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LINEAR AND NONLINEAR DEPENDENCE IN THE FINNISH FORWARD RATE AGREEMENT MARKETS

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ABSTRACT: This paper examines the statistical properties of the changes in the Finnish forward rate agreement (FRA) prices. The Finnish FRA is a standardized forward agreement. It has no daily settlement nor margin requirement features. We test both for linear and nonlinear dependence in the FRA log price changes. Only for two logarithmic return series the null hypothesis of no nonlinear dependence was not reject. We report evidence on ARCH-effects in our sample. The most interesting finding is, however, that we find also evidence on nonlinearities-in-mean. These results are acquired by testing for linearity against well specific alternative, namely smooth transition autoregressive (STAR) models. Furthermode, we use a selected simple trading rule in order to demonstrate the performance of such strategy during the observation period.

KEY WORDS: filter rule, forward rate agreement (FRA), nonlinear dependence, STAR models.

1. Introduction

This paper examines the statistical properties of the changes in the Finnish forward rate agreement (FRA) prices. We test for both linear and nonlinear dependence in the FRA price changes. Knowledge of the statistical properties of FRA prices can be utilized, for example, when constructing optimal hedging and trading strategies on the FRA markets.

The statistical properties of futures prices and the efficiency of future markets have received considerable interest among financial economists. Many of the studies have focused on the martingale property of the futures prices. We test for the null hypothesis that the logarithm of the prices follows a random walk. If prices follow a random walk, the returns should be uncorrelated. This has made the null hypothesis of random walk intuitively appealing. Note, however, that the random walk property is not a general implication of the theoretical asset pricing models.

Furthermore, we test for nonlinear dependence in price changes. Nonlinearities, especially ARCH-effects, in financial time series are well documented; see for example Bollerslev, Chou and Kroner (1992) and the many references therein. Murto (1992) reports GARCH-effects in the changes in one-month interest rates using Finnish money market data. However, ARCH-models (or GARCH models) are only one class of nonlinear model. Many other nonlinear models have also been developed. These include many models where nonlinearities enter through mean rather than through a variance as in ARCH-models.

Recently Hsieh (1989, 1991) and Praschnik (1991) and Scheinkman and LeBaron (1989) have tested for nonlinearities using data from the foreign exchange market, stock market and interest rate futures market. Hsieh (1989, 1991) as well as Praschnik (1991) concluded that most of the nonlinearities in the data are caused by conditional heteroscedasticity. The nonlinearity seems to enter through variances rather than through means.

We also report evidence on ARCH-effects in our sample. The most interesting finding is, however, that we obtain evidence of nonlinearities-in-mean. These results are accquired by testing linearity against well specific alternative, namely smooth transition autoregressive (STAR) models.

The remainder of the paper proceeds as follows: Section 2 describes the data used in this study. In Section 3 we test for randomness and in Section 4 for nonlinear dependence. In Section 5 we use a selected simple trading rule in order to demonstrate the performance of such strategy during the observation period. Conclusions are presented in Section 6.

2. Data and Markets

The market for FIM-denominated Forward Rate Agreements (FRA) started in November 1987. The market has grown quite substantially in recent years and the total nominal amount of all outstanding FRA agreements now typically exceeds FIM 100 billion.

The Finnish FRA is a **forward** agreement. It has no daily settlement nor margin requirement features. The use of forwards in this study make the results theoretically appealing. Namely, the theoretical literature is usually concerned with forward contracts, whereas empirical studies have mainly used data from the interest-rate future contracts.[1]

The standardized FRA is an agreement to buy or to sell a 3-month Certificate of Deposit (CD) at an agreed rate of interest on a certain date. These dates are March 15th, June 15th, September

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[1] For more on the difference between futures and forward rate contracts see Cox, Ingersoll and Ross (1981).

15th and December 15th each year.[2] If the banking day is a holiday, the following banking day is used as the expiration day. Because the FRA is based on calendar dates and not on a fixed number of days, the duration of the underlying security varies somewhat. Of the four FRA periods each year, the FRAs for only three periods are actively traded in the money market at any one time.

The nominal amount of the FRA is usually FIM 20 million. On the expiration date the forward rate is compared with the 3-month HELIBOR, which is the arithmetic average of the 3-month offer rates of the five major Finnish banks (KOP, UBF, Postipankki, Skopbank of Finland, Okobank). The difference is settled on a cash basis in the customer's account five banking days after the expiration. It should be noted that no real delivery of the underlying CD takes place on the expiration date. On the expiration date the difference between the forward rate and the 3-month HELIBOR is calculated as follows:

$$\left[\frac{\text{Nominal amount}}{1 + R_{FRA}\text{*days}/36500} \right] - \left[\frac{\text{Nominal amount}}{1 + R_{Hel}\text{*days}/36500} \right],$$

where: R_{FRA} = price of FRA expressed as discount rate,

R_{Hel} = 3-Month HELIBOR on the expiration date,

days = the number of days between the expiration date and the date three months after the expiration.

Interest rate forwards, like other Finnish money market instruments, are based on a 365-day year.

Banks are functioning in the FRA market as market makers and give on request both bid and ask prices to each other with a spread of five basis points for a contract of FIM 20 million. Currently there are seven market makers, KOP, UBF, Postipankki,

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[2] Currently each of these days have been altered to be the 3:rd Wednesday in the respective month.

