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INFORMATION IN T-BILL AUCTION BID DISTRIBUTIONS[1]

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Abstract

In this paper UK data is used to compare two potential sources of information regarding market uncertainty about future short interest rates. One is the so-called risk-neutral density function (RND) derived from interest rate option prices, the other is the distribution of bids submitted to an auction of short-term Treasury bills. More specifically, time series of RND standard deviations and auction bid standard deviations are compared. The results suggest that in some periods the auction bid standard deviations co-moved with those of the RNDs. Thus, in principle, auction bid standard deviations may be useful to get a picture of market uncertainty about future short rates even in the absence of well-developed interest rate options markets.

In the Supplement, encouraged by the above results, the author uses Hungarian T-bill auction data to check whether auction bid dispersion measures in Hungary make any sense as indicators of market uncertainty about future interest rates. Lacking any RND data for this country, this can only be done in indirect ways. These include looking at the correlations of auction dispersion measures of different T-bill maturities, comparing the time series of these measures and bid-ask spreads (another possible indicator of uncertainty) and conducting an intuitive consistency check for a certain time period.

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Introduction

In recent years, central banks in developed countries have been paying an increased attention to asset prices. In addition to being available with high frequency and in good quality, asset prices are determined in markets which are predominantly forward-looking. This characteristic of asset prices makes them straightforward candidates for indicators to be used by central banks, as market expectations and their responses to monetary policy measures are captured in these prices the most quickly and most reliably.

The market's expectations of future interest rates are of particular interest to central banks whose primary instrument is the short-run interest rate. These expectations can be viewed as subjective probability distributions formed by the market about possible future outcomes of an interest rate with a given maturity. In the absence of risk premia, interest rate forward and futures markets, or implied forward rates derived from the yield curve, provide some information about the mean of these probability distributions. More recently some techniques have been developed to estimate (assuming risk-neutrality) the whole subjective probability distribution of a future interest rate from the prices of options on that future interest rate. Higher moments of these distributions, such as variance, skewness and kurtosis provide useful information for central banks about the size and nature of uncertainty about a particular future interest rate, enhancing the information incorporated in the mean. Of course, these techniques are of little use in countries where interest rate option markets are illiquid or do not exist at all.

However, there is another occasion when the market may possibly reveal the whole of its subjective probability distribution of future interest rates: the auction of government securities. Bidders on these auctions try to guess the level of the spot interest rate the market as a whole considers to be "fair". In doing so they bid yields (or prices) together with amounts, thus attaching weights to different possibly "fair" levels of the interest rate. The auction bids therefore can be represented as a distribution.

Whether or not the level of the spot rate is "fair" depends on expected future rates, if the Expectation Hypothesis is supposed to hold in some form. Uncertainty about future rates will appear in bidders' subjective probability distributions about the "fair" value of the spot rate, affecting the second and higher moments. If the auction scheme is efficient in forcing bidders to bid "honestly", i.e. according to their subjective probability distributions, and if bidders can be taken as representing the market as a whole, then the auction bid distribution represents the market's subjective probability distribution about the "fair" value of the spot interest rate, which, as it was mentioned earlier, is related to the market's subjective probability distribution of future interest rates.

This suggested relationship rests upon many assumptions and needs to be tested empirically. For this purpose, U.K. data seemed especially appealing, since in this country detailed T-bill auction data and, more importantly, a sufficiently long time series of option-based estimates of risk-neutral density functions (RNDs) on future short rates were available. We used the U.K. data to examine whether auction bid distributions contain part of the information which is present in option-based RNDs. The main motivation behind this was to see whether it is sensible, in the absence of sophisticated interest rate option markets, to use auction data in order to get a picture about the market's uncertainty about future interest rates.

The structure of the paper is as follows: Section 1 outlines in what ways the probability density function of the short interest rate can be useful for the monetary authority. Section 2 describes how the market's subjective probability distribution of a future asset price is related to the option-based implied RND and how the latter can be estimated from option prices. Section 3 sets out a theoretical framework in order to demonstrate how T-bill auction bid distributions are related to the market's subjective probability distribution of short interest rates and, indirectly, to option-based implied RNDs. Section 4 uses U.K. data to investigate this relationship empirically, after making some assumptions about the term structure of interest rate uncertainty. Finally, Section 5 concludes.

In the Supplement, Hungarian T-bill auction data is used to check whether auction bid dispersion measures in Hungary make any sense as indicators of market uncertainty about future interest rates. Naturally, lacking any RND data for this country, this can only be done in indirect ways.

1. The relevance of probability density functions of the short interest rate for monetary policy

Option-based implied RNDs represent the probabilities the market attaches to different future short rates and, since short rates to some extent depend on monetary policy actions, to future monetary policy steps.

Therefore, these probability distributions may help to assess the market's possible reaction to a monetary policy action. In this respect, higher moments of the distribution have useful information in addition to that already incorporated in the mean (the forward rate). Suppose for example, that the central bank decides to raise the short interest rate. The market's reaction to this action may be different if its subjective probability distribution was narrow and symmetric around the present rate (i.e. if they did not expect any change) compared to the case in which the probability distribution had the same mean but was more disperse and/or skewed towards higher values (i.e. the market attached a higher weight to a rate hike).

Subjective probability distributions of the future short rate are also useful for the *ex post* analysis of monetary policy steps. If, for example, the shape of the distribution remains the same after a rate change by the central bank, then one can conclude that the market fully expected the action. In contrast, an unchanged mean in itself not necessarily means the same, since higher moments of the distribution (representing the market's uncertainty) may have changed significantly.

2. Option-based Risk-Neutral Densities (RNDs) as measures of interest rate uncertainty

The derivation of RND functions from option prices is based on a result by *Breeden* and *Litzenberger* (1978), which states that if a continuous probability density function exists for the price of the underlying asset at the time of expiry, S_T, then the second derivative of the European call pricing function[2] is equivalent to (the present value of) this distribution:

[2] A European call pricing function gives the prices of European call options on the same underlying asset and with the same expiry date (T in the formula above) as a function of the strike prices of the options.

$$\left.\frac{\partial^2 c(X)}{\partial X^2}\right|_{X=S_T} = e^{-r(T-t_0)} q(S_T)$$

where:

X is the exercise (strike) price
T is the expiry date
r is the risk-free interest rate over the period T-t_0
$q(S_T)$ is the (risk-neutral) probability density function of the underlying asset's price at expiry

Therefore, if we had call options on the same underlying asset with strike prices covering the whole range of $(0,\infty)$ we could derive the whole density function. Of course, in practice this is not the case as the number of strikes is limited. This means that the RND function has to be estimated from the observable option prices. There are different strategies to do this estimation. The first is to interpolate between the observable points on the call pricing function using an assumed parametric functional form (e.g. a cubic spline, see *Bates* (1991)). The same interpolation can be done after translating the call pricing function into an implied volatility function using the Black-Scholes formula (or an appropriate version of it) . The advantage of this procedure is that the Black-Scholes implied volatilities are supposed to be more smooth than the option prices themselves (see *Shimko* (1993)). Having estimated the implied volatility "smile" curve, it can be translated back to the call price domain, giving a continuous call pricing function, from which the (discounted) RND function can be derived as the second derivative.
Another strategy is to assume a parametric form directly for the RND function and estimate its parameters so that the distance between the observable option prices and those generated by the assumed RND is as small as possible. Knowing that asset price distributions can be crudely approximated with the lognormal density function, it seems reasonable to follow the procedure suggested by *Ritchey* (1990) and assume that $q(S_T)$ is the weighted sum of k lognormal density functions. A density function of the terminal price consisting of independent lognormal densities can be sufficiently flexible, in particular it can be skewed and leptokurtic, thus producing the empirically observable implied volatility "smile". *Melick* and *Thomas* (1994) apply this methodology using the mixture of three lognormal densities to extract implied RNDs from option prices on crude oil futures. However, for a number of financial option markets the range of strike prices is relatively small, which limits the number of distributional parameters that can be estimated from the data. *Bahra* (1997) suggests that the mixture of two lognormal densities, which requires the estimation of five parameters only (two means, two variances and a weight parameter), is still flexible enough to capture the main features of the terminal price distribution. He also mentions two other "tricks" to increase the degrees of freedom for the estimation. One is to take the observable forward price of the underlying asset to be the mean of the estimated RND, another is to use put options as well as calls. He then applies the two-lognormal methodology to equity, interest rate and foreign exchange markets. *Söderlind* and *Svensson* (1997) show that if the true distribution of the log stochastic discount factor and the log bond (or short instrument) price is a mixture of bivariate normal distributions, then the risk-neutral distribution is indeed a mixture of log-normals.

It is the two-lognormal methodology which is used by the Bank of England to recover implied RNDs from LIFFE option prices on the 3-month sterling interest rate. Part of the data we use in this paper is the daily time series of (the variances of) these estimated RNDs.
As their names suggest, these density functions are risk-neutral, i.e. they should coincide with the "true" subjective distribution of future interest rates (or any other underlying asset) if investors are risk-neutral, which sounds unrealistic. However, some results suggest (see *Rubinstein* (1994)) that, at least for some markets, implied RNDs and subjective probability distributions may not differ significantly. Moreover, if the degree of risk aversion is constant, then changes in the implied RND can be attributed to changes in the subjective probability distribution alone.

