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Measuring Changes in Liquidity Using the Bid-offer Price Proxy: Determinants of Liquidity in the United Kingdom Gilt Market

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Abstract: *Financial market liquidity is an important yardstick of value for investors and central monetary authorities. Secondary market liquidity itself cannot be observed directly and is instead measured using a number of different proxies. The most common proxy is the asset bid-off price spread. In this study we conduct time series analysis of the bid-offer spread in order to ascertain if the level of liquidity in a specified market has improved over a period of time. The market we select is the United Kingdom government bond market or gilt market. During the 1990s the UK monetary authorities introduced a number of structural reforms in the gilt market, designed to improve secondary market liquidity. We measure the success of the reforms by attempting to determine if liquidity levels improved in the post-reform period, via the examination of the bid-offer spread. We examine the determinants of this proxy measure, and estimate which of the explanatory variables carries the greatest weight in influencing liquidity levels. We conclude that a number of the independent variables that we examined, including bond issue size and maturity, are found to be significant determinants of liquidity. We conclude further that similar structural reforms should be considered by other central monetary authorities wishing to improve bond market liquidity levels, and that the determinant factors we cite should be reviewed during periods of market correction, when liquidity levels decrease.*

Keywords: financial markets, government bonds, liquidity, bid-offer spread, risk-free yield, repo, strips

Introduction

Participants in the financial markets emphasise the importance of secondary market liquidity as an indicator of value. This is also noted in the academic literature. During the late 1990s the United Kingdom (UK) central monetary authorities introduced a number of structural reforms in the sovereign bond market, designed to improve market liquidity. Liquidity itself cannot be measured directly; hence it is not possible to determine conclusively that the UK authorities succeeded in their goal. Instead it is necessary to assess changes in the level of liquidity with recourse to proxy measures. This paper attempts to measure liquidity in the UK gilt market by examining a common proxy for liquidity, the bid-offer price spread. The time period considered is the pre-reform and immediate post-reform period, to determine if these reforms succeeded in achieving their goal. We also investigate the significance of the relationship between the proxy measure of liquidity and its explanatory variables.

This work is motivated by the following factors: a desire to fill gaps in the existing literature with respect to sterling markets and UK government bonds; a response to an examination of gilt market liquidity by the sovereign issuing authority itself, which concluded with an observation that further research was required; to ascertain if results from earlier academic studies of the US Treasury market also hold in the UK market; and the wish to ascertain if the proxy liquidity measure we selected was effective in its purpose.

The time period under observation is the UK gilt market during the period 1993–2002. This period covers the timing of the structural reforms themselves, implemented during 1996-1998, as well as the period immediately before and after the implementation of the reforms. The market reforms, which were introduced by the Bank of England (BoE), were designed to improve liquidity and accessibility of the UK gilt market; they included the introduction of new products as well as technical changes to operating processes. The reforms are described Bank of England (1995) and Debt Management Office (1998, 2000).

We examine a proxy measure of liquidity, the bid-offer price spread. We consider the relationship between this spread and a number of explanatory variables, to assess the size and significance of their effect on the spread itself and indirectly liquidity itself. We determine the significance of the selected determinants, and also observe if any significant change in the relationship has taken place during the observation period. At the end of our examination we identified the significant determinants of bid-offer proxy measure of liquidity, and concluded that they had increased in level of significance and size of coefficient during the observation period. Additional tests indicated a strong likelihood that a structural break was evident in the time series data at an expected point in the period after the market reform implementation period.

Our study makes three contributions to the existing literature: it is the first such study using the bid-offer spread proxy indicator that looks specifically at the UK gilt market; secondly, it presents results of the interaction of various explanatory variables as drivers of liquidity, which will be of value to the UK sovereign monetary authority; and thirdly it suggests that the market reforms had indeed contributed to increased liquidity in the gilt market.

This paper is organised as follows: first, we review the literature and provide a discussion of market liquidity. This is followed by the background to empirical testing, the formulation and testing of the regression model for the bid-offer price spread, the tests of structural change, and a discussion of the results. The final section presents our conclusions.

Literature review

There are a number of definitions of financial market liquidity. Commonly accepted ones in the academic and practitioner literature include the following:

- a market where buyers and sellers may transact at any time (during opening hours) in size, at no extra cost, without this transaction causing prices to move;
- a case where two-way prices are available to market participants in wholesale market size, and where there is an openness in determining asset fair value.

The first definition has been suggested by O'Hara (1995) and Fleming (2001) among others, while the second was described by Mackintosh (1995).

The academic literature on the gilt market is sparse. A general introduction to the structure of the gilt market and its trading mechanisms is given in Choudhry et al (2003). Steeley and Ahmad (2002) conducted an investigation on the behaviour of prices in the gilt market during the period 1993 to 2001. This period takes in the time of the reforms to the gilt market, as well as periods of general market correction such as the Russian debt default in 1998 and the Long Term Capital Management (LTCM) hedge fund crisis of 1999. The authors consider the information and price action efficiency of gilts, specifically (i) the characteristics of the price movements during this period, testing the null hypothesis that price movements are random, and (ii) whether any non-random behaviour in prices could have been exploited. Their conclusion is that the gilt market exhibited returns persistence during the crisis period, due to its safe-haven status.

Repo is a common feature in many government bond markets, both in developed and emerging economies, being a standard money market secured loan instrument. A basic coverage and definition of repo is given in Blake (1990). The Bank for International Settlements (BIS) study group (2000) investigated the part played by repo in the maintenance of an orderly secondary market and concluded that it was a vital tool for such purposes.

The academic literature reviews financial market liquidity in some depth. O'Hara (1995) defines liquidity as the ability to trade a security quickly and with little cost. Sundaresan (1997) defines liquidity as a market where investors can buy and sell large amounts of stock with ease, at a narrow bid-offer spread and without an adverse price reaction. Gravelle (1998) defines liquidity as being the ease with which large-size bond transactions can be effected without market prices being impacted. He also reports that the central authorities desire the maintenance of a liquid market. Borio (2000, p.38) describes a liquid market as one where "…transactions can take place rapidly and with little impact on price".

Previous research into liquidity measures has concentrated mainly on the United States (US) government bond market and US corporate markets; these included Fleming (2001), Diaz and Skinner (2001) and Moulton (2004). Other studies considered the US dollar interest-rate swap market; these included Amihud and Mendelson (1991a, 1991b), Alexander, Edwards and Ferry (2000), Hong and Warga (2000), Brown, In and Fang (2002) and Kalimipalli and Warga (2002).

For our proxy measure of liquidity we use the bid-offer spread. Studies that looked specifically at this measure in the US dollar market included Schultz (1998), Chakravarty and Sarkar (1999), Hong and Warga (2000), Ap Gwilym, Trevino and Thomas (2002) and Brown, In and Fang (2002). All these studies were undertaken on the basis that a standard measure of liquidity is the extent of the bid-offer spread, and that as liquidity increases the spread will narrow. No such study has been conducted for the UK gilt market.

