attention to news releases about these variables. To the extent that they contain new information, the release of macroeconomic data prior to policy meetings should therefore cause more pronounced reactions in financial markets.

Previous research has attempted to empirically infer market views on transparency and credibility from the reaction of market interest rates to economic announcements. Clare and Courtenay (2001) and Lasaosa (2004) examined whether changes in the United Kingdom's monetary policy framework and the accompanying move towards greater transparency had indeed altered UK

markets' response to key economic announcements. They found that since the Bank of England's independence in 1997, prices reacted less to macroeconomic announcements. Moreover, they found no difference in the response of prices to policy rate changes. Asking a very similar question, Gravelle and Moessner (2001) found that Canadian interest rates in the period 1995-2000 reacted very little to domestic macroeconomic announcements, but substantially more to US releases. Moreover, they observed that in spite of greater transparency in the Bank of Canada's policy process, the sensitivity of interest rates to news about the domestic economy that would be expected to enter the Bank of Canada's policy decision had not increased over time.

In this paper, we examine whether the reaction of market interest rates to macroeconomic announcements has changed since Bank independence. Before stating the hypotheses for our empirical tests, it is useful to look at Figure 1, which describes the earlier-discussed relationships between market interest rates and monetary policy. In Figure 1, the reaction of interest rates to macroeconomic announcements is shown to be determined by a range of factors, including transparency, credibility and the macroeconomic environment. If the behaviour of interest rates has changed since 1997, then the figure shows that this could be attributed to one or several factors. The present paper designs a series of tests to quantify changes in the reaction of interest rates to macroeconomic news, and uses the results to make inferences about changes in transparency. Figure 1 shows, however, that in interpreting these test results, one needs to be mindful that they could have been affected by the remainder factors, not addressed in the paper.

**Figure 1**




With these observations in mind, we turn to the literature on transparency to formulate our theoretical priors. Our first hypothesis, based on Haldane and Read (2000), states that prices

should react *more* to macro announcements post-BI if market participants have a better understanding of the Bank's reaction function. It is also the main hypothesis in Clare and Courtenay (2001) and Lasaosa (2004). For future reference, we call this our *macro hypothesis.* Our second hypothesis is concerned with price volatility and trading volume. It again postulates that with increased transparency market participants are likely to be better informed about the monetary policy process. This in turn means that fewer investors are likely to have distinct information advantages. Drawing on market micro structure theory, we suggest that such a reduction in information asymmetries would lead to *larger* immediate volume and price volatility reactions on announcement days, and a quicker adjustment process thereafter. For convenience, we label this the *micro hypothesis.*

In a departure from Clare and Courtenay (2001) and Lasaosa (2004), but more in line with intraday event studies of US bond and equity markets, we use simple regression techniques to test for evidence of transparency. This allows us to incorporate additional information in the form of the size of the surprise in macro indicators, allows us to test for the reaction of indicators that are released at the same time and allows us to better control for other factors. In extensions of our regression specification, we consider the reaction to foreign indicators and control for the effect of uncertainty about future levels of interest rates on the reaction to macro indicators. We also extend our analysis to the United States, in an attempt to capture any wider changes in market behaviour which could bias our UK tests. The empirical tests are carried out for two UK futures contracts, Short Sterling and Long Gilt, for the period 1994-2003. We pay most attention to the reaction of Short Sterling futures contracts to macroeconomic indicator announcements, as this provides the clearest test of our two hypotheses. Finally, taking note of the existing event study literature, we focus on the surprise component of macroeconomic indicators, rather than looking at the simpler reaction that occurs at the time of the announcement.

Our main results are as follows. We find mild evidence of a reduction in price change of Short-Sterling futures to the surprise component of UK indicators post-BI. These results are unaltered when accounting for changes in market uncertainty. Yields also appear to react a little less to US indicators in the post-BI period. In terms of our hypotheses on transparency, this is evidence against the *macro* hypothesis. These results are consistent with Clare and Courtenay (2001) and Lasaosa (2004).

In our tests of the price adjustment dynamics, we observe that price volatility reacts a little more to domestic indicators post-BI in the minutes immediately following news releases, and that the

subsequent price adjustment appears to take place more quickly. These results are supportive of our *micro* hypothesis, as traders appear more inclined to respond to the arrival of news, and seem to reach an earlier consensus on the likely impact of this news on future monetary policy. The effects on trading volume, however, are more difficult to asses because of changes in trading practices during our sample period.

The remainder of the paper is organised as follows. Section 2 considers the Clare and Courtenay (2001) and Lasaosa (2004) hypothesis and sets out our own hypotheses. Section 3 offers a brief literature overview, considering studies of announcement effects as well as empirical microstructure work. We describe our data and methodology in Sections 4 and 5 respectively. We then present our results, first testing the *macro* hypothesis (Sections 6 and 7) and second the *micro* hypothesis (Section 8). Section 9 concludes.

## 2 The impact of news on financial assets and its relation to transparency

In this section, we outline the hypothesis used in Clare and Courtenay (2001) and Lasaosa (2004) and detail the revisions we make to it in testing for evidence of increased transparency. Clare and Courtenay (2001) and Lasaosa (2004) explicitly use the hypothesis that the reaction to macroeconomic indicators should be higher and the reaction to monetary policy lower post Bank independence (BI), if monetary policy has become more transparent. Both papers make clear, however, that their tests are not very powerful – the results of the tests may or may not reflect a change in transparency, but other factors. We attempt to improve the strength of the tests by refining the hypothesis.

*Macroeconomic news and interest rate reactions*

Clare and Courtenay (2001) and Lasaosa (2004) hypothesise that with increased transparency, market participants should pay more attention to macroeconomic announcements, thereby causing more pronounced reactions in asset prices. They argue that in a regime where market participants have a clear understanding of the central bank's objective function, they are also more capable of assessing which macroeconomic news is most likely to affect policy decisions. As a result, when macro data are released, market participants are likely to pay close attention to them and react strongly to information that they believe to be relevant. This hypothesis is consistent with theoretical signal extraction models, in that data releases can be viewed as public

signals of monetary policy. The policy maker's reaction function becoming clearer can be thought of as a reduction in noise in the signal, which causes agents to become more sensitive to the signal and thus to react more to the data announcements than previously.[1] In what follows, we refer to the above hypothesis as the *macro hypothesis*.

Lasaosa (2004) also presents a number of reasons why financial markets' reaction to macro announcements may *not* have increased, in spite of greater transparency. First, she argues that the MPC's committee structure may make it harder for market participants to predict policy rate decisions and correctly assess macroeconomic developments, as the weight committee members attach to these data may differ. Second, she suggests that interest rate decisions are made within a framework that allows for considerable flexibility on the part of the MPC (who can be said to use an implicit, rather than an explicit rule), and are for this reason harder to anticipate.

One could also argue that monetary policy has become more predictable because of the practice adopted by many central banks in the past decade to make small, gradual changes in their policy rates, rather than large and frequently-reversed changes. As a result market participants may be able to anticipate, not only the next policy rate change, but also the future path of short-term interest rates. This implies that when macroeconomic news arrives, market participants may assess its potential impact on the entire path of future rates, and may therefore choose to react less. Lildholdt and Vila Wetherilt (2004) show that UK monetary policy is indeed characterised by small frequent changes, and that predictability has increased, but these changes are shown to predate Bank independence.

The *macro hypothesis* is concerned with the amount of change in prices (or yields) observed following macro announcements, but remains silent on the dynamics of the actual adjustment. It is possible, however, that improvements in transparency significantly affect the manner in which prices incorporate the newly released information, in addition to their impact on the actual price change. For this reason, we bring an additional set of questions to the data. Do prices react more quickly than before? Are these price changes greater? Are there more trades and price changes for a given trade interval? In the next section, we outline a second hypothesis showing how the dynamics describing prices and trading volume can be related to advances in transparency.

---

[1] See e.g. Geraats (2001).

*Macroeconomic news, interest rate volatility and trading volume*

Market microstructure literature provides a number of interesting theories, relating information flows, price volatility and trading volume.[2] On the whole, this literature shows that the increased availability of information reduces both price volatility and volume, although there are models that contradict this result. In this section, we examine whether these theories could explain the volatility and volume reactions in the United Kingdom since Bank Independence.

The first model to tackle the volume-volatility relationship is the so-called Mixture of Distributions hypothesis. Introduced by Clark (1973), this statistical model states that both volume and price volatility are driven by the same latent variable, namely the number of so-called information arrivals. These information flows induce both trading and price volatility. For the Mixture of Distributions hypothesis to be relevant in our context, we must accept that agents perceive macro announcements to contain *more* information after Bank independence. If this were the case, then we would indeed expect trading volume and volatility to react *more* to macro announcements post Bank independence.

More recent market microstructure theory models the reaction of traders to information flows, thereby endogenising price volatility and trading volume. Two strands of this literature are worth considering. First, a large number of models consider the impact of asymmetric information on prices and trades. In the absence of information asymmetries, prices immediately and fully reflect all new information, and no trading occurs. In the presence of information asymmetries, however, informed and uninformed traders behave differently, generating distinct patterns in volume and volatility. Driving these patterns is the concept of adverse selection risk. This risk arises because less informed traders are at a distinct disadvantage when trading with more informed counterparts. But, since they are often unaware of their counterparty's identity, uninformed traders are at risk of trading with a better informed investor, thereby accepting to trade at the 'wrong' price. The greater the information asymmetry, the larger this adverse selection problem, and the lower the inclination of uninformed investors to trade.

The relationship between volume and price volatility is slightly more complex. One view is that when the number of traders in the market decreases, total volume is likely to fall as well, while price volatility increases, reflecting the deterioration of liquidity in the market. In an alternative

---

[2] See O'Hara (1995), chapter 6.

view, the volume-volatility relationship is positive, rather than negative. This could happen when traders have different beliefs. Differences in beliefs generate trading volume and create price volatility as investors may find it more difficult to agree amongst themselves on a price. The greater the disagreement, the larger the volume and the larger the price fluctuations required to reach this consensus price. In what follows, we use the positive volume-volatility relationship, as it has received strong support in empirical studies.[3]

Asymmetric information models provide a useful starting point for thinking about the impact of improvements in transparency. Informational asymmetries between different classes of market participants are likely to be less pronounced when the monetary policy process is more transparent. This could be either because fewer investors have a distinct information advantage or because the interpretation of the data is more straightforward. With improved transparency, adverse selection risk is lessened, and trading volume and price volatility are likely to be higher.

It follows that when traders are less concerned about adverse selection risk, we might expect public news arrivals to be met by a stronger initial response in both price volatility and trading volume. Consequently, we would expect *greater* volatility and volume reactions immediately following macro announcements post Bank independence. In a more transparent environment, we might also expect the subsequent adjustment process to be completed earlier, and both price and volume reactions to fade out more rapidly.

The opposite view is taken by Wang (1994), who shows that the volume responses to public announcements increase with the level of information asymmetry. Such a reaction in volume arises when traders respond differently to news releases. Wang argues that differences in interpretation arise because of information asymmetries, as uninformed traders use the public signal to update their expectations and correct previously made mistakes. The higher the degree of information asymmetry in the market, the greater the volume response. Hence, according to this view, trading volume should respond *less*, not more, to macroeconomic announcements if information asymmetries have diminished after Bank independence.

A second strand of the literature assumes that even if all traders have access to the same information, they may still differ in their interpretation. For example, traders could hold different opinions on the implications of macroeconomic news for future monetary policy. In a much cited

---

[3] See O'Hara (1995).

paper, Harris and Raviv (1993) show that if traders hold different opinions, then the arrival of new information could lead to larger price volatility and more trades. If transparency has promoted the debate about monetary policy, it could have lead to greater heterogeneity amongst traders.[4] In this case, we might expect volatility and volume to react *more* to macroeconomic announcements after Bank independence. [5] If, on the other hand, more market participants agree on the economic data, then much *less* price volatility or trade would follow.

In conclusion, we have outlined microstructure models that support both *greater* and *smaller* announcement effects post Bank independence. For future reference, we take the adverse selection view as our benchmark hypothesis, as it is arguable the best-known model. Hence, we will postulate that reduced information asymmetries post-Bank independence would lead to *larger initial* volume and price volatility reactions on announcements days, and a quicker adjustment process after the first few trades. We term this our *micro hypothesis*. But given the divergence of views in the literature, extra care will needed when interpreting the empirical evidence.

*Macroeconomic Indicators and the yield curve*

In our discussion so far, however, we have treated short and long interest rates in the same way (and so did Clare and Courtenay (2001) and Lasaosa (2004)). But rates of different maturities capture different aspects of monetary policy and are therefore likely to be affected differently by improvements in transparency. Changes at the longer end of the yield curve capture immediate monetary policy expectations, as well as medium-term monetary policy and inflation expectations. Haldane and Read (2000) suggest that the reaction of long rates to policy announcements reflects market participants' perceptions of both transparency and credibility.

This is illustrated by Joyce and Read (1999) who found that long bond yields responded strongly to inflation releases, particularly so after 1992, whereas three-month Libor rates did not. They suggested that the longer-rate movements reflected market participants' revisions of inflation expectations that might invite some policy change in the medium term. For this reason also, the

---

[4] This is the view taken in Amato et al (2002), who argue that one outcome of increased central bank transparency and credibility is that market participants attach too much weight to the central bank's views, and too little in their own beliefs. This will in turn affect the informational role of financial markets, as less private information will be aggregated and disseminated through prices and trading.

[5] A similar result is found in the announcement literature, for example, which has noted an increase in trading volume after corporate earnings announcements. See O'Hara (1995).

response of long rates diminished over the period 1992-1997, as inflation expectations became more firmly anchored over time. In the United States, Fleming and Remolona (1999) found that the reactions to macroeconomic announcements were the strongest for intermediate maturities (one to five years), creating a hump-shaped pattern in announcement effects. They attributed this result to the Federal Reserve's preference for interest rate smoothing, namely the practice to adjust interest rates in small steps towards a somewhat longer-term target.

In contrast, as suggested in the previous paragraph, changes in long rates could be affected by both improved transparency and credibility. Evidence suggests that medium-term inflation expectations in the United Kingdom were better anchored around the inflation target post-BI.[6] As a result, the reaction of the Long Gilt contract to news in macro indicators that affected medium-term inflation expectations could be smaller post-BI, rather than larger, even in the absence of reduced information asymmetries about the short-term.

For these reasons, the present paper focuses mostly on Short Sterling, as it best captures expectations of immediate monetary policy, being a three-month interest rate, deliverable 0 to 3 months forward. For comparison purposes, we will also present results of the Long Gilt. Clare and Courtenay (2001) and Lasaosa (2004) further examined the reaction of equity prices. But here again, there may be little reason to expect a greater reaction in the FTSE 100 contract post-BI, as macroeconomic news can affect both the expected future pay-off and discount factor components of equity, making the ex-ante expectation of a transparency effect ambiguous.

*Monetary Policy News*

Haldane and Read (2000) show that if market participants have a better understanding of monetary policy, they should be able to better predict future policy rate changes. Consequently, the reaction of market interest rates to monetary policy announcement should be smaller. But both theory and empirical evidence suggest that we should test the reaction of asset prices to the surprise component of monetary policy (see e.g. Kuttner (2001)). Clare and Courtenay (2001) and Lasaosa (2004) examine the actual announcement, rather than the unexpected component. Using only the surprise component has two implications for their hypothesis.

---

[6] There is strong evidence of increased credibility post-BI, see King (2002).

First, it is not clear whether we should indeed expect markets to react more to a given surprise in a monetary policy announcement post-BI. If financial markets better understand the MPC's reaction function, it is likely that the surprise component of monetary policy announcements will be smaller. It is less clear what should happen to the size of the markets' reaction to that surprise. In other words, if we find evidence of a decline in the surprise component of the monetary surprise variable, but a similar (or even larger) reaction in prices, do we interpret this as evidence for or against the hypothesis of Clare and Courtenay (2001)?

A second problem is that we do not have sufficient data to carry out the empirical tests properly. Survey expectations of MPC decisions are not available before Bank Independence. In addition, any analysis of the effect of monetary policy before Bank Independence would have to use a lower frequency of financial market data, as the time at which policy changes became clear to the market before was somewhat ambiguous[7]. For these reasons, this paper does not examine the reaction to monetary policy news.