3. The relationship between auction dispersion measures and interest rate uncertainty

The RND function implied by the prices of options on the short interest rate expresses the probabilities the market attaches to different future outcomes of the short interest rate. It is less clear, however, what the distribution of yield bids on a T-bill auction represents, since we tend to think of spot interest rates as non-random. In order to understand better what is going on at T-bill auctions it is useful to rely on a particular formulation of the pure expectations hypothesis (PEH) and introduce the concept of the *fair value* of the spot interest rate. The formulation of the PEH used here states that the *n*-period interest rate is the multiple of expected future one-period interest rates:

$$\left(1+i_{n,t}\right)^{n} = E_t\left[\left(1+i_{1,t}\right)\left(1+i_{1,t+1}\right)\ldots\ldots\left(1+i_{1,t+n-1}\right)\right] \qquad \text{Eq.1}$$

where:
$i_{n,t}$: the *n*-period nominal spot interest rate at time t
E_t: expectations conditional on information at time t

Or in log form, assuming away the terms generated by Jensen's inequality:

$$i_{n,t} \approx \frac{1}{n} E_t\left(i_{1,t} + i_{1,t+1} + \ldots\ldots + i_{1,t+n-1}\right) \qquad \text{Eq.1*}$$

Let us define $i^{f}_{n,t+n}$, the fair value of the spot *n*-period interest rate as the *n*-period rate which ex post equals the multiple of the *realized* 1-period rates, i.e.

$$\left(1+i^{f}_{n,t+n}\right)^{n} = \left(1+i_{1,t}\right)\left(1+i_{1,t+1}\right)\ldots\ldots\left(1+i_{1,t+n-1}\right) \qquad \text{Eq.2}$$

Or in log form (again neglecting Jensen's inequality):

$$i^{f}_{n,t+n} \approx \frac{1}{n}\left(i_{1,t} + i_{1,t+1} + \ldots\ldots + i_{1,t+n-1}\right) \qquad \text{Eq.2*}$$

It is obvious that the fair value of the spot *n-period* rate is not revealed before the maturity of the *n*-period bond (or *t+n-1*, if, as in Eq.1-2, we decompose the bond's maturity into periods of unit length). At any time before that, the fair value of the *n*-period spot rate is random and therefore has a probability distribution. It is this distribution (or more precisely, the market's view of it conditional on information at time *t*) which is revealed at an auction on the *n*-period bond. (Provided that bidders are risk-neutral, there are no transaction costs associated with taking part in the auction, and that the auction scheme successfully discourages strategic behaviour.) Assuming that bidders represent the market as a whole, the auction bid distribution is the market's view on the probability distribution of the fair value of the *n*-period interest rate.
Now take the expectations of both sides of Eq.2* and then use the PEH form Eq.1* to get

$$E_t\left(i^f_{n,t+n}\right) = i_{n,t} \qquad \text{Eq.3}$$

Thus, in a risk-neutral world, the current *n-period* market spot rate is the mean of the market's subjective probability distribution of the fair value of the *n*-period interest rate, and as such, is fixed (non-random). Although this distribution exists at any time *t*, the only occasions when the whole distribution (not only its mean) is revealed are the auctions.
The variance of the fair value $i^f_{n,t+n}$ is given by

$$\mathrm{var}\left(i^f_{n,t+n}\right) = a'Va \qquad \text{Eq.4}$$

where:

a is an *n* x 1 column vector of weights ($\left(\frac{1}{n}\right)$s in this case)

V is the *n* x *n* variance-covariance matrix of the one-period rates $i_{1,t}, i_{1,t+1} \ldots\ldots, i_{1,t+n-1}$

Taking expectations, we get the expected variance conditional on information at time *t*, which is the second moment of the market's subjective probability distribution of the fair value of the *n*-period spot interest rate.

$$\mathrm{var}_t\left(i^f_{n,t+n}\right) = a'V_t a \qquad \text{Eq.4*}$$

where:

V_t is the *n* x *n* matrix of expected variances and covariances of the one-period rates $i_{1,t}, i_{1,t+1} \ldots\ldots, i_{1,t+n-1}$

The decomposition of the *n*-period bond's maturity into periods of unit length is of course arbitrary: one may choose shorter or longer periods (or a mixture of them) and the resulting fair values will not necessarily be the same unless, during the life of the *n*-period bond, each spot rate was equal to its fair value. However, if the market indeed forms a view about the fair value of the spot *n-period* interest rate, then it also specifies an implicit (unobservable) decomposition of the *n-period* bond's life into subperiods.

It seems reasonable to argue that for small values of *n*, the spot rate is always quite close to its fair value. For example, the overnight market rate probably always equals its fair value, since there are few intraday investment opportunities. Let us assume that this is true for a bit longer maturity as well, i.e. that the 1-month rate always equals its fair value. In this case, since the 1-month spot rate is fixed, the probability distribution at time *t* of the fair value of the *n*-month spot interest rate is the probability distribution of the *n*-1-month interest rate 1 month ahead (assuming that the market subjectively decomposed the *n* periods into an initial 1-month and a subsequent *n*-1 month-long period).
The ultimate aim of this lengthy exposition was to demonstrate how the distribution of auction yield bids relates to uncertainty about future rates, and, in this indirect way, to the probability distributions derived from option prices, and now we are getting closer to it. Suppose we observe the distribution of the yield bids in an auction on the 3-month T-bill. This is the market's view at the time of the auction on the fair value of the 3-month interest rate in the form of a probability distribution. Under our previous assumptions (i.e. that the 1-month rate always equals its fair value and that the market subjectively divides the 3 months into an initial 1-month and a subsequent 2-month periods) this probability distribution will be the probability distribution the market attaches to the 2-month interest rate 1 month ahead. Now if there were options (with a sufficient number of strike prices) on the 2-month rate expiring exactly in 1 month, then we would expect the probability distribution derived from the prices of these options to be the same as the distribution of the auction yield bids[3].
The decomposition of the 3-month period in the above example is again arbitrary: in fact we are never able to tell which future rate's probability density function affects the auction bid distribution, while in the case of option-based RNDs both the maturity (2 months in the example) and the horizon (1 month ahead) of the underlying asset is exactly specified. However, even if we do not know exactly which future rate affects the auction bid distribution, in principle it still has some information about the market's judgement of future spot rates in general. One might say for example, that the standard deviation of the bid distribution of an auction on the 3-month T-bill today reflects the degree of uncertainty about short rates with an unspecified *n* months of maturity *m* months ahead, for which $n+m=3$. Although less precise than the standard deviation of the option-based RND, this information is still relevant.
Our implicit assumption, that the auction demand curves reflect the true demand curves of bidders is challenged by many studies from the field of auction theory. *Friedman* (1960) for example, suggested that multiple-price auctions (the type we will examine later), mainly because of the "winners' curse", discourage bidders from revealing their true demand curves and encourage collusion. However, adding the information linkage between auctions and the secondary market, which may induce signalling from bidders, as it was done by *Bikhchandani and Huang* (1989), makes these results less straightforward. Uniform price auctions in this framework yield higher revenues than multiple-price auctions only under certain conditions.
In our empirical analysis we did not rely on any of these ambiguous results from auction theory. All we wanted to see if the auction bid distributions contained part of the information which was present in the subjective distributions represented by the option-based RNDs.

[3] Assuming that the difference between default risk on interbank loans and that on Treasury securities is negligible.

4. The information content of auction dispersion measures in the UK

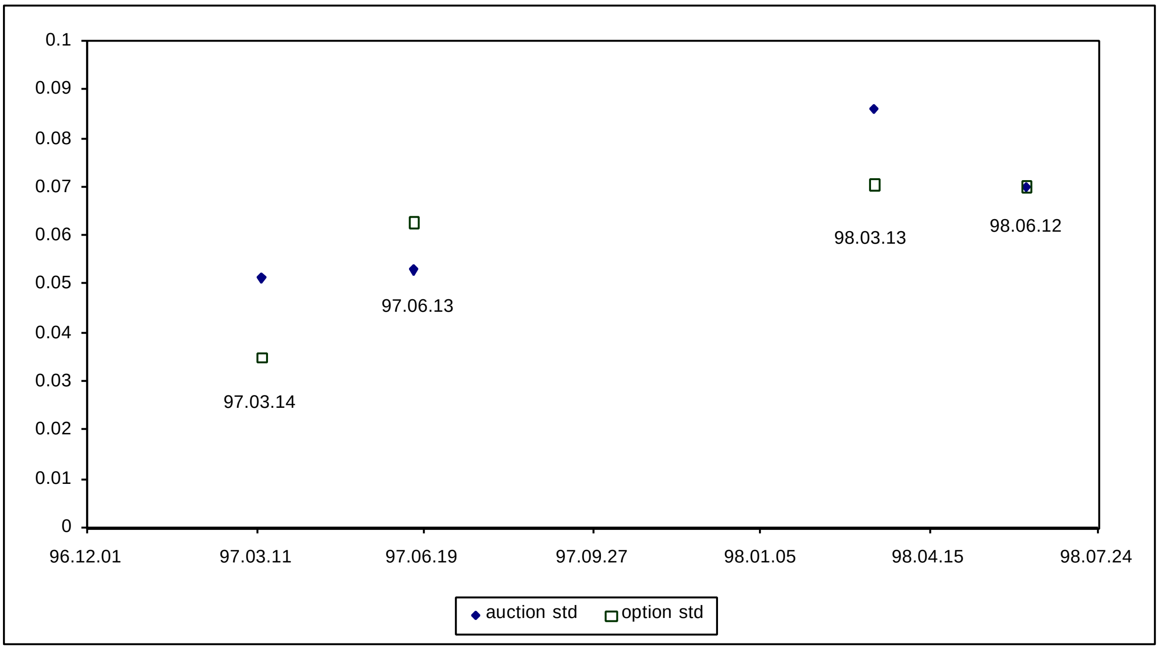
4.1 Data issues

In the previous section we have established a theoretical relationship between auction bid distributions and option-based RND functions of the future spot interest rate. In the next section we will examine whether there is any empirical evidence supporting this relationship. For this purpose, we use U.K. data, partly because the interest rate derivative markets in the U.K. are well-developed and quite liquid but mainly because by now the estimation of implied RND functions is fully operational in the Bank of England and data series with a daily frequency are readily available. When comparing auction bid distributions and option-based RNDs we restrict our focus to the standard deviations (and another measure in the case of the former, see below) of these distributions as the main measure of uncertainty about future rates. Higher moments such as skewness and kurtosis may contain useful information as well, but we wanted to focus on a first check of the hypothesised relationship between the two types of distributions, and left the examination of higher moments for future research.

As it is clear from the previous section, ideally one would compare the auction bid distribution on a particular maturity with the option-based RND on the same day, where the underlying asset of the options used matures exactly when the auctioned instrument matures. For instance, as in the example of the previous section, we would like to compare the auction bid distribution on the 3-month T-bill with the option-based RND derived from options with different strike prices written on the 2-month rate 1month ahead (or the 1-month rate 2 months ahead, etc.). However, such matching pairs are difficult to find. The options used to estimate implied RNDs by the Bank of England are all written on the 3-month short sterling rate with usually 4 or 5 expiry dates available, separated by 3 months and with the last one approximately 1 year ahead. At the same time, in recent years the weekly T-bill auctions in the Bank of England have involved the 1-month and 3-month maturities. There are auctions by the Debt Management Office for significantly longer instruments (gilts) maturing well beyond the expiry of the longest option contract. So recently there are no auctions on instruments with maturities between 3 months and 3 years, which poses a significant problem for us, since the 6-month and 1-year maturities would probably be the most suitable for our analysis.

