1997/9

Oslo
August 5, 1997

# Arbeidsnotat

Research Department

A Test of Uncovered Interest Rate Parity for Ten European Countries Based on Bootstrapping and Panel Data Models

by

Tom Bernhardsen

A Test of Uncovered Interest Rate Parity for Ten European Countries Based on Bootstrapping and Panel Data Models

1997/9

NORGES BANK

Notatet fås ved henvendelse
til Norges Bank, Biblioteket,
PB 1179 Sentrum, 0107 Oslo

Norges Banks arbeidsnotater inneholder
forskningsarbeider og utredninger som
vanligvis ikke har fått sin endelige form.
Hensikten er blant annet at forfatteren
kan motta kommentarer fra kolleger og
andre interesserte.

Synspunkter og konklusjoner
står for forfatterens regning.

*Copies of this Working Paper are obtainable
from Norges Bank, Library,
P.B. 1179, Sentrum, 0107 Oslo, Norway.*

*Norges Bank's Working papers present
research projects and reports
(not usually in their final form),
and are intended inter alia to enable
the author to benefit from the comments
of colleagues and other interested parties.*

*Views and conclusions expressed are the responsibility
of the author alone.*

**ISSN 0801-2504**
**ISBN 82-7553-114-4**

# A TEST OF UNCOVERED INTEREST RATE PARITY FOR TEN EUROPEAN COUNTRIES BASED ON BOOTSTRAPPING AND PANEL DATA MODELS

by

**Tom Bernhardsen**

**Research Department, Norges Bank**
**(The Central Bank of Norway)**

**Pb. 1179, Sentrum 0107 Oslo**
**phone: +47 22316490**
**fax: +47 22424062**
**E-mail: Tom.Bernhardsen@norges-bank.no**

**5. August 1997**

## Abstract

Based on both single country models and panel data models uncovered interest rate parity is tested for ten European countries relative to Germany by regressing exchange rate changes on interest rate differentials. The period is from March 1979 to February 1996 at one month, three, six and twelve months maturity. Since exchange rate changes follow a non-normal distribution, the distribution of the test-statistic is bootstrapped from the sample. The bootstrapped confidence intervals are wider with larger upper limits than the confidence intervals based on the normal distribution. The regression coefficients, all estimated to be less than one, are considerably lower for long term maturities than for short term maturities, and lower for countries outside the ERM than for ERM countries. This is explained by differences in the variance of exchange rate changes and thereby by the risk premium between long and short term maturities and the risk premium between these two groups of countries.

JEL Classification: C23 E43

Keywords: Uncovered interest rate parity, interest rate differentials, risk premium, bootstrapping, panel data models.

I am grateful to Farooq Akram, Steinar Holden, Bennett T. McCallum, Asbjørn Rødseth, Øistein Røisland, and Bent Vale for comments.

## 1) Introduction

In this paper I test uncovered interest rate parity (UIP) for the European countries France, Belgium, Denmark, Italy, the Netherlands, Austria, Switzerland, Great Britain, Norway, and Sweden, all relative to Germany. The data period is from March 1979 to February 1996 with monthly observations, and the maturity is at one month, three, six and twelve months[1]. The study extends existing literature in two directions. First, the distribution of the test statistic is estimated by bootstrapping from the sample. This may be necessary as exchange rate changes follow a non-regular distribution and inference based on the normal distribution may be misleading. Second, in addition to traditional single country models, I also estimate panel data models. By estimating panel data models it may be possible to reduce the Peso-problem, since we increase the number of countries and observations. Furthermore, panel data models make it possible to test UIP at longer maturities like six and twelve months with a relatively large number of degrees of freedom. In addition, panel data models give a tool for comparing differences between groups of countries, and models are estimated for both ERM-countries and countries outside the ERM separately. Since different coefficient estimates may reflect differences in the risk premium, one can analyse whether ERM countries differ from countries outside ERM with respect to the risk premium.

If UIP holds, the rate of return on domestic bonds is equal to the expected rate of return on foreign bonds. The standard way to test UIP is to regress exchange rate changes on interest rate differentials and test whether the regression coefficient is equal to unity. In most studies the regression coefficient is estimated to be considerably lower than unity. The reason for this may be 1) expectations are not rational, 2) there is a risk premium correlated with the interest rate differential, or 3) a peso problem exists.

A peso problem may arise if investors ex-ante expected exchange rate changes are not equal to the actual exchange rate changes. Then, non-systematic expectational errors may appear to be systematic in a small sample. Suppose that there is a large probability of a small change in the exchange rate (no realignment) and a small probability of a large change in the exchange rate (a realignment). Ex-ante the investors take the expected rate of devaluation rationally into account and this will be reflected in the interest rate differential. If no realignment has taken place ex-post, the

[1]Data on interest rates and exchange rates have been provided by BIS. Euro interest rates are quoted at 10 a. m., and exchange rates are quoted at 2:15 p. m. (Swiss time).

investors appear to have overestimated or underestimated the probability of a realignment systematically. To detect investors rationality from the sample, we need a longer sample period so that the actual number and size of realignments approach investors' realignment expectations.

Basically, the peso problem consists of two parts, a small sample problem and a problem caused by the fact that the distribution of exchange rate changes deviates considerably from the normal distribution. First, due to discrete changes in the central parity the exchange rate changes have an irregular probability distribution. This is particularly relevant for currencies in "fixed" exchange rate systems like the ERM, where realignments sometimes occur. In addition, it is well documented in the literature that the distribution of exchange rate changes, conditional upon no realignment, has thicker tails and a higher peak at zero than the normal distribution (Baillie and McMahon (1989) and Mundaca (1991)). Hence inference based on the normal distribution may be invalid[2].

Correct inference could be made if we either knew the true distribution of exchange rate changes or if the sample period and the number of observations were sufficiently large to estimate the true distribution of exchange rate changes correctly. The approach in this study is to estimate the distribution of the test statistic by bootstrapping from the sample. Bootstrapping is a resampling method where observations are drawn from the original sample with replacement. The test statistic is calculated on the bootstrapped samples. This is repeated many times. The resulting empirical distribution of the test statistics is taken to estimate the true distribution of the test statistic.

Basically, inference can be based on a-priori assumptions about the population distribution and the distribution of the test statistic, or inference can be based on the sample at hand generated from the population distribution. If there are good reasons to believe that traditional parametric inference based on a-priori assumptions is not valid, one should consider alternative inference methods based on the actual sample. Vikøren (1994) and Holden, Kolsrud and Vikøren (1993), who test the standard UIP hypothesis for respectively Norway and the Nordic countries, use Monte Carlo simulations from target zone models to estimate the distribution of the test-statistic. The models are calibrated on the basis of the sample. Vikøren show that the critical values of the simulated distribution are considerably greater than the critical values from the normal distribution. Holden,

[2]There is a large literature on UIP, see Engel (1996), Froot and Thaler (1990), Hodrick (1987), Lewis (1995), and MacDonald and Taylor (1992) for details and references. Vikøren (1994) tests UIP for the Nordic countries.

Kolsrud and Vikøren indicate that the rejection of UIP for target zone currencies is probably due to either the existence of a time variant risk premium or non-rational expectations, while it is less likely that a peso problem causes UIP to be rejected.

In the next section the standard UIP test is reviewed. I show how non-rational expectations, a risk premium and a peso problem can lead to rejection of the standard null hypothesis that the regression coefficient is equal to one. I also discuss how inference can be improved by bootstrapping from the sample. Then I discuss the estimation result from the standard test. The panel data models are discussed in section 3, and section 4 concludes.

**2) The standard UIP test**

In the absence of capital controls the expected rate of depreciation must be equal to the interest rate differential adjusted for a possible risk premium, i.e.,

(1) $$\Delta e^{e} = (i-i*) + rp ,$$

where i-i* is the interest rate differential relative to the foreign country, $\Delta e^e$ is the expected change of the logarithm of the exchange rate (e is the logarithm of the exchange rate, units of domestic currency per unit of foreign currency) and rp is the risk premium. UIP holds if and only if the risk premium is equal to zero. Since we do not observe the expected rate of depreciation, we impose rational expectations. In the literature UIP together with the auxiliary hypothesis of rational expectations has frequently been tested by estimating the model

(2) $$\Delta e = \alpha_1 + \beta_1(i-i*) + u ,$$

where $\alpha_1$ and $\beta_1$ are constants and u is an error term. Under the simultaneous null hypothesis that UIP holds and expectations are rational, are $\beta_1$=1 and u is free for autocorrelation[3]. In most of the

[3]Some also include $\alpha_1$=0 in the null hypothesis. However, in the recent literature the null hypothesis has been restricted to be $\beta_1$=1. This is because $\alpha_1$ may differ from zero due to the effect of Siegel's (1972) paradox (Froot (1990) and Lewis (1995)).

earlier studies $\beta_1$ is estimated to be lower than unity. There are at least three explanations for this: 1) expectations are not rational, 2) there exists a risk premium correlated with the interest rate differential, and 3) there is a Peso problem. Since we test a simultaneous hypothesis under the null, it is difficult to distinguish these explanations if the null hypothesis is rejected.