## 3 Literature review

This section provides a short overview of the empirical literature on announcement effects. For a thorough review of the issue of central bank transparency and its impact on financial prices, the reader is referred to the literature reviews in Geraats (2003), Lildholdt and Vila Wetherilt (2003), and Lasaosa (2004).

Early event studies of the reaction of asset prices to macroeconomic and monetary policy announcements, such as Cook and Hahn (1989) and Ederington and Lee (1995), sought to measure the effect of either policy rate changes or macroeconomic news on returns and return volatility. Later papers, such as Becker et al (1996), Fleming and Remolona (1997) and (1999), extended this approach by considering only the surprise component of news announcements. At around the same time, the literature also moved from using daily closing prices to intraday data,

[7] Between ERM and Bank independence, monetary policy was decided by the Chancellor of the Exchequer, following a monthly consultation with the Governor of the Bank of England. The timing of the implementation of the Chancellor's decision was then left at the discretion of the Bank of England, with the restriction that it had to be implemented before the next meeting of the Chancellor and Governor. The Bank generally indicated the decision to the markets by changing the rate at which it conducted its daily money market operations. The Bank conducted its money market operations at regular times throughout the day, so, for changes in the policy rate, newswire reports can be used to obtain the time at which the market became aware of the new policy rate. However, in the case of no-

ranging from one-hour to one-minute intervals. On the whole, the empirical literature on announcement effects has found that public announcements cause short-lived reactions in financial prices.

Broadly speaking, this literature consists of two types of studies. A first group is descriptive in nature and examines the patterns of intraday returns and return volatility before and after announcements. Ederington and Lee (1995) found that prices in the US Treasury and foreign exchange markets adjusted as early as ten seconds after a news release, with most of the price adjustment completed in the following 30 seconds. Fleming and Remolona (1999) found evidence of a two-stage reaction process in US Treasury markets: Volatility was found to rise sharply in the first two minutes following a macroeconomic announcement. This reaction was entirely explained by dealers widening their bid-ask spreads in response to the increased risk of rapid price changes and associated inventory losses. In the second stage, volatility declined, although it remained significantly higher for at least an hour after the news release. Fleming and Remolona (1999) attributed the relative length of the second stage to market participants taking a while to reach a consensus about the significance of the announcement. This two-stage adjustment process is confirmed by Balduzzi et al (2001) who studied Treasury bond prices too, but at a slightly lower frequency.

In the United Kingdom, ap Gwilym et al (1998) found both short sterling and FTSE 100 futures to react immediately to UK announcements. Short sterling prices took up to three minutes to fully adjust, whereas FTSE100 adjusted within 90 seconds. Clare et al (1999) observed that return volatility of Gilt futures prices was higher in the half hour following macroeconomic news releases. Clare and Courtenay (2001) conducted a similar analysis for a range of futures contracts and found volatility to peak in the first minute after news releases, and to decline sharply in the next two minutes. Compared to non-announcements days, return volatility remained higher for another 15 to 20 minutes. Clare and Courtenay (2001) subsequently used this descriptive approach to compare the price response to announcements before and after the Bank of England's independence in 1997, and so did Lasaosa (2004).

A second group of studies relies on regression analysis to assess the impact of announcements on returns and return volatility. This is done either via dummies corresponding to the individual announcements, or by including the surprise component of the announcement as an additional

---

change decisions, the decision not to change rates did not become clear at any discrete point in time, but rather it became clear slowly, as the Bank continued to conduct money market operations at the previous policy rate.

regressor. Ederington and Lee (1993), Becker et al (1996), Fleming and Remolona (1997) and Balduzzi et al (2001) used the regression approach to examine various US announcements. Chief amongst their findings was the observation that Treasury markets reacted differently to various announcements, with employment and inflation releases generating the greatest response, and GDP releases the weakest one. Moreover, the price response to a given announcement appeared to be stronger in periods of greater market uncertainty. Fleming and Remolona (1997) further established that using the surprise component of the announcements provided more accurate estimates of the reaction of prices to new information. Most subsequent studies have followed this approach. In the context of the United Kingdom, ap Gwilym and Thomas (1998) examined the impact of both US and UK macroeconomic announcements on Gilt futures. Like the US studies, they found inflation and labour market statistics to have the largest announcement effects.

Andersen and Bollerslev (1998) and Bollerslev et al (2000) extended the regression approach (usually a simple OLS equation) to account for the dynamics of intraday volatility, in particular its well-documented persistence. Using a carefully constructed GARCH specification to examine five-minute returns in foreign exchange and Treasury bond markets, respectively, they confirmed the presence of announcement effects documented in the earlier literature. Using a similar framework, Andersen et al (2003) found that in their response to macroeconomic surprises, exchange rates reacted immediately, but their volatility took up to an hour to adjust.

More recent papers analyse the joint reaction of different markets to the same public information arrivals. Faust et al (2003) examined the joint response of interest rates and foreign exchange rates to US macroeconomic and policy rate announcements. They were able to characterize the impact of news releases on both expected exchange rate dynamics and foreign exchange risk premia. Although they employed daily data, the work of Ehrmann and Fratzscher (2002) is interesting as they found evidence of changes in the nature of spill-over effects between the United States and the euro area. In particular, they observed that as markets have become more knowledgeable about the ECB's monetary policy, their reaction to euro area news has changed too, to become qualitatively similar to the response of German markets to domestic announcements prior to the creation of the ECB. Furthermore, they found that the response of euro area prices to US announcements had declined over time.

Event studies have also examined the reaction of trading volumes following announcements. As explained in the previous Section, the market microstructure literature relates trading volume to

the arrival of news.[8] If this news leads to different interpretations amongst market participants, or intensifies existing information asymmetries, then news releases could lead to increases in both trading volume and volatility. On the whole, event studies confirm this insight. Fleming and Remolona (1997) found that trading volumes in US Treasury bond markets increased after most macroeconomic announcements, and that this rise was more pronounced in periods of uncertainty about market prices. Balduzzi et al (2001) noted that this response was quite persistent, with Treasury bond trading volumes often remaining higher than normal for up to one hour. In contrast, their study showed the price response to be very short-lived. Using data of even higher frequency, Fleming and Remolona (1999) found that trading volume in Treasury bonds dropped immediately following economic announcements, reflecting, as explained above, dealers' reluctance to trade. This initial fall was followed by a sharp rise in trading volume. Volumes remained high for up to 90 minutes, longer than price volatility.

In the United Kingdom, ap Gwilym, et al (1998) observed that trading volumes in short Sterling futures rose almost instantaneously in response to news releases, and again in contrast to prices, remained higher for longer. Clare et al (1999) found evidence of a rise in the size of long Gilt futures trades following macroeconomic announcements, but did not detect a rise in the number of trades. Lasaosa (2004) found the rise in futures volumes (both number of trades and their size) following macroeconomic announcements to be more muted in the period after the Bank of England's independence.

Microstructure theory has also demonstrated that the relationship between prices and trading volumes itself may be affected by information asymmetries. In his seminal paper, Kyle (1985) showed that prices will be more sensitive to trades either when liquidity is low (buyers have difficulty finding sellers) and/or when the level of information asymmetry increases. Event studies have explored this idea by examining the price-volume relationship following macroeconomic announcements. Green (2001) found that the release of public information increased information asymmetries in Treasury bond markets, with prices becoming more sensitive to trading activity in the half hour following announcements. Although in his tests, trading volume remained high for several hours after the announcement, price volatility and sensitivity measures returned to normal within 15 minutes.

---

[8] See also Bessembinder and Seguin (1993) and O'Hara (1995).

## 4 Data description

### *4.1 Intraday prices, number of trades and trading volume*

This paper examines the intraday reaction of futures markets to announcements on all trading days between 1994 and 2003.[9] We use two futures contracts traded on LIFFE: Short Sterling and Long Gilt, though our hypotheses on transparency lead us to focus on the Short Sterling contract and look in less detail at the Long Gilt contract. For each trade recorded, the LIFFE dataset gives the time of the transaction to the nearest second, the price at which it occurred, the volume of contracts traded in the transaction and the maturity of the contract traded. Following Clare and Courtenay (2001), we use the closest to expiry Short Sterling contract until it expires and the closest to expiry Long Gilt contract until the next contract becomes more liquid.[10] From this data set, we construct return and number of trades series for non-overlapping intervals of a variety of lengths, for every working day between 1994 and 2003. The highest-frequency intervals we use are one minute in length, the lowest-frequency are 60 minutes. The lower-frequency series are used to test the *macro* hypothesis, whereas the higher-frequency ones are employed to provide evidence about the *micro* hypothesis.

To obtain a first idea about the intra-day behaviour of both price volatility and trading volume, Charts 1-4 show the mean absolute price changes (Charts 1 and 3) and mean deviation of the number of trades from a rolling mean level (Charts 2 and 4) for the Short Sterling and Long Gilt contracts for each five-minute period of the working day across the entire sample period. Two series are shown, one for days on which the UK macroeconomic indicators covered in this paper were announced and one for all other trading days[11].

Insert Charts 1-4 here

Price volatility and the number of trades appear to follow a similar pattern across both markets, with high volatility and heavy trading at the beginning and end of the day and at brief, but regular periods throughout the day. Trading and price volatility are heavier in the Long Gilt market in the afternoon, possibly due to the opening of US markets. In both markets, there are spikes in

---

[9] We were not able to take our sample back further, eg to include the introduction of inflation targeting, due to lack of data.
[10] We apply filters to the data to remove erroneously recorded trades. These are available from the authors.
[11] Note that, by construction, policy announcements, foreign macroeconomic announcements and UK macro announcements not covered in this paper can fall in either line.

both price volatility and number of trades at 09:30, 13:30 and 15:00. The first coincides with the release of the vast majority of UK macroeconomic indicators, while the latter two coincide with the two main release times of US macroeconomic indicators. The clarity of these three spikes in the Short Sterling market is encouraging, as they will form the focus of this paper.

Given the high frequency nature of our data, we can focus in more closely on the 09:30 spike. Charts 5 and 6 present the average one-minute absolute return in the minutes before and after the 09:30 announcement time on both announcement and non-announcement days. Chart 5 shows the period of 1994 – Bank Independence and Chart 6 shows the period Bank Independence – 2003. The latter appears to show a sharper reaction of volatility to macroeconomic announcements after 1997. We will investigate this further in Section 7. Charts 7 and 8 repeat the analysis using cumulative returns. They consider the period out to one hour and plot the average cumulative absolute return for announcement and non-announcement days in the period 1994 – BI and BI – 2003. The Charts appear to show evidence of a sharp initial reaction of prices to macroeconomic announcements, but also a continued price reaction over the first 30-40 minutes following an announcement.

Insert Charts 5-8 here

There is a potential structural break in the time series we construct from this data, when the trading of both contracts on LIFFE switched from pit trading to electronic trading. This occurred on 1 July 1999 for both contracts. When looking at the data, the break is more apparent in the trades series, which shows a marked increase in trading volume. This break must be accounted for in our tests of both hypotheses, and we explain how we go about this in Section 5.

### *4.2 Announcements and surprises*

For the United Kingdom, we use the same set of macroeconomic indicators as Clare and Courtenay (2001). Summary statistics are provided in Table 1. The data are taken from Money Market Services International (MMS), a database which includes the exact time of the indicator announcement, the period which the data covers, the outturn of the indicator and the median of a

survey of brokers' expectations of the outturn.[12] The survey is conducted the Friday before the indicator is announced and usually covers 30-40 money-market brokers.

Insert Table 1 here

The median survey expectation is of interest as it provides a proxy for the market's expectation of each outturn.[13] It allows us to construct for each indicator a surprise variable, $S_{kt}$, by subtracting the survey expectation of the outturn from the outturn and standardising by dividing by the standard deviation (this is done solely to make the regression coefficients comparable across indicators). It is equal to the expression below on days when an announcement is made, and zero otherwise:

$$S_{kt} = \frac{A_{kt} - E(A_{kt})}{\sigma_{s_k}}$$

We test the expectations series for each indicator for evidence of bias and test the reaction of both contracts to each indicator for evidence of efficiency.[14] We find some evidence of upwardly biased expectations in the Unemployment, Industrial Production and PPI Input indicators. Balduzzi et al (2001) argue that the resulting errors-in-variables problem will bias downwards the regression coefficients and upwards their standard errors. Noting in addition that the tests are based on the assumption of a symmetric distribution of surprises, when it seems plausible to suppose for our sample period that at least unemployment and perhaps inflation surprised more on the downside than the upside, we follow their approach, which is to run standard OLS regressions, but use care in the interpretation of the t-statistics, which will be downward biased.

---

[12] We checked to make sure that our release times did not coincide with other regular releases of information. DMO T-bill and gilt auctions do not pose a problem, beginning at 11:00 and 10:30 respectively, with results announced within 20 and 40 minutes respectively. The publication times of the CBI Distributive Trades and Monthly Trends surveys varied throughout the period, but they never coincided with the ONS macroeconomic releases. The CIPM and CIPS manufacturing and services surveys were published at 09:30 for much of the period. The CIPM coincide with 23 of the Provisional M0 announcement times and the CIPS coincide with 15 of the Provisional M0 and 9 of the Industrial Production announcement times. This could bias the M0 and Ind. Prod. Coefficients in the dummies regressions and to a lesser extent bias their standard errors in the surprises regression. The Minutes of the MPC meeting coincide with the Average Earnings/Unemployment indicator on 27 occasions and with a few other indicators on a few occasions. We tested whether including dummies for the MPC Minutes affected the coefficients on the Average Earnings/Unemployment indicators in our regressions and found that they did not.

[13] This, however, implied that we could only look at announcements for which a surprise variable can be calculated. For example, such a variable could not be constructed for survey announcements such as those conducted and published by the BCC or the CBI, which arguably contain valuable information for policy makers. As a result, the research in this area tends to look at historical data releases only.

[14] See Appendices A and B.

Our macroeconomic data for the United States is taken from the same source (MMS) and comes in exactly the same form. A description of the data is available in Appendix C. We ran the same tests for unbiasedness and efficiency and found the survey expectations to be unbiased and the market reactions to be efficient for all of the macroeconomic indicators we use[15].

## 5 How to measure announcement effects

The regression equations described in this section follow the standard format used in the event studies outlined in the literature review of Section 2.[16] Two main types of dependent variables are used, a measure of the change in price for the yield and volatility regressions and the deviation of the number of trades from a rolling average for the trades regressions.

### *5.1. Dependent variables*

For the yields regressions using the Short Sterling contract, we measure the change in yield on the contract as the basis point change in yield on the contract. We can do this because the yield is a simple function of the price of the contract.[17] For the Long Gilt, the contract is traded on the price of a bond of loosely specified maturity and coupon size, so we are forced to use for our change in yield series $-10{,}000 \times \left[\ln(P_{t+1}) - \ln(P_t)\right]$, where $P_t$ is the price of the contract immediately before a macroeconomic announcement and $P_{t+1}$ is the price five minutes after the announcement. The scale of the coefficients is only meaningful for comparison between coefficients in Long Gilt regressions. The separate measures of change in yields for the two contracts will both be referred to as 'yields' variables.

Each regression contains one observation per working day for each of the announcement times. Thus, in the regressions that measure the *x*-minute reaction of UK asset prices to UK announcements, all of which occur at 09:30, the yields series comprises the 2490 change in yields between 09:30 09:30+*x* (one for each working day from January 1994 to September 2003; 930 of these observations follow announcements, the other 1560 do not). In the regressions that test the

---

[15] The tests are available from the authors.
[16] One possible drawback of the regression method is that one needs to make distributional assumptions. The descriptive approach, outlined in Section 3, avoids this, by using non-parametric tests of differences. We opted for the regression approach, however, as it is more flexible and permits a more robust model specification. This will become clearer in the remainder of this Section.
[17] The price of the contract is given by the formula P=100-yield (in percentage points).

reaction of UK bonds to US announcements, which occur at 13:30 and 15:00, the yields series comprises 4980 change in yields (one for each of the *x*-minute periods following 13:30 and 15:00 for each working day from January 1994 to September 2003). As described below, the length of the reaction window, *x*, varies between tests.