Another problem stems from the fact that short sterling option expiry dates are fixed, therefore as the expiry gets closer, uncertainty about the short sterling rate, and therefore the standard deviation of the corresponding RND gets (ceteris paribus) smaller, while no such thing is expected to happen with the standard deviation of the weekly auction bid distributions, which reflect uncertainty on a constant horizon. All these problems mean that a direct comparison of the levels of the standard deviations of the auction bid distributions (on the longest possible T-bill, i.e. the 3-month) and the implied RNDs would only be possible once every 3 months, when an option gets very close to expiry. This would mean loosing roughly 11/12 of our auction observations, of which we have a limited amount anyway. Another problem is that the estimation procedure of the implied RNDs is rather unreliable with such a short time remaining until expiry. Bearing this in mind, we nevertheless plotted the standard deviations from both distributions for the 4 dates in our sample for which we could find both auction data and an option maturing in less than a week in Chart 1. As can be seen from the chart, the magnitudes of the standard deviations from the two sources are of the same order.

Chart1



One solution to these problems was to construct a time series of standard deviations which, using a simple linear interpolation between the standard deviations of RNDs corresponding to two consecutive option expiry dates, has a constant 6-month horizon, i.e. on each day it crudely reflects the standard deviation of an RND from hypothetical options on the 3-month rate, expiring exactly 6 months ahead[4].

The two option expiries chosen for the interpolation are always the second and the third in the order of expiry, ensuring that the end of the 6-month horizon always falls between these two. The formula for calculating this approximation of the 6-month constant-horizon RND standard deviation $(\overline{s})$ was:

$$\overline{s} = s_2 + (s_3 - s_2)\frac{126 - t_2}{t_3 - t_2} \qquad \text{Eq.5}$$

Where:

s_2, s_3 : the standard deviations of the RNDs from the 2nd and 3rd expiry dates, respectively
t_2 , t_3: distance of the 2nd and 3rd expiry dates from today's date in business days, respectively (126 was assumed to be the number of business days in 6 months)

[4] Theoretically, it would have been more appropriate to use linear interpolation between variances instead of standard deviations. However, this did not change the main empirical results, so the original formulation (using standard deviations) was kept.

Of course, a constant horizon shorter than 6 months would have suited better our 3-month T-bill auction data, but as it was mentioned earlier, the estimation of the standard deviations of the RNDs corresponding to the first expiry date tends to get rather unreliable as the expiry gets very close, therefore the RND of the first expiry date was omitted from the weighting scheme of Eq.3, and the 6-month horizon was chosen.

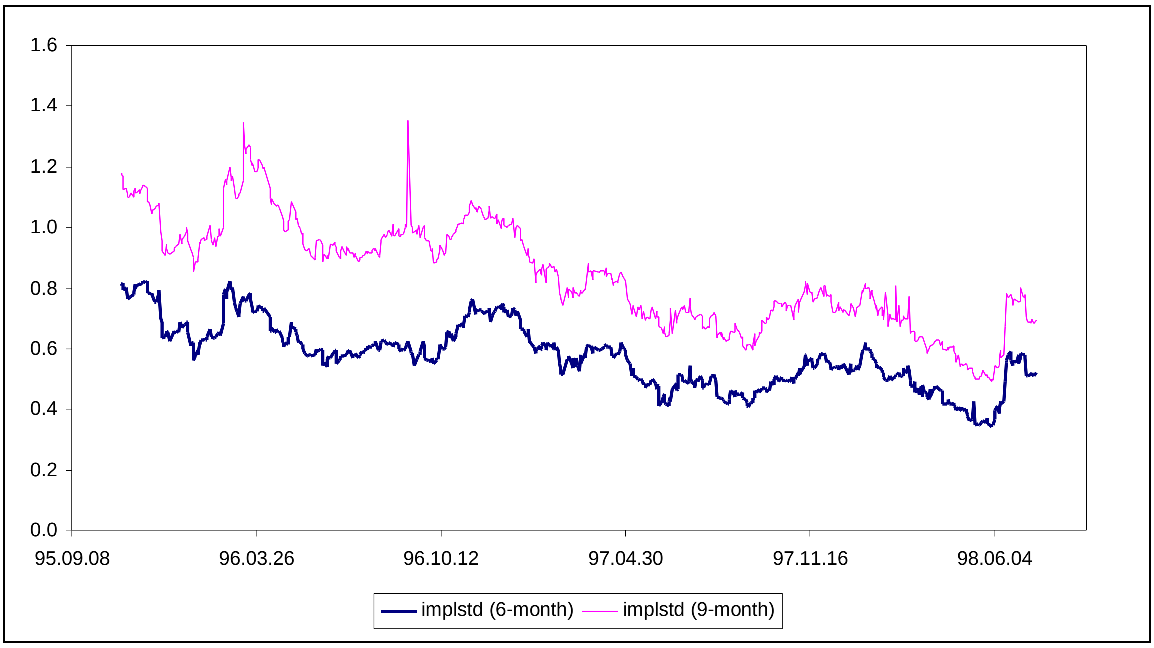
These limitations of the data however make it necessary to add some further assumptions to those underlying the basic relationship between auction bid distributions and option-based RNDs outlined in the previous section, in order to be able to carry out a meaningful test of this relationship.

The test of the relationship between auction bid distributions and option-based RNDs will involve an examination of the time series behaviour of the forward curve of uncertainty about the 3-month rate, or more precisely those two points on it that we can observe (at zero and 6-month horizons).

In order to be able to test this relationship, we will have to make an additional, and rather restrictive assumption, namely that movements in the forward curve of interest rate uncertainty are systematic, i.e. that measures of interest rate uncertainty on different horizons are cointegrated.

Some evidence supporting this assumption was found by constructing a 9-month constant-horizon implied RND standard deviation measure in a similar way as shown in Eq.3, and then checking if it is cointegrated with the 6-month measure described earlier. Chart 2 shows the two series for the whole sample period. Both series were I(1) in our sample, and a standard (Engle-Granger 2-step) test showed that they were cointegrated in the shorter sub-sample we used for testing cointegration between the auction dispersion measures and the 6-month implied RND standard deviation (see in Section 4.2).

Chart 2



If we can show that some measure of uncertainty (variance, standard deviation or the slope of the cumulated demand curve (see below)) calculated from the auction bids and that from an option-based RND are cointegrated, then, under the assumption of cointegrated uncertainty measures along the forward curve of interest rate uncertainty, we can conclude that the auction bid distribution (or its variance), similarly to that of an option-based RND contains some information about the market's sentiment of uncertainty of future interest rates.

4.2 Empirical tests

For the auction bid distribution, we have calculated two different measures of dispersion. The first one is the standard deviation of the auction yield bids (where the bid amounts were used as weights), the second is the slope coefficient of an OLS regression of the yield bids on the cumulated bid amounts. The idea behind the second measure is that the more disperse the auction bid distribution the steeper the slope of the cumulated demand curve. The difference between the two measures lies mainly in their weighting of individual bids: for example while the standard deviation is rather sensitive to an individual bid at an extreme yield and with a large amount, the slope measure tends to place a smaller weight on such a bid.

Our sample of weekly 3-month T-bill auction yield bids runs from January 5, 1996 to July 17, 1998, with one auction (January 24, 1997) missing because of incomplete bid data. The average number of bids per auction was 20.7 during this period, the average total amount bid for was 1,682 million pounds, while the allocated amounts varied between 200 million to 1,200 million pounds.

For the same period we have daily estimations of the standard deviation of the 6-month constant-horizon hypothetical implied RND, which are estimated from the prices of options on the 3-month short sterling rate.

Charts 3 and 4 show the time series of the standard deviation of the implied RND (**implstd**) together with the standard deviations (**aucstd**) and estimated slope coefficients (**beta**) of the weekly auction bid distributions, respectively, together with the 1-month moving averages of the latter two (**aucstdma** and **betama**).