The ordinary least squares estimator of $\beta_1$ is given by

(3) $$\hat{\beta}_1 = \frac{\sum (i-i*)\Delta e}{\sum (i-i*)^2} .$$

However, if a risk premium exists, equation (2) is misspecified in the sense that a relevant variable is not included in the model. The appropriate specification would be

(4) $$\Delta e = \alpha_2 + \beta_2(i-i*) + \lambda rp + \epsilon ,$$

where $\lambda$ is a constant and $\epsilon$ is an error term. Then, the estimate of $\beta_1$ is given by

(5) $$\hat{\beta}_1 = \beta_2 + \lambda \frac{\sum (i-i*)rp}{\sum (i-i*)^2} + \frac{\sum (i-i*)\epsilon}{\sum (i-i*)^2} .$$

Under the null hypothesis is $\beta_2$=1. The estimate of $\beta_1$ will deviate from $\beta_2$ to the extent that the interest rate differential is correlated with either of the risk premium or the error term $\epsilon$. If expectations are rational, the error term $\epsilon$ is unsystematic and uncorrelated with all variables in the information set, including the interest rate differential. Hence the third term on the right side of equation (5) will be zero. Furthermore, if the risk premium is zero or uncorrelated with the interest rate differential, the second part on the right side of equation (5) will be zero. Hence non-rational expectations and a risk premium correlated with the interest rate differential may both cause the estimate of $\beta_1$ to deviate from $\beta_2$, which is equal to one under the null hypothesis that UIP holds and expectations are rational.

However, even though expectations are rational and the risk premium is equal to zero, the null hypothesis that $\beta_1=1$ may be rejected if the actual exchange rate changes in the sample do not approach investors' expectations. The peso problem can be illustrated in the following way: The exchange rate can be decomposed into two parts; the central parity and the exchange rate's deviation from the central parity, *the exchange rate within the band.* Let e=c+x, where c and x denote respectively the logarithm of the central parity and the exchange rate within the band. Then, the estimator of $\beta_1$ is given by

$$\hat{\beta}_1 = \frac{\Sigma(i-i*)\Delta e}{\Sigma(i-i*)^2} = \frac{\Sigma(i-i*)\Delta c}{\Sigma(i-i*)^2} + \frac{\Sigma(i-i*)\Delta x}{\Sigma(i-i*)^2} . \tag{6}$$

Suppose that no realignment occurred during the period. Then, the first part on the right side of equation (6) is zero and the estimate of $\beta_1$ is determined solely by the interest rate differential and the exchange rate within the band. It is realistic to assume that the correlation between the interest rate differential and the expected rate of devaluation is positive. Then, if the investors' ex-ante expected rate of devaluation is positive, while no devaluation occurred during the period, $\beta_1$ will be underestimated. To avoid the underestimation of $\beta_1$ we need a longer sample period so that the number and the size of the central parity changes correspond to investors' ex-ante expectations.

To test the null hypothesis that $\beta_1=1$ we need to know the distribution of the test statistic. Equation (3) shows that the true distribution of the standard t-statistic of the coefficient $\beta_1$ depends linearly on the true distribution of the exchange rate changes. Hence if the true distribution of the exchange rate changes were normal, the true distribution of the t-statistic would also be normal. However, as noted in the introduction, we have a strong reason to believe that the true distribution of exchange rate changes is not normal. To discuss inference problems further, let f(Δe) denote the true probability distribution of exchange rate changes, g(Δe) the empirical distribution of the exchange rate changes in the sample, and $T(t\beta_1)$ the true distribution of the standard t-statistic of the coefficient $\beta_1$. To test the hypothesis that $\beta_1=1$, the t-statistic should be compared with the critical values from the true distribution of the t-statistic, $T(t\beta_1)$, and not with the critical values from the normal distribution. However, the true distribution of the t-statistic is unknown and must be estimated.

One way to estimate the true distribution of the t-statistic is to bootstrap the distribution from the sample. Bootstrapping is a resampling method where observations are drawn from the original sample with replacement. The test statistic of interest is calculated on the bootstrapped samples. This is repeated many times. The resulting empirical distribution of the test statistic is taken to estimate the true distribution of the test statistic. One can show that the empirical distribution of the exchange rate changes, $g(\Delta e)$, is the non-parametric maximum likelihood estimator of the true distribution $f(\Delta e)$ (Duval and Mooney (1993)). Given that we have no other information about the true distribution of the exchange rate changes, the sample is the best estimate of that distribution. Let $B(t\beta_1)$ denote the bootstrapped distribution of the t-statistic. Then, $B(t\beta_1)$ depends on $g(\Delta e)$ in the similar way as $T(t\beta_1)$ depends on $f(\Delta e)$ (see the appendix for more details on the bootstrap method).

Assume first for simplicity that the sample period is large enough to remove the peso problem, i. e., the number and the size of realignments corresponds to investors' ex-ante expectations. In this case only non-rational expectations and a non-zero risk premium can reject the null hypothesis. In a sufficiently large sample the empirical distribution of exchange rate changes approaches the true distribution of exchange rate changes. Then, the bootstrapped distribution of the t-statistic will approach the true distribution of the t-statistic and provide correct inference. Furthermore, in a large sample the central limit theorem ensures that the true distribution of the t-statistic converges to the normal distribution (Hamilton (1994)). Hence in a sufficiently large sample the bootstrapped distribution of the t-statistic should give approximately the same inference as the normal distribution.

However, this is not true for small samples. The more the distribution of the error term in the regression equation deviates from the normal distribution, the more observations are needed to justify the central limit theorem (Hamilton (1994)). Since we know a-priori that realignments cause the true distribution of exchange rate changes to deviate strongly from the normal distribution, the central limit theorem must be used with caution. In addition, we do not know how large sample period we need to remove the peso bias of the $\beta_1$-estimate. Therefore, in a realistic situation we must expect the $\beta_1$-estimate to be peso biased and the true distribution of the t-statistic to deviate from the normal distribution.

However, even in small samples with the peso problem present we could still make correct inference if we knew the true distribution of the test statistic. Unfortunately, in small samples the empirical distribution of exchange rate changes may not approach the true distribution of exchange rate changes, and therefore, the bootstrapped distribution of the test statistic may deviate from the true distribution of the test statistic. However, since we have strong arguments that $f(\Delta e)$ deviates considerably from the normal distribution, $g(\Delta e)$ may still approximate $f(\Delta e)$ better than the normal distribution. This may be the case if $g(\Delta e)$ contains outliers caused by realignments. In that case we would still expect $B(t\beta_1)$ to approximate $T(t\beta_1)$ better than the normal distribution. Hence when we test the simultaneous hypothesis that $\beta_1=1$, the bootstrapped distribution is in any case expected to give more correct inference than the normal distribution.

For the currencies considered, realignments have frequently been devaluations relative to the Deutsche mark, i.e., there have been discrete positive jumps in the exchange rate. Hence both the true distribution and the empirical distribution of the exchange rate changes are expected to be considerably more skewed to the right than the normal distribution. Therefore, the upper limits of the confidence intervals based on the bootstrapped distribution are expected to be greater than the upper limits of the confidence intervals based on the normal distribution.

Table (1) and (2) show the estimation results for model (2) at one month and three months maturity, respectively[4]. For the parameters $\alpha$ and $\beta$ the coefficient estimate is given in the first row, the confidence interval based on the normal distribution in the second row and the bootstrapped confidence interval in the third row (both at the 95 percent level). *DW* is the Durbin Watson test statistic for first order autocorrelation of the error term (an asterisk indicates autocorrelation), $R^2$ is the coefficient of determination, *N-test* shows the p-value of the null hypothesis that the error term is normally distributed (low p-values indicate that the distribution of the error term deviates from the normal distribution), *st.dev* shows the estimated standard deviation of the error term, *skewness* shows to what extent the distribution of the error term is skewed to the right (positive values) or to the left (negative values), *excess k.* shows excess kurtosis, i.e., to what extent the distribution of the error term has longer tails than the normal distribution. For the normal distribution both the

[4]At three months maturity I only use every third observation to avoid overlapping data caused by the fact that the maturity time exceeds the observation frequenzy. This would otherwise create autocorrelation in the error term (Baillie and McMahon (1989) and Hansen and Hodrick (1980)).

skewness and the excess kurtosis are equal to zero[5].

We first look at the inference results in that we compare the bootstrapped distribution with the normal distribution. The *N-test* in table (1) and (2) show that for most models the error term cannot be assumed to be normally distributed. This is confirmed by *skewness* and *excess k.*, which indicate that in most models the distribution of the error term is more skewed to the right and has longer tails than the normal distribution. Figure (1) and (2) show the residual distribution, which for most models clearly deviates from the normal distribution. Hence inference based on the normal distribution may be invalid. Therefore, the distribution of the t-statistic has been bootstrapped from the sample.