Blennerhassett and Bowman (1998) studied the New Zealand stock exchange and found that the bid-offer spread became more sensitive to changes in trade size, and this effect may mean that larger-size dealers carry disproportionate costs. Naik et al (1999) studied the impact of the change from the dealership market to an order book market in the London Stock Exchange (LSE) and found that investors had benefited from the change as they now faced narrower bid-offer spreads.

A number of other proxies for liquidity have been considered. Yield spreads were investigated in Sarig and Warga (1989), Blume, Keim and Patel (1991), Warga (1992) and Crabbe and Turner (1995). Trading volume was investigated in Kamara (1994) and Alexander, Edwards and Ferri (2000). Nunn, Hill and Schneeweis (1986) used a combination of three proxies for liquidity, which were (i) the bond's age (ii) the bid-offer spread and (iii) the amount of bonds outstanding. Mackintosh (1995) proposed a liquidity score for a bond based on an aggregate liquidity rating.

Some studies have tested the hypothesis that larger size bond issues are more liquid. For example Hong and Warga (2000) observed that larger-size issues had smaller bid-offer spreads. Alexander, Edwards and Ferri (2000) examined the determinants of trading volume for corporate bonds, and concluded that the larger-size issues were more liquid. Kalimipalli and Warga (2002), Chakravarty and Sarkar (2003) and Moulton (2004) all suggest that issue size is an important determinant of liquidity. Moulton studied the US Treasury repo market however, in which large issue size must be considered the norm and by implication a prime driver of liquidity. However

Crabbe and Turner (1995) and Fridson and Garman (1998) when studying yield spreads as proxies for liquidity found no backing for the issue size hypothesis.

The importance of liquidity to the smooth functioning of financial markets is emphasised frequently in the literature. Datar, Naik and Radcliffe (1998) suggest that liquidity has an impact on asset returns. Amihud and Mendelson (1986) concluded that investors allow for lower liquidity by demanding a higher return premium, which is the trade-off required for bearing the higher cost of trading in illiquid markets. The same authors (1991) also found that the difference in bid-offer spread between US Treasury bills and Treasury securities had an impact of yield-to-maturity. Amihud, Mendelson and Lauterbach (1997) observed that asset values on the Israeli stock exchange underwent changes when the equities began to be traded on a more liquid electronic system.

McCauley and Remolona (2000) reported how a number of Organisation for Economic Cooperation and Development (OECD) governments continued to maintain gross issuance in an effort to preserve market liquidity, despite budget surpluses removing the need to issue debt. This reflects the importance of the government bond market to all market participants, including investors, traders and brokers. The authors emphasise the importance of a liquid market in government bonds.

Researchers have identified various factors that are determinants of liquidity. Alexander, Edwards and Ferri (2000) and Sarig and Warga (1989) found that corporate bonds that were issued more recently were more actively traded, implying bond age as a liquidity factor. Among numerous studies that make this observation, Babbel et al (2001) showed that benchmark or "on-the-run" US Treasury securities were more actively traded than older Treasuries. This has implications for our own research. Fleming and Remolona (1999) found that macroeconomic announcements had a significant impact on the bid-offer spread. Another factor is the outstanding amount in issue for a bond; one expects this to be an influence on liquidity and Fisher (1959) observed this in a very early study. Garman (1976), Stoll (1978), Amihud and Mendelson (1980), and Ho and Stoll (1981) found that the bid-offer spread increases with the bond price and the credit risk of the bond, and also decreased with higher levels of trading activity.

In corporate bond markets credit ratings have an impact on bond liquidity, as shown by Fridson and Garman (1998). Kamara (1994) concluded that looking only at risk-free sovereign bonds removes liquidity issues arising from credit risk, because all the bonds in the sample are credit-risk-free. The same applies to UK gilts: all bonds in a sample of gilts have uniform tax, trading and settlement issues, and zero credit risk.

Testing Determinants of the Bid-offer Spread

In the UK gilt market the bid-offer spread is quoted by market primary dealers, known as gilt-edged market makers (GEMM). This spread quote is good for what is considered "normal market size" (NMS), a standard-sized order.[1] To eliminate the quantity aspect in our study, we consider only the bid-offer spread for the NMS bargain quantity.

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[1] The concept of normal market size is originally an equity market one, where live prices are displayed on electronic order boards. These prices will be good for all orders of size up to NMS quantity. Because bond markets are traded as over-the-counter (OTC) rather than on-exchange, live prices are not displayed on screens. However institutional investors assume, when obtaining a price from a GEMM, that this price is good for dealing up to a standard size,

We obtain the bid-offer spreads for the four benchmark bonds for each week of the period under study. These are market closing prices as at close of business each Friday, obtained from the Bloomberg trade system. The observation period is broken into three distinct periods:

- Period 1, the sample period before the implementation of the structural reforms (Jan 1993 – Dec 1995);
- Period 2, the period during which reforms were introduced (Jan 1996 – Jan 1998);
- Period 3, the sample period after the introduction of reforms (Feb 1998 – Dec 2002).

We assemble a database of weekly prices for each benchmark bond from price data available on Bloomberg for the period January 1993 to December 2002. Ignoring the end-year periods this is 51 weeks per year, hence a 510-week sample period.[2] Thus the dataset is, for each week, four benchmark bonds (the two-year, five-year, 10-year and long-bond) for each week in the sample period.

We wish to address the following key issues:

- using our proxy measure, has liquidity improved post the reform period?
- what are the determinants of the bid-offer spread?
- do any aspects of the bonds themselves influence this spread?
- how firm are the spreads themselves during the sample period, and how stable?

We consider a number of variables as determinants of the bid-offer spread. The selection of explanatory variables is influenced by the previous literature. Based on this literature, our model captures all the relevant explanatory variables except macro-level indicators. To date, with the exception of Chakravarty and Sarkar (2003), this is because previous studies have excluded macro-economic factors. Following Chakravarty and Sarkar (2003), we modify the model for a second test, which includes a macro-level factor. This is the base interest rate at each weekly observation period.[3] A third model that includes an extension to test for the impact and influence of macroeconomic news announcements is also tested. Our models also follow findings from Fleming and Remonola (1999) that concluded that macroeconomic data announcements had a significant impact on the bid-offer spread. We wish to see if the explanatory variables are significant without the impact of the macro-level indicator and announcements variable, hence we test three models.

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which is assumed to be the NMS quantity. This size ranges from GBP 10million to GBP 50 million depending on bond maturity.

[2] Note that the benchmark bond itself changes over time, as its maturity reduces and it is replaced by the new benchmark. For example the five-year benchmark during 1995 was the 8% Treasury 2000 stock, while the following year it was the 7% Treasury 2001 stock. Benchmarks are identified explicitly on Bloomberg. We use the relevant benchmark bond price for each weekly price we set up on the database.

[3] We consider the UK base interest rate, set by the Bank of England (BoE), to be an indicator as well as measure of overall macroeconomic conditions. As such it can be expected to have an influence on the bid-offer spread, so we model for it here. Note that both of these premises are part of the conclusions reached in Fleming and Remolona (1999) and Chakravarty and Sarkar (2003).