Chart 3

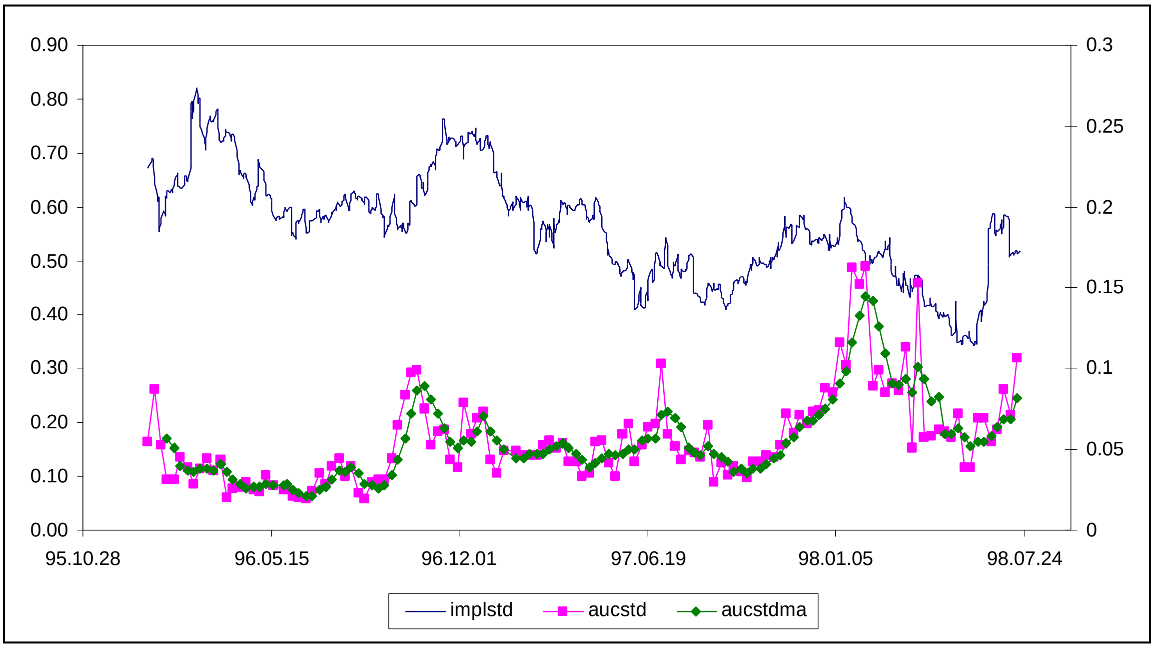


Chart 4

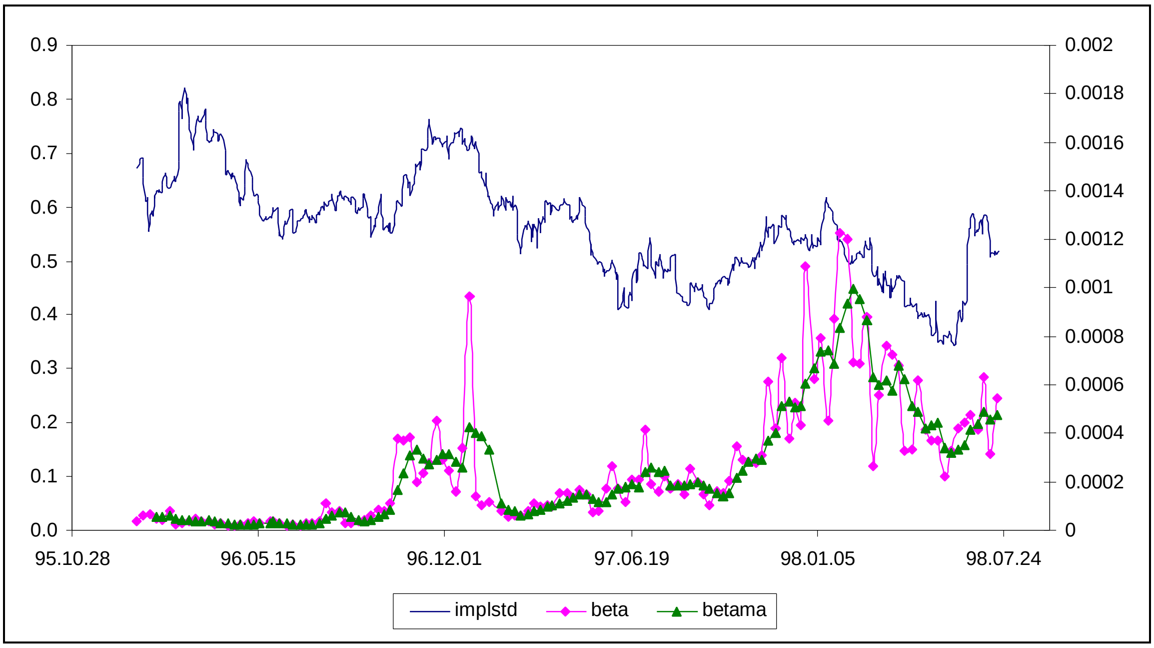
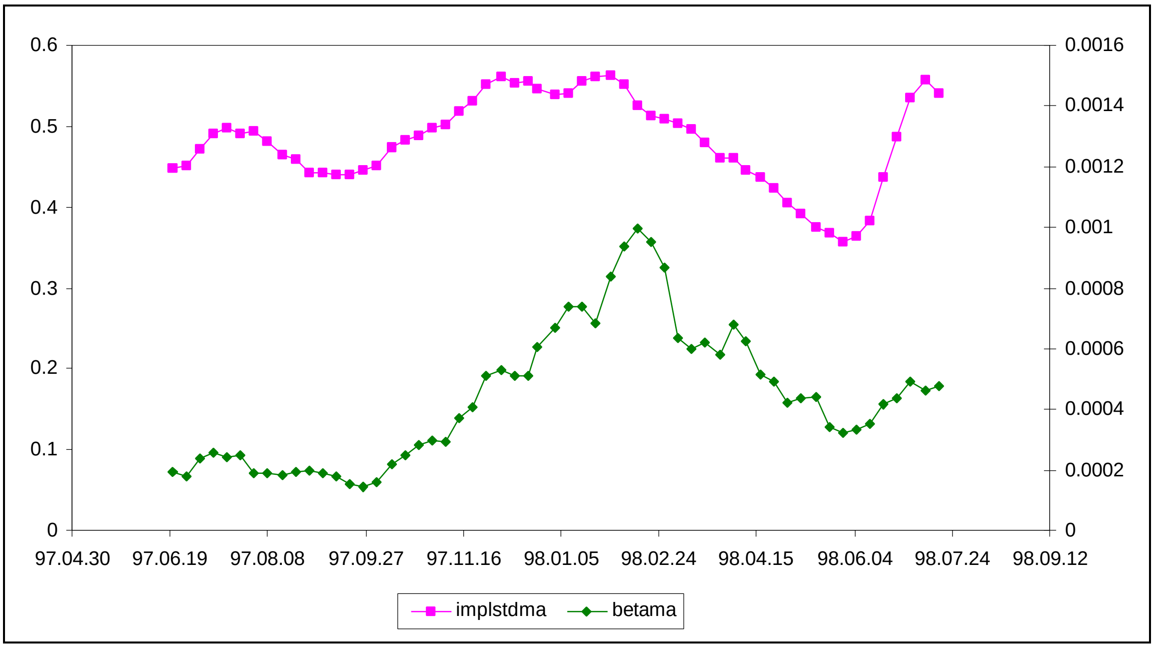


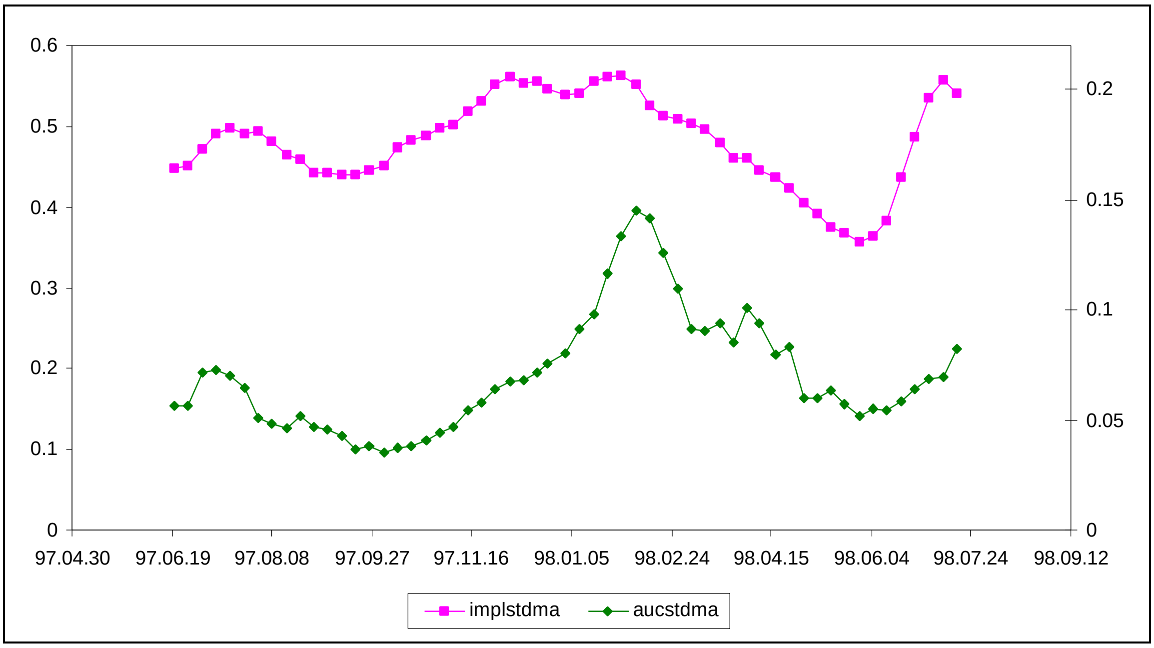
Chart 4 suggests that before October 1996, the slopes of the cumulated demand curves at the auctions were much closer to zero and varied much less than in the period afterwards. The period from October 1996 to January 1997 brought the first sign of co-movement of the auction slope with the implied RND standard deviation. After this "turbulent" period, auction slopes stabilised on a bit higher level than before, and beginning roughly from the end of May, 1997 the visual evidence of their co-movement with the implied RND standard deviations is quite remarkable[5]. The latter is better observable from Chart 5, which shows the 1-month moving averages of both the auction slopes (**betama**) and the standard deviations of the implied RNDs on the days corresponding to auction dates (**implstdma**) for this last period (first observation used to calculate monthly moving averages is May 30, 1997, last is July 17, 1998). The same pattern can be seen from Chart 6, which plots **aucstdma** and **implstdma** for this period. The starting point of the period of apparent co-movement seems to coincide with the granting of independence to the Bank of England and the subsequent change in the conduct of monetary policy.

Chart 5



[5] The fact that Chart 4 seems to tell a more detailed story than Chart 3, i.e. that it helps to distinguish the initial period with an apparent lack of co-movement stems from the property of the slope measure that it is more robust to outliers, i.e. obviously nonsensical bids than the standard deviation measure of Chart 3.

Chart 6



In the case of Chart 3 and 6, both of which plot together two standard deviation measures corresponding to different horizons, it makes sense to compare the magnitudes, i.e. to look at the difference in the left and right scales. What we find is that the implied RND standard deviations are much higher, in the shorter sample they are 3-9 times larger than the auction standard deviation. Earlier we argued that the auction standard deviation expresses the uncertainty about the fair value of the spot 3-month rate or, alternatively, the uncertainty about an *n*-month rate *m* months ahead, where $n+m=3$. The implied RND standard deviation on the other hand reflects the uncertainty about the 3-month rate (or its fair value) 6 months ahead. If the short (say, 3-month) interest rate is a highly persistent process (I(1) or close to I(1)) then it is perfectly plausible that uncertainty in a longer horizon is bigger. However if the short rate is mean-reverting, then this is not necessarily true, and the question on what horizon the mean reversion takes place becomes crucial. It seems plausible to assume, however, that mean-reversion does not take place so quickly and that the 3-month rate can be taken as an I(1) process in the 6-moth horizon.