As expected, for most models the bootstrapped confidence intervals are wider with greater upper limits than the confidence intervals based on the normal distribution. Exceptions are Italy and Great Britain at both maturities and Austria at three months maturity. For the Italian lira and the British pound the reason could be that both currencies to a larger extent have been floating compared to the other currencies analysed. Until the currency crisis in 1992 the currency band for all ERM currencies, except the Italian lira, was ±2.25 percent around the central parity. However, for the Italian lira the currency band was ±6 percent around the central parity (until 1990), and after the currency crisis Italy was not participating in the ERM. The wider is the currency band, the more the exchange rate behaves as in a pure float (Krugman (1991) and Froot and Obstfeld (1991)). The pound sterling has to a large extent been floating over the period. In a floating exchange rate system we expect the exchange rate to move more smoothly than in "fixed" exchange rate systems, where realignments sometimes occur. Hence in a floating exchange rate system there should be less outliers in the distribution of exchange rate changes than in a "fixed" exchange rate system. The Austrian shilling has been held fixed to the Deutsche mark with very small margins over a long period without large discrete jumps in the exchange rate. It also withstood the currency crisis in 1992 without being depreciated. As for the Italian lira and the British pound, the empirical distribution of changes in the Austrian shilling should not be expected to display large positive outliers as opposed to most ERM currencies, which from time to time have been devaluated.

For France (both maturities), Denmark (both maturities), the Netherlands (both maturities),

[5]The models have been estimated with the software package PCGIVE. For more details on the normality tests, see Doornik and Hendry (1994).

Switzerland (three months maturity), Norway (three months maturity) and Sweden (both maturities) the upper limits of the bootstrapped confidence intervals are noticeably greater than the upper limits of the confidence intervals based on the normal distribution. Hence it may be useful to bootstrap the distribution of the t-statistic to improve the inference. However, even though the bootstrapped confidence intervals for several models differ considerably from the confidence intervals based on the normal distribution, only in the case of Denmark do the two methods lead to different conclusions. For Denmark (both maturities), based on the normal distribution the null hypothesis is rejected, while the null hypothesis is not rejected if inference is based on the bootstrapped distribution.

In all models, except for Belgium and Sweden at one month maturity, the Durbin Watson test statistic indicates that the error term is free from autocorrelation. This is important, since autocorrelation in the error term leads to inefficient estimators (Hamilton (1994)). Furthermore, with autocorrelated disturbances the bootstrap method may not be valid. The structure of the error term is reflected in the test statistic based on the original.data. However, the resampling breaks up whatever dependence there may be in the original data, and the bootstrapping method cannot be relied on if such dependence is present (see the appendix for references).

We now turn to the coefficient estimates. The $\beta_1$-estimates differ considerably between the countries, from -1.48 (Switzerland) to 1.19 (France) at one month maturity and from -1.31 (Great Britain) to 0.95 (France) at three months maturity. In all models, except for France, $\beta_1$ is estimated to be lower than 0.5. Based on the bootstrapped distribution the null hypothesis that $\beta_1=1$ is rejected for Italy, the Netherlands, Austria and Great Britain at both maturities and for Switzerland at one month maturity. On the contrary, the null hypothesis is not rejected for France, Belgium, Denmark, Norway and Sweden at both maturities, and not for Switzerland at three months maturity. However, even though the null hypothesis is not rejected for these countries, the $\beta_1$-estimates are considerably below unity (except for France). Note that the $\beta_1$-estimates tend to be greater for ERM countries than for countries outside the ERM. This will be analysed below.

**3) Panel data models**

Below UIP will be tested on panel data models. Let

$$(7) \qquad \Delta e_{i,t} = \alpha_i D_i + \beta(i - i*)_{i,t} + u_{i,t} ,$$

where i=1,..., N is the number of countries, t=1,..., T is the time, $D_i$ is the country specific dummy variable for country i and $\alpha_i$ is the corresponding coefficient. Furthermore, $u_{i,t}$ is the error term with the assumptions that $E(u_{i,t})=0$, $E(u_{it}u_{j,s})=\sigma^2$ if i=j and s=t, and $E(u_{it}u_{j,s})=0$ otherwise. The dummy variables are supposed to reflect other country specific conditions than the interest rate differential, which influence the exchange rate changes. In addition, since the regression coefficient is imposed to be equal for all countries, the dummy variables may also capture the extent to which the interest rate differential has different effects on the exchange rate changes between the countries.

There are several reasons why it is interesting to estimate the panel data model (7). First, by estimating a panel data model we may reduce the peso bias of the β-estimate, since we increase the number of countries and observations. This argument follows Flood and Rose (1994), who also test UIP on panel data models. Using daily data for the period March 1979-March 1994 Flood and Rose find that for floating exchange rate regimes β is estimated to be negative. On the contrary, for fixed exchange rate regimes (ERM) the estimate of β is positive though below unity. Flood and Rose argue that by pooling data across countries the Peso problem may be removed. By estimating panel data models with and without realignments in the sample they estimate the "peso bias" to be about -0.35. If the use of panel data actually removes the peso bias of the β-estimate, we have ruled out one of the explanations under the alternative hypothesis. Then, only non-rational expectations and a risk premium can reject the null hypothesis.

Second, the empirical distribution of exchange rate changes from the panel data set may better approximate the true distribution of exchange rate changes compared to the data sets for each country separately. In this case the bootstrapped distribution of the t-statistic from the panel data model may better approximate the true distribution of the t-statistic compared to the single country models. Of course, with panel data and hence a large number of observations one could argue that the central limit theorem ensures the t-statistic to be approximately normally distributed. Then, the bootstrapped distribution of the t-statistic should approximate the normal distribution. The results will indicate whether this is the case.

Third, it is interesting to test UIP at both short and long term maturities. At one month and three months maturity the sample period covers enough observations to estimate models for each country separately. However, at 6 and 12 months maturity I only use every sixth and twelfth observation to avoid the overlapping data problem (see footnote 4). This reduces the number of observations considerably, and by restricting the β-coefficient to be equal for all countries one saves degrees of freedom. An additional argument for estimating panel data models is that the regression coefficients are equal for all countries under the null hypothesis of UIP. This is opposed to "standard cases" in panel data studies, where homogeneity assumptions and parameter restrictions may be more doubtful.

Fourth, by estimating panel data models one can compare different groups of countries. Below panel data models are estimated for all countries as one group and for the ERM countries and the countries outside the ERM separately. Since different coefficient estimates may reflect differences in the risk premium, one can analyse whether there are differences between the two groups of countries with respect to the risk premium.

Table (3) and (4) show the estimation results for model (7) at one month, three, six and twelve months maturity. For the regression coefficient the estimate is given in the first row, the confidence interval based on the normal distribution in the second row, and the bootstrapped confidence interval in the third row. To test for first order autocorrelation in the error term I estimate the model

$$\text{(8)} \qquad \hat{u}_{i,t} = \rho \hat{u}_{i,t-1} + v_{i,t} \quad where \quad \hat{\rho} = \frac{\sum_{i=1}^{N}\sum_{t=1}^{T} \hat{u}_{i,t}\hat{u}_{i,t-1}}{\sum_{i=1}^{N}\sum_{t=1}^{T} \hat{u}_{i,t-1}\hat{u}_{i,t-1}} .$$

The *AR-test* shows the estimate of ρ and the standard t-statistic in parentheses[6]. For all models at three, six and twelve months maturity the hypothesis that the error term is free for autocorrelation is not rejected. However, at one month maturity the hypothesis that the error term is free for autocorrelation is rejected[7].

[6]See Kmenta (1986) for more details.

[7]The autocorrelation test in equation (8) was performed with the software package GAUSS. The panel data models were estimated with PCGIVE (see footnote 5).

For all models the hypothesis that the error term is normally distributed is clearly rejected. As for the single country models, one should therefore bootstrap the distribution of the t-statistic on the basis of the actual sample. For all models, except for the model with all countries at three months maturity, the upper limit of the bootstrapped confidence interval is greater than the upper limit of the confidence interval based on the normal distribution. Generally, as for the single country models the bootstrapped confidence intervals are wider than the confidence interval based on the normal distribution. The reason is that the bootstrapped distribution of the t-statistic reflects the relatively large standard deviation of the exchange rate changes and "outliers" caused by realignments.

However, for all models the hypothesis that $\beta=1$ is rejected. Furthermore, the largest coefficient estimate is 0.52 (the ERM countries at one month maturity), i.e., all coefficient estimates are considerably below unity. Generally, the coefficient estimates are considerably greater for the ERM countries than for the countries outside the ERM, although this difference is not significant, since the confidence intervals for the two groups of countries overlap. For all models with only the ERM countries included the coefficient estimates are positive, while they are negative for all models with only the countries outside the ERM included. Furthermore, the coefficient estimates tend to be lower for long term maturities. In particular, the coefficient estimates at twelve month maturities are considerably lower than the coefficient estimates at one month, three and six months maturity. It is difficult to say to what extent the use of panel data in our analysis actually reduces the peso problem. But following the argument by Flood and Rose (1994), it is likely that non-rational expectations or a time varying risk premium are important explanations for the rejection of the null hypothesis, since the use of panel data is expected to reduce the peso problem in the first place.

The models are specified with a full set of country specific dummy variables and no general intercept term. The relative size between the coefficients can be interpreted as reflecting different policy between the countries. For all models with all countries included, Italy and Sweden have the highest dummy coefficients. These two countries also have the highest dummy coefficients in the models with respectively only the ERM countries and the countries outside the ERM included.