Researchers into corporate bond markets have identified other factors associated with the price spread that purport to explain liquidity. Sarig and Warga (1989) and Alexander, Edwards and Ferri (2000) found that younger corporate bonds were more actively traded, while for US Treasury securities Babbel *et al* (2001) showed that benchmark ("on-the-run") Treasuries were more liquid. Hence, *bond age* is significant to influencing the bid-offer spread. However it would not apply to our study, because we are considering *only* the relevant benchmark gilt at any one time. That is, there is only one bond being price-tested each week (one benchmark bond per maturity); there is no comparison to other bonds involved in the test. Therefore we do not need to include this variable in our analysis.

We estimate the following model:

$$Y_{bid-offerspread} = \alpha + \beta_1 C_{it} + \beta_2 M_{it} + \beta_3 I_{it} + \beta_4 V_{ti} + \varepsilon_{it} \quad (1)$$

where Y is the mean weekly bid-offer spread for bond i (i = 1, 4 for each benchmark) for each week over the given sample period under study. The explanatory variables are:

C	the bond coupon
M	the term to maturity, the time in years from price quote date and date of maturity
I	bond issue size
V	the volatility of the bond price over the sample period.
ε	the error term.

This is applied to a pooled cross-section and time-series analysis. We use the pooling technique when applying (1), as well as a separate test for each benchmark bond. Model (1) is estimated for each bond using our week-based dataset, with data collected for each benchmark bond for each week. Volatility is measured over the week using the close of week price changes.

Table I summarises the determinants in the model that are being tested, and their expected sign ahead of the tests.

The explanatory variable directly related to the structural reforms introduced by the BoE is issue size *I*. The benchmark programme instituted as part of these reforms lead to higher issue sizes for the selected benchmark bonds. Note that the issue size *I* is also taken to be a proxy measure of liquidity, as well as a determinant of the bid-offer spread, itself a proxy measure. This might compromise interpretation of the test results. However the previous literature, including Kalimipalli and Warga (2002), Chakravarty and Sarkar (2003) and Moulton (2004), suggests that it is an important determinant and therefore we test it here for the gilt market.

We also test the following extension, which introduces the base interest rate explanatory variable:

$$Y_{bid-offerspread} = \alpha + \beta_1 C_{it} + \beta_2 M_{it} + \beta_3 I_{it} + \beta_4 V_{it} + \beta_5 R_{it} + \varepsilon_{it} \quad (2)$$

where R_i is the base interest rate at any specific point in time for bond i.

Table I: Predicted Signs of Explanatory Variables

Explanatory Variable	Predicted Sign	Reasoning
Bond coupon	Not predicted (insignificant)	The coupon on issue sets the price to par or just below; therefore should not impact subsequent liqudity as the yield should set to the equilibrium fair value rate
Term to maturity	Positive	Higher term to maturity increases interest-rate risk (modified duration), therefore wider spread
Issue size	Negative	A larger issue size suggests greater liquidty and hence smaller spread
Volatility	Positive	Greater volatility translates into greater returns uncertainty, hence wider spread
Base rate	Postitive	High relative base rates indicate an economy requiring slowing-down of activity, hence wider spreads in illiquid markets
Announcement dummy	Positive	Any unexpected or negative announcment would result in lower prices, hence wider spreads to control risk exposure of the dealer

We do not consider the trade size variable in this model, which has been used in previous literature, because we assume that the bid-offer spread on the Bloomberg price quote is good for NMS bargains. As long as the trade is at or under the NMS quote size, there will be no change in the quote spread.

Our third model considers the impact of macroeconomic announcements on market prices and consequently the bid-offer spread. We add an additional variable for each week in the sample when a macroeconomic announcement was made. An announcement is considered as worthy of inclusion if it relates to the level of employment, inflation, retail sales or industrial production.[4] The third model incorporates the extension for the macroeconomic announcement variable:

$$Y_{bid-offerspread} = \alpha + \beta_1 C_{it} + \beta_2 M_{it} + \beta_3 I_{it} + \beta_4 V_{it} + \beta_5 R_{it} + \beta_6 A_{it} + \varepsilon_{it} \qquad (3)$$

where A represents the macroeconomic announcement variable.

A acts as a dummy variable: it takes the value of 1 if there is a macroeconomic announcement that week, and 0 if there is no announcement.

Based on our initial hypotheses we expect the following results. We expect coupon to be neither positively nor negatively related to the spread. We expect issue size to have a negative relationship, and we expect maturity and volatility to be positively related. The base interest rate, if it

[4] In the UK market, these announcements are generally made in the same week each month, so it is straightforward to isolate the price quotes that were recorded during a week of an announcement.

is at all statistically significant, is expected to have a positive relationship. We expect the announcement to have a widening impact on spread, hence a positive relationship. These sign predictions are summarised in Table I.

We conduct one additional test to determine the significance of the explanatory variables after the conclusion of the structural reforms. This adds a dummy variable for time to the model at (2), so that it becomes:

$$Y_{bid-offerspread} = \alpha + \beta_1 C_{it} + \beta_2 M_{it} + \beta_3 I_{it} + \beta_4 V_{it} + \beta_7 D_{it} + \varepsilon_{it} \qquad (4)$$

where D is the dummy variable. It is given a value of 0 (pre-reform) or 1 (post-reform). Hence D is set equal to 0 for the period January 1993–January 1998 and set at 1 for February 1998–December 2002, the period after the reforms were fully implemented.

The cut-off date for the conclusion of the reforms is January 1998. The β coefficient should be positive and statistically significant in order to provide evidence that the reforms have had the desired impact on the bid-offer spread and, by implication, the level of liquidity.

Volatility

The issue of volatility is a problematic one. We have included it in the model at (1) as the previous literature cites it as an independent variable. However liquidity itself can be viewed as influencing volatility, so the relationship, rather than being a causal one, is more likely to be a two-way one. In the first instance we believe that market volatility would be linked, negatively, to market liquidity. This is because *a priori*, we may expect the bid-offer spread to be widening with volatility as the market risk faced by primary dealers would increase simultaneously. Also, an illiquid market can be slow to benefit from new price action and updated information, which would lead to sudden changes in prices and hence increased volatility. This would be an instance of low levels of liquidity leading to higher volatility.

For this reason, we do not consider market volatility of itself to be an effective proxy measure of liquidity. However we include it in the main model to test the significance of its relationship with the bid-offer spread.

Empirical Results

We calculate the bid-offer spread for each benchmark bond for the period under study, given by

$$\frac{P_{Offer} - P_{Bid}}{P_{Offer}} \times 100.$$

Table II presents summary statistics on the mean bid-offer price spread for each benchmark bond. We note that the data is not unacceptably noisy, with standard deviation of the bid-offer spread lying between 0.1 to 0.6 basis points for the two-year benchmark and between 0.4 and 1.03 basis points for the long-bond.

Figure 1 is a plot of the average daily bid-offer spreads by month for each benchmark bond. We observe a decrease in bid-offer spread from Period 1 to Period 3. The noticeable difference in spreads between the 10-year bond and long-bond in Period 3 would appear to indicate that once price quotes had moved to decimals (another market reform), the market could differentiate between the two bonds better, in terms of price spread, than when ticks were used.

The lowest maturity bond always has the narrowest bid-offer spread, which would be expected because it represents the lowest interest-rate risk (measured by basis point value (BPV)), and hence the lowest cost inventory management for gilt primary dealers. The widest spread is observed for the longer-dated benchmark, as expected.