Skopbank of Finland, Okobank, Midland Montagu Finland and Nordbanken. There has been only minor changes in the market structure during the observation period. Non-market makers, i.e. other banks and companies, typically indicate their interest prior to dealing with the market makers. Some market makers have agreed to give each other bid and ask prices for a contract of 100 million with a spread of ten basis points. There are no brokers in the market.

The stock of outstanding contracts typically exceeds the outstanding stock of the underlying security, FIM-denominated CDs emitted by Finnish banks. For instance, the value of outstanding FRA contracts on 26th March 1991 was FIM 100 billion, whereas the value of outstanding CDs was FIM 78 billion. It should be noted, however, that the calculation of the stock of traded CD's is not equivalent to the calculation of the stock of traded FRA's. If a bank receives a CD emitted by itself, the bank is not able to sell it to a third party. Therefore, part of the traded CD's are not included in the outstanding stock. On the other hand, the stock of FRA-agreements is not reduced by such factors.

In this paper also the performance of one trading rule is examined. Some features of the Finnish domestic FRA market make the examination of technical trading rules especially appealing.

Firstly, the transaction costs in the FRA market are relatively small. The actual costs included in the bid/ask spread of the market maker are 5 basis points (0,05 percentage points). This is the transaction cost for market makers. Other parties usually have a transaction cost of 8 - 9 basis points or less.

Secondly, entering the market does not affect the liquidity position of the participant. Using a FRA does not require the opening of any margin or initial accounts. Instead, the credit risk between market participants is generally managed by using traditional credit limits. The absence of any marking to market feature, however, sets more requirements for managing the credit risk. Therefore, using limits as a means of managing credit risk can lead to exclusion of many potential market participants.

Thirdly, although the number of participants in the market is limited, the size of the FRA contract and the relatively large trading volume of market makers enable quite sizeable operations without prices being affected substantially.

The raw data used in this study consists of daily FRA prices of Kansallis-Osake-Pankki quated as nominal discount rates. The observations, which are extracted from the REUTER's page KORL at 1:00 pm every banking day, were provided by KOP Economic Research Department. The use of one quotation instead of an aggregate of quotations removes some of the problems encountered if the unit of analysis is an index of quotations. The plots of raw data and 3-month HELIBOR-rates are presented in the Appendix I.

The standarized market for FRAs was established on November 16th 1987. The first unbroken FRA period starts December 15th 1987. The periods included in this study are all the unbroken FRA periods between 15.12.1987 and 15.03.1991 totalling 2040 daily observations. The market for FRAs was closed between January 17th 1990 and March 7th 1990 because of a banking strike and a moratorium on payments between banking groups. This period is excluded from the study.

It should be noted that observations in consecutive FRA periods can not be assumed to be uncorrelated nor independent, because three periods of FRAs are always simultaneously traded. This dependence is obvious because shocks are likely to affect all these three FRA periods. The correlation coefficient between the daily changes of nominal rates of return of the nearest FRA period and the second nearest FRA period was 0.6836 during the observation period. The correlation coefficient between daily changes of returns of the second nearest and the most distant FRA period was 0.8818 and 0.6368 between the nearest and the most distant FRA period.

We base our inference on the logarithmic returns: $\ln(P_t)-\ln(P_{t-1})$, where P_t is the daily price at time t calculated from the nominal discount rates. In table 1 we report a range of descriptive statistics for the logarithmic returns series. For example, S881 denotes for the series starting from 16th March in 1988 and ending at 15th December in 1988. S882 states for the series starting from 16th June in same year and so for.

The largest daily return is 1.1 % per day, but the average returns are negative in most series. In three cases the mean is significantly different from zero according to the t-statistics. The standard deviation of the mean is increasing toward the end of the period. The coefficients of skewness and excess kurtosis present evidence on the non-normality of unconditional distribution of log price changes. Expecially the coefficients of excess kurtosis are clearly different from zero for all series.

Table 1. Summary statistics for daily price change

Contracts	obs[1] / skewness	mean / kurtosis[3]	SD / minimum	t-stat[2] / maximum
S874	196	$-0.91*10^{-5}$	0.00014	-0.93
	-0.45	0.89	-0.00039	0.00034
S881	197	$-0.22*10^{-4}$	0.00014	-2.10
	-0.79	3.58	-0.00070	0.00046
S882	195	$-0.77*10^{-5}$	0.00018	-0.61
	0.33	1.62	-0.00056	0.00059
S883	194	$-0.23*10^{-4}$	0.00022	-1.47
	-0.60	3.06	-0.00091	0.00069
S884	195	$-0.17*10^{-4}$	0.00020	-1.15
	-0.38	2.38	-0.00092	0.00068
S891	197	$-0.60*10^{-4}$	0.00023	-3.57
	-0.83	1.76	-0.00080	0.00051
S892	159	$-0.56*10^{-4}$	0.00021	-3.29
	-0.28	1.10	-0.00071	0.00061
S893	158	$-0.21*10^{-4}$	0.00026	1.02
	0.42	1.02	-0.00062	0.00085
S894	161	$-0.23*10^{-4}$	0.00024	1.20
	-0.29	4.44	-0.00012	0.00082
S901	196	$-0.74*10^{-5}$	0.00025	0.41
	0.01	1.16	-0.00065	0.00080
S902	194	$-0.10*10^{-4}$	0.00027	-0.53
	0.44	1.96	-0.00089	0.0011

[1]: number of observations
[2]: t-statistics for the null hypothesis that the mean is zero.
[3]: The coefficient of excess kurtosis, which is zero under the null hypothesis of normal distribution.

