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The long-run exchange rate for NOK: a BEER approach*

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Abstract

This paper investigates a long-run relation for the trade weighted NOK exchange rate. I find that the NOK Trade Weighted Index (TWI) cointegrates with the real oil price, the price differential and the real interest differential. The paper documents a long-run solution for the TWI. The paper's main contribution is that the analysis is based on a test for cointegration that is robust to mixed orders of integration in the data.

The estimated long-run relation can be considered a benchmark for the nominal exchange rate. This interpretation allows the model to be used when analysing deviations of the nominal exchange rate from the model consistent level. The model is part of the suit of simple cross check models used when analysing the exchange rate in Norges Bank. I also find that the long-run relation is robust to the recent problems in the financial markets.

Keywords: *Time-Series Models, Financial Econometrics, Foreign Exchange*

JEL-Classification: *C32, C58, F31*

1 Introduction

Understanding the dynamics of the exchange rate is important because of the interactions between the (real) exchange rate and the rate of inflation and productivity in the short- to medium-run. Furthermore, there is a evidence that in

*I am grateful to Q. Farooq Akram, Tom Bernhardsen, Arne Kloster, Francesco Ravazzolo and my colleagues in the Market Operations and Analysis Department for helpful comments and discussion. The author alone is responsible for the views expressed in the paper and for any errors that may remain. Correspondence: geir.engesland.alstad@norges-bank.no. Address: Norges Bank, Bankplassen 2, P.O. Box 1179 Sentrum, Norway. Tel: + 47 22 31 60 35.

an environment of low and stable inflation, increasing the interest rate leads to a appreciation of the exchange rate.

This paper examines a specific type of model for the krone exchange rate based on macroeconic fundamentals and documents that the model forms a basis for the exchange over the medium- to long-run.[1] In order to make meaningful analyses of short-run exchange rate dynamics, it is necessary to establish a benchmark or long-run level. With an estimated long-run level for the exchange rate at hand, short-run deviations from this model can be quantified.

There is no consensus on how to estimate the long-run exchange rate. For example, the literature is split between whether the long-run exchange rate is constant or time varying. The most common assumption in the constant case is that purchasing power parity (PPP) forms the basis for the nominal exchange rate. In this case shocks to the real exchange rate (RER) will dissipate over time.

Empirical evidence on the validity of the long-run PPP hypothesis has been less negative for Norway than for other developed economies. In particular, it is often found that the krone RER has lower volatility and lower persistence in shocks relative to economic theory and many other currencies. Moreover, several authors have found that the krone real exchange is stationary. In general, these results point to the absence of the PPP-puzzle from the krone RER, see e.g. Habib and Kalamova (2007) and Akram et al. (2004) and the references therein.

The Fundamental Equilibrium Exchange Rate (FEER) methodology also assumes a constant equilibrium exchange rate albeit with adjustment towards the long-run along the steady state paths of the explanatory variables. An example of this, and the approach I will follow in this paper, is the Behavioural Equilibrium Exchange Rate (BEER). The BEER approach to exchange rate modelling assumes that deviations from the long-run (a weighted average of macroeconomic variables)

[1] Given certain assumptions, the model can also be used to discuss deviations from the long-run model consistent exchange rate.

is transitory. Over time, the nominal exchange rate will fluctuate around this weighted average.

The BEER methodology has been used frequently for analysing the long-run krone RER, see e.g. Akram (2000a), Bjørnland and Hungnes (2008) or Bjørnstad and Jansen (2007). Following the Johansen and Juselius (1992) (JJ) approach, these papers all assume joint long-run equilibrium in the goods market and capital market when estimating the long-run relation. Following the JJ methodology, the authors test for cointegration using the Johansen-test. These papers' results differ in that Bjørnland and Hungnes (2008) and Bjørnstad and Jansen (2007) find a single long-run relation between the RER and other macroeconomic variables, whereas Akram (2000a) find several relations between the variables considered.