We performed cointegration tests both on the whole sample and on the shorter sub-sample May 30, 1997 – July 17, 1998, using 132 and 60 observations, respectively. The standard deviation from the implied RNDs turned out to be I(1) in both samples, while unit root tests on the auction slope coefficient and standard deviation series yielded more ambiguous results. Although the ADF tests suggested that both these series are I(1) in both samples, the Phillips-Perron tests implied that they are I(0) in both samples. Therefore we used the technique described in *Pesaran, Shin and Smith* (1996), which was especially designed for testing for the existence of a long-run relationship when the orders of integration of the underlying variables are not known with certainty. The test statistic is the F statistic for testing the joint significance of the lagged level terms in an unrestricted error-correction specification. Pesaran, Shin and Smith present two sets of critical values for this test: one of them assuming that all the variables are I(0) and another assuming that all of them are I(1). These two sets of critical values provide critical values bounds which cover all

possible classifications of the variables into I(*d*) processes, where $0<=d<=1$. If the F statistic is greater then the upper (I(1)) bound, then the null hypothesis of no long-run relationship between the regressors can be rejected without the need to know the exact order of integration of the regressors. If the F statistic is below the I(0) bound, the null of no long-run relationship cannot be rejected, but again, we do not need to know the order of integration of the variables to draw this conclusion. However, if the F statistic falls in the region between the bounds, one cannot make conclusive inference without knowing the order of integration of the variables.
The unrestricted ECMs we estimated had the following form:

$$\Delta y_t = a + f y_{t-1} + d x_{t-1} + \sum_{i=1}^{p-1} y_i \Delta y_{t-i} + \sum_{i=0}^{p-1} j_i \Delta x_{t-i} + e_t \qquad \text{Eq.6}$$

Where:
y_t : **aucstd** or **beta**
x_t : **implstd**

Since the number of lagged differenced variables in the ECM can have a significant effect on the test results, lag specifications $p=q=2,3,4$ were estimated for each pairs of variables.
The F statistic for testing the joint null hypothesis that $f=0$ and $d=0$ are reported for both sample periods in Tables 1 and 2.

Table 1
Test results for sample: 1996.01.05-1998.07.17

	F statistic	
p,q	**aucstd**	**beta**
2,2	3.802	4.200
3,3	3.557	3.135
4,4	3.378	1.464

Table 2
Test results for sample: 1997.05.30-1998.07.17

	F statistic	
p,q	**aucstd**	**beta**
2,2	4.979	4.824
3,3	3.993	3.663
4,4	3.127	3.484

Table 3
Critical Value Bounds of the F Statistic*
Testing for the Existence of A Long-Run Relationship between 2 variables
Intercept and No Trend in the Unrestricted ECM

90%		95%		97.5%		99%	
I(0)	*I(1)*	*I(0)*	*I(1)*	*I(0)*	*I(1)*	*I(0)*	*I(1)*
4.042	4.788	4.934	5.764	5.776	6.732	7.057	7.815

*Tabulated in Pesaran, Shin and Smith (1996)

As it is clear from Table 1, we can not conclude that any of the two dispersion measures were cointegrated with the implied RND standard deviation using the whole sample. However, there is some week evidence of the existence of a long-run relationship for the shorter sample: in the case of both the auction slope and the standard deviation, the F statistic is above the 90% upper critical value. This result is sensitive to the number of lagged differenced terms included, as it only holds for $p=q=2$ in the case of both variables. It turns out however that when searching for the proper lag order based on, for example, the Schwarz information criterion, the results suggest the $p=q=2$ specification for both the auction standard deviation and the slope coefficient[6].
Based on these results, for the shorter sample we have estimated the long-run relationship between the auction dispersion measures and the implied RND standard deviation, as well as the preferred ($p=q=2$) ECM specification.
The parameter estimates and their standard errors (in brackets) are presented in Table 4 and 5.
The point estimates of the coefficients in the long-run relationship are interpretable only in the case of the equation including the auction standard deviation. The point estimate of the long-run multiplier has the correct sign and its size is plausible.

Table 4

Dependent variable	Intercept	Long-run coefficient of **implstd**
aucstd	-0.025 (0.029)	0.198 (0.061)
beta	-0.000393 (0.000249)	0.001716 (0.000512)

[6] Given that both the auction dispersion series and the implied RND standard deviation series can plausibly be assumed noisy, we have tried cointegration tests with smoothed (moving average and Hodrick-Prescott filtered) versions of these series, but the results did not improve.

Table 5

$\Delta y_t =$	Δy_{t-1}	Δx_t	Δx_{t-1}	u_{t-1}	R^2
aucstd	-0.439 (0.122)	0.065 (0.079)	0.00390 (0.083)	-0.303 (0.118)	0.422
beta	-0.26642 (0.12918)	0.000869 (0.00066)	-0.00032 (0.00069)	-0.32668 (0.11479)	0.315

u_{t-1} denotes the lagged values of residuals from the levels equation

The adjustment coefficients are around -0.3 in the case of both the auction slope and the standard deviation ECMs, suggesting that a deviation of the auction dispersion measures from their long-run values dictated by the implied RND standard deviation is corrected for approximately in 3 weeks.
We have found some, although weak, evidence that there existed a long-run relationship between the 3-month auction dispersion measures and the standard deviation of the 6-month constant-horizon implied RND in a sub-sample which started roughly at the same time as the Bank of England was granted independence and changed the conduct of its monetary policy. If one believes that the 3-month auction dispersion measures are indeed good approximations of uncertainty about short rates, then Chart 5 and 6 reveal some changes in the slope of the forward curve of uncertainty about short rates, in particular, a temporary tilt (flattening) at the end of January-beginning of February, 1998 period. All in all, the fact that we were able to find some evidence of cointegration between noisy data form such different sources, coupled with the appealing visual evidence in Charts 5 and 6, makes us conclude that there is some information in auction dispersion measures regarding the market's uncertainty about future short rates.

5. Conclusion

In this paper we tried to examine whether dispersion measures of an auction bid distribution such as the slope of the cumulated demand curve and the standard deviation represent well the market's uncertainty about the short interest rate on a short horizon. We took the standard deviation of an option-based RND as a direct measure of this uncertainty, corresponding to a particular maturity and a particular horizon. Ideally we would have wanted to see if the standard deviation of an auction bid distribution can be viewed as an approximation of that of a properly chosen implied RND, i.e. one which is derived from the prices of options on the *n*-month rate with *m*-months to expiry, where $m+n$ equals the maturity of the auctioned government security. The available U.K. data proved to be much more limited than what is needed for this purpose. The longest regularly auctioned maturity was the 3-month, while the shortest available horizon for a constant-horizon implied RND series on the 3-month rate was 6 months. Thus what we had was two observed points on the rough equivalent of the term structure (more precisely the forward curve) of interest rate uncertainty, and we wanted to see whether one of these points (the auction standard deviation/slope) indeed reflects uncertainty on a short horizon. Therefore we had to make a rather restrictive assumption, namely that uncertainty measures at different horizons on the forward curve of uncertainty are cointegrated at least up to the 6-month horizon. Some empirical support for this assumption was provided by the cointegration tests of the 6-month and 9-month constant-horizon implied RND standard deviation series.

We have found some, although weak, evidence that the 3-month auction dispersion measures were cointegrated with the standard deviation of the 6-month constant-horizon implied RND in a sub-sample which started roughly at the same time as the Bank of England was granted independence and changed the conduct of its monetary policy. Under our assumption of shift-dominated movements in the forward curve of uncertainty, this evidence suggests that the auction dispersion measures are good approximations of the uncertainty about short rates a short period ahead. One limitation of these measures is that, unlike in the case of implied RND standard deviations, it is not possible to tell exactly which future short rate's uncertainty is reflected in them and on what horizon, i.e. they correspond to any n-month rate m months ahead for which $m+n$ equals the auctioned maturity (3 months in our case). However, even this general information about "short rate uncertainty on a short horizon" can be relevant for the monetary authority.

The motivation of this paper was to show that measures of interest rate uncertainty derived from interest rate option prices can be roughly approximated by variables which are available even in the absence of highly sophisticated financial markets. In emerging and developing countries, where the interest rate option market is illiquid or does not exist at all, data on government securities auctions is usually available for the monetary authority. Of course, the fact that we have found some evidence that auction dispersion measures contain information about interest rate uncertainty in the U.K. does not mean that the same is necessarily true for other countries. Since the lack of a liquid interest rate option market means that the information content of auction dispersion measures cannot be tested in these countries, one have to resort to less exact methods such as comparing the time series of these measures with the timing of monetary policy steps and "news" that could have possibly effected interest rate uncertainty, and check whether the message in the auction data made any sense historically in this respect. By providing some evidence of information in auction dispersion measures at least for one country, this paper suggests that such efforts may be justified and may yield some useful supplementary information for monetary policy.

Supplement: An application to Hungarian data

Data issues

In the following analysis we have used two measures of the dispersion of bid distributions of 3-, 6- and 12-month T-bill auctions organised by the Government Debt Management Agency (GDMA).[7] The first one is the variance of the auction yield bids (where the bid amounts were used as weights); the second is the slope coefficient of an OLS regression of the yield bids on the cumulated bid amounts. In the following sections we refer to the latter as *beta* for simplicity.

In our sample period the GDMA organised auctions for the 3-month T-bill on every Tuesdays, while there were biweekly auctions for the 6- and 12-month T-bills (on every second Wednesdays and Thursdays, respectively). Until January 1, 1998 banks, investment banks and other financial institutions as well as primary dealers could participate T-bill auctions. After this date, only primary dealers were allowed to bid. GDMA auctions are multiple-price, i.e. each successful bidder gets the amount he/she bid for at the price he/she submitted. One well-observable feature of

[7] The data was provided by GDMA.

bidding strategies is that instead of bidding at a single price, bidders submit multiple bids, i.e. they bid along their individual demand curves.

We have a sample of dispersion measures of the biweekly 6- and 12-month T-bill auctions covering 3 years (79 auctions). Unfortunately, the sample period of the dispersion measures of the weekly 3-month T-bill auctions is shorter, covering 20 months (85 observations) only. In Table 1, we presented some summary statistics of the auctions calculated over a sample period in which we had data about all the three maturities.

Table 1. Summary statistics of auctions in the common sample period

Maturity	3-month	6-month	12-month
First auction in sample	97.07.09	97.07.10	97.07.11
Last auction in sample	99.02.16	99.02.17	99.02.18
No. of auctions in sample	85	43	43
Average number of bids	92	162	176
Avg. amount on offer (bn HUF)	5.0	16.4	17.3
Avg. total amount bid (bn HUF)	11.9	34.6	42.2
Avg. yield (%)	18.14	18.12	18.11
Avg. variance of yield bids	0.027	0.026	0.021
Avg. beta	6.88E-07	1.61E-07	1.27E-07
Avg. beta (normalised)	1.72E-07	8.03E-08	1.27E-07
Avg. beta (normalised, w/o Sept., 1998)	1.15E-07	7.24E-08	1.05E-07

The reason why we used the beta measure in addition to variance in order to represent the dispersion of the auction bids was that in the calculation of the variance, the bid amounts were used as weights. This might cause a problem, since bidders in principle might submit bids for large quantities at implausibly high yields hoping that total bid amount would be less than the amount on offer, thus increasing the variance. Therefore it is not the market uncertainty about the future course of the short interest rate which is responsible for this part of the variance but rather the ‘tactical behaviour’ of some bidders. The beta measure is less sensitive to this ‘tactical behaviour’, i.e. large-amount bids at high yields, since when fitting a line on the auction demand curve, each bid gets the same weight regardless of the amount.