Table II: Summary Statistics on Weekly Percentage Bid-offer Spreads (Basis points 0.01), 1993-2002

Period 1
1993 - 1995
Panel A: Descriptive Statistics

Bond	Mean	Median	Standard Deviation	Lowest Spread	Highest Spread
Two-year	3.066	3.05	0.188	2.800	4.000
Five-year	6.3	6.25	0.289	6.000	7.800
10-year	9.502	9.4	0.516	9.375	12.500
Long-bond	9.476	9.45	0.486	9.000	12.000

Period 2
1996 - 1998
Panel A: Descriptive Statistics

Bond	Mean	Median	Standard Deviation	Lowest Spread	Highest Spread
Two-year	3.024	3	0.2197	2.700	3.900
Five-year	4.885	4.68	0.694	4.000	6.900
10-year	9.349	9.375	0.659	8.000	12.000
Long-bond	9.706	9.6	0.586	9.500	12.500

Period 3
1998 - 2002
Panel A: Descriptive Statistics

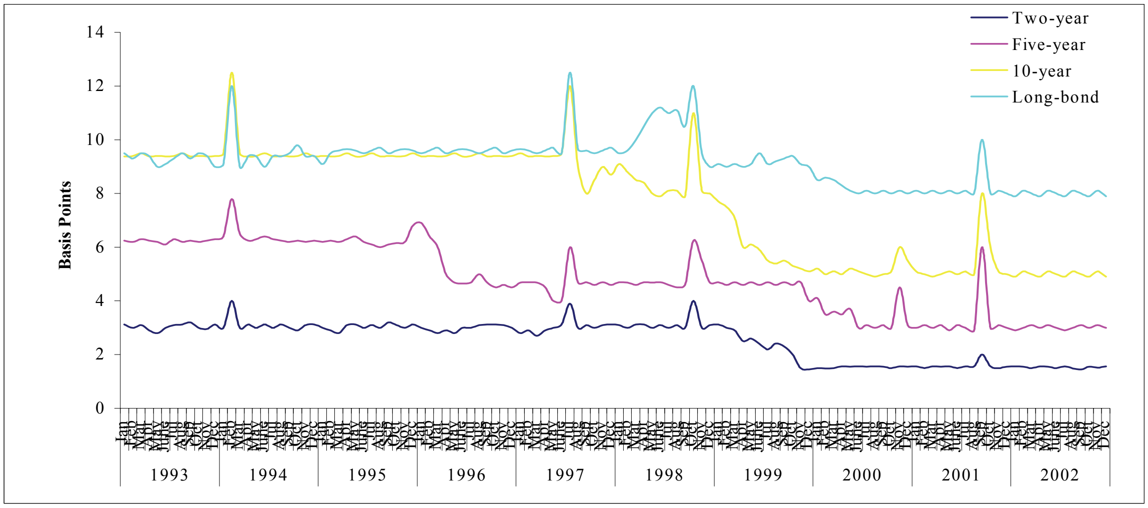
Bond	Mean	Median	Standard Deviation	Lowest Spread	Highest Spread
Two-year	2.01	1.56	0.688	1.450	4.000
Five-year	3.82	3.5	0.896	2.900	6.250
10-year	5.96	5.1	1.422	4.900	11.000
Long-bond	8.78	8.1	1.032	7.900	12.000

Test results

We examine the empirical determinants of the bid-offer spread. Initially, this is done be considering the characteristics of each bond. We subsequently consider the influence of base interest rates and macroeconomic announcements.

Figure 1: Average Bid-offer Spreads (Basis points 0.01), 1993-2002

(Price source: Bloomberg L.P.)



Note: Prices quotes were reported in "ticks" (1/32nds) before October 1998, when price format changed to basis points (0.01) Tick prices for this period have been converted to basis points.

Durbin-Watson (*DW*) test statistics from previous literature on financial markets suggest significant serial correlation in the error terms when the regressions for models similar to (2), (3) and (4) are estimated using OLS. In fact the *DW* test results in our case are between the upper value (d_U) and ($4 - d_U$), so we do not reject the null hypothesis; there is no evidence of autocorrelation. A Lagrange test is also conducted to assess for heteroscedasticity in the OLS error terms.[5]

If necessary, we adjust the *t*-statistic for autocorrelation using the Newey-West method.[6] However given the sufficient degrees of freedom it is still appropriate to use standard normal critical values.

Table III Panel A shows the regression results, using the pooling technique. This is given for each of the three sample periods during 1993 – 2002. We compare the coefficient values with our predicted values (see Table I) and observe that the results in Table III are generally in line with prediction.

[5] There is precedent for using the *DW* test in a pooled cross-section time-series analysis, most notably in Diaz and Skinner (2001).

[6] The Newey-West (1987) variance-covariance estimator is consistent in the presence of autocorrelation as well as heteroscedasticity.

Table III: Test Regression Results: Pooled Results
Dependent Variable is the Bid-offer Spread for Benchmark Bond

Panel A

Coefficients	Period 1 (1993-95)	Period 2 (1996-98)	Period 3 (1998-2002)	Full period (1993-2002)
Intercept α	0.00785[a] (4.16)	0.00687[a] (3.98)	0.00754[a] (4.12)	0.00694[a] (3.97)
Coupon β_1	0.0086 (1.43)	0.0009 (0.48)	0.0011 (0.78)	0.002 (0.19)
Term to maturity β_2	0.021[b] (3.23)	0.025[b] (3.94)	0.027[b] (2.96)	0.023 (1.99)
Issue size β_3	-0.015 (-0.32)	-0.009 (-0.57)	-0.014 (-0.54)	-0.005 (-0.64)
Volatility β_4	0.0891[a] (4.23)	0.0945[a] (5.42)	0.1027[a] (4.86)	0.1253[a] (4.33)
Base rate β_5	0.0001[a] (6.41)	0.00004[a] (4.67)	0.00001[a] (5.28)	0.00001[a] (4.82)
Macro announcement β_6	0.081[b] (3.13)	0.075[b] (3.72)	0.067[b] (3.26)	0.053[b] (3.47)
Dummy variable β_7	-	-	-	0.062[a] (4.45)
Adjusted R^2	0.54	0.53	0.49	0.55
Adjusted R^2 (model 7.4)	0.61	0.58	0.57	0.58
D-W statistic	2.142761	2.090534	1.994753	2.013697
White test $\chi^2_{9\text{df}}$	16.83	15.87	16.41	15.38
N	153	106	251	510

t-test in brackets, using Newey-West standard errors
[a] Significant at 1% level
[b] Significant at 5% level
Number of observations is per observation period. 51 weekly observations per year

Panel B Period 1 (1993-95): Time series
Dependent variable is recorded weekly bid-offer spread for benchmark bonds