3. Testing for randomness

In this section we report tests for the departuness for randomness.

Table 2a reports the Box-Pierce statistic, Q, for the price changes. The null hypothesis is that the first K autocorrelation coefficients are zero. It is well known that the Box-Pierce statistic rejects the null hypothesis too often under the heteroscedasticity. For this reason, we present also adjusted Box-Pierce Q* statistics, which are adjusted to allow heteroscedasticity as discussed by Diebold (1988). The adjusted

Box-Pierce statistic is $\sum_{\tau=1}^{K} [\rho(\tau)/S(\tau)]^2$, where $\rho(\tau)$ is the τth autocorrelation coefficient and $S(\tau) = \{(1/T)(1 + \gamma(\tau)/\sigma^4)\}^{1/2}$, where T is the number of observations, $\gamma(\tau)$ is the τth sample autocovariance based of the squared data, and σ is the sample standard deviation of the data. The Q^* and Q statistics are asymptotically chi-square distributions with K degrees of freedom. The adjusted Box-Pierce statistics are reported in table 2b.

According to the results presented in table 2a the null hypothesis of serial independence is rejected at the 1 percent significance level for S882, S891 and S892 and at the 5 percent level for S893 and S894 with 5, 10 or 20 lags. The null hypothesis is not rejected for six series at the conventional significance levels. These results indicate that there is serial correlation present in the data.

By using the adjusted Box-Pierce statistics we reject the null hypothesis at the 5 percent level for the series S882, S892 and S894, but do not reject it for the series S891 and S893 as in table 2a. In that respect results change compared to those in table 2a. These results indicate that some of the rejections of serial independence, as presented in table 2a, can be caused by heteroscedasticity in the data. Indeed, when testing for no-ARCH we reject the null hypothesis of no-ARCH for the series S891. This result was robust as to whether or not we include a linear term in the mean equation.[3] We conclude that according to the Box-Pierce statistics we can reject the null hypothesis of serial independence for series S882, S892, S893 and S894.

In tables 3a and 3b we report results for testing the null hypothesis that log prices follow a random walk. We use the variance ratio test of Lo and MacKinlay (1988, 1989).

The variance-ratio test is motivated by the notion that if the

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[3] In section 3 we present results from ARCH-tests.

returns are serially random with constant variance, then the variance of the return over k periods is simply $k\sigma$, where σ is the variance of the one-period return. Based on this notion Lo and MacKinley (1988) shows that variance ratio, $VR(k) = [Var(r^{k}_{t})/ Var(r^{1}_{t}) \times 1/k]$, has the following limiting distribution under the null hypothesis of random walk:

$$(1.1) \qquad T^{1/2}\ (VR(k) - 1)\ \{2(2k-1)(k-1)/3k\}^{-1/2} \sim N(0,\ 1).$$

The null hypothesis is that prices follow a random walk with a possible drift and errors are independently and identically distributed Gaussian increments.

Furthermore Lo and MacKinley (1988) derive a version of the variance ratio test which is robust with respect to heteroscedasticity. In this case the null hypothesis is the same as above with the exception that errors can be heteroscedastic. We present results using both versions of the variance ratio test.

Tables 3a and 3b show the estimates of the variance ratio, VR(k), and variance ratio test statistics. We report test statistics using 5 and 10 days return intervals. In table 3a the null hypothesis is rejected at the 1 percent significance level for S882, S901 and S902 and at the 5 percent level for S893 and S894 using either 5 or 10 days return intervals. The null hypothesis is not rejected for six series at conventional significance levels. As with Box-Pierce statistics the results change when we adjust for possible heteroscedasticity. In that case we reject the null hypothesis at the 1 or 5 percent significance levels only for series S882 and S902. However, the p-values for series S893, S894 and S901 are 0.09, 0.05 and 0.06 respectively. Furthermore in these series we find little evidence on ARCH-effects, see for example tables 4a and 4b below. We conclude that according to variance ratio test statistics we can reject the null hypothesis for series S882, S893, S894, S901 and S902.

Based on the above-presented results, we conclude that the null hypothesis of the random walk of prices is not rejected for series: S874, S881, S883, S884 and S891. This is based on the results using either adjusted Box-Pierce statistics or variance ratio test statistics. In other series we reject the null hypothesis. In the next section we investigate possible nonlinear dependence.

Table 2a. Tests for randomness. Box-Pierce statistics

contract	Q(5)	Q(10)	Q(20)
S874	5.35	11.83	19.09
S881	2.12	11.32	19.04
S882	15.45**	19.96*	39.35**
S883	4.71	9.84	14.65
S884	6.19	7.37	11.41
S891	12.34*	23.25**	33.19*
S892	17.09**	22.01*	53.90**
S893	13.99*	16.94	29.49
S894	10.96	18.79*	33.85*
S901	4.55	7.80	15.72
S902	10.64	13.97	23.36

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of significance respectively.