In the two charts below we have plotted the demand curve (cumulated yield bids) and the corresponding line fitted by OLS as well as the histogram of yield bids observed on a 1997 auction of the 12-month T-bill.[8] Some ‘tactical’ bids are well observable as well as a few bids at very low yields. These charts illustrate the need for the beta measure, which places less weight on bids at extreme yields and large quantities, thus representing the dispersion of the subjective probability distributions (about the future level of the short rate) presumably better than the variance.

[8] Auction demand curves are classified as confidential information by the GDMA, so we could not indicate the exact date of the auction.

Chart 1. Auction demand curve and the OLS regression line fitted on it

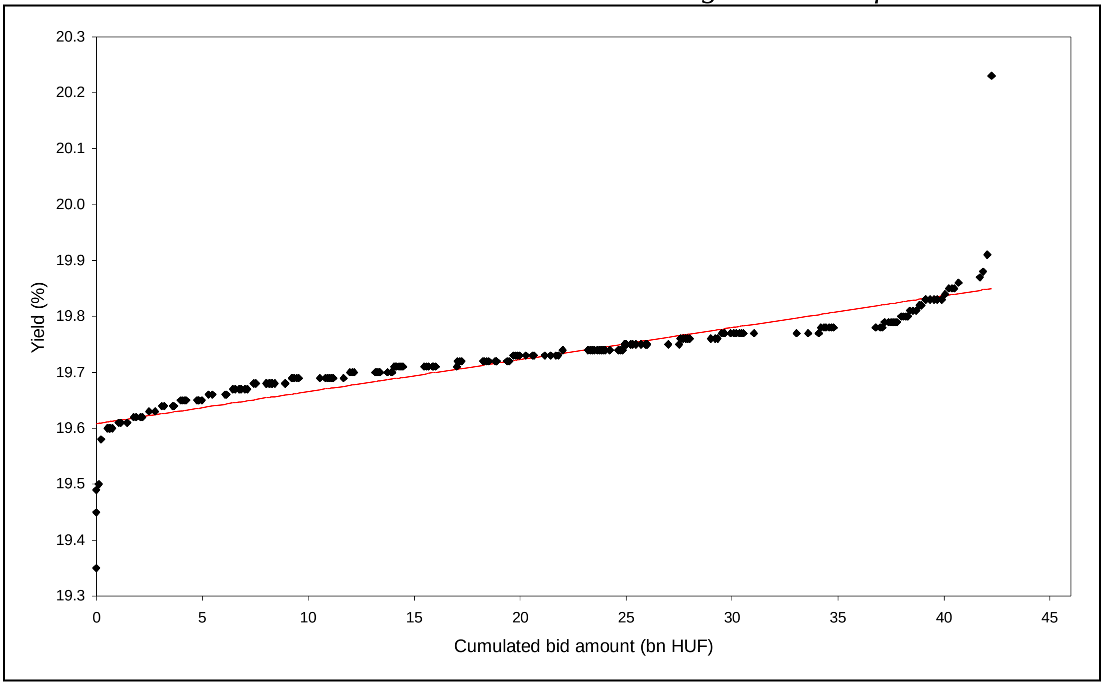
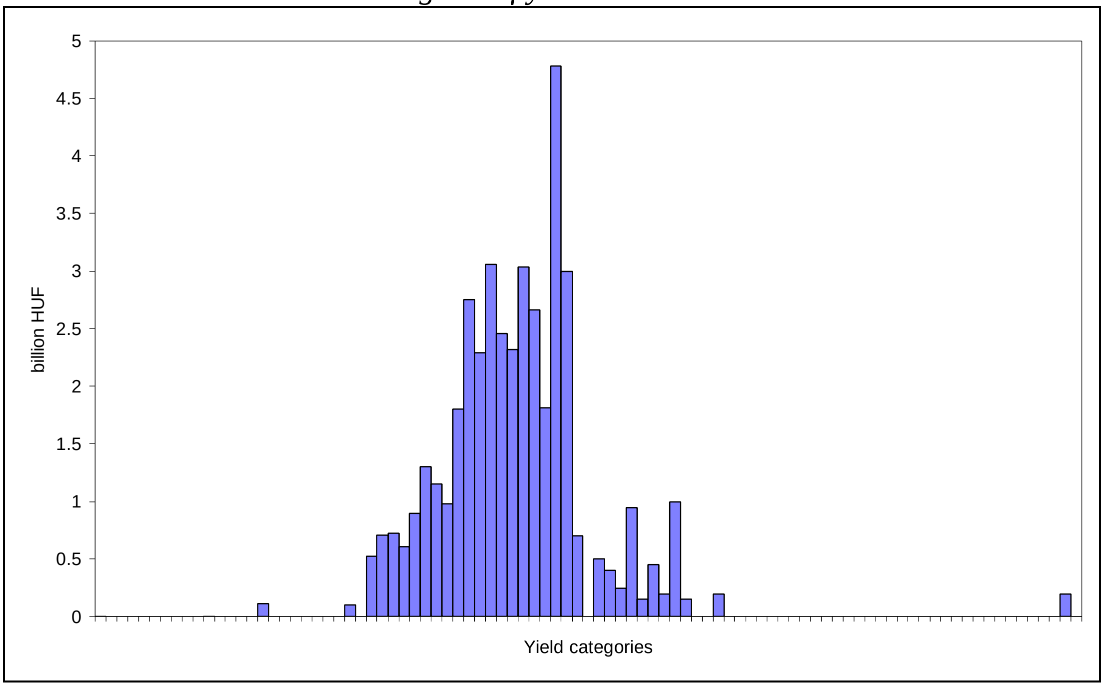


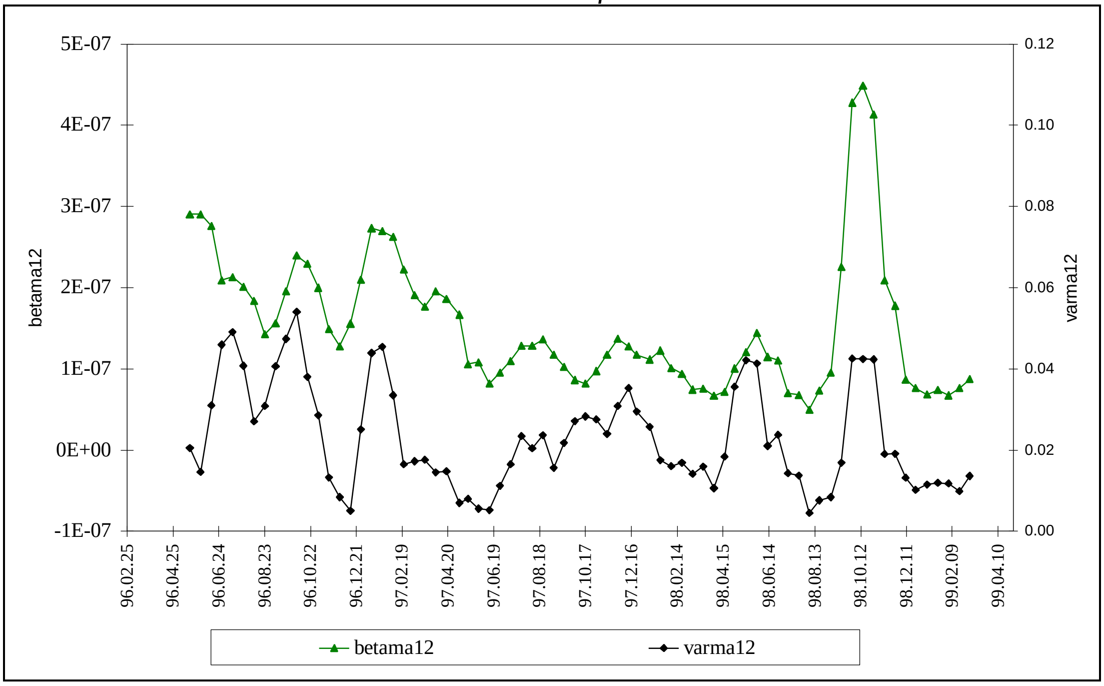
Chart 2. The histogram of yield bids on the same auction



Because of these problems, we used both dispersion measures in the time series analysis below. In Chart 3 the evolution of the beta and the variance of 12-month T-bill auctions are plotted. The two measures usually reflect the same tendencies, but there are periods (Sept-Oct 1997, for example) when the two measures move in different directions. If the beta indeed places less weight on ’tactical bids’, then different movements in the two dispersion measures indicate an increased

’tactical’ activity of bidders. However, since the GDMA provided whole demand curves for only a few selected auctions, we were not able to check this hypothesis.

Chart 3. The variance and beta of 12-month T-bill auctions



In order to be able to compare the betas of the three different maturities, we had to normalise them, since the slopes of the auction demand curves were originally estimated in the yield-quantity plane. To map these betas into the price-quantity plane we used the following formula:

$$\frac{dp}{dQ} = -m\frac{dy}{dQ} , \qquad (1)$$

where p is the natural logarithm of the T-bill price, m is maturity (in years), y is the yield-to-maturity (YTM) of the T-bill and Q is the amount demanded.[9] $\frac{dy}{dQ}$ in (1) is the beta measure we estimated in the yield-quantity plane, while $\frac{dp}{dQ}$ is its counterpart in the price-quantity plane, to which we refer to as *normalised beta* in the following sections. According to (1), we expect the original (non-normalised) beta of a 3-month T-bill auction to be roughly four times bigger, and that of a 6-month auction to be twice as big as the beta of the 12-month auction on the same date, provided that uncertainty about the short interest rate is the same on these three horizons. It can be seen from Table 1 that among the average betas, roughly these expected proportions prevailed in our sample period, especially if we leave out September 1998 from the sample. Since auctions

[9] In order to get (1) we used that for small ys the log price of a discount bond is $p = -my$. Differentiating both sides with respect to Q , we got (1).

for the 3-, 6- and 12-month T-bills are not held on the same day (even on weeks when all three maturities are auctioned, the auctions are on 3 consecutive days), the information sets of investors may change between two auctions on the same week. We assumed that these day-by-day changes in the information sets were usually not too big and treated the auctions on consecutive days as if they were simultaneous observations. Clearly, there were periods in our sample when this assumption was wrong. In September, 1998 for example, the escalation of the Russian crisis and the events on the Hungarian currency futures markets suggested that one day may had mattered a lot in terms of changes in the information sets of investors. We also calculated average betas leaving out this period from our sample, since we wanted to illustrate the differences of average betas due to differences in maturities, not to differences in information sets.