A contribution of this paper is to apply a non-standard empirical test for cointegration and extend the set of explanatory variables used in these earlier studies. The cointegration test is based on Pesaran et al. (1999) (PSS). An advantage of the PSS-test relative to more common cointegration tests, e.g. the Johansen test or Engle and Granger test, is that the PSS-test is robust to the variables' order of integration. The existing empirical literature on the analysis of the krone RER has shown that there is considerable uncertainty surrounding the order of integration. In this case the traditional cointegration tests are invalid. Another advantage of the test relative to other single equation cointegration tests is that it estimates the long-run coefficients directly from the test regression.

Another contribution of the paper is that the model is estimated on monthly data. As such the model is better suited for analysing current events compared to models estimated on quarterly data since the data is updated frequently. A final contribution, is that the data set includes the recent financial crisis.

Using the PSS-test, I find evidence of a stable long-run relationship between the TWI, the real oil price and the interest differential.[2] As in previous studies, I find

[2] Technically, I reject non-cointegration between the variables.

that the implicit half-life for the NOK (given the model) is short. The estimated model is part of the "suit of models" used to analyse the exchange rate in Norges Bank, see Flatner et al. (2010) for a discussion.

2 The theoretical framework

The BEER approach, see e.g. MacDonald and Stein (1999) and Akram et al. (2004), is based on the real uncovered interest rate parity (UIP) ignoring possible risk premia. The real UIP hypothesis can be written as follows:

$$\mathbb{E}_k(q_{t+k}) - q_t = r_{t,k} - r^*_{t,k} \tag{1}$$

where q is the (log) RER, $r_{t,k}$ is the k-period real (domestic) interest rate and $r^*_{t,k}$ is the k-period real foreign interest rate. In other words, the *expected* k-period real rate of depreciation equals the current k-period real interest rate differential. Assuming that a linear long-run relation between the exchange rate and other macro variables exists, the expected real interest rate can be written as:

$$\mathbb{E}(q) = \mathbb{E}\left(\alpha + \boldsymbol{\beta}'\boldsymbol{M}\right) \tag{2}$$

where $\boldsymbol{M}$ represents a vector of macro variables, α is a constant term and β is a vector of coefficients.[3] If $\boldsymbol{\beta} = 0$, the long-run RER is constant (in expectation) and the model collapses to the PPP hypothesis. From equation (2) we see that the model-consistent exchange rate will be constant when the long-run forcing variables are at their equilibrium levels. In the BEER terminology, the deviation from this constant level is referred to as the *total* misalignment. Deviations from (2) at any point in time is called the *current* misalignment.[4]

By inserting equation (2) into (1), and using the law of iterated expectations, real

[3] I suppress the time indices for brevity.

[4] For a detailed description, see Akram (2004) and MacDonald and Stein (1999).

UIP can be stated as follows:

$$q = \alpha + \boldsymbol{\beta}'\mathbb{E}(\boldsymbol{M}) - (r - r^*) \tag{3}$$

Thus, based on the BEER methodology, the estimated RER will, in general, be time varying. However, given cointegration, misalignment from the model consistent level tends to dissipate over time. Which variables to include in $\boldsymbol{M}$ is largely an empirical question. Here, the real oil price (in NOK) and the unemployment gap are included.

Habib and Kalamova (2007) and Amano and van Norden (1998) argue that the real oil price is a leading indicator for the terms of trade in countries that are net exporters of oil. In this case, the currency tends to appreciate when the oil price increases and vice versa. They refer to such currencies as “petro-currencies”. The Norwegian krone is often considered to be a “petro-currency”.

The unemployment gap is included since it is often considered a proxy for the output gap, via Okun's law.[5] Thus, in our framework, the real oil price and unemployment gap act as anchors to the real economy in the long-run relation.