As noted above, the rejection of the null hypothesis that $\beta=1$ may be caused by 1) non-rational expectations, 2) the existence of a risk premium correlated with the interest rate differential, or 3) a peso problem. By estimating a panel data model we can hope that the peso problem is reduced, even though we cannot be sure of that. Non-rational expectations may certainly cause the estimates of the regression coefficients to be considerably lower than unity. However, it is less plausible that non-rational expectations explain the difference in the coefficient estimates between the ERM countries and the countries outside the ERM, and the fact that the coefficient estimates are lower for long term maturities. It is more likely that these issues are explained by differences with respect to the risk premium, which, in addition to non-rational expectations, may also explain why the levels of the estimates are considerably below unity.

To explain these results further we need to know more precisely what determines the risk premium, which, I will argue, may explain why the β-estimates differ between the ERM countries and the countries outside the ERM, and why the β-estimates are lower for longer maturities. This can be shown within a mean-variance analysis (Dornbusch (1983) and Rødseth (1996)), where an investor maximizes a utility function with respect to expected real return and risk, the latter measured by the variance of the real return. The resulting optimal share of foreign bonds in the portfolio can be divided into two parts, the minimum variance portfolio and the speculative portfolio. The minimum variance portfolio is the share of foreign bonds in the portfolio which minimizes risk. The speculative portfolio is proportional to the difference in the expected return of foreign and domestic bonds. Dornbusch and Rødseth show that in equilibrium we must have

$$(9) \qquad i - i* - \Delta e^{e} = R\sigma_e^2[\alpha* - \alpha] \; .$$

The *relative expected return*, $i-i^*-\Delta e^e$, is equal to the risk premium. The risk premium consists of three parts, the coefficient of relative risk aversion, $R>0$, the variance of the exchange rate changes $\sigma^2_e>0$, and the difference between the minimum variance portfolio, $\alpha^*$, and the actual share of foreign currency in the market, $\alpha$. Suppose that the share of foreign currency in the market is equal to the minimum variance portfolio. Then the risk premium is equal to zero. In some intuitive sense the investor's actual share of foreign currency in the portfolio is then equal to the share of foreign currency investors want to hold. If the investors in this situation are "forced" to increase their share of domestic bonds in the portfolio, domestic bonds become more "risky". To accept this the investors demand a higher relative expected return on domestic bonds. Hence in the new equilibrium the domestic interest rate is higher, or the expected rate of depreciation is lower (assume $i^*$ exogenous). Similarly, if the investors are forced to increase the share of foreign bonds in the portfolio, foreign bonds become more risky, and the expected relative return must decrease. Hence risk is relative to deviations from the minimum variance portfolio. It is clear that the sign of the risk premium depends on the proportion of foreign currency in the market relative to the minimum variance portfolio. In addition, the magnitude of the risk premium also depends on the coefficient of relative risk aversion and the variance of the exchange rate changes. Changes in either of these variables, or changes in investor's preferences, may all change the risk premium and the expected relative return.

If there is no peso problem and expectations are rational, an estimate of $\beta$ in model (7) less than unity implies that a one percent increase in the interest rate differential is associated with a less than one percent increase in the rate of depreciation. This means that the risk premium, defined by equation (9), must rise with the interest rate differential. The larger is the risk premium, and the more correlated it is with the interest rate differential, the lower is the estimate of $\beta$ expected to be.

The coefficient of relative risk aversion and the share of foreign currency in the market relative to the minimum variance portfolio are unknown. Hence we do not observe the risk premium and to what extent it is correlated with the interest rate differential. However, we observe the variance of the exchange rate changes, and, if we are lucky, it might happen that the variance of the exchange rate changes may throw some light on the problem. This could be the case if the variance of the exchange rate changes is important for the determination of the risk premium. Table (5) shows the standard deviation of the one month, three, six and twelve months percentage changes of the exchange rate for both ERM countries and countries outside the ERM. The standard deviation is considerably higher for countries outside the ERM. This is not surprising, since the ERM countries participate in a fixed exchange rate system and exchange rate changes are measured relative to an ERM currency, the Deutsche mark. Furthermore, for both groups of countries the standard deviation of the exchange rate changes increases with the differential horizon.

This can be interpreted as consistent with the estimates of $\beta$ from model (7). The $\beta$ estimates are considerably lower for the countries outside the ERM, which may be explained by the relatively large standard deviation of the exchange rate changes for these countries. In addition, the fact that the $\beta$ estimates are lower for long term maturities may be explained by the relatively large standard deviation for the exchange rate changes at long differential horizons. Therefore, interpreted with some caution, the estimation results for model (7) indicate that differences with respect to the risk premium may explain the different $\beta$ estimates for the two group of countries and the fact that the $\beta$ estimates are lower for long term maturities.

## 4) Conclusion

In this paper I have used both single country models and panel data models to test UIP by regressing exchange rate changes on interest rate differentials. Under the joint null hypothesis of UIP and rational expectations the regression coefficient is equal to unity. In almost all single country models the regression coefficient is estimated to be lower than unity, and in about half of the models the regression coefficient is significantly lower than unity. The reason for this may be non-rational expectations, a non-zero risk premium, and a peso problem.

Since exchange rate changes follow a non-regular distribution caused by realignments, inference based on the normal distribution may be incorrect. Therefore, the distribution of the t-statistic has been estimated by bootstrapping from the sample. The bootstrapped confidence intervals are wider with larger upper limits than the confidence intervals based on the normal distribution. This reflects outliers in the sample distribution of exchange rate changes (realignments). However, the bootstrapped distributions give almost the same test results as the normal distribution. Only in the case of Denmark do the two methods lead to different conclusions: based on the normal distribution the null hypothesis is rejected, while the null hypothesis is not rejected if inference is based on the bootstrapped distribution.

By testing UIP on panel data the peso problem may be reduced, since we increase the number of observations considerably. In addition, compared to the single country models the bootstrapped distribution of the t-statistic may give an even better approximation to the true distribution of the t-statistic. However, also for the panel data models the null hypothesis is rejected. To the extent that the large number of observations actually reduces the peso problem, this indicates that non-rational expectations or a risk premium are the explanations for the rejection of the null hypothesis.

Furthermore, the regression coefficients are considerably lower for long term maturities than for short term maturities and lower for countries outside the ERM than for the ERM countries. This may be explained by differences in the risk premium between long and short term maturities and differences in the risk premium between these two groups of countries. The standard deviation of exchange rate changes, which influence the risk premium, are considerably larger for countries outside the ERM than for the ERM countries, and for both group of countries the standard deviation of exchange rate changes increases with the differential horizon.

## Appendix: The bootstrapping method

Traditional parametric inferences rely on a-priori assumptions about the distribution of the test statistic or some underlying error component. For example, in standard regression analysis the normality of the OLS estimator follows from the assumption that the error term is normally distributed. Alternatively, for non-normal disturbances the central limit theorem ensures the estimator to be asymptotically normally distributed. However, in small samples the distribution of the test statistic may deviate considerably from the normal distribution. Then, rather than making unrealistic assumptions about the distribution of the test statistic, it may be better to estimate the distribution from the sample at hand.

Bootstrapping relies on the analogy between the sample and the distribution from which the sample was drawn. Suppose that the sample consists of T observations, $x_1, x_2,..., x_T$, which are generated from an unknown probability distribution $F(x,\alpha)$, where x is the vector of observations and $\alpha$ is the population parameter of interest. Let a denote the estimator (and the estimate) of $\alpha$. Construct an empirical distribution function (EDF) from the sample by placing the probability of 1/T at each point, $x_1, x_2,..., x_T$. From the EDF one draws a simple random sample of size T with replacement. Thus the bootstrap sample will contain some of the original observations more than once, and others of them not at all, in a completely random order. Then, $\alpha$ is estimated from the bootstrap sample, and the estimate is denoted $a_1$. Repeat this N times, where N is a large number, for example 10.000. The true distribution of the test statistic is estimated by placing a probability of 1/N at each point $a_1, a_2,..., a_N$. This is the bootstrap distribution of the test statistic a, which can be used to make inferences about the underlying population parameter $\alpha$. By cumulating the bootstrap distribution one can calculate upper and lower critical values at the one-sided 2.5 percent level to obtain a 95 percent confidence interval for the population parameter. This is the method used to calculate bootstrap confidence intervals in tables (1)-(4).

The justification of bootstrapping rests on two analogies; first, the sample EDF with the population distribution that generated the data, $F(x,\alpha)$, and second, the random resampling mechanism with the random component of the function $F(x,\alpha)$. It can be shown that the empirical distribution function, EDF, is the non-parametric maximum likelihood estimate of the population function, $F(x,\alpha)$. Given that we have no other information about the population function, the sample is the best estimate of that population. In this sense the sample is treated as the population. Then, the resamples are analogous to independent random samples from the population distribution, $F(x,\alpha)$. The sampling distribution of the tests statistic, $a_1, a_2,..., a_T$, is supposed to reflect the error component of the population distribution $F(x,\alpha)$. For more details on bootstrapping and proofs, see Davidson and MacKinnon (1993), Duval and Mooney (1993) and Li and Maddala (1996) and their references.

The bootstrap method is not valid if the error term of the model is serially dependent. The structure of the error term is reflected in the test statistic based on the original data. However, the resampling breaks up whatever dependence there may be in the original data, and the bootstrapped results cannot be relied on if such dependence is present.