Coefficients	Two-year	Five-year	10-year	Long-bond
Intercept α	0.00741[a] (4.39)	0.00717[a] (4.18)	0.00619[a] (4.37)	0.00724[a] (3.68)
Coupon β_1	0.0075 (2.05)	0.0005 (0.58)	0.0009 (0.91)	0.008 (0.35)
Term to maturity β_2	0.031[b] (3.41)	0.031 (2.94)	0.049[b] (3.15)	0.053[b] (3.95)
Issue size β_3	-0.018 (-0.67)	-0.012 (-0.61)	-0.014 (-0.45)	-0.002 (-0.56)
Volatility β_4	0.0828[a] (4.23)	0.0961[a] (5.42)	0.1027[a] (4.86)	0.1253[a] (4.33)
Base rate β_5	0.0001[b] (3.62)	0.00001[b] (3.71)	0.00001[b] (3.65)	0.00001[b] (3.47)
Macro announcement β_6	0.071[b] (3.73)	0.062[b] (3.22)	0.051[b] (3.89)	0.047 (3.09)
Adjusted R^2	0.59	0.54	0.51	0.55

t-test in brackets, using Newey-West standard errors
[a] Significant at 1% level
[b] Significant at 5% level

Panel C Period 3 (1998-2002): Time series
Dependent variable is recorded weekly bid-offer spread for benchmark bonds

Coefficients	Two-year	Five-year	10-year	Long-bond
Intercept $\boxed{\alpha}$	0.00811[a] (4.85)	0.00789[a] (5.21)	0.00723[a] (3.97)	0.00854[a] (3.88)
Coupon $\boxed{\beta_1}$	0.0064 (1.05)	0.0017 (0.61)	0.0013 (0.94)	0.005 (0.39)
Term to maturity $\boxed{\beta_2}$	0.042[b] (3.91)	0.045[b] (3.17)	0.035 (1.99)	0.042[b] (3.15)
Issue size $\boxed{\beta_3}$	-0.026 (-0.73)	-0.022 (-0.34)	-0.0015 (-0.67)	-0.0019 (-0.51)
Volatility $\boxed{\beta_4}$	0.0921[b] (3.86)	0.0885[a] (4.52)	0.1137[b] (3.69)	0.1289[a] (3.96)
Base rate β_5	0.0001[b] (3.57)	0.00001[b] (3.58)	0.00001[b] (3.95)	0.00001[b] (3.53)
Macro announcement β_6	0.065[a] (2.13)	0.047[a] (3.82)	0.072[a] (2.19)	0.059[a] (5.09)
Adjusted R^2	0.62	0.57	0.53	0.59

t-test in brackets, using Newey-West standard errors
[a] Significant at 1% level
[b] Significant at 5% level

The dependent variable is the recorded weekly bid-offer price spread (per £100 par value) of the benchmark bond, calculated as shown in paragraph 7.3.3. The lower and upper 5% critical values for DW test are 1.46 and 1.63 (T = 153, 106 and 251; k = 4 (model 7.2) or 5 (7.5), and the value of (4 – d_U*) is 2.37. The critical values of the Newey-West test are 3.96, 3.19 and 2.16 at the 1%, 5% and 10% level respectively. Note there are sufficient degrees of freedom, for example for model (7.2) for Period 1 there are 51 observations per year per bond, which is 153 observations, so df = (n – k – 1) = 148. For each year there is one observation per bond per week, which is 204 observations per year for all four bonds.*

To check if there is any difference overall between the various benchmarks, that is, any different sign or significance according to maturity of benchmark, for Period 1 we conduct a separate test for each benchmark bond. In other words, we differentiate according to maturity date of each bond. This test means that there are a much smaller number of bonds to run the test for, for example the actual long-bond benchmark bond remained unchanged during January 1993 through to December 1995. As such, for each week the independent variables C_i and I_i remained unchanged ones for the long-bond, because the same bond is represented each week and its coupon and issue size would be unchanged.

The results for this test are shown in Table III Panel B, with the results for Period 3 given at Panel C, which also indicate that the maturity of the particular benchmark is not significant. There are some detailed differences which we discuss next, but the sign of the coefficients are the same as for the pooled results at Panel A and so we consider these results to be valid.

We apply the White (1980) test for heteroscedasticity running an auxiliary regression of (2). This has nine regressors so there are 9 degrees of freedom. The critical value of χ^2_{9df} at the 5% level is 16.9190. From Table III we see that the test statistic does not exceed the critical value for any of the observation periods. We conclude that there is no evidence of heteroscedasticity in the time series data.

Evaluation of results

In general, results observed are as expected with regard to statistical significance and magnitude, and sign of coefficient.

The intercept fluctuates around 0.007 for the entire period, thus the predicted value of the bid-offer spread in the absence of *all* other explanatory factors is 0.7 basis points. In other words, this suggests that a reasonable estimation of "fair value" in the two-way bond price spread is 0.7 basis points, the value of the spread in the absence of influence from any of the explanatory variables.

We find no significant relationship between bond coupon and bid-offer spread. However the relationship *is* significant for the two-year bond. We believe that this reflects the different price sensitivity of short-dated securities compared to longer-dated securities. When a bond becomes short-dated (deemed as such once it has two years or less remaining to maturity), it starts to trade more like a money market instrument than a capital market instrument. This change in behaviour is subtle and gradual, but in essence it results in a tighter bid-offer spread. A bond with a high coupon, which has a more sensitive "pull-to-par" effect as it approaches maturity, will be priced above par for longer and is viewed as a less liquid instrument by investors. For this reason, a high coupon may be positively related to price spread, and this is borne out by the results in Table III.

A positive relationship exists with time to maturity. This contrasts with Ap Gwilym *et al* (2002), but is similar to the findings of Hong and Warga (2000) and Chakravarty and Sarkar (2003) who reported that US government and corporate bond spreads are significantly positively related to time to maturity.

Similar to Ap Gwilym *et al* (2002) and Kalimipalli and Warga (2002) for US dollar bond markets, we found a statistically significant positive relationship between price spread and volatility. That is, a wider price spread exists for bonds with higher uncertainty of return. Volatility has the largest impact on the magnitude of the bid-offer spread.

We find a statistically significant negative relationship between spread and issue size. The sign of the coefficient is as predicted. While a bond with a small issue size would be expected to have a lower liquidity, benchmark bonds all tend to exist in large size. Hence, the result appears to suggest that where a certain issue exists in relatively smaller size, this will impact on the price spread.

The base rate is statistically significant even at the 1% level but is not *practically* significant. This confirms that as the base interest rate value for all other market rates, it is has no practical impact on bid-offer spread: the implication is that it affects all bond prices, as well as other market prices, in the same manner.

The macroeconomic announcement week variable coefficient is positive and significant. The magnitude of the coefficient is large, indicating that this variable has a large practical impact on the bid-offer spread. This is perhaps surprising, because the week of the announcement is known in advance and it might be expected to be priced-in already into market makers' deal quotes. Nevertheless the regression suggests a widening of spread during this week. This mirrors results reported by Chakravarty and Sarkar (2003).

In summary, the bond bid-ask spread increases with time to maturity, on announcement weeks, and during times of higher market volatility. The nature of the relationship of the dependent variable to the independent variables is unchanged across all three sample time periods. Note that the bid-offer spread is not considered by market practitioners to be a proxy for the credit risk or the interest-rate risk of a bond, so we do not need to consider whether the independent variables are risk factors.