Table 2b. Tests for randomness. Adjusted Box-Pierce statistics

contract	Q*(5)	Q*(10)	Q*(20)
S874	3.86	8.70	15.70
S881	1.51	9.73	18.36
S882	11.40*	14.54	31.08
S883	3.96	9.00	14.52
S884	4.05	4.89	7.11
S891	6.53	12.74	24.18
S892	10.42	13.82	42.60**
S893	10.06	13.14	24.75
S894	11.29*	17.12	30.89
S901	3.61	7.09	16.20
S902	7.97	10.82	21.59

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of significance respectively.

Table 3a. Tests for random walk in logarithmic prices. The variance ratio is reported in first row. Numbers in paranthesis are variance ratio test statistics.

contract	V(5)	V(10)
S874	1.35	1.42
	(2.17)	(1.70)
S881	1.03	1.06
	(0.20)	(0.25)
S882	1.55	1.74
	(3.44)**	(2.98)**
S883	1.29	1.43
	(1.78)	(1.72)
S884	1.02	1.07
	(0.13)	(0.29)
S891	1.09	1.33
	(0.56)	(1.32)
S892	0.80	0.87
	(-1.07)	(-0.48)
S893	1.40	1.54
	(2.18)*	(1.89)
S894	1.41	1.58
	(2.26)*	(2.07)*
S901	1.33	1.33
	(2.09)**	(1.34)
S902	1.47	1.51
	(2.90)**	(2.06)*

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of significance respectively.

Table 3b. Tests for random walk in logarithmic prices allowing heteroscedasticity. The variance ratio is reported in first row. Numbers in paranthesis are variance ratio test statistics.

contract	V*(5)	V*(10)
S874	1.35	1.42
	(1.84)	(1.46)
S881	1.03	1.06
	(0.17)	(0.23)
S882	1.55	1.74
	(3.24)**	(2.78)**
S883	1.29	1.43
	(1.31)	(1.43)
S884	1.02	1.07
	(0.11)	(0.25)
S891	1.09	1.33
	(0.43)	(1.04)
S892	0.80	0.87
	(-0.92)	(-0.41)
S893	1.40	1.54
	(1.69)	(1.54)
S894	1.41	1.58
	(1.95)	(1.87)
S901	1.33	1.33
	(1.90)	(1.28)
S902	1.47	1.51
	(2.59)**	(1.87)

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of significance respectively.

4. Testing for Nonlinear Dependence

In this section we test for nonlinear dependence. First, we report results from testing for the no-ARCH. Second, we test linearity against the smooth transition autoregressive (STAR) model. The null hypothesis is that the prices follow a random walk or that returns follow a linear autoregressive process.

By far the most popular nonlinear model in empirical econometric work has been ARCH, which have been used to describe heteroscedasticity in many financial time series.

Tables 4a and 4b present Box-Pierce test statistics based on the squared log price changes or the squared residuals. The null hypothesis depends whether or not we rejected the random walk hypothesis. We use squared price changes in the series in which we did not reject the null hypothesis of the random walk in the Section 3. These results are presented in table 4a. In the case of series in which the random walk hypothesis was rejected, we use linear autoregressive model, where the lags are selected using AIC model selection criteria. These results are presented in table 4b.

The squared data exhibit substantially more autocorrelation than the raw data, which is indicative of conditional heteroscedasticity. Only series that do not seem to have conditional heteroscedasticity are S881 and the four latest series: S893, S894, S901 and S902. We also carried out Engle's (1982) ARCH-test. These results were essentially the same as for the Box-Pierce statistic and are not presented here.

Table 4a. Tests for no-ARCH based on the random walk model for prices.

contract	Q(5)	Q(10)	Q(20)
S874	19.08**	30.87**	45.96**
S881	4.76	5.71	10.08
S883	24.28**	25.16**	27.05
S884	12.05*	19.01*	21.58
S891	41.52**	68.18**	71.76**

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of significance respectively.

Table 4b. Tests for AR-model with homoscedastic errors against AR-model with ARCH-errors.

contract	Q(5)	Q(10)	Q(20)
S882	11.70*	21.19*	33.27*
S892	16.40**	41.72**	52.83**
S893	2.91	8.95	18.25
S894	3.40	13.14	29.65
S901	5.74	7.45	15.56
S902	5.28	10.54	18.62

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of significance respectively.

In addition to ARCH, we consider also other nonlinear models. Especially we test linearity against smooth transition autoregressive models (STAR).

Consider the following STAR model of the order p

$$(2.1) \qquad r_t = \pi_{10} + \pi_1' w_t + (\pi_{20} + \pi_2' w_t)F(r_{t-d}) + u_t$$

where $u_t \sim nid(0,\sigma^2), \pi_j = (\pi_{j1}, \ldots, \pi_{jp})', j=1,2, w_t=(r_{t-1}, \ldots, r_{t-p})'$. The term $F(r_{t-d})$ denotes a transition function, where d is the delay parameter and r_t is the logarithmic daily return at time t. Teräsvirta (1990) and Teräsvirta and Anderson (1991) present two possible transition functions. The first one is the following:

$$(2.2) \qquad F(r_{t-d}) = [1 - \exp(-\gamma(r_{t-d}-c)^2)], \quad \gamma > 0$$

The second is the logistic function:

(2.3) $F(r_{t-d}) = [1 + \exp(-\gamma(r_{t-d}-c))]^{-1}, \quad \gamma > 0$

Model (2.1) with (2.2) is called an exponential STAR (ESTAR) model. An ESTAR model implies that high positive or negative returns have similar local dynamic structures, whereas the returns in the mid-region have different local dynamics. An ESTAR model can present return dynamics in which returns go back to towards "normal" returns in the same manner both from high negative or positive returns. The ESTAR model is a slight generalization of the exponential autoregressive model of Haggan and Ozaki (1981).