One should note that even if the sample distribution is the non-parametric maximum likelihood estimator of the population distribution, the bootstrapped distribution does not necessarily approximate the true distribution of the test statistic. The bootstrapped distribution will approximate the true distribution of the test statistic only if the sample is a good approximation of the population. Lack of congruence between the sample distribution and the population distribution could arise, either because of a small sample or just because of bad luck. Hence bootstrapping does not necessarily solve all problems caused by the fact that the population distribution is unknown. But bootstrapping should be regarded as an alternative to standard parametric methods, in particular in situations where we a-priori expect the test statistic not to have a standard distribution or where we a-posteriori rejects the distribution of the test statistic to be standard.

**References**

Baillie R. and McMahon P. (1989) "The Foreign Exchange Market: Theory and Econometric Evidence", Cambridge University Press.

Davidson, R. and MacKinnon, J. G. (1993) "Estimation and Inference in Econometrics", Oxford University Press, Oxford.

Doornik and Hendry (1994) "PCGIVE Profesional 8.0", International Thomson Publishing, London.

Dornbusch, R. (1983) "Exchange Rate Risk and the Macroeconomics of Exchange Rate Determination", Printed in Hawkins, R., Levich, R., and Wihlborg, C. G. (eds.) "The Internationalization of Financial Markets and National Economic Policy, vol. 3", Greenwich Press. Also Printed in Dornbusch, R. (1993) "Exchange Rates and Inflation", The MIT Press, Massascusetts.

Duval, R. D. and Mooney, C. Z. (1993) "Bootstrapping: A Nonparametric Approach to Statistical Inference", In the series: Quantitative Applications in the Social Sciences, Sage Publications, London.

Engel, C. (1996) "The Forward Discount Anomaly and the Risk Premium: A Survey of Recent Evidence", *Journal of Empirical Finance* 3, 123-192.

Flood, R. P. and Rose, A. K. (1994) "Fixes: Of the Forward Discount Puzzle", *NBER Working Paper* 4928.

Froot, K. A. (1990) "On the Efficiency of Foreign Exchange Market", Unpublished manuscript, Cambridge: Harward Business School.

Froot, K. A and Thaler, R. H. (1990) "Anomalies, Foreign Exchange", *Journal of Economic Perspectives* 4, 179-192.

Froot, K. A. and Obstfeld, M. (1991) "Exchange Rate Dynamics Under Stochastic Regime Shifts: a Unified Approach", *Journal of International Economics* 31, 203-229.

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

Hansen, L. P. and Hodrick, R. J. (1980) "Forward Exchange Rates as Optimal Predictors of Future Spot Rates", *Journal of Political Economy* 8, 829-853.

Hodrick, R. J. (1987) "The Empirical Evidence on the Efficiency of Forward and Future Foreign Exchange Markets", Fundamentals of Pure and Applied Economics no.4, Chicago: Northwestern University.

Holden, S., Kolsrud, D. and Vikøren, B. (1993) "Testing Uncovered Interest Parity: Evidence and some Monte Carlo Experiments", *Working Paper, 93/2 Norges Bank,* (The Central Bank of Norway).

Kmenta, J. (1986) "Elements of Econometrics", Macmillan Publishing Company, New York.

Krugman, P. (1991) "Target Zones and Exchange Rate Dynamics", *Quarterly Journal of Economics* 106, 669-82.

Lewis, K. K. (1995) "Puzzles in International Financial Markets", in Grossman, G. M. and Rogoff, K. (eds.) "Handbook of International Economics, vol. 3, North-Holland.

Li, H. and Maddala, G. S. (1996) "Bootstrapping Time Series Models", *Econometric Reviews* 15, 115-158.

MacDonald, R. and Taylor, M. P. (1992) "Exchange Rate Economics: A Survey", *IMF Staff Papers* 39, 1-57.

Mundaca, G. (1991) "The Volatility of the Norwegian Currency Baskets", *The Scandinavian Journal of Economics* 93, 53-73.

Rødseth, A. (1996) "Open Economy Macro Economics", University of Oslo.

Siegel, J. (1972) "Risk, Interest Rates and the Forward Exchange", *Quarterly Journel of Economics* 86, 303-339.

Vikøren, B. (1994) "Interest Rate Differential, Exchange Rate Expectations and Capital Mobility: Norwegian Evidence", Dr. Polit. dissertation, Sosialøkonomisk Institutt, University of Oslo.

Tab (1): Regression results from model (2), $\Delta e=\alpha_1+\beta_1(i-i^*)+u$, one month maturity. For each country the number of observations is 204.

| | France | Belgium | Denmark | Italy | The Netherlands |
|---|---|---|---|---|---|
| $\alpha$ | -0.0020<br>(-0.0036, -0.0004)<br>(-0.0042, 0.0006) | 0.0003<br>(-0.0013, 0.0020)<br>(-0.0006, 0.0013) | 0.0002<br>(-0.0016, 0.0020)<br>(-0.0013, 0.0016) | 0.0037<br>(-0.0011, 0.0086)<br>(-0.0012, 0.0088) | 0.0005<br>(-0.0005, 0.0010)<br>(-0.0004, 0.0010) |
| $\beta$ | 1.1941<br>(0.8463, 1.5419)<br>(0.2276, 1.9700) | 0.4295<br>(-0.1413, 1.0003)<br>(-0.0490, 1.0100) | 0.4429<br>(-0.0133, 0.9000)<br>(-0.0372, 1.0227) | 0.0782<br>(-0.6332, 0.7897)<br>(-0.5524, 0.7226) | -0.8549<br>(-1.4557, -0.2540)<br>(-1.6700, 0.0234) |
| DW | 2.06 | 1.54* | 1.73 | 1.82 | 2.06 |
| $R^2$ | 0.1839 | 0.0107 | 0.0177 | 0.0002 | 0.0372 |
| N-test | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 |
| st.dev | 0.0082 | 0.0075 | 0.0077 | 0.0180 | 0.0034 |
| skewness | 1.7745 | 4.4469 | 2.5227 | 2.6850 | 1.8017 |
| excess k. | 6.8671 | 39.9729 | 12.5079 | 17.3819 | 8.3863 |
| | | | | | |
| | Austria | Switzerland | Great Britain | Norway | Sweden |
| $\alpha$ | -0.0002<br>(-0.0004, 0.00001)<br>(-0.0005, 0.00001) | -0.0025<br>(-0.0052, 0.0002)<br>(-0.0050, 0.0001) | 0.0072<br>(0.0010, 0.0133)<br>0.0016, 0.0135) | 0.0006<br>(-0.0025, 0.0036)<br>(-0.0016, 0.0028) | 0.0017<br>(-0.0052, 0.0087)<br>(-0.0056, 0.0091) |
| $\beta$ | 0.1039<br>(-0.2194, 0.4272)<br>(-0.2581, 0.5049) | -1.4834<br>(-2.9000, -0.0668)<br>(-3.0000, 0.0041) | -1.2897<br>(-2.7251, 0.1458)<br>(-2.8100, 0.1382) | 0.4532<br>(-0.1762, 1.0825)<br>(-0.1751, 1.1109) | 0.4043<br>(-1.1795, 1.9882)<br>(-1.4153, 2.3807) |
| DW | 1.98 | 1.76 | 1.76 | 1.81 | 1.61* |
| $R^2$ | 0.0020 | 0.0205 | 0.0152 | 0.0098 | 0.0012 |
| N-test | 0.0000 | 0.1090 | 0.0001 | 0.0001 | 0.0000 |
| st.dev | 0.0016 | 0.0137 | 0.0258 | 0.0141 | 0.0229 |
| skewness | -2.9114 | -0.0976 | 0.3845 | 0.4342 | 2.8130 |
| excess k. | 16.5695 | 0.5960 | 1.7329 | 1.6958 | 15.4407 |

For the parameters $\alpha$ and $\beta$ the coefficient estimate is given in the first row, the confidence interval based on the normal distribution in the second row and the bootstrapped confidence interval in the third row (both at the 95 percent level). *DW* is the Durbin Watson test statistic, $R^2$ is the coefficient of determination, *N-test* shows the p-value of the null hypothesis that the error term is normally distributed (low p-values indicate that the distribution of the error term deviates from the normal distribution), *st.dev* shows the estimated standard deviation of the error term, *skewness* shows to what extent the distribution of the error term is skewed to the right (positive values) or to the left (negative values), *excess k.* shows excess kurtosis, i.e., to what extent the distribution of the error term has longer tails than the normal distribution. For the normal distribution both the skewness and the excess kurtosis are equal to zero.