By substituting the real oil price and the unemployment gap into equation (3), the long-run deviation of the RER from its benchmark can be written as:

$$q = \alpha + \beta_1\mathbb{E}(op) + \beta_2\mathbb{E}(\tilde{u}) - (r - r^*) \tag{4}$$

where op is the real oil price expressed in kroner and $\tilde{u}$ represents the unemployment gap. By assuming rational expectations, using the definition of the RER ($q \equiv s - p + p^*$), and redefining the coefficients, I obtain the following expression

[5] Okun's law (in gap form) states that if output is below its potential level, the involuntary unemployment rate will tend to be positive, see e.g. Knotek (2007) for policy implications and practical uses of Okun's law.

for the deviation of the TWI from it's nominal benchmark

$$s = \gamma_0 + \gamma_1(p - p^*) + \gamma_2 op + \gamma_3 \tilde{u} + \gamma_4(r - r^*) + \omega \qquad (5)$$

where s is the log nominal exchange rate (an increase indicates a depreciation of the domestic currency), $p - p^*$ is the log price differential and ω is the long-run error term. ω is a function of expectational errors and effects of linearisation. If we cannot reject cointegration, we have have found evidence that the ω is orthogonal to the long-run solution. Hence, we can base analyses on the current misalignment of the nominal exchange rate on equation (5). If equation (5) defines a long-run relation for the nominal exchange rate, then one would expect that $\gamma_1 > 0$, $\gamma_2 < 0$, $\gamma_3 > 0$, and $\gamma_4 < 0$.

3 The econometric framework

The Pesaran et al. (1999) test for cointegration is essentially an application of an Autoregressive Distributed Lag model (ADL) model. An advantage of the PSS test relative to standard cointegration tests, e.g. the Johansen test, is that the long- and short-run coefficients are robustly estimated even in the case of different orders of integration among the variables. Given mixed orders of integration, the power of the traditional tests is essentially zero.

To my knowledge the PSS-test has not been employed when analysing the long-run krone exchange rate. As such, the current paper's main contribution is documenting a long-run relation for the krone exchange rate using a robust estimation technique and not relying on the non-credible assumption that the interest differential is non-stationary. Nor do I rely on the empirically ambiguous assumption of the krone RER being non-stationary.

The testing procedure can be described more formally as follows. Consider the general VEqCM

$$\Delta \boldsymbol{z}_t = \boldsymbol{\Pi} \boldsymbol{z}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{z}_{t-i} + \boldsymbol{\epsilon}_t \tag{6}$$

where $\boldsymbol{\Pi}$ is the long-run multiplier. Partition the vector $\boldsymbol{z}_t = (y_t, \boldsymbol{x}_t')'$ and assume that $\boldsymbol{\Pi}$ is upper triangular. This ensures that there is *at most* one long-run relation among the variables $\boldsymbol{z}_t$ that includes *both* y_t and $\boldsymbol{x}_t$ irrespective of the level of integration of $\boldsymbol{x}_t$. Under this assumption $\boldsymbol{x}_t$ is said to be *long-run forcing* for y_t.[6]

Under these assumptions, we may write the conditional equilibrium correction model (EqCM) as follows (see Pesaran et al. (1999) for details):

$$\Delta y_t = c_0 + \pi_{yy} y_{t-1} + \boldsymbol{\pi}_{yx.x} \boldsymbol{x}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\psi}_i' \Delta \boldsymbol{z}_{t-i} + \boldsymbol{\omega}' \Delta \boldsymbol{x}_t + u_t \tag{7}$$

where the subscript on $\boldsymbol{\pi}_{yx.x}$ indicates that all regressors are weakly exogenous to the long-run coefficients. Thus, we specify the null hypothesis as $H_0 : \pi_{yy} = \boldsymbol{\pi}_{yx.x} = \boldsymbol{0}'$, i.e. there exists no long-run relation between the variables. Pesaran et al. (1999) tabulate results from Monte Carlo experiments showing that their test is less prone to falsely *not* rejecting the finding of a stable long-run relation than the standard tests.[7]

In this specific application, the null hypothesis may be written as $H_0 : \gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = 0$. If the null is rejected, we can estimate the long-run coefficients and standard errors using the Baardsen semi-parametric method.[8]

The test essentially proposes estimating an unrestricted ADL model with a convenient lag length, and then testing whether the level variables are jointly zero, using critical bounds. The test calls for critical bounds in order to compensate for the lack of knowledge of the regressors' level of integration prior to the test. Just as the Phillips-Perron test, the null hypothesis is no long-run relation. If the test

[6] This does not preclude the y_t being Granger causal for $\boldsymbol{x}_t$ in the short-run.