Model (2.1) with (2.3) is called an logistic STAR (LSTAR) model. In this case the high positive and negative returns can have different local dynamics, and a transition from high positive (negative) returns to the high negative (positive) returns may be smooth. The LSTAR model is a generalization of a two-regime threshold autoregressive model.

In testing and specifiying the STAR model we use the following steps suggested by Teräsvirta (1990). First we specify a linear model as a base for testing linearity. Second we test for linearity against the STAR model using the linear model under the null hypothesis. In this step we also determine the value of the delay parameter d. Finally we choose between LSTAR and ESTAR models keeping d fixed and using a sequence of test of the nested hypothesis.

When determining the maximum lag p in the linear model, we used the AIC model selection criterion. However, in every series we select p to be at least two, even when model selection criteria implies otherwise. If we choose too low value of p, the estimated AR model may have autocorrelated residuals. In this case the linearity test may be biased towards rejection if the true model is linear. That is because, as Teräsvirta (1990) shows, the test also has power against serially correlated errors.

Teräsvirta (1990) shows that a Lagrance multiplier (LM) type test of linearity against STAR (both LSTAR and ESTAR) assuming d is known is equivalent to the test of

$$H_o: \beta_{2j} = \beta_{3j} = \beta_{4j} = 0, \; j=1...p$$

against H_1: H_0 is not valid in the artificial regression:

$$(2.4) \qquad r_t = \beta_0 + \beta_1 ' w_t + \sum_{j=1}^{p} \beta_{2j} r_{t-j} r_{t-d} + \sum_{j=1}^{p} \beta_{3j} r_{t-j} r^2_{t-d} + \sum_{j=1}^{p} \beta_{4j} r_{t-j} r^3_{t-d} + v_t,$$

where v_t is an error term in the regression. We choose the value of d by carrying tests for different values of d. If we reject linearity for more than one value of d, we choose the value of d where the p-value of the selected test is the lowest.

The third stage is to choose between LSTAR and ESTAR. This can be done by a sequence of test within (2.4). Teräsvirta (1990) suggests that the following sequence of hypothesis can used to distinguish between LSTAR and ESTAR models:

$$H_{04}: \beta_{4j} = 0, \; j=1...p$$
$$H_{03}: \beta_{3j} = 0 \mid \beta_{4j} = 0, \; j=1...p$$
$$H_{02}: \beta_{2j} = 0 \mid \beta_{3j} = \beta_{4j} = 0, \; j=1...p.$$

If we reject H_{04}, we choose the LSTAR model. We also choose the LSTAR model, if we reject H_{02}, but accept H_{04} and H_{03}. The ESTAR model is chosen if we accept H_{04}, reject H_{03} and accept H_{02}. The definite choice can not be made if both H_{03} and H_{02} are rejected.

Table 5 reports results from the test for linearity against STAR model. We consider only the series in which the null hypothesis that prices follow a random walk was rejected. Thus we test for linearity against STAR-model for series S882, S892, S893, S894, S901 and S902.

According to our results, the linearity is rejected for three series: S893, S901 and S902. The interesting finding is that the series in which we reject the linearity in favour of STAR-model, we did not find any evidence on the ARCH.

The next step is to choose between the LSTAR- and ESTAR-models. Table 6 reports the results from testing that $\beta_{4j} = 0$, $j=1...p$ in the equation (2.4). We reject the null hypothesis at the 5 percent significance level for all series. Thus we choose the LSTAR model in all three cases.

Table 5. Tests for linearity against STAR-model.

contract	d=1	d=2	d=3	d=4	d=5
S882	1.56	1.57	1.49	1.29	1.77
S892	0.06	0.62	-	-	-
S893	2.30*	1.50	0.98	-	-
S894	0.70	1.00	0.45	-	-
S901	1.59	3.75**	-	-	-
S902	2.11	2.71*	-	-	-

The linear model under the null hypothesis is:

S882: $r_t = \alpha_0 + \alpha_1 r_{t-1} + \alpha_2 r_{t-4} + \alpha_3 r_{t-5} + u_t$

S892: $r_t = \alpha_0 + \alpha_1 r_{t-1} + \alpha_2 r_{t-2} + u_t$,

S893: $r_t = \alpha_0 + \alpha_1 r_{t-1} + \alpha_2 r_{t-2} + \alpha_3 r_{t-3} + u_t$

S894: $r_t = \alpha_0 + \alpha_1 r_{t-1} + \alpha_2 r_{t-2} + \alpha_3 r_{t-3} + u_t$

S901: $r_t = \alpha_0 + \alpha_1 r_{t-1} + \alpha_2 r_{t-2} + u_t$

S902: $r_t = \alpha_0 + \alpha_1 r_{t-1} + \alpha_2 r_{t-2} + u_t$,

Table 6. Choosing between LSTAR and ESTAR model. The null hypothesis is that $\beta_{4j} = 0$, $j=1...p$.

contract	d=1	d=2
S893	3.30*	-
S901	-	3.71*
S902	-	3.40*

Table 7 summarizes our results thus far. Only for one series, S881, the null hypothesis that logarithmic returns are white noise was not rejected. In all other series we find evidence on either linear or nonlinear dependence. Especially nonlinear dependence in the sample was evident. Only with two logarithmic return series the null hypothesis of no nonlinear dependence was not rejected. These series were S881 and S894, where the former was linearly dependent.