There is an equivalent trades series for each of the announcement times. As the variance of trading volume varies widely over time, and its mean also appears to be non-stationary, we choose to look at the deviation of trades from a normalised rolling average, rather than the simpler number of trades. For this reason, we define the reaction of trading volume at time *a* on day *t* as the number of trades *N* in the five-minute period following *a* minus the rolling mean number of trades $\mu$, all divided by the rolling standard deviation of the number of trades $\sigma$. For our rolling means and standard deviations, we use all non-overlapping five-minute periods of the trading day for the ten trading days before day *t* (note that there are 120 non-overlapping five-minute periods per trading day).

$$Y_t^{trade} = \frac{N_{t,a} - \mu_t}{\sigma_t} \quad \text{where } \mu_t = \frac{1}{120 \times 10} \sum^{A} \sum_{t-10}^{t-1} N_{t,a} \quad \text{and } \sigma_t = \sqrt{\frac{1}{120 \times 10} \sum^{A} \sum_{t-10}^{t-1} (N_{t,a} - \mu_t)^2}$$

*5.2. Tests of the change in yield following announcements*

We first consider the regressions used to test the change in price following macro announcements. These tests ignore the price adjustment dynamics and compare the price before an announcement with the price after the information in the announcement has been assimilated. The tests of the price adjustment dynamics are very similar and can be most easily explained as simple extensions of the price change tests. We make inferences about our *macro* hypothesis from the first set of tests and about our *micro* hypothesis from the second set of tests.

*Reaction to announcement time dummies*

First, we regress the yields variables on dummies for the time at which the announcement was made, ignoring the quantitative information in the size of the surprise. We use the absolute change in yield because the dummy variable is unsigned. The equation specified looks as follows:

$$(1) \qquad Y_t^j = a_1^j + \sum_{k=1}^{K} \beta_{1,k}^j \times D_{kt} + e_t^j$$

In equation (1), $Y_t^j$ is the absolute change in yield on security *j* from 09:30 to 10:30 on day *t*. $D_{kt}$ is a dummy which equals 1 when macroeconomic indicator *k* is announced at 09:30 on day *t* and which equals 0 at all other times. The regression is run for both contracts; Short Sterling and Long Gilt. The regression permits the following test; that the absolute reaction in yield on security *j* to announcement *k* differs significantly from the absolute change in yield on non-announcement days between 09:30 and 10:30 (ie, $\beta_k^{j,i} \neq 0$).

To address the question of whether the reaction has changed since Bank independence, we add post-independence multiplicative dummies, $D_t^{BI}$, to Equation (1)[18]. This permits a test of whether the yield reaction of security *j* to indicator *k* is significantly different post-independence compared to pre-independence:

$$(2) \quad Y_t^j = a_1^j + \sum_{k=1}^{K} \beta_{1,k}^j \times D_{kt} + \sum_{k=1}^{K} \beta_{2,k}^j \times D_{kt} \times D_t^{BI} + e_t^j$$

*Reaction to the surprise component of announcements*

A second set of regressions relates the yield response to the size of the surprise in the announcement. If markets are rational and forward looking, then the two securities would only be expected to react to macroeconomic indicators to the extent that the indicators contained news not already accounted for in the price. This should lead to a better-specified regression equation. Using quantifiable surprises has further advantages; it gives us a sign for the reaction to indicators and it allows us to include separate regressors for announcements made at the same time. This means that we can examine the reactions to PPI Input and Output, RPI and RPIX and Average Earnings and Unemployment individually.

Replacing the dummy variable in Equation (1) with the earlier-defined surprise variable, $S_{kt}$, permits a test of the reaction of yields ($Y_t^j$) to the surprise component of announcements:

---

[18] Using dummies to capture the break point may be sub-optimal if the effects of any increase in transparency occurred slowly, rather than manifesting themselves as a discrete shift in behaviour. However, our small sample size (recall that our macro announcements are monthly and quarterly) makes the simple, but parsimonious, dummy variable approach attractive.

(3) $$Y_t^{j,i} = a_1^j + \sum_{k=1}^{K} \beta_{1,k}^j \times S_{kt} + e_t^j$$

As with Equation (1), adding dummies for the post-independence period allows for a test of whether the reaction of indicators has changed over time (Equation 4).

(4) $$Y_t^j = a_1^j + \sum_{k=1}^{K} \beta_{1,k}^j \times S_{kt}^i + \sum_{k=1}^{K} \beta_{2,k}^j S_{kt} \times D_t^{BI} + e_t^j$$

For ease of interpretation, we report regressions with the surprise variables standardised by the sample standard deviation of the surprise, as detailed above, although this does not allow for the distribution of surprises to have changed during the time period. A simple visual inspection of a plot of the non-standardised surprise data suggests that the distribution of only the Unemployment surprises changed much throughout the sample period, with the surprises becoming smaller throughout the period, presumably as market participants lowered their estimates of the NAIRU. We confirmed this by running two alternate regressions to test the sensitivity of our results to the standardisation; one without standardised surprises and one with surprises standardised by the pre- and post-BI standard deviations separately. We found that only the coefficients on the Unemployment indicator were much affected, with the coefficients on the other indicators having the same sign and similar quantitative ranking as previously. The difference in the reaction of yield to the Unemployment indicator post-BI appears to be understated in our standardised regressions as a result of the tightening of the distribution of Unemployment surprises.

*5.3. Tests of the price adjustment dynamics*

We test the price adjustment dynamics by analysing both price volatility and trading volume, using high-frequency data. We first wish to test whether the initial increase in volatility following an announcement is larger post-BI and whether volatility dies down more quickly thereafter. Following Ederington and Lee (1993), this is done by running a series of regressions similar to Equation (1) on consecutive one-minute absolute price changes. We then repeat this exercise for a series of consecutive five-minute absolute price changes. We will present regression results for twelve five-minute intervals, and five one-minute intervals. For ease of interpretation, we run separate regressions for the pre- and post-BI periods, and compare the coefficients.

We employ a similar methodology to describe the behaviour of trading volume following macro announcements in the pre- and post-BI periods. We again estimate a version of Equation (1), this time using as dependent variable the number of transactions that occur in five-minute trading intervals. This regression is run for twelve successive five-minute intervals, in order to test the total volume adjustment during the full hour after the macro announcement. To further investigate any possible changes in behaviour, we also examine and compare the distribution of trades on announcement and other days.

As mentioned in Section 4, we have to account for the structural break corresponding to the start of electronic trading. We do this by splitting our sample period in three: the pre-BI period, the post-BI to electronic trading period, and the post-electronic trading period. We run separate regressions for each of the three periods.

Our third set of tests is on cumulative returns in the pre- and post-BI periods; describing the path of prices as information is assimilated following a macro announcement. The methodology is similar to that of Equation (3), but with a series of cumulative return variables as the dependent variable. The tests are again run for one- and five-minute periods, but with cumulative returns being defined as the change in yield between 09:30 and each of consecutive one- and five-minute intervals (i.e. 09:30-09:31, 09:30-09:32, etc.). Note that Equation (3) is a regression on surprises, and so it gives us a better fit of equation and allows us to measure separately the effect of announcements made at the same time.

## 6 The effect of macroeconomic announcements on yields (the *macro* hypothesis)

In this Section, we test the change in price that occurs following macro announcements. Our tests are conducted on the 60-minute change in yield. The hypothesis is tested using the Short Sterling contract, but we also report results for the Long Gilt contract. Our intentions in looking at the Long Gilt contract were largely exploratory; we had no hypothesis to test with a view to saying something about transparency, but rather thought that doing the tests on another market may provide an interesting perspective for our work on Short Sterling. The results are reported in this light. For the purposes of this paper, they do not merit as much attention as the Short Sterling results and our treatment of them in the text that follows reflects this view.

*6.1 The effect of UK macro announcements on Short Sterling and Long Gilt*

We first consider the simplest test of the impact of UK macroeconomic indicator announcements on UK yields by estimating Equation (1) for the full sample period. The results are given in Table 2.

Insert Table 2 here

The results show the yields on Short Sterling to react significantly to five of the eleven macroeconomic indicators. Average Earnings/Unemployment, Retail Sales and RPI/RPIX appear to provoke the strongest and most highly significant reaction. The preliminary estimate of GDP appears to provoke a stronger reaction than the two revisions to GDP, which is encouraging as it is the first of the three GDP releases and is thus likely to contain more new information than the other two. The results for the Long Gilt contract are similar, with seven of the eleven indicators eliciting significant reactions and the Average Earnings/Unemployment, Retail Sales and RPI/RPIX indicators again appearing to attract the strongest reaction.

Table 3 shows the results of the price regressions when the surprise component of announcements is incorporated, instead of the timing dummy (Equation (3)). As explained earlier, if markets are rational and reasonably efficient, we may expect a stronger relationship between the surprise component of announcements and change in yields than between simple announcement dummies and change in yields, because dummy variables do not allow us to differentiate between fully anticipated news, which should exact no reaction, and unanticipated news, which should exact a reaction. Note that the use of quantitative variable allows us to test separately for the reaction to indicators which are released at the same time, so the regression now includes separate variables for Average Earnings, Unemployment, PPI Input, PPI Output, RPI and RPIX.

Insert Table 3 here

Table 3 shows that Short Sterling yields react significantly to eight of the 14 indicator announcements; namely, Average Earnings, Preliminary GDP, Industrial Production, Provisional M4, PPI input, Retail Sales, RPI and Unemployment. A positive coefficient on indicators, such as the coefficients on Preliminary GDP and Retail Sales, suggests a larger than expected outturn in those indicators leads to a rise in yields. The negative coefficient on Unemployment suggests that a larger than expected outturn leads to a fall in yields. Larger than expected outturns in all

the indicators would generally be seen as leading to a rise in short rates, except for Unemployment, which would be expected to be associated with a fall in short rates. Thus, the signs of all the significant indicators are as would be expected. The rank of the magnitude of the coefficients is similar to that in the dummies regression, with Average Earnings, RPI and Retail Sales again eliciting the strongest reactions. The results for the Long Gilt contract are similar, though, surprisingly, the Long Gilt contract does not appear to react significantly to any of the GDP releases at the 5% level.

Comparing the specifications of our two Short Sterling regressions, Equations (1) and (3) from Tables 2 and 3 respectively, reveals a higher $R^2$ for the surprises regression (Table 4), and a larger number of variables significant at the 5% level[19]. We argued in Section 5 and above that for theoretical reasons, it was preferable to examine the reaction of financial markets to unexpected macro announcements. With the statistical evidence from our two regressions supporting this view, we concentrate on macroeconomic surprises for the rest of our macro tests.

### *6.2 The Effect of Bank Independence*

Table 4 shows the results of the regression equation (4), which tests for a differential response to indicator announcements post-BI. For presentational purposes, we drop indicators that were not significant in Equation (3) and which continue to be not significant in either the pre- or post-BI period.

Insert Table 4 here

Table 4 shows that before Bank independence (the top half of the table), the Short Sterling contract responded significantly and with the expected sign to surprises in seven indicators; the same as in Table 4, but with the exception of Preliminary GDP. Post-BI (the bottom half of the table), only the reaction to Retail Sales was significantly lower than it had been in the pre-BI period. The reaction to all the other indicators was not significantly different (at the 5% level) to the reaction pre-BI. Furthermore, of the coefficients that were not significant, all but one were of the sign that would suggest a smaller reaction post-BI. With respect to our hypotheses on transparency, this constitutes reasonably strong evidence against the *macro* hypothesis.

---

[19] The $R^2$ statistic does not have the standard interpretation, as no attempt is made in the regressions to explain the variation in change in yields on non-announcement days, other than the inclusion of a constant. Thus $R^2$ statistics are only directly comparable across regressions done on the same contract and with the same sample.

In the next two sub-sections, we provide further evidence on transparency by considering several extensions of the regression models used in this Section.

### *6.3 International macro indicators*

In making their interest rate decision, the MPC consider not only developments in the UK economy, but also developments in other large economies, such as the euro-area and the United States, which are likely to impact on the United Kingdom over the forecast horizon. If markets understand this, they will revise their expectations of UK monetary policy in line with information about aspects of other economies which they think may feature in the MPC's reaction function. If this is the case, it is important to try to include relevant foreign macroeconomic indicators in our tests of transparency, as the hypothesis that markets should react more to macroeconomic news post-BI applies equally to domestic and international macroeconomic news.

The impact of foreign economies on UK inflation will generally come through trade, making the countries with which the United Kingdom trades most the most relevant to UK monetary policy decisions. The euro-area is by far the United Kingdom's largest trade partner, but unfortunately the advent of monetary union makes it unclear to which indicators monetary policy was most likely to react throughout the sample period; those of the largest euro-area nations, Germany, France and Italy, or the euro-area-wide economic indicators introduced in the late 90's. This ambiguity, in addition to most of the old German indicators being published outside of LIFFE opening hours, makes euro-area macroeconomic announcements prohibitively difficult to examine. Macroeconomic indicators for the United States, the UK's second largest trade partner, are easier to incorporate. Table 5 show the estimation results for equation (4), with US indicators replacing UK ones. As in the previous section, some of the consistently non-significant indicators are omitted.

Insert Table 5 here

The Short Sterling contract reacts significantly, to only three of the US macroeconomic indicators in our dataset: Preliminary GDP, Housing Starts and Non-Farm Payrolls. The significant reaction to Preliminary GDP (the second GDP release) is perhaps surprising, given that the reaction to Advance GDP (the first release) is not significant. It should be noted that the sample size in each

case is quite small, as there are only a third as many observations for the quarterly indicators as there are for the monthly indicators. Of the post-BI indicators, only Initial Claims shows a significantly different reaction post-BI, the reaction being greater than in the pre-BI period (note that a greater than expected outturn in Initial Claims would generally be associated with weaker than expected US output). On the whole, the lack of a more general increased reaction to indicators post-BI can be seen as further evidence against the *macro* hypothesis.

The Long Gilt contract only responds significantly to three US announcements prior to 1997; Non-farm Payrolls Hourly Earnings and Preliminary GDP. The reaction to US Non-Farm Payrolls in the pre-BI period is stronger than the reaction to any of the UK indicators. Only two indicators elicit a significantly different reaction post-BI, with both Preliminary GDP and Non-Farm Payrolls provoking a smaller reaction than pre-BI.

That the Long Gilt contract reacts more strongly to US indicators than the Short Sterling contract likely reflects the difference in maturity of the contracts, as discussed in Section 2. The Short Sterling contract is most likely to be affected by influences on immediate monetary policy, and thus react more to direct indicators of UK economic prospects. The Long Gilt contract will be affected by information about longer-term economic prospects, and thus react more to indirect indicators of UK economic prospects, as business cycles are not perfectly synchronised and pass-through from foreign supply and demand to the domestic economy is slow.

*Controlling for uncertainty*

The tests carried out so far have not allowed for the possibility that markets may react differently to a given macroeconomic surprise in different macroeconomic climates. The significance of a given surprise to agents attempting to predict future monetary policy may depend on the amount of uncertainty surrounding future states of the economy. If there is very strong consensus about the immediate path of monetary policy, say, because the economy has received a strong shock to move it away from equilibrium and the policy maker's likely reaction is clearly understood, the value of additional information from macroeconomic indicator surprises may be relatively low. But if the immediate path of monetary policy is very uncertain, the amount of information in a surprise outturn may be relatively high. Failing to account for this could lead to poorly specified regressions. Worse still, if the balance of periods of certainty and uncertainty were greatly different in the pre- and post-BI periods, the tests of transparency used so far may suffer from bias.

Uncertainty about the immediate path of monetary policy can be proxied for using implied volatility from options prices. When implied volatility is low (high), uncertainty about the likely path of interest rates, relative to that implied by interest rate futures prices, is low (high). We use basis point implied volatility to calculate a dummy variable, $\beta_t^{IVOL}$, that equals 1 when the implied volatility of a three-month option on Short Sterling exceeds 80 basis points[20]. We introduce this to Equation (4) as a multiplicative slope dummy to test for a differential reaction to the surprise component of indicators in times of greater disagreement amongst market participants:

$$(5)\ Y_t^{j,i} = a_1^{c,i} + \sum_{k=1}^{K} \beta_{1,k}^{j,i} S_{kt}^{i} + \sum_{k=1}^{K} \beta_{2,k}^{j,i} S_{kt}^{i} \times D_t^{BI} + \sum_{k=1}^{K} \beta_{3,k}^{j,i} S_{kt}^{i} \times D_t^{IVOL} + e_t^{j,i}$$

The results are shown in Table 6. The top panel shows the various regression coefficients prior to Bank independence and under low uncertainty. The bottom panel examines whether markets responded more post BI, under low uncertainty. Finally, the middle panel reports whether the responses of prices and trades were more pronounced under high uncertainty (both before and after 1997).