The results of the additional test featuring the dummy variable D_1 are encouraging. The β value is 0.062 which is positive and the result is significant at the 5% level. This may suggest that the explanatory variables had greater significance on market prices after the reforms had been completed. The magnitude is also of practical significance, suggesting that the impact of the reforms themselves was of the order of 6 basis points in price spread, on average for all bonds during the period after the reforms were completed. The coefficient value for the D variable may be taken to be a proxy for the quantitative impact of the market reforms.

The results reported in Panels B and C can be studied to determine the possible effect on bid-offer price for specific benchmarks during the period before and after the reforms. The volatility variable V has, as expected, greater impact on the longer-dated bonds compared to the shorter-dated bonds, reflecting their increased interest-rate risk sensitivity to market movements. Conversely, the macro-announcement variable A has greater practical significance for the shorter-dated bonds, which is to be expected because news announcements have more impact in the immediate term than over the longer term. Perhaps surprisingly, the magnitude of the coefficient does not diminish in the later period (Panel C), suggesting that despite the reforms it continued to have an impact on bid-offer spread. On the other hand the intercept value increases in the later period, implying that after the reforms had been completed the impact of all of the variables, when taken together, was reduced.

Overall the results indicate a consistency in the relationships in each regression. The models for each period appear to explain over half of the variation in bid-offer spread, with adjusted R^2 lying above 50% in each case. This is satisfying; it is similar to results reported by Ap Gwilym *et al* (2002) and Chakravarty and Sarkar (2003), and contrasts favourably with the lower R^2 reported in Hong and Warga (2000). Based on the adjusted R^2 there is a slight preference for the model containing the macro announcement variable, although we note that model (4) also has an increased number of variables and hence we expect, and observe, a higher adjusted R^2 value.

The Bid-Offer Spread: Tests of Structural Change

Observation of the results suggests a change in the average bid-offer spread over the time period under observation. We wish to test if the bid-offer function has undergone a structural change at any point during the entire observation period. The tests undertaken to test for structural change in the relationships are the Chow test, the Wald test, the Quandt test and the CUSUM squares test.

Chow test: results and evaluation

The application of the F-test in tests of structural change is common in the financial economics literature, as suggested in Greene (2000). In this context it is named the Chow test after Chow (1960) for tests at a known or suspected known break date. The model we test is the one shown at equation (4). We split the sample period 1993-2002 into two sub-sample periods, which are sub-sample 1 for 1993-1997 and sub-sample 2 for 1998-2002. Thus we obtain three estimated regressions, shown in Table IV. The first regression is that shown already for the complete period in Table III Panel A.

The Chow F-statistic obtained is $F[7,496] = 17.386$.

If the constant in (4) is fixed at the 5% level, the critical value $F_{7,\,496}$ is 2.01. As the Chow test statistic exceeds this value, we reject the null hypothesis that the bid-offer function in both sub-sample periods is the same. Thus, there has been a structural change in the relationship between the variables, and the coefficients are indeed different in the two periods. The test statistic also exceeds the 1% critical value $F_{7,\,496}$ which is 2.64. We suggest that this implies that much of the difference in the model over the two periods is explained by changes in the constant and one or more of the coefficients. This implies a break in the relationships in the period after the market reforms were introduced.

The Chow test does not indicate explicitly which coefficient (slope or intercept) is different or whether both are different in both periods. Because of this limitation with we conduct further tests.

Table IV: Bid-offer Regression Equation: Chow Test
Sub-sample periods and Wald Statistic

Coefficients	Full period (1993-2002)	Sub-sample period 1 (1993-1997)	Sub-sample period 2 (1998-2002)
Constant α	0.00694	0.00717	0.00753
Coupon β_1	0.002	0.0007	0.0011
Term to maturity β_2	0.023	0.021	0.028
Issue size β_3	-0.005	-0.0008	-0.0147
Volatility β_4	0.1253	0.0912	0.1024
Base rate β_5	0.00001	0.00003	0.00001
Macro announcement β_6	0.053	0.076	0.0674
Adjusted R^2	0.58	0.51	0.57
Sum of squares	0.2812	0.0153	0.0731
Wald statistic	17.987	-	-
N	510		
N_1		255	
N_2			255
k	7	7	7

Wald test

We test for possibility of a "Type I" error using the Wald test for unequal variances, valid with large sample sizes.

The Wald test statistic for the regression shown in Table IV gives a value of 17.987. The χ-squared 5% critical value for 7 degrees of freedom is 14.07. Thus under the Wald test we reject the null hypothesis that the same coefficients apply in both sub-sample periods 1 and 2.

The QLR test for break at an unknown date

When conducting the Chow test we selected the date at which to split the dataset into two as the point at which implementation of the market reforms was complete. This pre-supposes that this was the date at which the structural break occurred. More realistically, the date of the possible break could lie at any point during the period the reforms were being undertaken – a time frame of 25 months – or on a point shortly after the reform process was completed. Because we cannot state the date of a possible break with certainty, we apply a modified Chow test which is the Quandt likelihood ratio (QLR) statistic.

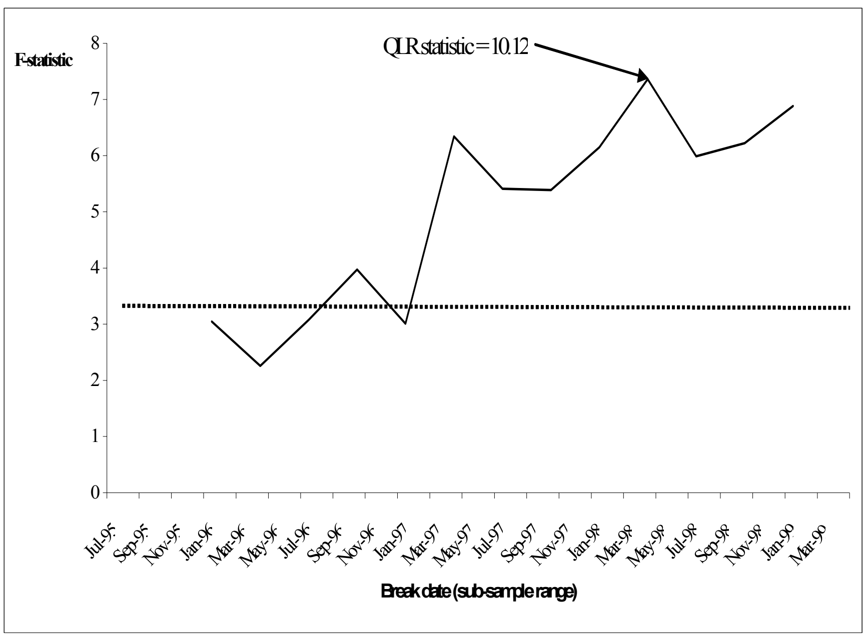
Our supposition is that the break lies during the period from the start of the reforms through to a point up to 12 months following their conclusion. We therefore apply the QLR test for breaks at all possible dates from January 1996 [t_1] to January 1999 [t_2] which is a sub-sample range within the full range t_0 to T of January 1993-December 2002.

We compute the *F*-statistic for possible break dates within the sub-sample range, for which we select quarterly periods in the range. The number of restrictions is 7 which is equivalent to the number of degrees of freedom in the standard *F*-test. Results are shown in Table V.