Tests for no-ARCH and for linearity against STAR-model were non-nested. This did not cause, however, any problems in inference because as noted before in the series in which we find evidence on STAR-model we did not find any evidence on ARCH-effects and vice versa.

Table 7. Summary of the Test Results.

contract	Test for random walk[1]	Test for r.w. against ARCH	Test for AR-model against AR-model with ARCH-residuals	Test for AR-model against STAR-model	Interpretation
S874[2]	-	+	na	na	ARCH
S881	-	-	na	na	white noise
S882	+	na	+	-	AR-model with ARCH-residuals
S883	-	+	na	na	ARCH
S884	-	+	na	na	ARCH
S891	-	+	na	na	ARCH
S892	+	na	+	-	AR-model with ARCH-residuals
S893	+	na	-	+	(L)STAR-model
S894	+	na	-	-	AR-model
S901	+	na	-	+	(L)STAR-model
S902	+	na	-	+	(L)STAR-model

1) Results are based on the heteroscedasticity adjusted Box-Pierce statistics and variance ratio tests.
2) +: rejection of the null hypothesis
-: acception of the null hypothesis
na: test not conducted

5. An example of the performance of a trading rule

The fact that there may exist some dependence or predictability in daily changes of logarithmic FRA prices does not mean that the information in historical prices could be used in earning abnormal returns. It is often argued that minor autocorrelations do not necessarily contradict the efficient market hypothesis because the trading costs could easily exceed any trading profits. Furthermore the random walk property is not a general implication of the theoretical asset pricing models as was noted in the introduction.

In this paper only one trading rule was selected to show the performance of a such method in the Finnish FRA market. It should be noted that we do not attempt to show with systematic method that the Finnish FRA market during the observation period was not efficient. The purpose is to show the magnitude of returns that can be achieved using a simple trading rule.

Numerous technical trading rules have been devised to uncover dependencies. Maybe the most commonly used technique is the filter rule introduced by Alexander (1961). In this paper a simple filter rule based on moving averages is tested. Filter rules and other technical trading rules have been tested by many academics. Some of the most recent studies are Cornell and Dietrich (1980), Bird (1985), Taylor (1985) and Drinka et al. (1991).

Cornell and Dietrich (1980) used moving averages to examine whether deviations from a moving average of some length could be used to systematically predict future price behaviour in the foreign exchange markets. Taylor (1985) used daily futures prices for eight commodities and currencies in various markets. Daily observations for at least eight years were used. He concluded that it can not be claimed that it is impossible to find actively traded contracts permitting profitable speculation. He concludes, however, that the observation period of approximately

10 years is not necessarily sufficiently long to justify the findings.

Bird (1985) used filter rules to test for dependency on cash and future contracts for copper, lead, tin and zinc on the London Metal Exchange. He used daily cash and futures prices for the period 1972 - 1982 and tested 25 different trading rules. He found positive dependency in price changes for some of the commodities periods. Inefficiency was indicated for over two thirds for cash and futures copper. For lead and zinc at least twelve of the rules indicated inefficiency of the market, although for tin Bird did not found any sign of inefficiency. Drinka et al. (1991) used a total of 34 technical trading rules to test their performance in the IMM Japanese Yen futures market. They used daily prices and their observation period was from March 1978 to September 1987. They found some strategies, which produced positive returns after adjusting for risk and concluded that the hypothesis of weak form efficiency could not be accepted.

The following steps were taken to test performance of the trading rule. First a moving average of FRA rates of the last 6 days was calculated. When the FRA rate moved 10 basis points above the moving average the FRA was sold short. Equivalently, when the FRA moved 10 basis points below the moving average the FRA was bought. The contract size of the first transaction was FIM 10 million. The position was held until the next counteractive signal. After the following signal the open position was closed and a new opposite position was opened. At expiration day the open position of FIM 10 million was closed. The net profit or loss for each period after adjusting for a transaction cost of 5 basis points at expiration date are shown in table 8.

Table 8. Net Profit/Loss of Each Period with 0.05 % Bid/Ask Spread, Finnish Marks

Period	Net Profit / Loss
P874	-8945.63
P881	+28685.27
P882	+44918.05
P883	+44942.51
P884	-3220.72
P891	+57921.06
P892	+75077.94
P893	+55979.47
P894	+63876.50
P901	+20758.85
P902	+14890.30
Mean profit/Loss	+35898.51
Standard deviation	+26342.28

The number of buy and sell triggers as well the average rates of returns for both sell and buy triggers are provided in Table 9.