Insert Table 6 here

Most strikingly in the Short Sterling regression, none of the uncertainty dummies are significant. A joint test of the uncertainty coefficients shows that they are jointly not significant and thus suggests that the specification of the regression is inferior to that of Equation (4). Given this, it makes little sense to put much weight on the interpretation of the coefficients on the other regressors. Doing so shows that they are little different to those in Table 4.[21]

Considering the Long Gilt contract leads to the same conclusion – there appears to be no role for uncertainty about the immediate path of monetary policy in explaining the reaction of the Long Gilt contract to the surprise component of macroeconomic indicators. This is not overly surprising, as ten-year interest rates should not be very sensitive to the immediate path of

---

[20] The choice of 80 basis points is set to cause the dummy to equal 1 (high volatility) 25% of the time. The choice of 25% is necessarily arbitrary. The results are reasonably robust to different ratios of high-to-low uncertainty. We use implied volatility measured in basis points because it is insensitive to the level of Short Sterling, which varies widely over the sample.

[21] In contrast, Fleming and Remolona (1997) find that the reaction of Treasury futures to macro announcements is greater during periods of increased uncertainty.

monetary policy, except in so much as it contains information about the manner in which monetary policy is likely to be conducted in the future.

### *6.5 Concluding remarks*

All our evidence on the *macro hypothesis* is in favour of its rejection. One hour after major macroeconomic releases, prices have reacted less to the new information in the post BI period than before. Although we use a different sample period and a different methodology, our results are in line with Clare and Courtenay (2001) and Lasaosa (2004). Hence, at first sight there appears to be little evidence of increased transparency since Bank independence. We suggested in Section 2, however, that this result need not be seen as incompatible with greater transparency. In particular, we explained that other changes in the conduct of monetary policy (committee decision making, forward-looking rules, and gradualism), put in place at the same time as increases in transparency, could have led to smaller, rather than larger responses to macroeconomic news. We explore this further in the next section.

## 7 Testing the *macro* hypothesis for the United States

One concern with our approach is that it may be influenced by broader trends in the behaviour of central banks or the workings of the Short Sterling market. As noted above, a particular concern is that we may fail to account for the effect of a more forward-looking approach in monetary policy strategy. In addition, as noted in Section 3, monetary policy in the United Kingdom and United States has become increasingly characterised by a series of small, positively correlated changes in policy rate. Martin and Salmon (1999) find evidence of increased gradualism for the United Kingdom and Sack (2000) for the United States. Both increased gradualism and the use of forward-looking policy rules might explain why the path of monetary policy might have become increasingly predictable, as empirically demonstrated by Lange et al (2003), Swanson (2003) and Lildholdt and Vila Wetherilt (2004), for the United States and the United Kingdom, respectively.[22] Increased predictability could in turn have led to smaller market reactions for a given amount of macroeconomic news. In this Section, we rerun our tests for the United States, to see whether the somewhat similar set of circumstances has led to a similar change in market

[22] These papers also demonstrate that predictability changes slowly over time. Lildholdt and Vila Wetherilt( 2004) further show that it is difficult to find discrete shifts in the data corresponding to policy changes, and that much of the

reactions. Admittedly, this is a rather weak test, as we are not able to exactly attribute the results to either increased transparency, or to increased predictability. Nonetheless, it might be instructive to compare the two countries' experiences.

To this end, we estimate equation (4) using US yields and announcements data. We use the EuroDollar contract for the dependant variable, as it is the US equivalent of the Short Sterling contract. Table 7 shows the results.

Insert Table 7 here

At first pass, using our standard measure of the change in yields in Equation (4), it appears that there is a decline in the reaction of yields to five of the most prominent indicators in the post-BI period. However, Equation (4) is likely inappropriate for the EuroDollar yield series because it contains structural breaks in 1997 and 1998, when the tick size decreases from one basis point to one half of, and then one quarter of, a basis point. The decline in tick size was accompanied by a decline in intra-day volatility, which makes it desirable to look not at the reaction to a surprise at time *t*, but at the reaction at time *t* relative to the general level of intra-day volatility at time *t*. To do this, we construct a series similar to that constructed for trading volume for the United Kingdom.[23] Using this series, reported in the third column of Table 8 as equation *(4)*, it appears that the reaction of EuroDollar yields to the core macroeconomic indicators is little different in the post-BI period compared to the pre-BI period, with only one of the post-BI coefficients being significant and with them being mixed in sign.

The evidence for the United States shows a similar picture to that seen in the United Kingdom; little difference in price reaction to macro news after 1997. We earlier suggested that the lack of a marked change in market reactions to UK announcements did not necessarily constitute a rejection of increased transparency, as changes in monetary policy might have had a counterbalancing effect. In other words, our results leave us undecided as to whether there is evidence of increased transparency. The fact that we do not find any strong patterns in the US data means that our US data are not useful in redressing our indecision.

---

increase in predictability in the United Kingdom took place before 1997. See also our earlier footnote on the drawback of using breakpoints in event studies.

[23] There are some subtle differences. An explanation of the construction of the series, and a justification for giving it precedence over the standard yields series, are given in Appendix D.

## 8 The effect of macro announcements on UK volatility and trades (the *micro* hypothesis)

Having examined whether there is evidence of a change in prices one hour after a macroeconomic announcement, we now analyse how this price adjustment occurs during the hour following the news release. Referring back to our *micro* hypothesis, we are interested in both the immediate reaction of price volatility (say, in the first minute following an announcement), and the reaction over a longer period (the first hour following an announcement). To address these questions, Section 8.1.considers the reaction of volatility, over successive one- and five-minute intervals out to an hour after announcements. Section 8.2.analyses the reaction of trading volume to new information and completes our tests of the *micro* hypothesis.

### *8.1 The effects on volatility*

Price volatility is defined as the absolute return over a short trading interval, either one or five minutes. Our regressions relate the reaction of this volatility measure to macro news via simple announcement dummies.[24] The regressions are estimated for the first five one-minute and the first twelve five-minute periods following announcements. Doing so, we can obtain a profile of price volatility over the entire hour following a macroeconomic announcement. Instead of using post-BI dummies (as in Sections 6 and 7), we run separate regressions for the pre- and post-BI periods. Table 8 shows the pre- and post-BI coefficients alongside each other, with the results for the one-minute regressions in Panel A and the results for the five-minute regressions in Panel B. We show results only for indicators to which there was a significant reaction at the 5% level in the first five-minute period.

Insert Table 8 here

Consider first the higher-frequency results for the first five one-minute periods (Panel A), which contain information about the immediate market reaction to an announcement. For Preliminary GDP, Retail Sales and RPI/RPIX, price volatility appears to rise more quickly in the first minute. Moreover, most of the price adjustment occurs in the first minute, thereafter it fades out rapidly, as shown by the much smaller coefficients in the post BI sample. In contrast, Average Earnings/Unemployment announcements see a larger rise in the first minute post BI, but volatility

[24] In considering the results of the price volatility regressions, we should note that the regressions are only done for dummy variables, and thus do not account for the differing size of the news content of the announcements. This is undesirable, but unavoidable given the necessity of using absolute returns.

remains elevated in the next few minutes. There is no evidence that the price adjustment is occurs more quickly post-BI. Price volatility following the Industrial Production seems generally smaller post-BI, consistent with the smaller change in price observed in the macro tests of Section 6, but the reaction, albeit smaller, appears no slower than in the pre-BI period. The reaction of PPI Input/Output is barely significant pre-BI and not significant post-BI. Taken together, the one-minute tests show a sharper immediate reaction in the post-BI period for most indicators, but present a more heterogeneous picture regarding the price adjustment after the first minute.

The five-minute return regressions illustrate the price adjustment that occurs over the hour following the macro announcements. Except for Preliminary GDP and Retail Sales, the reaction of price volatility over the first five minutes following an announcement is smaller post BI. Panel B of Table 8 further shows that the response of price volatility over the remainder of the hour is mixed. In most cases, it drops sharply in the second five-minute interval. But it remains elevated for the Average Earnings/Unemployment announcement, with volatility post BI significant in each five-minute period out to 35 minutes. There is also evidence of higher price volatility following the Preliminary GDP announcement. There is, however, less sustained price volatility in the post-BI period following the Industrial Production and RPI/RPIX announcements. Finally, the evidence on volatility following the PPI Input/Output and Retail Sales announcements is ambiguous, with volatility being significant and of similar size for similar periods after each. Overall, the picture obtained from the five-minute data is quite unclear, with five-minute volatility not appearing to decline any quicker or slower in the hour following macro announcements in the post-BI period. As such, there is little support for the hypothesis that greater transparency may have contributed to a faster price adjustment, and thus a quicker decline in five-minute volatility following macro announcements.

So far, our tests are supportive of a sharper initial reaction post BI, but show no strong evidence of a faster price adjustment occurring thereafter, as postulated in the *micro* hypothesis. At this point, it is worth recalling the cumulative returns shown in Charts 7 and 8, which suggested that, on aggregate across all the indicators, the total price change was completed more quickly in the post-BI period. How does one reconcile the evidence on cumulative returns with that of the five-minute regression results reported in Table 8? One possible answer is that, as suggested by Charts 7 and 8, the price level reaches its new 'equilibrium' quicker (because the information is assimilated more rapidly), but that high-frequency volatility around the new level continues for as long as it did pre-BI.

We test this idea by repeating the one- and five-minute tests using cumulative returns rather than one- or five-minute absolute returns.[25] Rather than showing the results in tabular form, we present them in graphic form in Charts 9-16. Each chart shows pre- and post-BI lines, with one-standard error bands, which are plotted using the coefficient values on the indicator from different horizon pre- and post-BI regressions.

Insert Charts 9-14 here

For all six of the indicators, the pre-BI cumulative reaction line suggests a sharp price adjustment in the first five minutes, with further, small price adjustments occurring more gradually across much of the remainder of the first hour. This gives the lines an asymptotic appearance. For the first four indicators, Average Earnings, GDP (preliminary), Industrial Production and Retail Sales, the post-BI lines differ markedly from the pre-BI lines, with almost all of the change in price appearing to occur in the first minute following an announcement and the lines remaining largely flat thereafter[26]. For RPI, the differences in reaction pre- and post-BI are partly obscured by the much smaller overall reaction post-BI. The post-BI reaction does not seem as pronounced as for the other indicators, but still looks a little sharper than the pre-BI reaction. Finally, the Unemployment indicator appears to elicit the slowest reaction in both periods, perhaps suggesting that the market finds it most difficult to interpret.[27] The reaction to this indicator in the post-BI period does not seem sharper than in the pre-BI period, but nor does it seem obviously slower.[28]

To conclude, the cumulative-returns regressions confirm our earlier impression of a sharper initial reaction to announcements post BI. Of all the tests, they provide the strongest evidence that the subsequent price change occurs more rapidly than in the pre-BI period, although this is

---

[25] Since these regressions use cumulative, rather than absolute returns, the independent variable is the surprise variable. As a result, the effects of Average Earnings/Unemployment and RPI/RPIX can be disentangled.
[26] The results for Retail Sales may appear at first to contradict the volatility results, with a smaller price adjustment occurring in the first minute post-BI than pre-BI (in the volatility regressions, volatility in the first minute is higher in the post-BI period). This can be explained by recalling that the volatility variable is regressed on a simple dummy for Retail Sales announcement days, and thus does not account for the size of the surprise, as does the surprise variable in the cumulative return regressions. The distribution of surprises in the Retail Sales announcement does not appear to change in the post-BI period.
[27] Some caution should be exercised in interpreting the results for the Unemployment indicator, as the expectations series appeared to be a poor proxy for market expectations (see Section 4).
[28] Considering again the volatility result for the Average Earnings/Unemployment indicator, Chart 14 suggests that the prolonged significant volatility in the post-BI period is most likely a result of the smaller error with which volatility is measured in the post-BI period.

more pronounced for some indicators than for others. In the next Section, we test whether a similar pattern is seen in the trading volume data.

### *8.2 The effect on trades*

We now consider the reaction of trades to macro announcements. Recall that our *micro* hypothesis postulated a sharper immediate response in trading volume on announcement days post-BI, and a faster levelling off in subsequent trading rounds. To test this assertion, we estimate Equation (1) for successive five-minute intervals over the period of one hour. As stated earlier, our dependent variable is the total number of trades in a five-minute interval, expressed as a deviation from its ten-day moving average, and standardised by its rolling standard deviation. Separate regressions are run for the pre- and post-BI periods, with an additional break in July 1999 to account for the introduction of electronic trading on LIFFE.[29] The results are reported in Table 9, with row 1 reporting the pre-BI results, row 2 the post-BI results prior to electronic trading, and row 3 the post-BI, post-electronic results.

Insert Table 9 here

Focusing first on the results prior to electronic trading, volume in the first five minutes following an announcement is highest in the first five minutes after an announcement and tapers off slowly thereafter. Not much changes between 1997 and 1999. The immediate volume response rises in the post-BI period in only two out of six cases: preliminary GDP and RPI/RPIX. It is slightly lower in response to Average Earnings/Unemployment, Industrial Production, and PPI input/output releases, and roughly equal following Retail Sales news. When looking at the next 55 minutes, it is not apparent that volatility declines either more quickly or more slowly in the post-BI period. Taken together, these trade regressions provide neither support for, nor evidence against, our *micro* hypothesis.

Table 9 also presents the results for the final four years of our sample period when all trading was done electronically. When comparing these with the post-BI results prior to electronic trading, we observe a marked increase in the volume response during the first five minutes of trading on

[29] The results were robust across the two other specifications of trades regressions that we tried; one using absolute surprises (similar to Equation 3) and one using binary surprises (in which we separated surprises into small and large surprises). The results are available from the authors upon request.

announcement days. Moreover, in all six cases, the estimated coefficients are larger than in the pre-BI period. The subsequent adjustment process does not appear to have changed markedly.

So does this mean that we do not find support for our *micro* hypothesis until after July 1999? Prior to that date, the increased transparency of monetary policy and accompanying reduction in adverse selection risk did not translate into higher trading volumes immediately following macroeconomic announcements, yet it did after the introduction of electronic trading. One interpretation is that the results after July 1999 are a reflection of the changed nature of trading conditions. Markets with electronic trading facilities are generally believed to be more transparent and allow for more rapid dissemination of new information into traded prices (CGFS, (2001)). Hence, electronic trading may have led to a reduction in adverse selection risk, similar to that associated with increased policy transparency.[30]

Although our regression tests do not allow us to distinguish between the effects of increased transparency and the impact of electronic trading, a closer look at the distribution of trades might shed some light on the issue. Table 10 presents summary statistics for each of the three periods.

Insert Table 10 here

Considering first only the pre-BI and post-BI, pre-electronic trading periods, the distributions seem quite similar. The means are very similar, while the standard deviation is a little higher in the latter period. In addition, some of the probability density seems to have moved away from the upper tail in the latter period. The change in the distribution after electronic trading is introduced is much more marked. Both the mean and the standard deviation roughly double. The weight of the upper tail, however, remains similar to that in the post-BI, pre-electronic trading period.

At first pass, these summary statistics suggest that there may be no reason to expect larger coefficients post-BI in the post-electronic trading period. Although trading volume doubles, so too does the standard deviation, meaning that if the number of trades in a given period following an announcement also doubled in the post-electronic period, our trades variable (defined in Section 5.1) would take the same value as in the pre-electronic trading period. This suggests that the increase in trading volume alone cannot account for the larger coefficients observed in the first five minute, post-electronic trading period regressions.

*8.3 Concluding remarks*

The price volatility and cumulative return regressions show a sharper immediate reaction post-BI, with almost all of the price adjustment taking place within the first minute. There is also some evidence of price volatility declining more rapidly subsequently. The trades regressions do not pick up a sharper initial reaction of trading volume until after the switch to electronic trading. There is no evidence of trading volume subsiding quicker in the post-BI period after the first five minutes.

Considered together, the results provide some support for our benchmark *micro* hypothesis, based on the information asymmetry faced by uninformed traders. Prices react more sharply initially, as the more transparent environment reduces the cost to uninformed traders of trading. Information is then assimilated into price more quickly, though more clearly so for some indicators, with the result that prices adjust more quickly in the post-BI period. However, contrary to our expectations, we find little corroborative evidence in our trading volume tests.