Tab (2): Regression results from model (2), $\Delta e=\alpha_1+\beta_1(i-i^*)+u$, three months maturity. For each country the number of observations is 68.

| | France | Belgium | Denmark | Italy | The Netherlands |
|---|---|---|---|---|---|
| $\alpha$ | -0.0043<br>(-0.0093, 0.0008)<br>(-0.0098, 0.0028) | 0.0013<br>(-0.0051, 0.0076)<br>(-0.0032, 0.0068) | 0.0008<br>(-0.0057, 0.0073)<br>(-0.0066, 0.0083) | 0.0131<br>(-0.0041, 0.0307)<br>(-0.0055, 0.0352) | 0.0007<br>(-0.0008, 0.0028)<br>(-0.0007, 0.0022) |
| $\beta$ | 0.9511<br>(0.5948, 1.3074)<br>(0.1676, 1.4006) | 0.3791<br>(-0.3400, 1.0983)<br>(-0.3056, 1.2200) | 0.4051<br>(-0.1225, 0.9327)<br>(-0.2368, 1.0982) | 0.0160<br>(-0.7818, 0.8137)<br>(-0.8370, 0.6840) | -0.1329<br>(-0.7859, 0.5202)<br>(-0.8516, 0.8154) |
| DW | 1.97 | 1.64 | 1.88 | 2.20 | 1.88 |
| $R^2$ | 0.2963 | 0.0162 | 0.0337 | 0.0001 | 0.0024 |
| N-test | 0.0210 | 0.0000 | 0.0013 | 0.0000 | 0.0005 |
| st.dev | 0.0135 | 0.0160 | 0.0146 | 0.0359 | 0.0056 |
| skewness | 0.7841 | 3.3022 | 1.1698 | 1.6749 | 1.2403 |
| excess k. | 0.5639 | 17.1195 | 3.0832 | 6.4088 | 3.8309 |
| | | | | | |
| | Austria | Switzerland | Great Britain | Norway | Sweden |
| $\alpha$ | -0.0005<br>(-0.0013, 0.0003)<br>(-0.0012, 0.00008) | -0.0038<br>(-0.0128, 0.0051)<br>(-0.0124, 0.0057) | 0.0221<br>(0.0022, 0.0419)<br>(0.0027, 0.0452) | 0.0025<br>(-0.0090, 0.0140)<br>(-0.0063, 0.0106) | 0.0077<br>(-0.0181, 0.0335)<br>(-0.0194, 00346) |
| $\beta$ | -0.1379<br>(-0.5393, 0.2634)<br>(-0.7350, 0.1899) | -0.5343<br>(-2.2441, 1.1754)<br>(-2.6200, 1.7019) | -1.3135<br>(-2.8776, 0.2506)<br>(-3.0667, 0.2173) | 0.4074<br>(-0.3915, 1.2063)<br>(-0.3960, 1.3820) | 0.2128<br>(-1.8028, 2.2284)<br>(-2.1770, 2.6759) |
| DW | 1.88 | 2.10 | 1.99 | 1.65 | 1.88 |
| $R^2$ | 0.0069 | 0.0057 | 0.04 | 0.0151 | 0.0007 |
| N-test | 0.0000 | 0.6242 | 0.0317 | 0.1463 | 0.0000 |
| st.dev | 0.0030 | 0.0255 | 0.0454 | 0.0290 | 0.0474 |
| skewness | -4.2086 | -0.0945 | 0.7897 | 0.4930 | 1.8332 |
| excess k. | 24.3793 | 0.1471 | 0.9574 | 0.7040 | 4.8000 |

For the parameters $\alpha$ and $\beta$ the coefficient estimate is given in the first row, the confidence interval based on the normal distribution in the second row and the bootstrapped confidence interval in the third row (both at the 95 percent level). *DW* is the Durbin Watson test statistic, $R^2$ is the coefficient of determination, *N-test* shows the p-value of the null hypothesis that the error term is normally distributed (low p-values indicate that the distribution of the error term deviates from the normal distribution), *st.dev* shows the estimated standard deviation of the error term, *skewness* shows to what extent the distribution of the error term is skewed to the right (positive values) or to the left (negative values), *excess k.* shows excess kurtosis, i.e., to what extent the distribution of the error term has longer tails than the normal distribution. For the normal distribution both the skewness and the excess kurtosis are equal to zero.

Tab (3): Regression results from model (7), $\Delta e_{i,t}=\alpha_i D_i+\beta(i-i^*)+u_{i,t}$, one month and three months maturity

| | One month maturity | | | Three months maturity | | |
|---|---|---|---|---|---|---|
| | All countries N=10, T=204 | ERM N=5, T=204 | Non-ERM N=5, T=204 | All countries N=10, T=68 | ERM N=5, T=68 | Non-ERM N=5, T=68 |
| Interest rate differential | 0.2256<br>(-0.048, 0.499)<br>(-0.137, 0.590) | 0.5234<br>(0.274, 0.773)<br>(0.084, 0.940) | -0.2095<br>(-0.735, 0.316)<br>(-0.808, 0.417) | 0.1669 (1.03)<br>(0.206, 0.841)<br>(-0.239, 0.551) | 0.3923<br>(0.104, 0.681)<br>(-0.039, 0.751) | -0.1936<br>(-0.818, 0.430)<br>(-0.940, 0.570) |
| dfrf | 0.0012 (1.07)<br>(-0.0001, 0.003) | 0.0002 (0.26)<br>(-0.001, 0.002) | | 0.0041 (1.06)<br>(-0.0007, 0.009) | 0.0017 (0.59)<br>(-0.003, 0.006) | |
| dbef | 0.0008 (0.73)<br>(-0.0004, 0.002) | 0.0001 (0.16)<br>(-0.001, 0.002) | | 0.0027 (0.75)<br>(-0.0016, 0.007) | 0.0012 (0.44)<br>(-0.003, 0.006) | |
| ddkk | 0.0009 (0.80)<br>(-0.0006, 0.003) | -0.0001 (-0.06)<br>(-0.002, 0.002) | | 0.0032 (0.84)<br>(-0.002, 0.009) | 0.0009 (0.32)<br>(-0.004, 0.006) | |
| ditl | 0.0028 (2.18)<br>(-0.0004, 0.007) | 0.0011 (1.06)<br>(-0.003, 0.005) | | 0.0103 (2.24)<br>(-0.002, 0.002) | 0.0061 (1.66)<br>(-0.007, 0.021) | |
| dnlg | 0.0001 (0.11)<br>(-0.0004, 0.001) | 0.0001 (0.02)<br>(-0.0005,0.0006) | | 0.0004 (0.12)<br>(-0.001, 0.002) | 0.0002 (0.08)<br>(-0.001, 0.002) | |
| dats | -0.0002 (-0.23)<br>-0.0005, 0.0001) | | -0.0002 (-0.14)<br>(-0.0004,0.0001) | -0.0007 (-0.21)<br>(-0.002, 0.0000) | | -0.0005 (-0.11)<br>(-0.001, 0.0004) |
| dchf | -0.0002 (-0.20)<br>(-0.002, 0.002) | | -0.0008 (-0.61)<br>(-0.003, 0.001) | -0.0011 (-0.33)<br>(-0.007, 0.005) | | -0.0025 (-0.58)<br>(-0.009, 0.004) |
| dgbp | 0.0018 (1.65)<br>(-0.002, 0.006) | | 0.0033 (2.16)<br>(-0.0006, 0.007) | 0.0065 (1.68)<br>(-0.005, 0.019) | | 0.0103 (1.91)<br>(-0.003, 0.025) |
| dnok | 0.0014 (1.26)<br>(-0.0008, 0.004) | | 0.0031 (1.91)<br>(0.0004, 0.006) | 0.0052 (1.33)<br>(-0.0008, 0.004) | | 0.0094 (1.68)<br>(-0.0003, 0.02) |
| dsek | 0.0024 (2.09)<br>(-0.0007, 0.006) | | 0.0041 (2.53)<br>(0.0004, 0.008) | 0.0082 (2.09)<br>(-0.0007, 0.006) | | 0.0123 (2.21)<br>(-0.0005, 0.025) |
| std.dev | 0.015 | 0.010 | 0.018 | 0.028 | 0.020 | 0.034 |
| $R^2$ | 0.02 | 0.06 | 0.02 | 0.06 | 0.13 | 0.04 |
| AR-test | 0.13 (6.03) | 0.09 (2.77) | 0.15 (4.71) | 0.02 (0.57) | -0.04 (-0.71) | 0.04 (0.69) |
| N-test | 0.0000 | 0.0000 | 0.0000 | 0.000 | 0.0000 | 0.0000 |
| skewness | 2.00 | 3.85 | 1.47 | 1.75 | 2.68 | 1.46 |
| excess k. | 19.54 | 41.05 | 12.31 | 9.84 | 20.02 | 6.17 |

For the interest rate differential the estimated coefficient is given in the first row, the confidence interval for the coefficient based on the normal distribution in the second row and the bootstrapped confidence interval in the third row. For the dummy variables the estimated coefficient and the standard t-statistic are given in the first row and the bootstrapped confidence interval for the coefficient in the second row. The *Ar-test* shows the regression coefficient of the residuals on the first order lag of the residuals with the standard t-statistic in pharenthesis, $R^2$ is the coefficient of determination, *N-test* shows the p-value of the null hypothesis that the error term is normally distributed (low p-values indicate that the distribution of the error term deviates from the normal distribution), *st.dev* shows the estimated standard deviation of the error term, *skewness* shows to what extent the distribution of the error term is skewed to the right (positive values) or to the left (negative values), *excess k.* shows excess kurtosis, i.e., to what extent the distribution of the error term has longer tails than the normal distribution. For the normal distribution both the skewness and the excess kurtosis are equal to zero.