Table 9. The number of buy and sell transactions

Period	Buy Triggers	Sell triggers
P874	4	5
P881	2	3
P882	4	4
P883	5	5
P884	5	5
P891	3	4
P892	1	2
P893	4	5
P894	3	4
P901	6	6
P902	7	7

The two biggest profits produced by our trading rule were achieved, when the changes in logarithmic FRA prices follow AR-process. Furhermore, the largest four profits were achieved for the series where the Box-Pierce test statistics reject the null hypothesis that first K autocorrelation coefficients are zero. (See also table A1 in the Appendix II.) In our sample the Box-Pierce statistics appear to be qood indicator for profitability of the selected filter rule.

One of the findings in the previous sections was that the time series properities of different series vary. This indicates that investors should use different trading stategies for different series in order to maximize profits. Here we have considered only one trading rule.

The number of transactions generally increases as we move from the earliest to the latest periods. This could be partly explained by the increasing volatility in daily price changes as shown in Table 1. The increasing volatility combined with the rule that is based on absolute deviation instead of relative deviation could lead to more triggers in the last two periods.

It should be noted that because three FRA periods are traded simultaneously there exists dependence and similar patterns can be found in three consecutive time series. Therefore the technical rule used in this study can produce almost simultaneous buy and sell triggers in all three FRA periods traded at the same time. Although the expected return from using the rule is positive it varies considerably over time and it is not adjusted for this risk.

6. Conclusion

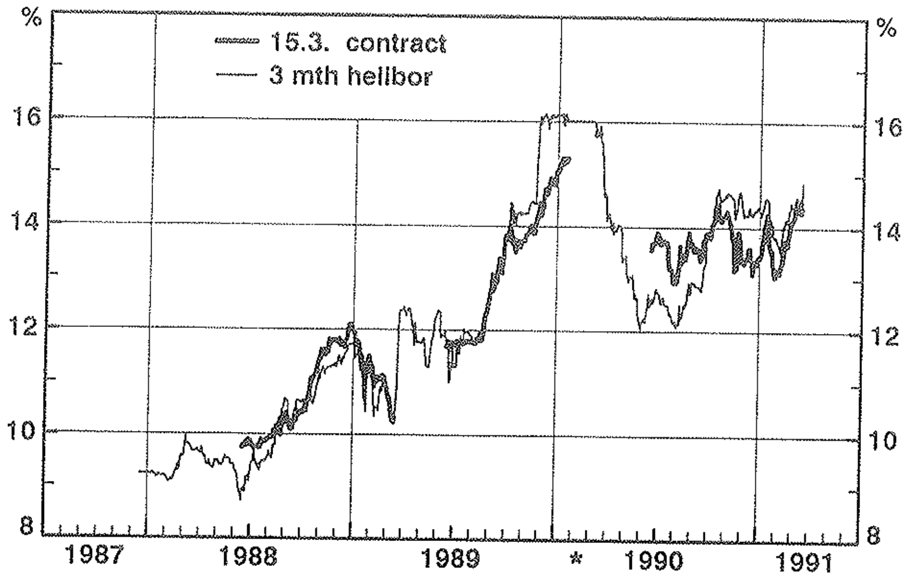
In this study we tested for linear and nonlinear dependences in the FRA price changes. We found that in almost every series there was evidence on either linear or nonlinear dependences. Especially we found evidence on ARCH-effects, which are well documented for many financial time series. The most interesting finding, however, was that all the nonlinearities can not be explained by the ARCH-model in our sample. We found evidence on the nonlinearities-in-mean by testing linearity against STAR-models. Furthermore nonlinearities-in-mean, presented by STAR models, appear only for series in which we did not find any evidence on ARCH-effects.

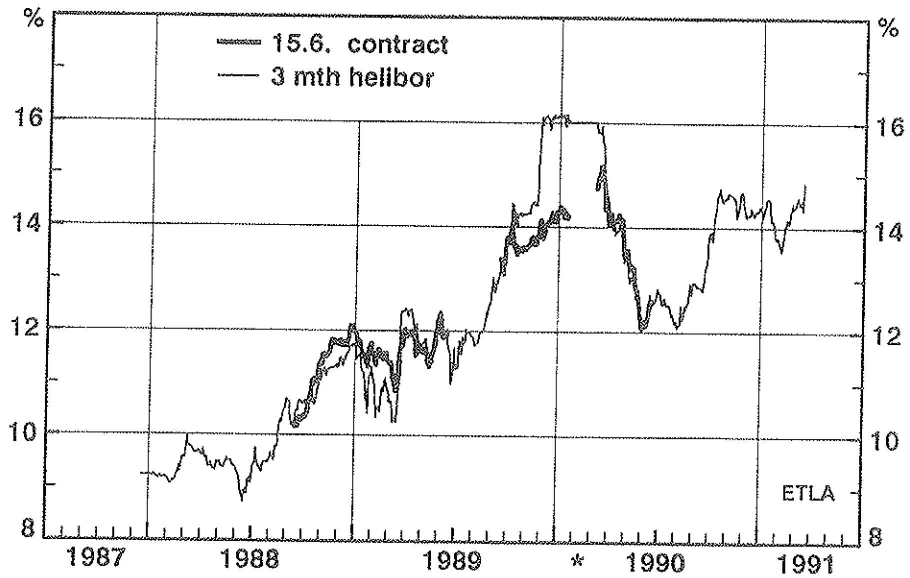
We also demonstrated the performance of the selected trading rule in our sample. As a whole the perfomance of the trading rule does not contradict the results achieved by testing for linear and nonlinear dependences in the FRA price changes. One of the findings was that the largest profits were achieved for the series where the null hypothesis of no autocorrelation was rejected using Box-Pierce test statistics.