The results further indicate that the introduction of electronic trading on LIFFE may have affected price and trade dynamics, as it too may have contributed to a reduction in adverse selection risk. More generally, trading practices changed vastly during our sample period, as market participants acquired more sophisticated trading and information processing tools. Hence, our empirical tests may reflect broader market developments, and not just greater monetary policy transparency. A similar question is addressed by Swanson (2003) who tests whether increased predictability of US policy rates during the 1990s was the result of increased monetary policy transparency, or increased sophistication of financial market participants, explained by better econometric techniques and/or increased computing power. Comparing market forecasts of short-term interest rates and real variables, he finds that while the former had improved in accuracy, the latter had actually deteriorated. Hence, he concludes that improved transparency had indeed helped predictability.

[30] Allen and Hawkins (2002) point out, however, that the effects of electronic trading on transparency are controversial, and depend on specific market settings. See also Ganley et al (1998).

## 9 Conclusion

In this paper, we test the high-frequency reaction of market interest rates to the release of macroeconomic indicators, in an attempt to find evidence of increased monetary policy transparency since the Bank of England was granted independence in June 1997. We put forward two hypotheses; first, that increased transparency should have led to a larger change in price following macroeconomic news, and, second, that increased transparency should lead to prices incorporating macroeconomic news more quickly.

Using our empirical tests to make inference about our two hypotheses, we find evidence to reject our *macro* hypothesis, as the change in yield following macro announcements does not appear significantly different post-BI. But we find partial support for our *micro* hypothesis, as there is evidence of prices adjusting more quickly, perhaps as traders are more willing to trade.

We conclude that the changes in the behaviour of interest rates following macro announcements since 1997 could be consistent with increased transparency. We are mindful, however, of the complexity of the mapping between interest rate reactions and monetary policy. We can not rule out (and have not tested) that the reaction of market interest rates to macroeconomic announcements may have been affected by aspects of monetary policy other than transparency, by changes in the macroeconomic environment itself, or by other, unidentified factors.

Indeed, monetary policy in the United Kingdom changed in more than one aspect during the past decade. The introduction of the inflation target in 1992 falls outside our sample period, and so did the adoption of a more gradualist approach towards rate setting.[31] Previous research by Lildholdt and Vila Wetherilt (2003) has shown that the resulting increase in interest rate predictability dates back to 1992, but was a gradual process, rather than a one-off jump coinciding with the adoption of the new framework. When the Bank of England was granted operational independence in 1997, the nature of decision making as well as the amount of information available on this process changed. The empirical evidence in the present paper shows that the adjustment in financial markets cannot easily be captured by a few simple hypotheses.

## Appendices

### Appendix A: Tests of unbiasedness of UK expectations series

Our reliance upon MMS survey expectations merits a brief consideration of their properties. For our measure of surprises to be valid, we would hope that the expectations are both unbiased and efficient. Recent papers have found MMS expectations to be unbiased and more efficient than regression-based forecasts for most US and euro area releases[32].

Following Balduzzi et al (2001) and Gravelle and Moessner (2001), we perform two simple tests. First, we test for unbiasedness in the MMS expectations series by running the following regressions:

$$A_t = \alpha + \beta E_t + \varepsilon_t ,$$

where $A$ is the actual announcement and $E$ the survey expectation. The unbiasedness test is a Wald test of the joint hypothesis that the constant $\alpha$ equals 0 and the coefficient on the expectation $\beta$ equals 1. Table 1 shows the regression coefficients together with the F-statistics associated with the Wald test. The results show that unbiasedness is rejected at the 5% level for three out of the 16 indicators only: Industrial Production, M0 and Unemployment.

**Table A1: tests for unbiasedness**

| | alpha* | | beta* | | Wald** test |
|---|---|---|---|---|---|
| Average Earnings | 0.31% | **(0.026)** | 0.93 | **(0.031)** | 0.06 |
| Ex-EU trade | -£0.21bn | **(0.049)** | 0.89 | (0.092) | 0.26 |
| GDP prel (Q/Q) | 0.04% | (0.116) | 0.93 | (0.143) | 0.98 |
| GDP prov (Q/Q) | -0.01% | (0.811) | 1.01 | (0.838) | 0.86 |
| GDP final (Q/Q) | -0.01% | (0.688) | 1.02 | (0.771) | 0.27 |
| Global Trade | -£0.01bn | (0.823) | 1.01 | (0.808) | 0.54 |
| Ind. Prod. (M/M) | -0.21% | **(0.010)** | 1.07 | (0.720) | **0.00** |
| M0 prov (Y/Y) | -1.89% | (0.312) | 1.33 | (0.283) | 0.11 |
| M4 prov (Y/Y) | 0.33% | (0.261) | 0.97 | (0.347) | 0.33 |
| PPI Input (M/M) | -0.13% | (0.070) | 1.43 | **(0.000)** | **0.00** |
| PPI Output (M/M) | -0.01% | (0.540) | 0.90 | (0.390) | 0.12 |
| PSBR | -£0.16bn | (0.360) | 1.06 | (0.060) | 0.11 |
| Retail Sales (M/M) | -0.04% | (0.651) | 1.14 | (0.501) | 0.69 |
| RPI (M/M) | 0.02% | (0.569) | 0.97 | (0.594) | 0.11 |
| RPIX (Y/Y) | 0.13% | (0.253) | 0.96 | (0.317) | 0.23 |
| Unemployment | -5k | **(0.000)** | 0.95 | (0.728) | **0.00** |
| Base rate (Δ in b.p.) | -0.02 | (0.13) | 1.31 | **(0.048)** | 0.09 |

* coefficients are given in their original units, p-values of t-tests are given in brackets
** p-values of the F-test are given

[31] Lildholdt and Vila Wetherilt (2003) show that both the number of large rate changes and the occurrence of rate reversals declined markedly after 1992.

**Appendix B: Tests of efficiency of reaction of Short Sterling and Long Gilt contracts to macro news**

The price of each of the assets should react only to the surprise component of macroeconomic announcements if the markets are semi-strong efficient.[33] We follow Moessner et al. (2000) in testing this by regressing change in yields on surprises and the median expectation from the MMS survery (Equation X).

$$(x) \qquad Y_t^{c,i} = \alpha^{c,i} + \sum^{K} \beta_{1,k}^{c,i} S_{i,k} + \sum^{K} \beta_{2,k}^{c,i} E_{i,k} + e_t^{c,i} + \beta_3^{break,c}$$

If markets react efficiently and significantly to macroeconomic announcements, and the MMS expectations series are unbiased, we would expect the $\beta_1^{c,i}$ coefficients to be significantly different to zero and the $\beta_2^{c,i}$ coefficients not to be significantly different to zero. This appears to be the case for the reaction of the Short Sterling contract to all of the indicators (none of the $\beta_2$ s are significant at the 5% level and only the PSBR is significant at the 10% level), but the Long Gilt contract appears to react significantly to the expectation of the Unemployment outturn (Table B1). Recalling that the Unemployment expectations series failed the test of unbiasedness, as survey respondents on average over-estimated unemployment over the sample period, the test of efficiency may not be valid for the Unemployment indicator.

Given that the Short Sterling and Long Gilt futures markets appear to react efficiently to the news contained in the macroeconomic indicators, we can conclude that it is more appropriate to test the reaction of yields to surprises than to dummies in both markets and thus that we are justified in focusing on the surprise regressions for each in our paper.

[32] See for example Balduzzi, Elton and Green (2001) and Ehrmann and Fratzscher (2002).
[33] Add ref.

## Table B1: test of market efficiency

| announcement | SS | | LG | |
|---|---|---|---|---|
| constant | 0.00 | | 0.00 | |
| (surprise) Average Earnings | -0.06 | (0.000) | -0.33 | (0.000) |
| (surprise) Ex-EU Trade | 0.00 | (0.036) | 0.00 | (0.644) |
| (surprise) GDP (final) | -0.03 | (0.185) | -0.09 | (0.182) |
| (surprise) GDP (prel) | -0.10 | (0.019) | -0.26 | (0.156) |
| (surprise) GDP (prov) | -0.11 | (0.000) | -0.25 | (0.000) |
| (surprise) Global Trade | 0.00 | (0.004) | 0.00 | (0.502) |
| (surprise) Industrial Production | -0.02 | (0.000) | -0.07 | (0.000) |
| (surprise) M0 (prov) | 0.00 | (0.178) | -0.03 | (0.022) |
| (surprise) PPI Input | -0.01 | (0.000) | -0.04 | (0.000) |
| (surprise) PPI Output | 0.00 | (0.568) | 0.00 | (0.561) |
| (surprise) PSBR | 0.00 | (0.702) | 0.00 | (0.052) |
| (surprise) Retail Sales | -0.04 | (0.000) | -0.13 | (0.000) |
| (surprise) RPI/RPIX | -0.10 | (0.000) | -0.57 | (0.000) |
| (surprise) RPIX | -0.03 | (0.146) | -0.04 | (0.765) |
| (surprise) Unemployment | 0.00 | (0.001) | 0.00 | (0.000) |
| (expectation) Average Earnings | 0.00 | (0.625) | 0.00 | (0.345) |
| (expectation) Ex-EU Trade | 0.00 | (0.821) | 0.00 | (0.908) |
| (expectation) GDP (final) | 0.00 | (0.140) | -0.01 | (0.230) |
| (expectation) GDP (prel) | 0.00 | (0.684) | 0.00 | (0.886) |
| (expectation) GDP (prov) | 0.00 | (0.735) | -0.03 | (0.061) |
| (expectation) Global Trade | 0.00 | (0.982) | 0.00 | (0.930) |
| (expectation) Industrial Production | 0.01 | (0.092) | 0.02 | (0.240) |
| (expectation) M0 (prov) | 0.00 | (0.406) | 0.00 | (0.849) |
| (expectation) PPI Input | 0.00 | (0.777) | 0.01 | (0.192) |
| (expectation) PPI Output | 0.01 | (0.363) | 0.00 | (0.985) |
| (expectation) PSBR | 0.00 | (0.071) | 0.00 | (0.557) |
| (expectation) Retail Sales | 0.00 | (0.462) | 0.02 | (0.456) |
| (expectation) RPI/RPIX | 0.00 | (0.791) | -0.03 | (0.315) |
| (expectation) RPIX | 0.00 | (0.141) | 0.00 | (0.761) |
| (expectation) Unemployment | 0.00 | (0.148) | 0.00 | (0.011) |
| $R^2$ | 0.34 | | 0.27 | |
| sample size | 2490 | | 2490 | |

## Appendix C: Summary statistics for US macroeconomic indicators

### Table C1: Macroeconomic and monetary policy announcements for the United States

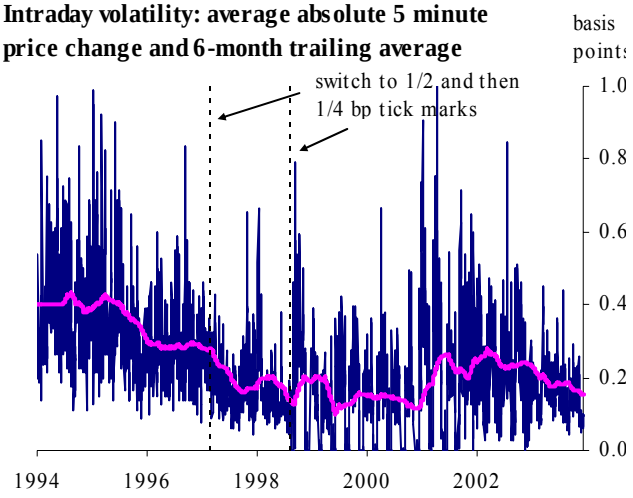
| | | | | | | | *outturn* | | *expectation* | | *surprise* | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| indicator | time | from | | to | frequency | number | mean | s.d. | mean | s.d. | mean | s.d. |
| CPI | 13:30 | Jan 94 | - | Dec 03 | monthly | 106 | 0.2% | 0.2% | 0.2% | 0.1% | 0.0% | 0.1% |
| Current Account | 15:00 | Mar 94 | - | Dec 98 | quarterly | 15 | -£35.9bn | £23.9bn | n/a | n/a | n/a | n/a |
| GDP Advance | 13:30 | Jan 94 | - | Oct 03 | quarterly | 36 | 3.2% | 1.9% | 2.7% | 1.7% | 0.4% | 0.8% |
| GDP Preliminary | 13:30 | Mar 94 | - | Nov 03 | quarterly | 35 | 3.3% | 2.0% | 3.2% | 2.0% | 0.2% | 0.3% |
| GDP Final | 13:30 | Mar 94 | - | Sep 03 | quarterly | 34 | 3.1% | 2.1% | 3.1% | 2.0% | 0.0% | 0.3% |
| Hourly Earnings | 13:30 | Jan 94 | - | Dec 03 | monthly | 106 | 0.3% | 0.2% | 0.3% | 0.1% | 0.0% | 0.2% |
| Housing Starts | 13:30 | Jan 94 | | Dec 03 | monthly | 107 | 1583k | 161k | 1563k | 136k | 20k | 73k |
| Industrial Prod. | 14:15 | Jan 94 | - | Dec 03 | monthly | 106 | 0.2% | 0.5% | 0.1% | 0.3% | 0.0% | 0.3% |
| Initial Claims | 13:30 | Jan 94 | | Sep 99 | weekly | 235 | 330k | 30k | 331k | 27k | -2k | 15k |
| NAPM | 15:00 | Jan 94 | - | Nov 01 | monthly | 81 | 51.0 | 4.6 | 51.3 | 4.3 | -0.3 | 2.0 |
| Nonfarm Payrolls | 13:30 | Jan 94 | - | Dec 03 | monthly | 106 | 108.7 | 186.5 | 130.8 | 120.9 | -22.1 | 116.4 |
| PPI | 13:30 | Jan 94 | - | Dec 03 | monthly | 105 | 0.1% | 0.4% | 0.2% | 0.2% | 0.0% | 0.3% |
| Retail Sales | 13:30 | Jan 94 | - | Dec 03 | monthly | 105 | 0.4% | 0.9% | 0.3% | 0.5% | 0.0% | 0.7% |
| Unemployment Rate | 13:30 | Jan 94 | - | Dec 03 | monthly | 106 | 5.0% | 0.7% | 5.1% | 0.7% | 0.0% | 0.1% |
| policy rate | | | | | | | | | | | | |

## Appendix D: relative price volatility series for the EuroDollar contract

In March 1997 and August 1998, the tick size in which the contract is traded was reduced from 1 basis point to ½ and then ¼ of a basis point, allowing traders a finer denomination in which to react to news than previously. It is possible that this led to a reduction in intra-day volatility, when measured in return space rather than tick space. A simple visual inspection of the EuroDollar price volatility series fails to give evidence for or against this possibility, with a marked decline around 1997 coming after a period of slighter decline over the period 1994-1997, making it hard to attributed the decline to the change in tick size (Chart D.1).

**Chart D.1**

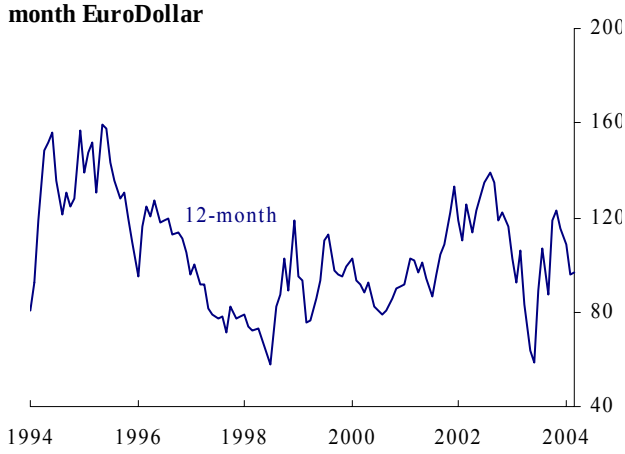
**Intraday volatility: average absolute 5 minute price change and 6-month trailing average**




Intraday volatility is the mean absolute 5-minute change in yield, calculated for individual working days. The 6-month average is a 6-month trailing mean of these figures.

**Chart D.2**

**US rate uncertainty: implied volatility 12-month EuroDollar**




An obvious approach to the decline in volatility would be to treat it as a simple structural break and account for it with dummies. This option has the disadvantage of the break, or breaks, coming at a very similar time to Bank Independence, which would prohibit us from testing for a smaller reaction to indicators in the post-BI period. It has a further disadvantage that we may be mistaking the apparent structural break for an entirely exogenous fall in volatility, which would make the dummy approach invalid. The option implied volatility on a 12-month EuroDollar contract suggests that this case should be given some consideration (Chart D.2).