Tab (4): Regression results from model (7), $\Delta e_{i,t}=\alpha_i D_i+\beta(i-i^*)+u_{i,t}$, six and twelve months maturity

| | Six months maturity | | | Twelve months maturity | | |
|---|---|---|---|---|---|---|
| | All countries N=10, T=34 | ERM N=5, T=34 | Non-ERM N=5, T=34 | All countries N=10, T=17 | ERM N=5, T=17 | Non-ERM N=5, T=17 |
| Interest rate differential | 0.2124<br>(-0.130, 0.555)<br>(-0.181, 0.570) | 0.4424<br>(0.127, 0.758)<br>(0.178, 0.801) | -0.1917<br>(-0.877, 0.493)<br>(-0.888, 0.503) | -0.1755<br>(-0.583, 0.232)<br>(-0.666, 0.298) | 0.0556<br>(-0.346, 0.457)<br>(-0.548, 0.631) | -0.5370<br>(-1.309, 0.235)<br>(-1.422, 0.343) |
| dfrf | 0.0077 (0.94)<br>(-0.002, 0.018) | 0.0027 (0.42)<br>(-0.007, 0.013) | | 0.0336 (1.89)<br>(0.008, 0.063) | 0.0241 (1.64)<br>(-0.002, 0.054) | |
| dbef | 0.0049 (0.65)<br>(-0.005, 0.047) | 0.0016 (0.28)<br>(-0.008, 0.014) | | 0.0216 (1.28)<br>(0.001, 0.048) | 0.0147 (1.09)<br>(-0.007, 0.043) | |
| ddkk | 0.0056 (0.69)<br>(-0.004, 0.017) | 0.0006 (0.10)<br>(-0.009, 0.012) | | 0.0297 (1.65)<br>(0.004, 0.059) | 0.0199 (1.34)<br>(-0.008, 0.052) | |
| ditl | 0.0200 (2.05)<br>(-0.005, 0.049) | 0.0113 (1.41)<br>(-0.015, 0.043) | | 0.0758 (3.51)<br>(0.015, 0.141) | 0.0592 (3.12)<br>(-0.013, 0.138) | |
| dnlg | 0.0007 (0.10)<br>(-0.002, 0.004) | 0.0003 (0.06)<br>(-0.003, 0.004) | | 0.0029 (0.19)<br>(-0.001, 0.008) | 0.0021 (0.18)<br>(-0.001, 0.006) | |
| dats | -0.0016 (-0.22)<br>(-0.003, -0.001) | | -0.0009 (-0.11)<br>(-0.003, 0.001) | -0.0022 (-0.14)<br>(-0.006, 0.001) | | -0.0014 (-0.08)<br>(-0.006, 0.003) |
| dchf | -0.0019 (-0.25)<br>(-0.015, 0.011) | | -0.0051 (-0.56)<br>(-0.019, 0.008) | -0.0089 (-0.55)<br>(-0.038, 0.018) | | -0.0152 (-0.77)<br>(-0.047, 0.013) |
| dgbp | 0.0120 (1.49)<br>(-0.012, 0.037) | | 0.0204 (1.80)<br>(-0.006, 0.047) | 0.0415 (2.33)<br>(-0.010, 0.095) | | 0.0562 (2.29)<br>(0.002, 0.109) |
| dnok | 0.0095 (1.14)<br>(-0.004, 0.024) | | 0.0190 (1.59)<br>(0.002, 0.037) | 0.0391 (2.12)<br>(0.005, 0.078) | | 0.0559 (2.14)<br>(0.017, 0.105) |
| dsek | 0.0176 (2.14)<br>(-0.005, 0.043) | | 0.0265 (2.28)<br>(0.025, 0.054) | 0.0591 (3.30)<br>(0.006, 0.115) | | 0.0742 (3.00)<br>(0.008, 0.141) |
| std.dev | 0.041 | 0.030 | 0.049 | 0.063 | 0.048 | 0.071 |
| $R^2$ | 0.12 | 0.23 | 0.08 | 2.23 | 0.3372 | 0.18 |
| AR-test | -0.08 (-1.37) | -0.04 (-0.41) | -0.10 (-1.21) | -0.13 (-1.63) | -0.11 (-0.94) | -0.15 (-1.35) |
| N-test | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0000 | 0.0270 |
| skewness | 1.54 | 2.52 | 1.30 | 1.03 | 2.35 | 0.71 |
| excess k. | 6.84 | 13.89 | 4.35 | 2.04 | 7.53 | 0.50 |

For the interest rate differential the estimated coefficient is given in the first row, the confidence interval for the coefficient based on the normal distribution in the second row and the bootstrapped confidence interval in the third row. For the dummy variables the estimated coefficient and the standard t-statistic are given in the first row and the bootstrapped confidence interval for the coefficient in the second row. The *Ar-test* shows the regression coefficient of the residuals on the first order lag of the residuals with the standard t-statistic in pharenthesis, $R^2$ is the coefficient of determination, *N-test* shows the p-value of the null hypothesis that the error term is normally distributed (low p-values indicate that the distribution of the error term deviates from the normal distribution), *st.dev* shows the estimated standard deviation of the error term, *skewness* shows to what extent the distribution of the error term is skewed to the right (positive values) or to the left (negative values), *excess k.* shows excess kurtosis, i.e., to what extent the distribution of the error term has longer tails than the normal distribution. For the normal distribution both the skewness and the excess kurtosis are equal to zero.

Tab (5): Average standard deviation of the percentage change of the exchange rate

| | ERM countries | Non-ERM countries |
|---|---|---|
| $\Delta e_{t-1}$ | 0.0092 | 0.016 |
| $\Delta e_{t-3}$ | 0.017 | 0.03 |
| $\Delta e_{t-6}$ | 0.025 | 0.044 |
| $\Delta e_{t-12}$ | 0.035 | 0.059 |

The table shows the standard deviation of $\Delta e_{t-s}$, where $\Delta e_{t-s}=\ln(E_t)-\ln(E_{t-s})$ for s=1, 3, 6, and 12 months, where E is the level of the nominal exchange rate.

Figure (1): Residuals from model (2), $\Delta e=\alpha_1+\beta_1(i-i^*)+u$, one month maturity