We observe that at the largest of the F-values on April 1998 we exceed the 5% critical value, suggesting that this is an estimator of the break date. However the value is also exceeded for the remainder of the sub-sample period suggesting that there is a gradual evolution of the regression function from this date to the end of the sub-sample period. Thus we conclude that at least one of the coefficients in model (4) has changed during this sub-sample period.

Table V: QLR Test for Break in Equation (4) during Sub-sample Period January 1996-January 1999 (*F*-statistics)

Date t_n	*F*-statistic
Jul-95	
Oct-95	
Jan-96	3.05
Apr-96	2.26
Jul-96	3.08
Oct-96	3.97
Jan-97	3.01
Apr-97	6.34
Jul-97	5.41
Oct-97	5.39
Jan-98	6.15
Apr-98	7.36
Jul-98	5.99
Oct-98	6.22
Jan-99	6.88
Apr-99	



F-statistic tests H_0 of a break in one or more of the coefficients or the intercept in Equation (7.4)

QLR statistic is the largest of the results obtained

The 5% critical value with 7 restrictions is 3.15

CUSUM of squares (CUSUMQ) test

A further test of model stability is the CUSUM test suggested by Brown, Durbin and Evans (1975). We apply the similar but alternative test CUSUM of squares (CUSUMQ) based on the cumulative sum of squared residuals. Under this technique we run the regression with no pre-determined break point, adding one period at a time, to see if the results indicate a coefficient change.

We calculate the CUSUMQ test statistic and plot this and the upper and lower confidence bounds (at the 5% significance level, with k equal to 7) against time t. The results are shown at Figure 2.

Figure 2: CUMSUM of Squares Test, Bid-offer Spread

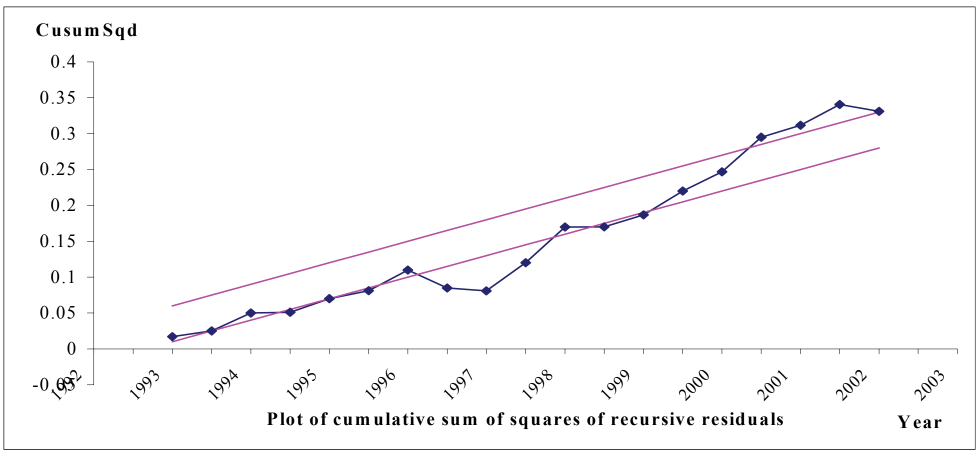


Figure 2 suggests that the model experiences instability at a number of points during the observation period, where the test statistic moves outside the confidence limits. This indicates model instability, and that we should reject the null hypothesis of parameter stability.

We note however that the test statistic indicates instability occurring both *during* the period of reforms as well as *after* their full implementation. That is, the statistic lies outside the confidence bounds during what we have labelled as "Period 2", and then during the second half of Period 3. This is broadly consistent with the results obtained earlier, although the dates of the implied break are different. For the Chow test our selected cut-off date was January 1998; for the QLR test the time of the most significant score was April 1998; and for the CUMUSQ test the test score was significant at January 1998 but not at June 1998. It was continuously significant in the second half of Period 3.

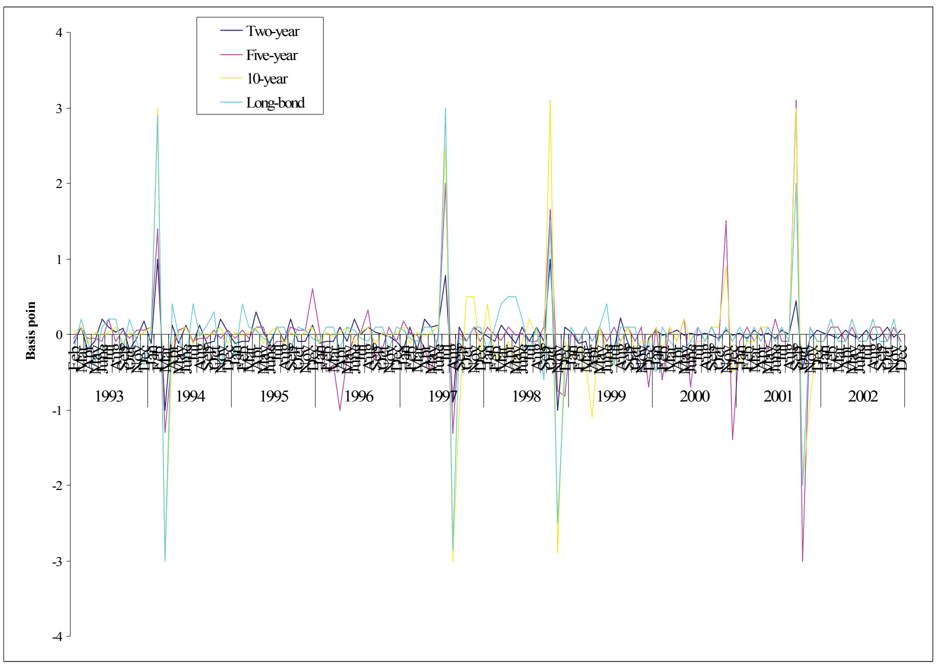
All the tests so far indicate an element of structural break in the model, but not necessarily at the same point in time. This may reflect other factors at work influencing market liquidity and the magnitude of the bid-offer spread, which are not specified in our model. It may also reflect the anticipatory effect of the BoE's reform announcement, which has been observed elsewhere in our study.

The Bid-Offer Spread: Test for Cointegration

The econometric analysis we apply here can be relied on to give valid results if the time series data is stationary. In practice, time series that have linear and stochastic trends and are I(1) are often observed to move together. This suggests that their difference is I(0) or stationary; for example see Hayashi (2000). We test now for this property in our bid-offer spread time series dataset.

Gujarati (1995) and Stock and Watson (2003) suggest that visual observation of a plot of the time series is a valid initial test of cointegration. Hatanaka (1996) states that "…visual inspection of time series charts is indispensable to empirical studies as it suggests what kind of formal tests should be applied." (1996, p.9) Figure 3 shows the first-differenced bid-offer spread data. We note that the differenced data does not exhibit any trend.