An alternative approach, which we adopt, is to focus our regressions on the reaction of the EuroDollar contract to an indicator at time *t*, above and beyond the prevailing volatility at time *t*. To do this, we construct a change in yield series similar to the trades series used for the Short Sterling contract. If the change in yield at time *t* is positive, we take the change in yield minus the absolute average change in yield, if it is negative, we take the change in yield plus the average absolute change in yield and if it is zero, we leave it at zero. In all 3 cases, we scale the variable

by dividing by the average absolute change in yield. We utilise this slightly different approach because the change in yield series has a symmetric distribution around zero, rather than being always non-negative.

## Annex 1: Tables

*For all tables, * denotes significance at the 5% level and ** denotes significance at the 1% level.*

## Table 1

| | | | | | | | *outturn* | | *survey* | | *surprise* | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| indicator | time | from | | to | frequency | number | mean | st. dev. | mean | sd | mean | sd |
| Average Earnings | 09:30 | Jan 94 | - | Sep 03 | monthly | 113 | 4.1% | 0.7% | 4.1% | 0.7% | 0.0% | 0.2% |
| Ex-EU trade | 09:30 | Oct 96 | - | Sep 03 | quarterly | 83 | -£1.7bn | £0.8bn | -£1.7bn | £0.7bn | £0.0bn | £0.4bn |
| GDP prel (Q/Q) | 09:30 | Jan 94 | - | Jul 03 | quarterly | 40 | 0.6% | 0.3% | 0.6% | 0.2% | 0.0% | 0.1% |
| GDP prov (Q/Q) | 09:30 | Feb 94 | - | Aug 03 | quarterly | 39 | 0.6% | 0.3% | 0.6% | 0.3% | 0.0% | 0.1% |
| GDP final (Q/Q) | 09:30 | Mar 94 | - | Jun 03 | quarterly | 39 | 0.6% | 0.3% | 0.5% | 0.3% | 0.0% | 0.1% |
| Global Trade | 09:30 | Jan 94 | - | Sep 03 | monthly | 118 | -£1.8bn | £1.0bn | -£1.8bn | £0.9bn | £0.0bn | £0.3bn |
| Ind. Prod. (M/M) | 09:30 | Jan 94 | - | Sep 03 | monthly | 117 | 0.0% | 0.8% | 0.2% | 0.4% | -0.2% | 0.6% |
| M0 prov (Y/Y) | 09:30 | Jan 94 | - | Oct 01 | monthly | 85 | 6.7% | 1.3% | 6.4% | 0.7% | 0.2% | 0.6% |
| M4 prov (Y/Y) | 09:30 | Jan 94 | - | Sep 03 | monthly | 111 | 7.7% | 2.3% | 7.8% | 2.4% | 0.1% | 0.7% |
| PPI Input (M/M) | 09:30 | Jan 94 | - | Sep 03 | monthly | 117 | 0.1% | 1.3% | 0.1% | 0.7% | -0.1% | 0.8% |
| PPI Output (M/M) | 09:30 | Jan 94 | - | Sep 03 | monthly | 117 | 0.1% | 0.2% | 0.2% | 0.2% | 0.0% | 0.2% |
| PSBR | 09:30 | Jan 94 | - | Sep 03 | monthly | 110 | £0.8bn | £6.2bn | £0.9bn | £5.6bn | -£0.1bn | £1.7bn |
| Retail Sales (M/M) | 09:30 | Jan 94 | - | Sep 03 | monthly | 117 | 0.3% | 0.7% | 0.3% | 0.3% | 0.0% | 0.6% |
| RPI (M/M) | 09:30 | Jan 94 | - | Sep 03 | monthly | 111 | 0.2% | 0.4% | 0.2% | 0.4% | 0.0% | 0.1% |
| RPIX (Y/Y) | 09:30 | Jan 94 | - | Sep 03 | monthly | 111 | 2.5% | 0.4% | 2.5% | 0.4% | 0.0% | 0.1% |
| Unemployment | 09:30 | Jan 94 | - | Sep 03 | monthly | 110 | -15k | 18k | -10k | 12k | -5k | 14k |

## Table 2: Macro tests - UK yield reaction to UK dummies

| | *Short Sterling* | | *Long Gilt* | |
|---|---|---|---|---|
| | *Eq. (1)* | | *Eq. (1)* | |
| C | 0.76 | ** | 2.68 | ** |
| Average Earnings/Unemployment | 1.47 | ** | 6.87 | ** |
| GDP (final) | 0.12 | | 0.06 | |
| GDP (prel) | 0.66 | * | 2.27 | ** |
| GDP (prov) | 0.19 | | 1.53 | * |
| Global Trade | -0.15 | | 1.17 | |
| Industrial Production | 0.70 | ** | 3.11 | ** |
| M0 (prov) | -0.07 | | 0.66 | |
| PPI Input/Ouput | 0.28 | | 2.02 | ** |
| PSBR | -0.03 | | 1.16 | |
| Retail Sales | 1.52 | ** | 6.61 | ** |
| RPI/RPIX | 1.63 | ** | 5.61 | ** |
| $R^2$ | 0.11 | | 0.20 | |
| Adjusted $R^2$ | 0.11 | | 0.19 | |
| sample size | 2489 | | 2489 | |

The coefficients on the indicators measure the difference between the absolute change in yield after a given indicator announcement and the absolute change in yield generally observed in the hour following the 09:30 ONS announcement time. A positive coefficient denotes a larger change in yield than is generally observed, so a positive and significant coefficient suggests there is a significant reaction to a given indicator. Coefficients in the Short Sterling regression are in basis points. Coefficients in the Long Gilt regression are 10,000 times the absolute change in log price.

## Table 3: Macro tests - UK yield reaction to UK surprises

| announcement surprises | *Short Sterling* *Eq. (2)* | | *Long Gilt* *Eq. (2)* | |
|---|---|---|---|---|
| C | -0.09 | ** | -0.11 | |
| Average Earnings | 1.69 | ** | 9.12 | ** |
| GDP (final) | 0.07 | | -0.42 | |
| GDP (prel) | 1.26 | ** | 2.66 | |
| GDP (prov) | 0.68 | | 4.70 | |
| Global Trade | -0.17 | | 1.36 | |
| Industrial Production | 1.30 | ** | 4.68 | ** |
| M4 (prov) | 0.34 | ** | 0.61 | |
| PPI Input | 0.32 | ** | 2.14 | ** |
| PPI Output | 0.04 | | -0.55 | |
| PSBR | -0.05 | | 3.61 | * |
| Retail Sales | 1.97 | ** | 6.45 | ** |
| RPI | 1.72 | ** | 13.35 | ** |
| RPIX | 0.64 | | 0.71 | |
| Unemployment | -0.80 | ** | -5.42 | ** |
| $R^2$ | 0.30 | | 0.09 | |
| Adjusted $R^2$ | 0.29 | | 0.08 | |
| sample size | 2489 | | 2489 | |

The coefficients on the indicators measure the change in yield associated with a surprise in a given indicator. A positive indicator coefficient suggests that a larger than expected outturn in a given indicator causes the yield on either Short Sterling or Long Gilt to increase in the 60 minutes following the indicator's announcement. Note that a positive surprise in the Unemployment indicator would generally be expected to lead to a fall in yields, so the negative coefficient on the Unemployment indicator is encouraging. All the other indicators would generally be expected to be yield positive. The scale of the Short Sterling coefficients is basis points per 1 standard deviation surprise. The scale of the Long Gilt contract can be thought of as a rough approximation to basis points per 1 standard deviation. It is comparable across indicators and across Long Gilt regressions on surprises. It is -10,000 times the log change in the price of the contract per 1 standard deviation surprise in the indicator.

## Table 4: Macro tests - UK yield reaction to UK surprises, with post-BI dummies

| announcement surprises | *Short Sterling* *Eq. (4)* | | *Long Gilt* *Eq. (4)* | |
|---|---|---|---|---|
| constant | 0.01 | | 0.32 | |
| Average Earnings | 1.95 | ** | 17.60 | ** |
| GDP (prel) | 1.73 | | 6.76 | * |
| Industrial Production | 2.48 | * | 8.26 | |
| M4 (provisional) | 0.71 | * | 2.00 | |
| PPI Input | 0.78 | * | 8.04 | * |
| PPI Output | 1.56 | | 1.88 | |
| Retail Sales | 3.04 | ** | 10.94 | * |
| RPI | 2.70 | * | 21.71 | ** |
| RPIX | -0.40 | | 3.32 | |
| Unemployment | -1.01 | ** | -6.32 | ** |
| PBI x Average Earnings | -0.44 | | -12.02 | ** |
| PBI x GDP (prel) | -0.64 | | -5.73 | |
| PBI x Industrial Production | -1.78 | | -5.30 | |
| PBI x M4 (provisional) | -0.50 | | -1.55 | |
| PBI x PPI Input | -0.57 | | -7.35 | * |
| PBI x PPI Output | -1.54 | | -2.44 | |
| PBI x Retail Sales | -1.45 | * | -5.58 | |
| PBI x RPI | -1.55 | | -13.47 | |
| PBI x RPIX | 1.64 | | -1.87 | |
| PBI x Unemployment | 0.64 | | 6.28 | |
| $R^2$ | 0.31 | | 0.11 | |
| sample size | 2489 | | 2489 | |

The yields regressions are equivalent to those in table 4, with the addition of post-BI multiplicative slope dummies for each of the indicators. A positive indicator coefficient in the yields regressions now suggests that a larger than expected outturn in a given indicator pre-BI caused the yield on either Short Sterling or Long Gilt to increase in the 60 minutes following the indicator 's announcement. A positive coefficient on the post-BI dummy coefficients suggests that yields increased by more for a given surprise post-BI than pre-BI, a negative coefficient suggests that yields increased by less and a coefficient that is not significant suggests that there was not a significantly different change in yields for a given surprise post-BI.

## Table 5: Macro tests - UK yield reaction to US surprises, with post-BI dummies

| | *Short Sterling* | | *Long Gilt* | |
|---|---|---|---|---|
| announcement surprises | *Eq. (4)* | | *Eq. (4)* | |
| constant | 0.00 | | 0.14 | |
| CPI | 0.25 | | 3.03 | |
| GDP (adv) | 0.44 | | 3.93 | |
| GDP (prel) | 0.51 | ** | 5.65 | * |
| Hourly Earnings | 0.31 | | 8.63 | ** |
| Housing Starts | 0.54 | * | 6.19 | |
| Initial Claims | 0.11 | | -0.44 | |
| Non-Farm Payrolls | 0.97 | ** | 28.03 | ** |
| PPI | 0.16 | | 2.65 | |
| Retail Sales | 0.73 | | 6.08 | |
| Unemployment | 0.07 | | -2.72 | |
| PBI x CPI | -0.24 | | -0.29 | |
| PBI x GDP (adv) | 0.05 | | -1.14 | |
| PBI x GDP (prel) | -0.19 | | -7.13 | * |
| PBI x Hourly Earnings | -0.44 | | -4.10 | |
| PBI x Housing Starts | -0.47 | | -4.70 | |
| PBI x Initial Claims | -0.20 | * | -0.75 | |
| PBI x Non-Farm Payrolls | -0.50 | | -19.53 | ** |
| PBI x PPI | -0.23 | | -1.33 | |
| PBI x Retail Sales | -0.10 | | 0.23 | |
| PBI x Unemployment | -0.35 | | -5.29 | |
| $R^2$ | 0.01 | | 0.04 | |
| sample size | 4978 | | 4978 | |

The regressions can be interpreted in the same manner as the regressions on UK indicators in Table 5. Note that a positive surprise in the Initial Claims and Unemployment indicators would normally be regarded as lowering yields.

## Table 6: Macro tests - UK yield reaction to UK surprises, with post-BI dummies and implied volatility dummies

| | *Short Sterling* | | *Long Gilt* | |
|---|---|---|---|---|
| announcement surprises/dummies | *Eq. (9)* | | *Eq. (9)* | |
| constant | -0.01 | | 0.06 | |
| Average Earnings | 1.75 | ** | 14.30 | ** |
| GDP (prel) | 1.08 | ** | 4.88 | ** |
| Industrial Production | 1.96 | ** | 10.20 | ** |
| PPI Input | 1.13 | ** | 4.42 | |
| PPI Output | 3.00 | ** | 30.24 | ** |
| Retail Sales | 1.32 | | 9.43 | |
| RPI | 1.73 | ** | 17.79 | ** |
| Unemployment | -0.33 | | -1.49 | |
| PBI x Average Earnings | -0.31 | | -8.60 | ** |
| PBI x GDP (prel) | 0.31 | | -3.54 | ** |
| PBI x Industrial Production | -1.24 | ** | -6.20 | ** |
| PBI x PPI Input | -0.75 | ** | -3.83 | |
| PBI x PPI Output | -2.27 | * | -27.02 | ** |
| PBI x Retail Sales | 0.45 | | -2.42 | |
| PBI x RPI | -0.16 | | -10.82 | ** |
| PBI x Unemployment | 0.10 | | 1.38 | |
| IVOL x Average Earnings | -0.29 | | 0.06 | |
| IVOL x GDP (prel) | -0.05 | | 1.22 | |
| IVOL x Industrial Production | 0.09 | | -0.92 | |
| IVOL x PPI Input | -0.12 | | 1.72 | |
| IVOL x PPI Output | -0.82 | | -3.65 | |
| IVOL x Retail Sales | 0.88 | | 0.47 | |
| IVOL x RPI | 0.29 | | -3.22 | |
| IVOL x Unemployment | -0.62 | | -3.82 | |
| $R^2$ | 0.35 | | 0.32 | |
| sample size | 2489 | | 2489 | |

The regressions can be interpreted similarly to the regressions in Table 5. For the yields regressions, the indicator dummies now show the reaction of yields to a 1 sd surprise in the pre-BI period when forecast uncertainty was not high, the PBI dummies show the differential reaction when uncertainty was not high post-BI and the IVOL (implied volatility) dummies show the differential reaction throughout the whole sample period when forecast uncertainty was high.