Until September, 1998 another institutional feature of GDMA T-bill auctions was that yield bids were bounded: bidders could not submit yield bids that were 100 basis points higher or 200 basis points lower than the average yield at the previous auction of the given T-bill maturity. In the third week of September 1998, partly because of the Russian crisis, yields on each maturities increased dramatically, and the upper bound on yield bids became effective on the September 22 3-month T-bill auction. Accordingly, only 8 bids were submitted instead of the usual number of 70-80. As a result, the GDMA abolished the yield bid limits on September 25.

Even if the auction bid distributions, as we assumed, represented the subjective probability distributions formed by the market about the future short rate, the yield bid limits would mean that only a trimmed version of the latter is captured in the former. Nevertheless, we assumed that this trimming did not divert the dispersion measures of the auction bid distributions from those of the subjective probability distributions, i.e. that in most of our sample period, only a negligible probability mass fell outside the institutional limits of yield bids.

Empirical analysis

In this section we try to prove our hypothesis, namely that dispersion measures of T-bill auction bid distributions roughly represent uncertainty about the future short rate, in three indirect ways.

Our first proof is based on the plausible assumption that short-term interest rate uncertainty is positively related to uncertainty about longer-term rates. If, for example, the uncertainty about the level of the 3-month rate in 1 month increases, then uncertainty about the level of the 6-month rate in 1 month also increases. In other words we assume that movements in the term structure of interest rate uncertainty (similarly to the yield curve) are dominated by parallel shifts.

Therefore if the time series analysis of the auction dispersion measures corresponding to different T-bill maturities show that there is a significant positive correlation between them, we take it as an indirect proof of our hypothesis.

The second method examines the relationship between auction dispersion measures and average bid-ask spreads observable on the OTC market of government securities. The idea here is that, although they are affected by some other factors, bid-ask spreads may also reflect uncertainty about future rates.

Our third, more qualitative method is a consistency check: by plotting the time series of dispersion measures we examine if their changes seem consistent with monetary policy steps and money market developments in the sample period that could have influenced market uncertainty about short interest rates.

Correlation of dispersion measures corresponding to different maturities

On Charts 4 and 5 we plotted the smoothed time series of 3-, 6- and 12-month T-bill auction variances and normalised betas.

Chart 4. Auction bid variances on 3-, 6- and 12-month auctions (1-month moving averages)

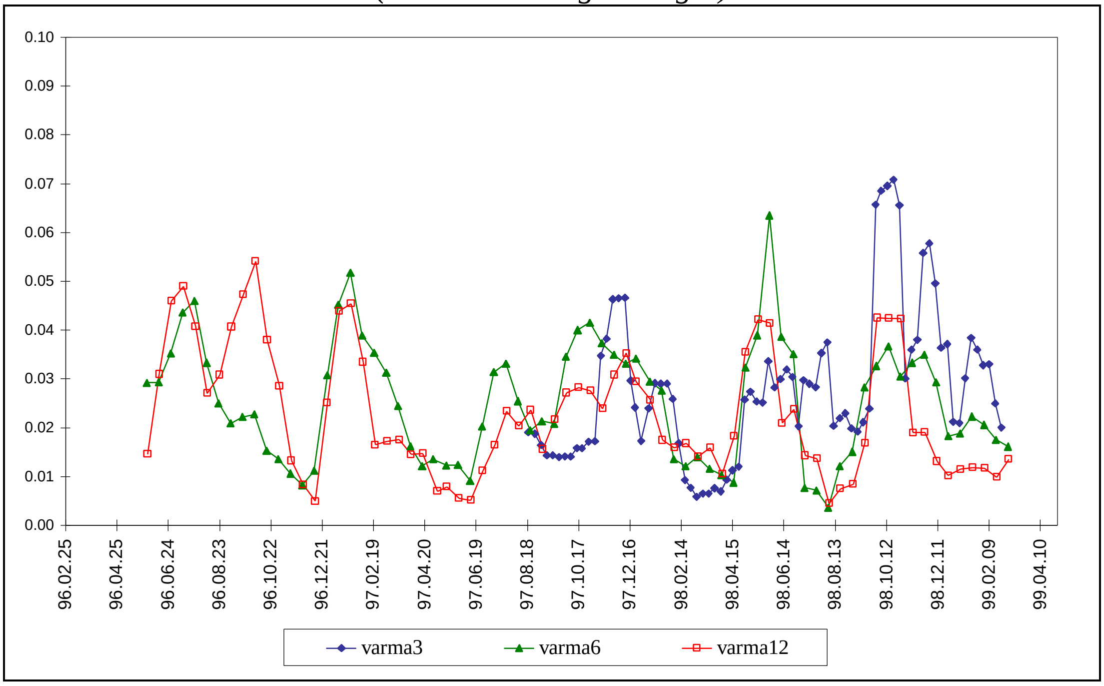
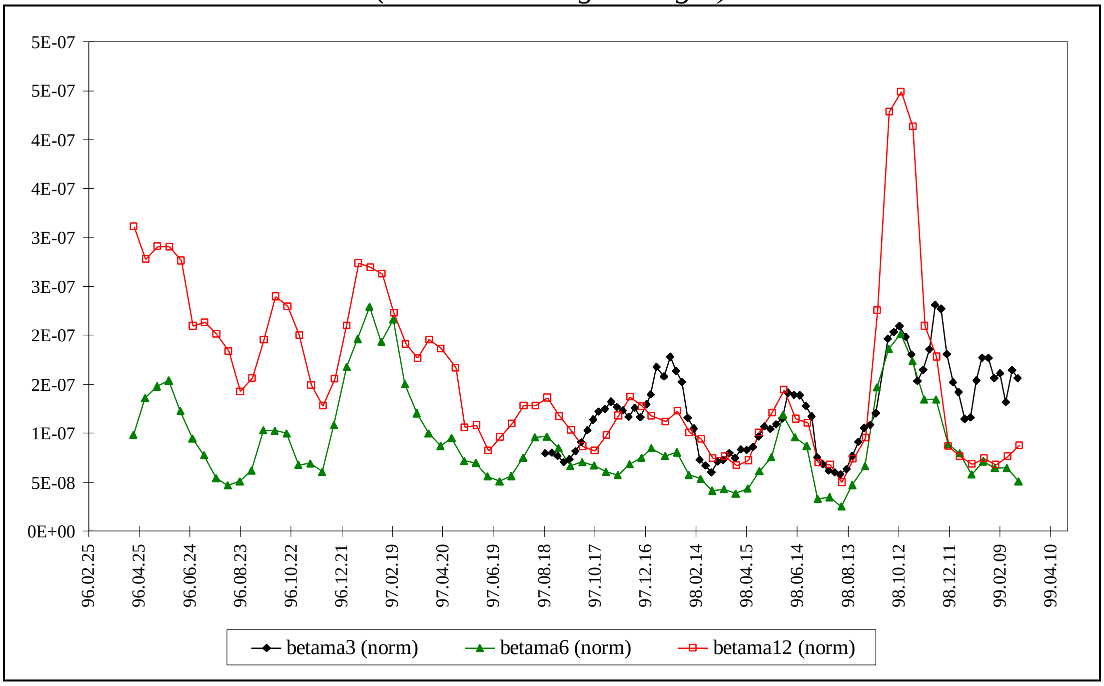


Chart 5. Normalised betas of 3-, 6- and 12-month auctions
(1-month moving averages)



Charts 4 and 5 suggest that there was some co-movement between dispersion measures of different maturities. Since unit root tests showed that both the variance and the normalised beta series were I(0), simple correlation is an appropriate way to measure how closely they tracked each other. Correlation coefficients for the common sample period are reported in Tables 2 and 3 below. When calculating these, we took auctions organised on the same week as simultaneous observations. Since 6- and 12-month auctions were held every second week, this meant that the weeks when only 3-month auctions took place had to be left out from the sample.

Table 2. Correlation of normalised betas of
3-, 6- and 12-month auctions
(sample:1997.07.09-1999.02.18)

	beta 3	beta 6	beta 12
beta 3	1		
beta 6	0.56	1	
beta 12	0.63	0.80	1

Table 3. Correlation of variances of
3-, 6- and 12-month auctions
(sample:1997.07.09-1999.02.18)

	var 3	var 6	var 12
var 3	1		
var 6	0.34	1	
var 12	0.53	0.53	1

The correlation coefficients for both dispersion measures were positive, significant and, with one exception bigger than 0.5. An interesting result was that for both dispersion measures, the correlation between the 3- and the 6-month maturities was less than that between the 3- and the 12-month maturities, despite the fact that observations of the former two are separated by only one day while there are two days between observations of the latter two. The information set of bidders may change more in two days than in one, so one would expect a smaller correlation between the 3- and the 12-month auction dispersion measures.

In the longer sample containing 6- and 12-month auctions only, the correlation coefficients were 0.74 and 0.65 for variances and normalised betas, respectively.[10]

If movements in the term structure of interest rate uncertainty are predominantly parallel shifts, then the strong positive correlation between auction dispersion measures on different maturities suggest that these measures reflect uncertainty about the future short interest rates.

Auction dispersion measures and average bid-ask spreads

Beginning with October, 1996, daily time series of OTC market best bid and ask quotes for government securities are available, from which bid-ask spreads for individual papers can be calculated. Presumably, interest rate uncertainty has an effect on bid-ask spreads, higher uncertainty leading to bigger spreads. However, it is obvious that other factors, first of all relative liquidity, may affect the bid-ask spread of an individual paper. On the basis of remaining maturity, we constructed maturity bands for government papers (3-6 months, 6-12 months, etc.) across which we calculated average bid-ask spreads. By averaging across the maturity band we probably got rid of relative liquidity and other instrument-specific factors influencing the bid-ask spreads of individual papers. Nevertheless, average bid-ask spreads may still reflect many other non-instrument-specific factors besides interest rate uncertainty.