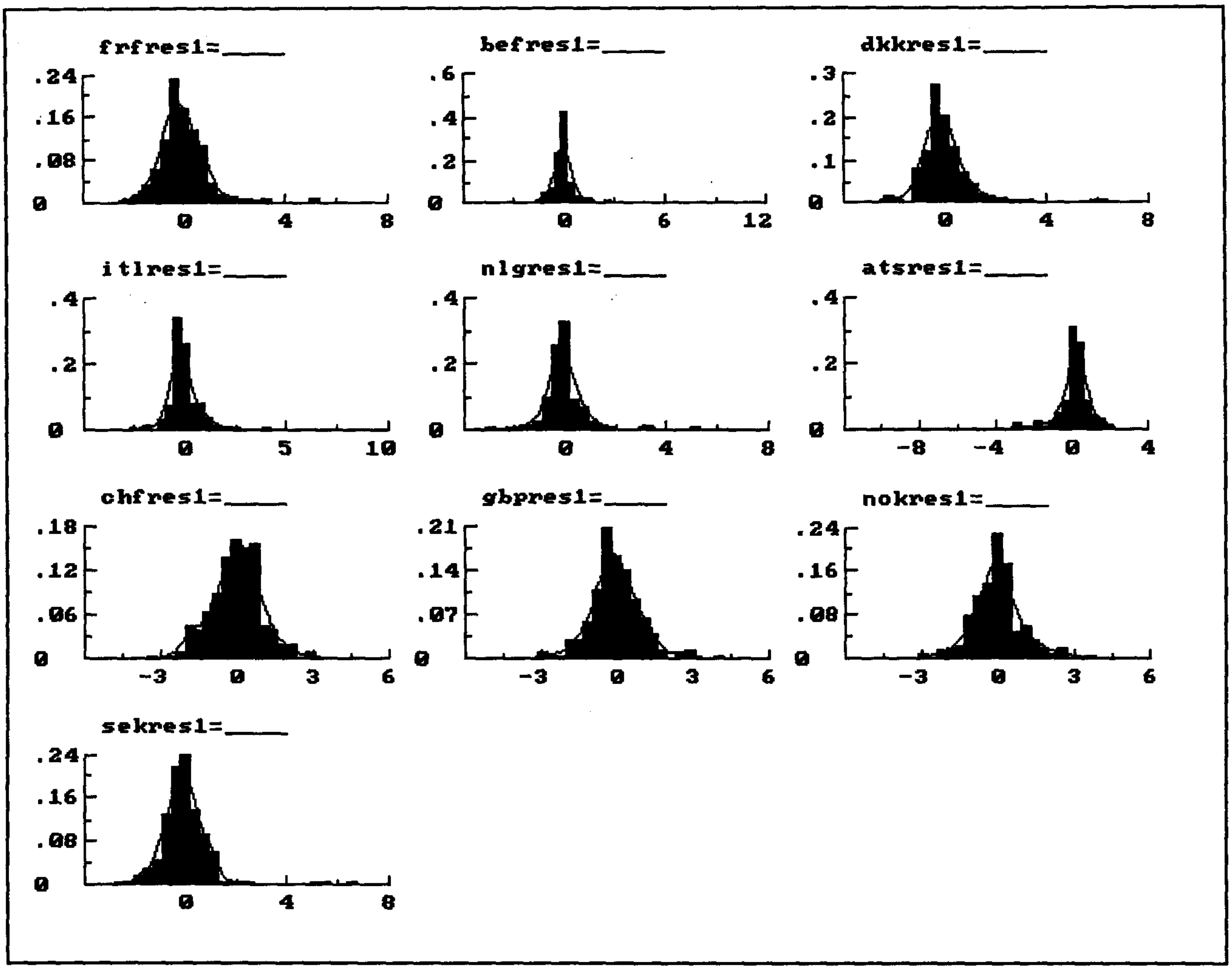
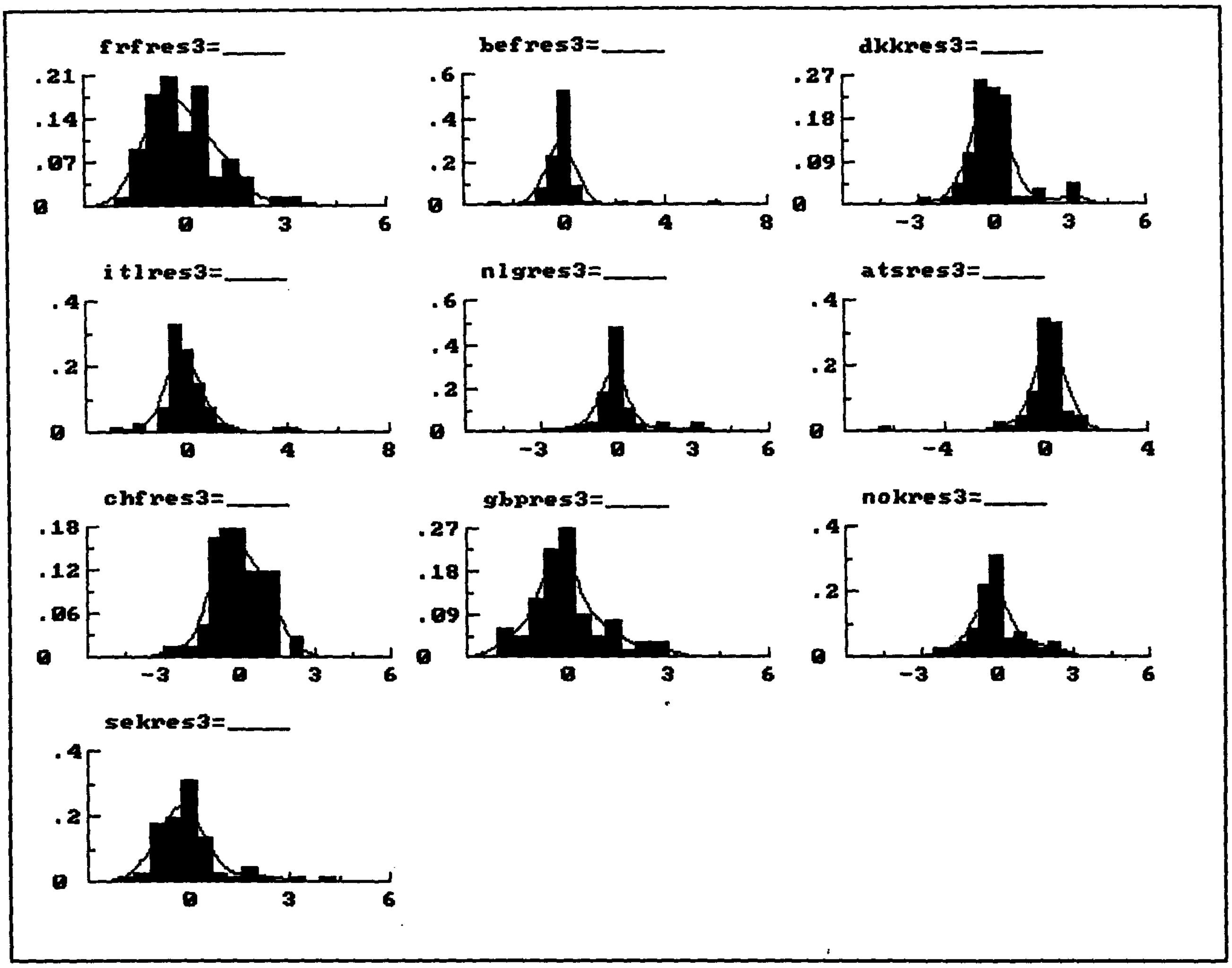


Figure (2): Residuals from model (2), $\Delta e=\alpha_1+\beta_1(i-i^*)+u$, three months maturity



ISSUED IN THE SERIES ARBEIDSNOTAT FROM NORGES BANK 1995-1997

1995/1 Bårdsen, Gunnar, Paul G. Fisher and Ragnar Nymoen
Business Cycles: Real facts of fallacies?
Research Department, 1995, 26 p.

1995/2 Bårdsen, Gunnar and Paul G. Fisher
The importance of being structured.
Research Department, 1995, 28 p.

1995/3 Oftedal, Knut Olav
Modellering av dagpengeordningen i en makroøkonomisk modell.
Utredningsavdelingen, 1995, 42 s.

1995/4 Mundaca,B. Gabriela
Probabilities of Realignments in a Currency Band: A Switching-Regime Approach.
Research Department, 1995, 48 p.

1995/5 Nilssen, Tore
On the Consistency of Merger Policy.
Research Department, 1995, 26 p

1995/6 Jansen, Eilev S. and Timo Teräsvirta
Testing Parameter Constancy and Super Exogeneity in Econometric Equations.
Research Department, 1995, 40 p.

1995/7 Bukh, Per Nikolaj D., Sigbjørn Atle Berg and Finn Førsund
Banking Efficiency in the Nordic Countries: A Four-Country Malmquist Index Analysis.
Research Department, 1995, 38 p.

1995/8 Eitrheim, Øyvind
The demand for broad money and tests for neglected monetary effects on inflations.
Empirical evidence for Norway 1969 to 1993.
Research Department, 1995, 128 p.

1996/1 Hammersland, Roger
The Structure of Exports.
An empirical analysis on Norwegian data.
Research Department, 1996, 65 p.

1996/2 Berg, Sigbjørn Atle
Central Bank Auctions of Deposit Certificates.
Research Department, 1996, 28 p.

1996/3 Ovesen, Vidar
Valutaspekulasjon innenfor et valutakursbåndregime og et styrt flytende kursregime.
Utredningsavdelingen, 1996, 70 s.

1996/4 Vale, Bent
Firm's Inventory Investments, Financial Conditions and the Banking Crisis in Norway.
Research Department, 1996, 27 p

1996/5 Bernhardsen, Tom
Devaluation Expectations and Macroeconomic Variables: A Critical Evaluation of the Literature.
Research Department, 1996, 38 p.

1996/6 Bernhardsen, Tom
A Test of Uncovered Interest Rate Parity for some EMS countries.
Research Department, 1996, 26 p.

1996/7 Mundaca, Gabriela B., Ole Bjørn Røste and Siri Valseth
The Dynamics of the Correlations between the short- and long-run Interest Rates in the Nordic Counties and Germany.
Research Department, 1996, 31 p.

1996/8 Mundaca, Gabriela B.
A Drift of the "Drift Adjustment Method".
Research Department, 1996, 13 p.

1996/9 Eika, Kari, Neil R. Ericsson and Ragnar Nymoen
Hazards in Implementing a Monetary Conditions Index
Research Department, 1996, 36 p.

1997/1 Akram, Qaisar Farooq og Espen Frøyland
Empirisk modellering av norske pengemarkeds- og obligasjonsrenter.
Forskningsavdelingen /Økonomisk avdeling, 1997, 34 s.

1997/2 Evjen, Snorre og Ragnar Nymoen
Har solidaritetsalternativet bidratt til lav lønnsvekst i industrien?
Forskningsavdelingen/Avdeling for finansielle instrumenter og betalingssystemer, 1997, 20 s.

1997/3 Nergård, Anita
Livsforsikringsselskaper - Regulering og risiko
Avdelingen for finansiell analyse og struktur (FIAS), 1997, 46 s.

1997/4 Lønning, Ingunn M.
Norsk penge- og valutapolitikk i lys av store oljeinntekter.
Forskningsavdelingen, 1997, 34 s.

1997/5 Mundaca, B. Gabriela and Jon Strand
Optimal exchange rate fluctuations under risk aversion and short-run wage rigidity.
Forskningsavdelingen, University of Oslo, 1997, 22 p.

1997/6 Berhnhardsen, Tom
The Relationship Between Interest Rate Differentials and Macroeconomic Variables: A Panel Data Study for European Countries.
Research Department, 1997, 46 p.

1997/7 Kim, Moshe and Bent Vale
Branch Banking in Dynamic Oligopoly.
Research Department, 1997, 26 p.

1997/8 Jenssen, Åsmund
Financial Market Imperfections and Inventory Investmand
An Empirical Study of Norwegian Manufacturing Firms 1991-1994
Research Department, 1997, 81 p.

1977/9 Bernhardsen, Tom
A Test of Uncovered Interest Rate Parity for Ten European Countries based on Bootstrapping and Panel Data Models.
Research Department, 1997, 31 p.