## Table 7: US yield reaction to US surprises, with post-BI dummies

| | *EuroDollar* | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | *Eq. (4)* | | *Eq. (4)* | | *Eq. (4)** | | *Eq. (9)* | |
| announcement surprises/dummies | prices | | normalised trades | | normalised prices | | normalised trades | |
| constant | -0.01 | | 0.52 | ** | -0.05 | | 0.35 | ** |
| Business Inventories | -0.25 | | 0.10 | | -0.40 | | 0.89 | ** |
| Consumer Confidence | 0.54 | ** | 1.50 | ** | 1.84 | ** | 0.83 | ** |
| Construction | -0.12 | | 0.25 | | -0.25 | | 0.35 | |
| CPI | 1.03 | | 3.52 | ** | 3.18 | | 2.88 | ** |
| GDP (adv) | 1.36 | * | 2.01 | ** | 3.96 | ** | 3.72 | ** |
| GDP (final) | -0.09 | | 0.70 | | -0.51 | | 1.22 | * |
| GDP (provisional) | 0.24 | | 1.44 | * | 0.49 | | 1.85 | * |
| Hourly Earnings | 1.42 | ** | 3.10 | ** | 4.48 | ** | | |
| Housing Starts | 0.74 | ** | 1.57 | ** | 1.96 | ** | 1.41 | ** |
| Initial Claims | -0.57 | ** | 1.44 | ** | -1.45 | ** | 1.53 | ** |
| Industrial Production | 1.00 | ** | 2.25 | ** | 2.49 | ** | 2.16 | ** |
| NAPM | 0.86 | * | 1.49 | * | 3.95 | ** | 1.70 | ** |
| Non-Farm Payrolls | 3.62 | ** | 1.93 | ** | 10.71 | ** | 8.25 | ** |
| PPI | 1.17 | ** | 3.81 | ** | 3.29 | ** | 3.59 | ** |
| Retail Sales | 3.15 | ** | 6.60 | ** | 8.18 | * | 2.31 | ** |
| Unemployment | -1.70 | ** | 1.62 | | -4.47 | * | | |
| PBI x Business Inventories | -0.01 | | 0.73 | | -0.58 | | 0.10 | |
| PBI x Consumer Confidence | 0.41 | | 0.24 | | 3.32 | | 0.62 | |
| PBI x Construction | 0.33 | | 0.63 | | 0.93 | | 1.25 | * |
| PBI x CPI | -0.96 | | -1.64 | | -2.46 | | -0.57 | |
| PBI x GDP (adv) | -0.65 | | 1.48 | * | -0.89 | | 0.13 | |
| PBI x GDP (final) | 0.38 | | -0.69 | | 2.05 | | -0.99 | |
| PBI x GDP (provisional) | -0.19 | | -0.80 | | -0.31 | | -0.84 | |
| PBI x Hourly Earnings | -0.85 | | -0.85 | | -0.67 | | | |
| PBI x Housing Starts | -0.62 | ** | -0.48 | | -1.43 | * | -0.22 | |
| PBI x Initial Claims | 0.35 | ** | -0.46 | | 0.52 | | -0.43 | |
| PBI x Industrial Production | -0.66 | | -1.18 | * | -0.78 | | -1.10 | * |
| PBI x NAPM | 0.37 | | 0.39 | | 1.98 | | -0.72 | |
| PBI x Non-Farm Payrolls | -1.73 | ** | 0.64 | | -1.20 | | -2.82 | ** |
| PBI x PPI | -0.70 | * | -1.49 | | -0.91 | | -0.62 | |
| PBI x Retail Sales | -2.36 | * | -5.39 | ** | -4.46 | | 0.02 | |
| PBI x Unemployment | 0.53 | | 0.05 | | -1.25 | | | |
| $R^2$ | 0.23 | | 0.26 | | 0.23 | | 0.38 | |
| sample size | 7446 | | 7446 | | 7446 | | 7446 | |

## Table 8: Panel A

return period (1-minute absolute returns)

| | | 1 | 2 | 3 | 4 | 5 |
|---|---|---|---|---|---|---|
| Av. Earn. / Unemploy. | pre | 0.27 | 1.26 ** | 0.52 ** | 0.22 | 0.25 |
| | post | 0.74 ** | 0.74 ** | 0.51 ** | 0.28 ** | 0.17 |
| GDP (prel.) | pre | 0.36 | 0.28 | 0.08 | 0.20 | 0.19 |
| | post | 1.05 ** | 0.37 ** | 0.19 * | 0.22 * | 0.12 |
| Ind. Prod. | pre | 0.85 ** | 0.47 ** | 0.37 * | 0.31 ** | 0.09 |
| | post | 0.57 ** | 0.20 ** | 0.17 ** | 0.12 ** | 0.07 |
| PPI Input / Ouput | pre | 0.28 | 0.39 ** | 0.15 | 0.20 ** | 0.18 |
| | post | 0.11 | 0.08 | -0.03 | 0.10 * | 0.06 |
| Retail Sales | pre | 0.73 ** | 0.78 ** | 0.55 ** | 0.42 ** | 0.16 |
| | post | 1.48 ** | 0.67 ** | 0.31 ** | 0.23 ** | 0.17 |
| RPI / RPIX | pre | 1.00 ** | 0.98 ** | 0.10 | 0.52 ** | 0.09 |
| | post | 1.16 ** | 0.44 ** | 0.15 ** | 0.28 ** | 0.17 |

## Table 8: Panel B

return period (5-mininute absolute returns)

| | | 5 | 10 | 15 | 20 | 25 | 30 | 35 | 40 | 45 | 50 | 55 | 60 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Av. Earn. / Unemployme | pre- | 1.87 ** | 0.31 * | 0.49 ** | 0.20 | 0.17 | 0.19 | 0.13 | 0.01 | 0.06 | 0.24 * | 0.18 * | 0.28 ** |
| | post- | 1.06 ** | 0.30 ** | 0.28 ** | 0.50 ** | 0.23 ** | 0.21 ** | 0.20 ** | 0.10 | 0.02 | 0.05 | 0.06 | 0.18 ** |
| GDP (prel.) | pre- | 0.24 | 0.00 | -0.03 | -0.09 | -0.12 | 0.12 | 0.12 | 0.05 | 0.05 | 0.06 | 0.11 | 0.03 |
| | post- | 1.21 ** | -0.08 | 0.26 * | 0.17 | 0.22 * | 0.07 | 0.06 | 0.15 | 0.01 | 0.07 | -0.08 | -0.01 |
| Ind. Prod. | pre- | 1.44 ** | 0.29 * | 0.37 * | 0.34 * | 0.21 | 0.23 | 0.06 | 0.11 | 0.05 | 0.20 | -0.01 | 0.14 |
| | post- | 0.55 ** | -0.01 | 0.08 | 0.11 | 0.08 | 0.07 | 0.05 | 0.02 | -0.04 | -0.02 | -0.02 | -0.06 |
| PPI Input / Ouput | pre- | 0.55 ** | 0.16 | 0.07 | -0.01 | 0.92 | -0.04 | 0.06 | 0.03 | -0.06 | -0.07 | 0.05 | -0.10 |
| | post- | 0.10 | 0.11 | -0.06 | -0.09 * | -0.01 | 0.03 | -0.12 ** | -0.02 | -0.08 * | -0.12 ** | -0.05 | -0.02 |
| Retail Sales | pre- | 1.14 ** | 0.48 ** | 0.12 | 0.36 | 0.30 ** | 0.22 * | 0.25 ** | 0.31 ** | 0.18 | 0.08 | 0.07 | 0.10 |
| | post- | 1.74 ** | 0.47 ** | 0.31 ** | 0.21 ** | 0.19 ** | 0.18 ** | 0.19 ** | 0.03 | 0.12 * | 0.13 * | 0.07 | 0.04 |
| RPI / RPIX | pre- | 1.26 ** | 0.43 * | 0.49 ** | 0.42 | 0.25 * | 0.17 | 0.10 | 0.16 | 0.09 | 0.07 | 0.25 * | 0.17 |
| | post- | 1.09 ** | 0.31 ** | 0.19 * | 0.23 * | 0.15 * | 0.13 * | 0.11 | 0.16 * | 0.07 | 0.11 | 0.19 ** | 0.04 |

## Table 9

| announcement | | volume (5 minute periods) | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | 5 | 10 | 15 | 20 | 25 | 30 | 35 | 40 | 45 | 50 | 55 | 60 |
| Av. Earn. / Unemploy. | pre- | 2.35 ** | 1.41 ** | 1.03 ** | 1.26 ** | 0.93 ** | 0.44 * | 0.50 ** | 0.18 | 0.59 ** | 0.69 ** | 0.45 ** | 0.47 ** |
| | post- | 1.78 ** | 1.05 ** | 0.55 * | 0.45 * | 0.74 ** | 0.32 | 0.71 ** | 0.58 ** | 0.57 ** | 0.48 ** | 0.27 | 0.34 * |
| | elec | 4.15 ** | 1.65 ** | 1.20 ** | 1.11 ** | 1.00 ** | 0.75 ** | 0.89 ** | 0.34 | 0.02 | 0.25 * | 0.39 * | 0.10 |
| GDP (prel.) | pre- | 1.55 ** | 0.47 | 0.60 * | 0.24 | -0.32 | -0.01 | 0.09 | 0.45 | 0.18 | 0.28 | -0.07 | 0.21 |
| | post- | 2.46 ** | 0.63 * | 0.64 * | 0.87 ** | 0.16 | -0.26 * | 0.18 | 0.07 | -0.06 | -0.08 | -0.27 ** | -0.08 |
| | elec | 3.77 ** | 0.53 | 0.69 | 1.00 * | 0.44 | 0.13 | 0.27 | 0.29 | 0.03 | 0.17 | 0.89 | 0.26 |
| Ind. Prod. | pre- | 1.96 ** | 0.91 ** | 0.95 ** | 0.79 ** | 0.60 ** | 0.28 | 0.18 | 0.44 ** | 0.25 | 0.19 | 0.10 | 0.19 |
| | post- | 1.06 ** | 0.55 | 0.40 * | 0.28 * | 0.04 | 0.11 | 0.28 * | 0.09 | -0.01 | 0.05 | -0.05 | -0.02 |
| | elec | 2.96 ** | 0.96 ** | 0.28 | 0.33 ** | 0.21 | -0.03 | 0.13 | 0.13 | -0.02 | -0.04 | 0.16 | -0.09 |
| Retail Sales | pre- | 2.40 ** | 1.38 ** | 0.99 ** | 0.88 ** | 0.71 ** | 0.57 ** | 0.46 ** | 0.99 ** | 0.62 ** | 0.66 ** | 0.35 ** | 0.45 ** |
| | post- | 2.36 ** | 1.64 ** | 1.22 ** | 0.69 ** | 0.29 | 0.55 ** | 0.36 * | 0.67 ** | 0.53 ** | 0.42 * | 0.12 | 0.43 ** |
| | elec | 6.09 ** | 1.92 ** | 0.90 ** | 1.16 ** | 0.62 ** | 0.50 * | 0.35 | 0.24 | 0.08 | 0.24 | 0.39 ** | 0.21 |
| RPI / RPIX | pre- | 1.68 ** | 1.14 ** | 1.29 ** | 1.08 ** | 1.11 ** | 0.61 ** | 0.65 ** | 0.68 ** | 0.73 ** | 0.35 | 0.27 | 0.74 ** |
| | post- | 1.88 ** | 1.34 ** | 0.87 ** | 0.36 | 0.35 | 0.47 * | 0.30 | 0.49 * | 0.36 | 0.33 * | 0.09 | -0.09 |
| | elec | 4.21 ** | 1.63 ** | 1.08 ** | 0.94 ** | 0.66 ** | 0.62 ** | 0.67 ** | 0.18 | 0.53 * | 0.49 ** | 0.33 * | 0.45 * |

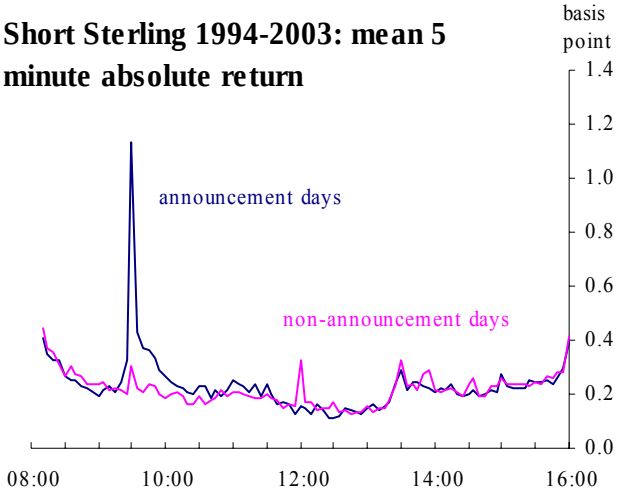
## Table 10: summary statistics for distribution of trading volume

| | pre-BI | post-BI, pre-elec. | post-elec. |
|---|---|---|---|
| mean | 2.6 | 2.5 | 5.0 |
| median | 0 | 0 | 2 |
| max | 126 | 232 | 338 |
| sd | 4.4 | 5.6 | 9.8 |
| distribution: | | | |
| +/- <1 s.d. | 88.7% | 93.4% | 91.2% |
| +/- <2 s.d. | 6.7% | 3.5% | 5.1% |
| +/- >2 s.d. | 4.6% | 3.1% | 3.7% |