What we want to see is whether time series of auction dispersion measures and average bid-ask spreads, being two potential sources of information about interest rate uncertainty, move together or not. In Chart 6, 1-month moving averages of average daily bid-ask spreads of government securities with 3-6 months of remaining maturities are plotted together with 1-month moving averages of variances corresponding to 6-month T-bill auctions.

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[10] The longer sample covered the period 1996.03.07-1999.03.04 and contained 79 pairs of auctions.

Chart 6. 6-month auction variances and average bid-ask spreads of government securities with 3-6 months of remaining maturity (1-month moving averages)

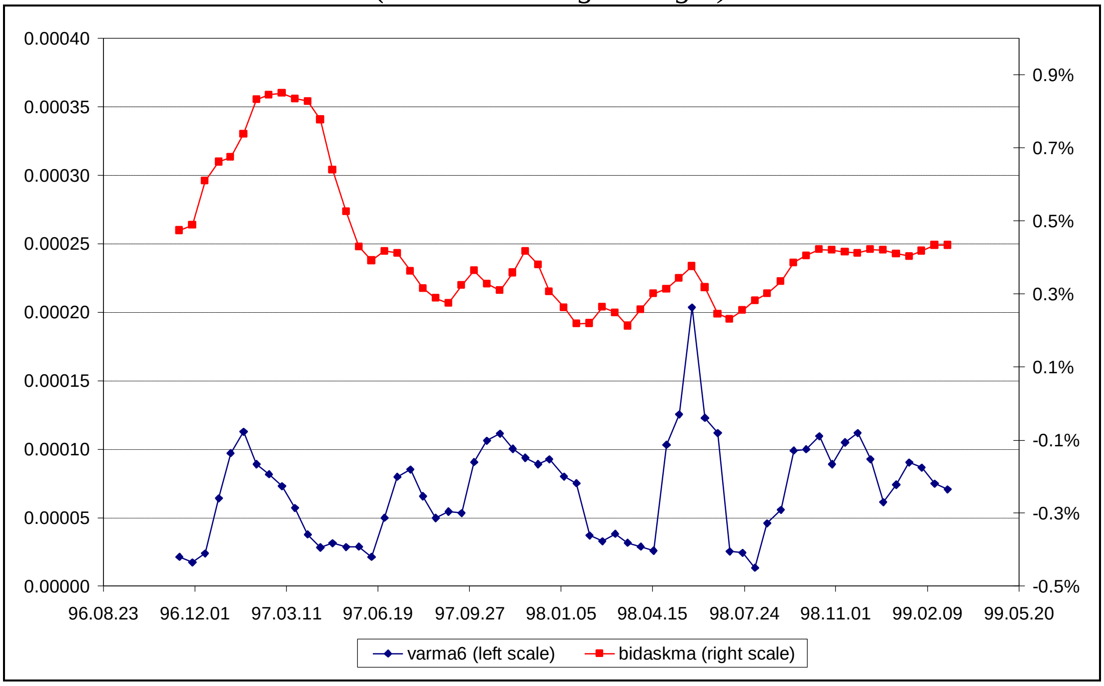


Chart 6 suggests co-movement to some extent. However, this is hard to quantify since the two series in this case have different orders of integration. While the series of average bid-ask spreads contains a unit root, that of the auction variances is I(0) according to either ADF or Phillips-Perron tests, though the short sample size seriously limits the reliability of these tests. In the presence of a unit root in one of the series, there is the danger that the calculation of correlation or regression coefficients would reflect spurious correlation to some extent.

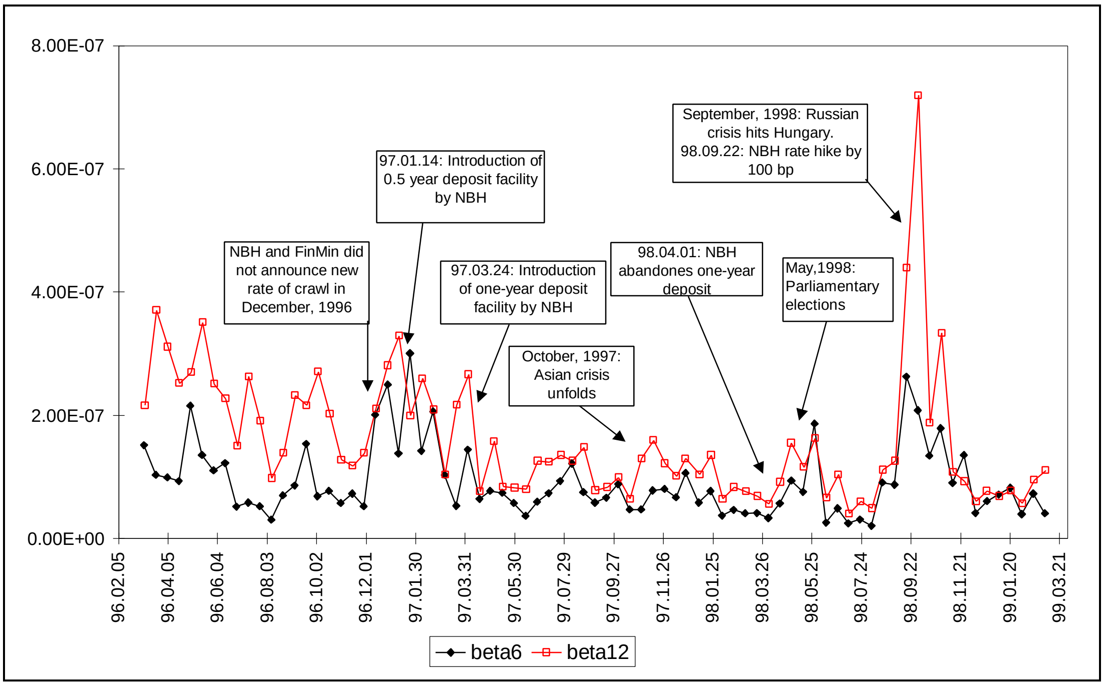
Another problem is that there was a significant institutional change in the OTC market quoting regulation in January, 1998, when the GDMA reduced the maximum bid-ask spread primary dealers can use from a 100 yield basis points to 50.

Because of these problems, we could not go further than reporting co-movement based on visual inspection.

An intuitive consistency check

In Chart 7, beside normalised betas for 6- and 12-month auctions, we showed those events (starting December, 1996) which we could plausibly assume to have had an effect on uncertainty about the future course of short interest rates. In order to show the exact timing of events and reactions by normalised betas we did not use the moving-average smoothing here.

Chart 7. 6- and 12-month normalised betas



At the beginning of December 1996 market participants expected the announcement of a decrease in the rate of crawl, effective on January 1997.[11] This announcement did not take place but the central bank decreased the interest rate on its main instrument, the 1-month reverse repo by an unusually large 50 basis points. These unexpected and somewhat contradictory steps from policymakers may have increased uncertainty about future short interest rates. The slopes of auction demand curves (the normalised betas) indeed increased significantly in this period.

In the first quarter of 1997 the National Bank of Hungary (NBH) lengthened the segment of the yield curve which it wanted to control directly in two steps. On January 17, it introduced a 6-month NBH deposit and on March 24 a 12-month NBH-bond. These were issued on tap, with a yield fixed and occasionally modified by the NBH. One would expect that the presence of central bank instruments with such long maturities would decrease market uncertainty about the future course of the short rate.[12] Normalised betas decreased by May and a relatively stable period (in terms of the evolution of betas) had started.

The normalised betas did not indicate unambiguously the increased interest rate uncertainty that the onset of the Southeast Asian crisis (October 1997) probably brought about. Only the 12-month normalised beta increased after the first serious shock occurred on the Budapest Stock Exchange on October 28.

One would expect that the abolishment of the NBH's 1-year bond as a standing deposit facility on April 1, 1998 would have increased interest rate uncertainty. The normalised betas indeed

[11] Since March 1995, a crawling band exchange rate regime has been operating in Hungary. The Forint is pegged to a basket of foreign currencies; the central parity is devalued daily according to a pre-announced 'rate of crawl'. The Forint is allowed to fluctuate within a +/- 2.25% band around the central parity. Changes in the rate of crawl are announced occasionally, roughly twice a year and always approximately 1 month in advance.

[12] The forward rate implied by the relative yields of the two NBH instruments may have operated as a message to the market about the intentions of the NBH with future interest rate policy.

increased at that time. However it may have been the general elections due to start in May that had affected interest rate uncertainty in this period.
The effect of the Russian crisis on the Hungarian debt market became full-fledged in September 1998. After the September 16 expiry on the currency futures market, an extremely strong pressure was on the Forint. On September 22, the NBH increased its 1-month reverse repo rate by 100 basis points. All these events suggest that uncertainty about future interest rates may have increased in this period. Again, this was reflected by the increase in normalised betas in the same period.
For the sake of simplicity we did not plot auction variances in Chart 7, but they were also consistent with the chronology of assumed interest rate uncertainty presented here.
Nevertheless, there were some changes in the time series of auction dispersion measures which we were not able to explain on the basis of our knowledge of monetary policy steps or money market events during the sample period.

Conclusion

In the previous sections we tried to examine the extent to which dispersion measures of Hungarian T-bill auctions reflected market uncertainty about future short rates. Since a direct measure of this uncertainty – the variance of an RND derived from interest-rate options would be such measure – is not available in Hungary; we had to revert to indirect ways to prove the relationship between auction dispersion measures and interest rate uncertainty.
The significant positive correlation between the dispersion measures of 3-, 6- and 12-month auctions suggest that there is a common underlying factor influencing dispersion on all these maturities. If movements in the term structure of interest rate uncertainty (similarly to yield curve movements) are dominated by parallel shifts, then this underlying factor may well be interest rate uncertainty. Another measure of interest rate uncertainty, which is probably affected by many other factors as well, is the average bid-ask spread on the OTC market of government securities. The auction dispersion measures to some extent seemed to co-move with this alternative measure of uncertainty as well. Last, the evolution of the auction dispersion measures seemed to be consistent with our intuition about possible changes in interest rate uncertainty, based on the knowledge of monetary policy steps and global and domestic money market events.
Based on these results we suggest the construction of a composite index, appropriately combining dispersion measures of T-bill auctions and, perhaps, average bid-ask spreads. The evolution of this index would provide policymakers a broad picture of current movements in market uncertainty about future short interest rates. This piece of information could be a useful supplement to information about the expected value of future short rates, which is available in the form of implied forwards.

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