The relative strong dependency found in some of the FRA periods and the performance of the example trading rule could be partly explained with the limited number and the homogeneity of the actors in the market. A majority of the transactions are made among five market making banks. For most of the observation period the Finnish FRA market was limited to only Finnish participants. The use of Finnish derivative securities was freed in January 1991. After the abolishment of the limitation the foreign participants have shown only minor interests in Finnish FRAs. Similar explanations were also given by Bird (1985). He proposed that the differences in the commodity markets dependency properties could be explained by the presence of the foreign actors in the market.

APPENDIX I

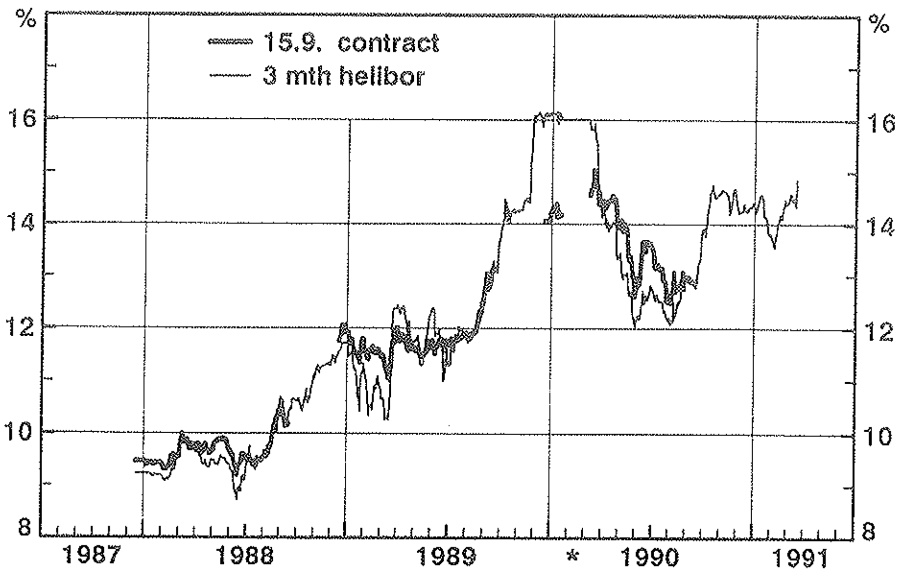
<u>Figures 1a and 1b.</u> Daily observations of FRA-contracts ending 15.3. and 15.6 respectively and three-month Helibor-rates

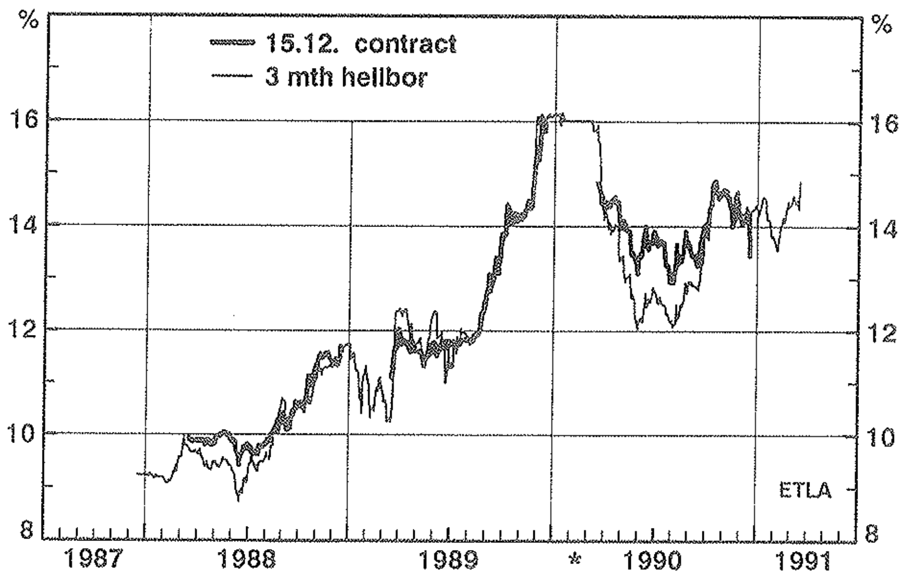




* bank strike

Figures 1c and 1d. Daily observations of FRA-contracts ending 15.9. and 15.12. respectively and three-month Helibor-rates





* bank strike

APPENDIX II

Table A1. Net Profit/Loss and Box-Pierce statistics

Period	Net Profit/Loss	Q(5)	Q(10)	Q(20)
P874	-8945.63			
P881	+28685.27			
P882	+44918.05	**	**	**
P883	+44942.51			
P884	-3220.72			
P891	+57921.06	*	**	*
P892	+75077.94	**	*	**
P893	+55979.47	*		
P894	+63876.50		*	*
P901	+20758.85			
P902	+14890.30			

* and ** denote a rejection of the null hypothesis at the 5 % and 1 % levels of sifnificancel respectively.

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