## Annex 2: Charts

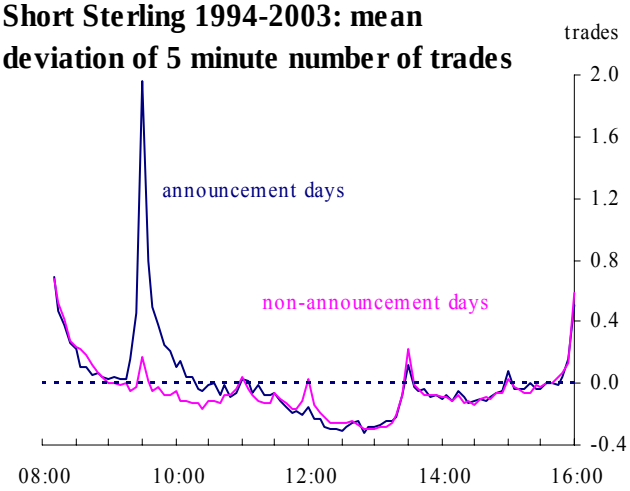
### Chart 1

**Short Sterling 1994-2003: mean 5 minute absolute return**



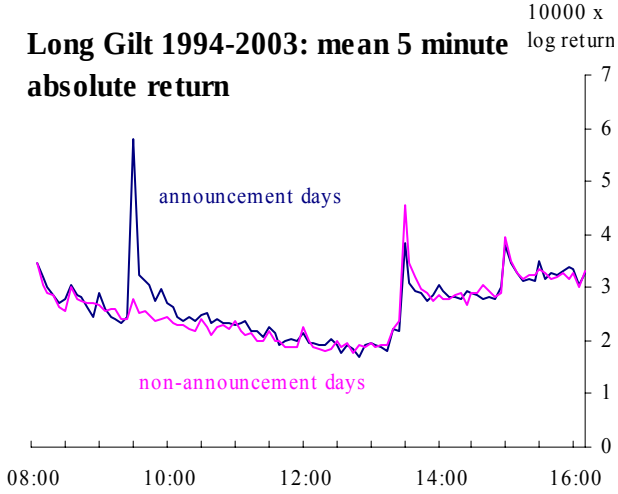

### Chart 2

**Short Sterling 1994-2003: mean deviation of 5 minute number of trades**



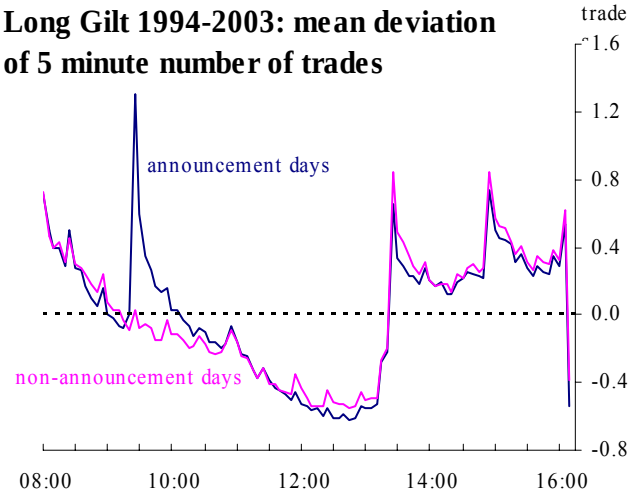

### Chart 3

**Long Gilt 1994-2003: mean 5 minute absolute return**



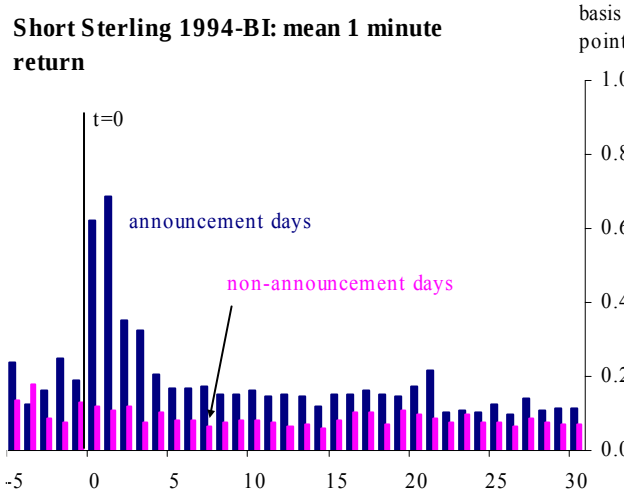

### Chart 4

**Long Gilt 1994-2003: mean deviation of 5 minute number of trades**



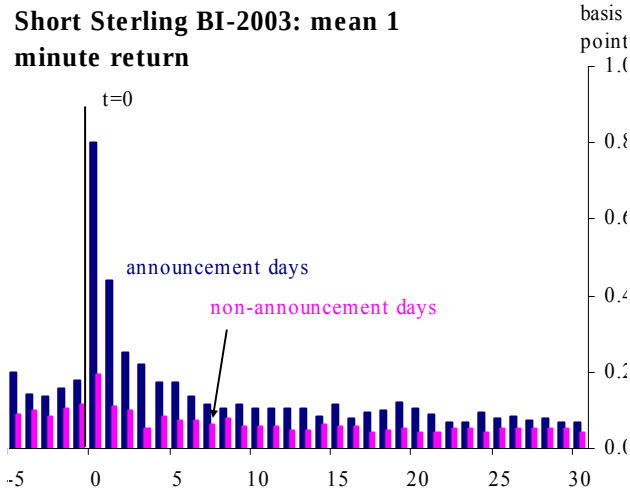
## Chart 5

**Short Sterling 1994-BI: mean 1 minute return**



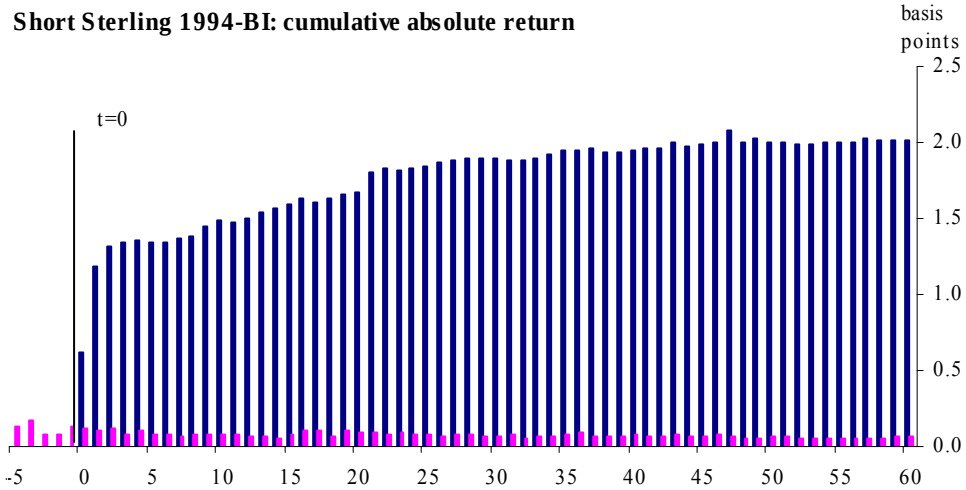

## Chart 6

**Short Sterling BI-2003: mean 1 minute return**



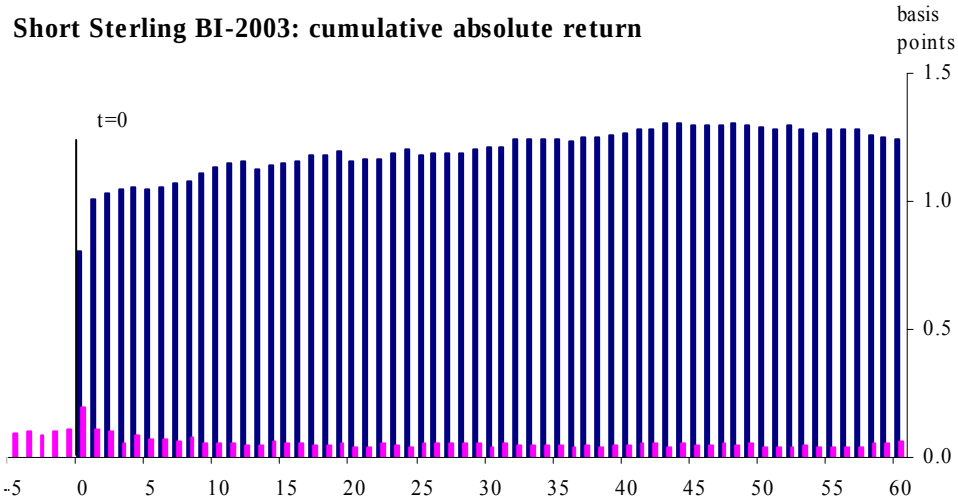

## Chart 7

**Short Sterling 1994-BI: cumulative absolute return**

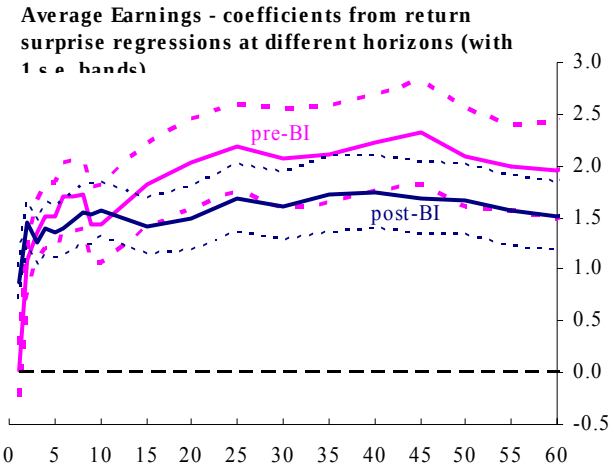



## Chart 8

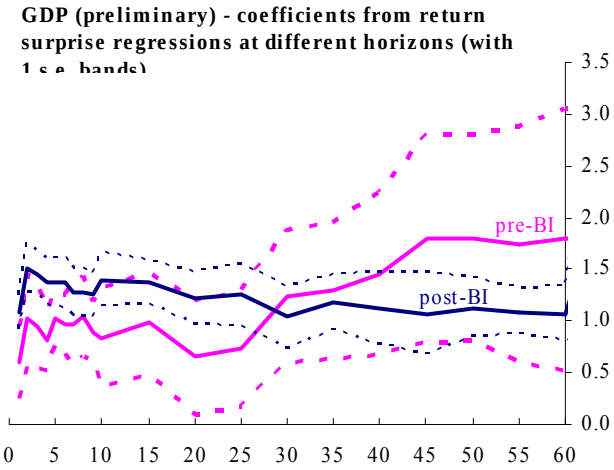
**Short Sterling BI-2003: cumulative absolute return**



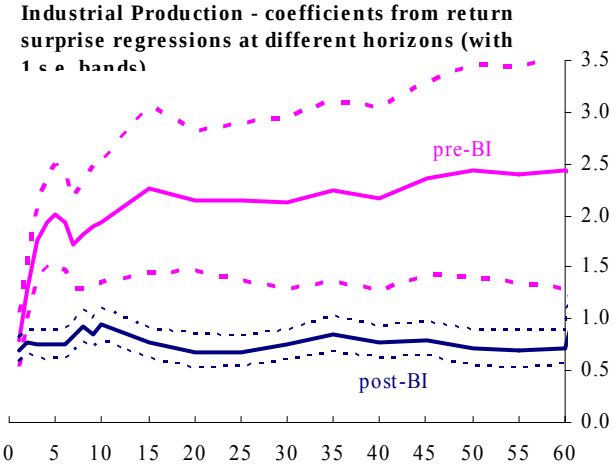
## Table 9



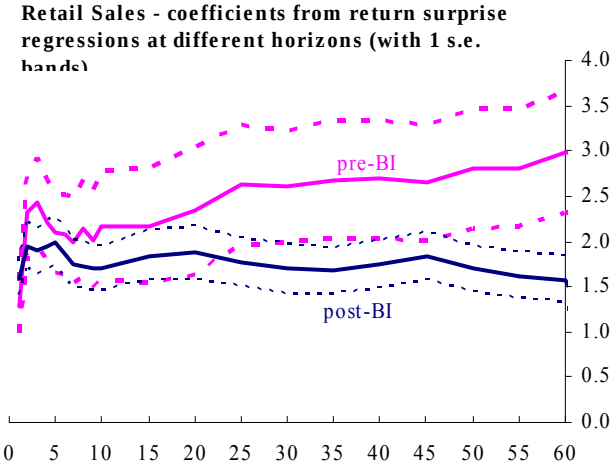

## Table 10



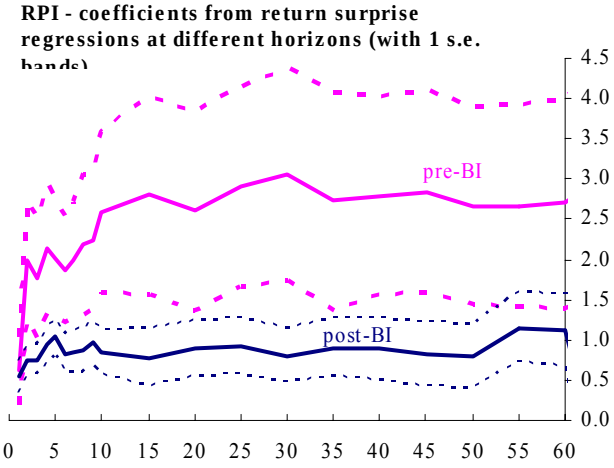

## Table 11



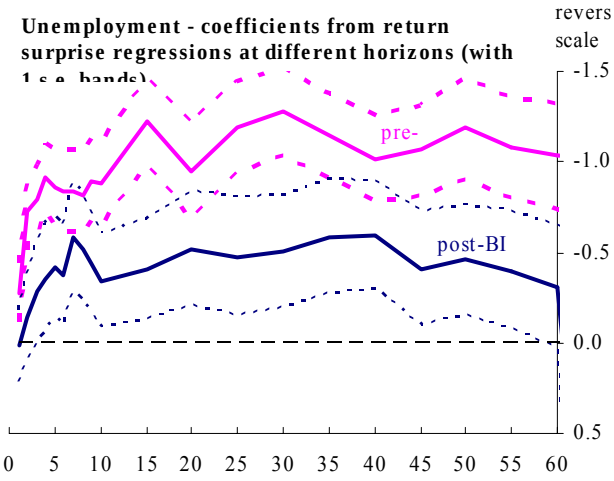

## Table 12




## Table 13




## Table 14



**References**

**Allen, H and Hawkins, J (2002)** 'Electronic trading in wholesale financial markets: its wider impact and policy issues,' *Bank of England Quarterly Bulletin*, Spring, p50-58.

**Amato, J.D, Morris, S and Shin, H.S (2002)**, Communication and Monetary Policy, mimeo, London School of Economics.

**Andersen, T and Bollerslev, T (1998)** 'Deutsche Mark–Dollar Volatility: Intraday Activity Patterns, Macroeconomic Announcements, and Longer Run Dependencies, *Journal of Finance*, Vol. 53, p219-265.

**Andersen, T, Bollerslev, T, Diebold, F X and Vega, A (2003)** 'Micro effects of macro announcements: real-time price discovery in foreign exchange', *American Economic Review*, 1, p38-62.

**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*, p7-17.

**Ap Gwilym, O, Thomas, S (1998)** 'Which macroeconomic announcements have most impact on Long Gilt futures market volatility?', *Derivatives Use*, *Trading and Regulation,* Vol. 4 No. 2 p164-73

**Balduzzi, P, Elton, E J and Green, T F (2001)** 'Economic news and bond prices: evidence from the U.S. Treasury market', *Journal of Financial and Quantitative Analysis*, Vol. 36, No. 4 p523-543

**Becker, K G, Finnerty, J E, and Kopecky, K J (1993)** 'Economic News and Intraday Volatility in International Bond Markets' *Financial Analysts Journal,* May/June 1993, p81-5

**Becker, K G, Finnerty, J E, and Kopecky, K J (1996)** 'Macroeconomic News and the Efficiency of International Bond Futures Markets' *Journal of Futures Markets,* Vol. 16, No. 2, p131-45

**Bessembinder, H, Seguin, PJ (1993)** 'Price Volatility, Trading Volume, and Market Depth: Evidence from Futures Markets' *Journal of Financial and Quantitative Analysis,* Vol. 28, No. 1, p21-39

**Bollerslev, T, Cai, J and Song, F (2000)**, 'Intraday periodicity, long memory volatility, and macroeconomic announcements in the US Treasury bond market, Journal of Empirical Finance, Vol. 7, p37-55.

**Chadha, JS, Nolan, C (2001)** 'Inflation Targeting, Transparency and Interest Rate Volatility: Ditching "Monetary Mystique" in the U.K.' *Journal of Macroeconomics*, Vol. 23, No. 3, p349-366

**Clare, A and Courtenay, R (2001)** 'Assessing the impact of macroeconomic news announcements on securities prices under different monetary policy regimes', *Bank of England Working Paper, No 125*

**Clare, A, Johnson, M, Proudman, J and Saporta, V (1999)** 'The impact of macroeconomic news announcements on the market for gilts', in Market liquidity: research findings and selected policy implications, Committee on the Global Financial System, The Bank for International Settlements, Basle, Switzerland.

**Committee on the Global Financial System (CGFS) (2001),** The Implications of Electronic Trading in Financial Markets, The Bank for International Settlements, Basle, Switzerland.

**Cook, T and Hahn, T (1989),** 'The Effect of Changes in the Federal Funds Rate Target on Market Interest Rates in the 1970s, *Journal of Monetary Economics*, Vol. 24, p331-51.

**Ederington, L H and Lee, J H (1995)** 'Short Run Dynamics of the Price Adjustment to New Information', *Journal of Financial and Quantitative Analysis*, Vol. 30, p117-134.

**Ederington, L H and Lee, J H (1993)** 'How Markets Process Information: News Releases and Volatility', *Journal of Finance*, Vol. 48, p1161-91.

**Ehrmann, M and Fratzscher, M (2002)** 'Interdependence between the Euro area and the US: What role for EMU?', *European Central Bank Working Paper, No. 200*

**Faust, J, Rogers, J H, Shing-Yi, B W, and Wright, J H (2003)** 'The High-Frequency Response of Exchange Rates and Interest Rates to Macroeconomic Announcements' *Board of Governors International Finance Discussion Paper No. 784*

**Fleming, M J and Remolona, E M (1997)** 'What moves the bond market?, *FRBNY Economic Policy Review,* p31-50.

**Fleming, M J and Remolona, EM (1999)** 'Price Formation and Liquidity in the U.S. Treasury Market: The Response to Public Information', *International Economic Review* Vol. LIV No 5, p1901-15.

**Fleming, M J and Remolona, EM (1999)** 'The Term Structure of Announcement Effects', *FRBNY Staff Reports,* No. 76.

**Fleming, M J (2001)** 'Measuring Treasury Market Liquidity', *mimeo*, Federal Reserve Bank of New York

**Ganley, J, Holland, A, Saporta, V and Vila, A (1998)**, 'Transparency and the design of securities markets,' *Bank of England Financial Stability Review*, No. 4, p8-17.

**Geraats, P M (2002)** 'Central Bank Transparency', *Economic Journal 2002,* p532-65

**Gravelle, T, and Moessner, R (2001)** 'Reactions of Canadian Interest Rates to Macroeconomic Announcements: Implications for Monetary Policy Transparency', *Bank of Canada Working Paper*

**Green, T C (2001)** 'Economic News and the Impact of Trading on Bond Prices' *Journal of Finance,* Forthcoming

**Haldane, A and Read, V (2000)** 'Monetary Policy Surprises and the Yield Curve', *Bank of England Working Paper, No 106*

**Harris, M and Raviv, A (1993)**, 'Differences of Opinion Make a Horse Race,' *Review of Financial Studies*, Vol. 6, p473-506.

**Joyce, M, Read, V (1999)** 'Asset Price Reactions to RPI Announcements', *Bank of England Working Paper, No 94*

**King, M, (2002)** 'The inflation target 10 years on', Speech at The London School of Economics.

**Kuttner, K (2001)** 'Monetary policy surprises and interest rates: Evidence from the Fed funds futures market', *Journal of Monetary Economics,* Vol. 47, p523-44

**Lange, J, Sack, B and Whitesell, B (2003)**, Anticipations of monetary policy in financial markets,' *Journal of Money, Credit and Banking*, Vol. 35, p889-909.

**Lasaosa, M (2003)** 'Learning the rules of the new game? Comparing the reactions in financial markets to announcements before and after the Bank of England's operational independence', *mimeo*, Bank of England

**Lildholdt, P and Vila Wetherilt, A (2004)**, Anticipation of Monetary Policy in UK Financial Markets, *mimeo*, Bank of England.

**Moessner, R, Gravelle, T and Sinclair, P (2000)** 'Measures of Monetary Policy Transparency and the Transmission Mechanism', *mimeo*, Bank of England

**O'Hara, M (1995)** *'Market Microstructure Theory'* Cambridge, Massachusetts: Blackwell.

**Sack, B (2003),** 'Does the Fed act gradually? A VAR analysis,' *Journal of Money, Credit and Banking*, Vol.46, p229-56.

**Wang, J (1994)**, 'A Model of Competitive Stock Trading Volume,' *Journal of Political Economy*, Vol. 102, p.127-68.