Figure 3: First Differences of Bid-offer Spread



Because the original bid-offer time series is non-stationary, we conduct the Johansen test for cointegration in a multivariate environment. We ignore the A macroeconomic announcement variable, which leaves five variables in the system; that is, $g = 5$. Therefore there can be at most 4 linearly independent cointegrating vectors, so that $r \leq 4$. We use the trace statistic formula shown in Brooks (2002, p.404).

Results are shown in Table VI. Critical values are given in Osterwald-Lenum (1992). We see that the test statistic is larger than the critical value for the first two rows, so we reject the null hypothesis for $r = 0$ and $r = 1$ at the 5% level. The null cannot be rejected for higher r. Therefore we conclude that there is at least one cointegrating vector in the series. This weakly supports the conclusions reached in the previous section.

Table VI: Results of Johansen Test for Co-integration between Bid-offer Spread Explanatory Variables

r (number of cointegrating vectors under null hypothesis)	Test statistic	5% critical value
0	41.6	38.6
1	24.7	23.8
2	9.3	12.1
3	1.51	4.2
4	0.06	3.4

Summary and Conclusions

The test results suggest that market liquidity, as measured by the bid-offer proxy, had increased in the post-reform period. Therefore the structural reforms introduced in the UK gilt market by the BoE were associated with a period of increased liquidity in the market. Our conclusions rely on the values of the proxy measure itself over the time period under observation, the increase in magnitude and significance of the independent variables, and the evidence of a structural break in the time-series data after the reforms had been completed. We provide detailed conclusions and policy recommendations below.

Conclusions from statistical tests

In this study we report the findings of econometric testing of the explanatory variables behind a proxy measure of liquidity, the bid-offer spread, in order to determine (i) if these variables are significant, (ii) whether the relationships have changed over the time of the observation period, and (iii) whether the implied level of liquidity has increased during the time period under observation.

In the first instance we conclude that market liquidity levels had improved in the period following the introduction of the structural reforms.

We tested the determinants of the benchmark bond bid-offer spread. We observed the mean weekly bid-offer spread for each benchmark gilt for the sample period under study. Bid-offer spreads for all the benchmark bonds had narrowed during the period under study. Measured by this indicator, gilt market liquidity had improved during the time pre-reform to post-reform. We are not able to verify this with complete certainty, because other factors may also have contributed to the reduction in the quote spread, but it is a reasonable conclusion to infer.

The statistical analysis we conducted accounted for over 50% of the total variation of the price spread. We determined from model testing that the bid-offer price spread is influenced as follows:

- longer time to maturity (interest-rate risk), and higher volatility: wider bid-offer spread;
- larger issue size: narrower bid-offer spread.

One of the Bank of England's market reforms had created larger size benchmark issues, so we conclude that this reform was a key factor in improved liquidity.

Volatility was shown to be a significant causal factor, with higher volatility influencing wider price spreads and hence lower liquidity. That the *overall* level of liquidity was improving during a period of high volatility, the latter in part connected with several market crises during 1994-2001, implies a degree of success associated with the BoE reforms in helping to create a more orderly market.

We observed that there was a structural break in the bid-offer spread time series data, both during and after the reform period. This implies a change in the relationship between the variables. The results of the CUSUMQ test were strongly consistent with the earlier results and suggested that bid-offer spread experienced instability at a number of points during the observation period. We concluded that the bid-offer spread is cointegrated with four of its explanatory variables, these being bond coupon, bond term-to-maturity, bond volatility and the central bank base rate. Although they may individually exhibit a random walk feature, the evidence appears to suggest a stable long-run relationship between the bid-offer spread and the four variables.

This result suggests with a reasonable degree of certainty that market liquidity had improved during this period.

Impact of market reforms and policy recommendations

Our preliminary conclusion, that gilt market liquidity as measured by the proxy indicator had increased during the observation period, leads to a second conclusion that the market structural reforms introduced by the BoE led to improved liquidity in the gilt market. This has further implications for monetary authorities worldwide, as it suggests that they should introduce repo and strips markets, and a uniform trading infrastructure, in order to improve liquidity in their respective markets.

Our results show also that the main explanatory variables are significant determinants of the bid-offer spread, in itself a proxy measure of liquidity. This is a key measure of the effect of the reforms in market efficiency: monitoring these variables and understanding the relationship between them will assist in the maintenance of an orderly, liquid market.

One further conclusion we draw from the results is that they suggest how market liquidity can be maintained through most trading conditions. For example, we see that a relatively narrow bid-offer spread, implying a liquid market, is influenced by the level of market volatility, benchmark bond issuance and other factors. In times of market correction or instability, central banks and regulatory authorities may wish to consider these factors when addressing market policy.

Finally, the suggested increase in liquidity levels reflects positively on the value of an open repo market, the most significant of the BoE's reforms in the gilt market. Therefore sovereign debt agencies should consider introducing this reform as an aid to increasing and maintaining market liquidity.

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Impact of Electronic Tax Registers on VAT Compliance: A Study of Kenyan Private Business Firms

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Abstract: *The purpose of this study was to assess the impact of use of Electronic Tax Registers (ETRs) on Value Added Tax (VAT) compliance among private business firms in Kisumu city, Kenya. A sample of 233 private firms was selected from a population of 590 private firms using stratified sampling technique. The primary data was gathered using structured questionnaires and analysed by use of correlation and descriptive statistics. Empirical results of the study indicate that effective and regular use of ETR has a significant positive impact on the VAT compliance. The regular frequency of inspection of private businesses by the tax authorities (staff from Kenya Revenue Authority) has a slight impact on VAT compliance however; sales of Private Business Firms show insignificant relationship with VAT compliance. From these findings, the study concludes that inspection of businesses by tax authorizes as well as use of ETRs are major determinants of VAT compliance among private Business Firms in Kenya.*

Keywords: Value added tax, electronic tax registers, private business firms

Introduction

Value Added Tax, (VAT) on consumer expenditure was introduced in Kenya in 1990 in order to replace sales tax, which had been in operation since 1973. It was introduced as a measure to increase government revenue through expansion of tax base, which hitherto was confined to income tax and sales tax. VAT is levied on consumption of taxable goods and services supplied in Kenya or imported into Kenya. Registered persons acting as agents of government of Kenya collect VAT at designated points and then submit to the Kenya Revenue Authority (KRA) (Simiyu 2003). Previous empirical study conducted by Moyi and Ronge (2006), indicates that VAT contribution is estimated to an average of 5.4% of GDP between the year of its introduction (1990) and the year 2005. The average of total tax contribution to GDP for the same period was 19.8%. This clearly indicates that in Kenya, VAT contributes substantially to the growth of the economy (table 1).

Another study conducted by Waris et al, (2009) reveals that despite the importance of VAT in the national budget, the period between the year 2000-2003 showed that VAT had the highest share of total tax (above 30%). However, VAT contribution trend declined to total taxes collected from the year 2003 onwards as given in table 1 which captures the composition of various taxes to total taxes in Kenya (1996-2008). This trend is worrying and calls for intervention reforms. Kenya revenue Authority (KRA) has since introduced several reforms in its revenue collection system including the introduction of Electronic Tax Registers.

ETRs were first introduced to Kenya in 2004, through a gazette notice no. 47 issued in October 22, 2004. According to this notice, ETR or printer is defined as any device approved by the