[7] The distribution for the bounds test is non-standard and the critical values are calculated numerically.

[8] See Doornik and Hendry (2006) for details.

value falls below the bound, we cannot reject the null. On the other hand, if the test falls above the bound, we reject the null. Conversely, if the test statistic falls within the bounds, the test is inconclusive.

The underlying assumption is that the *level* of the regressand is weakly exogenous for the other equations in the system. In the language of Pesaran et al. (1999) the right-hand side variables are considered to be *long-run forcing.* The procedure is also robust to more than one cointegrating relations, provided that there is a single long-run relation including *all* variables.

Table 1 summarises the results of Augmented Dickey–Fuller tests for the variables. Similarly to Bjørnland and Hungnes (2008) and Bjørnstad and Jansen (2007), I cannot reject non-stationarity for the real interest differential. This could arise due to a non-stationary risk premium or non-stationary expected depreciation rate.[9] Any expected gain over and beyond the pure UIP hypothesis is defined as a risk premium. However, a non-stationarity risk premium or expected depreciation rate seems implausible from standard economic theory.[10] It is the speculative behaviour of these agents that drive the RER away from its historical benchmark values. This, according to Juselius (2006), invalidates univariate analysis of stationarity for the RER or real interest rate differential.

Furthermore, a time varying (non-stationary) risk premium implies that significant excess returns exist in the foreign exchange market and that these can be predicted using current information, see Sarno and Taylor (2001). Concerning the RER and oil price, these variables are often found to be non-stationary and table 1 indicates that they are in this sample as well. We note, however, that the krone RER has been found to be stationary in some papers.

Thus, the variables I consider appear to be a mix of $I(0)$ and $I(1)$ variables and

[9]For a discussion of the Norwegian krone and the risk premium, see Alendal (2010).

[10]In their 2007 paper, Bjørnstad and Jansen (2007) offer an interpretation of the interest differential as a proxy for a non-stationary risk premium. Frydman and Goldberg (2006) develop theoretical models in which this may occur due to risk averse and myopic agents basing forecasts on imperfect knowledge.

the standard tests do apply. Consequentially, the Pesaran et al. (1999) test seems appropriate. None of the combinations appear to be $I(2)$. Also, I reject non-stationarity for the unemployment gap at the 5 percent level. I conclude that the necessary conditions for the validity of the PSS procedure are satisfied.

Table 1. *Unit root (ADF) tests*

Variable	$I(2)$	$I(1)$	lags
Nominal TWI	-4.745**	-2.526	11
$p - p^*$	-4.091**	-2.314	8
$r - r^*$	-5.096**	-2.189	8
$oil^\dagger$	-7.705**	-2.320	7
$\tilde{u}$.	-5.647**	-2.874*	6
Real TWI	-5.065**	-2.598	10

† indicates a trend adjusted ADF test. ** and * indicate rejection of the null at the 1 and 5 percent level respectively.

4 The data

Our data set consists of 342 monthly observations from January 1982 to June 2010 inclusive. The set of regressors include the Norwegian exchange rate TWI, Norwegian CPI, a trade weighted measure of foreign CPI, the real Norwegian interest rate (based on the 12 month NIBOR rate), trade weighted interest rate, and the real oil price (1-month future contract, brent blend) expressed in Kroner. Unemployment is measured by the labour force survey (LFS).[11] All variables are in logs, except unemployment which is expressed in gap form using a Hodrick-Prescott (HP) filter with $\lambda = 180000$.

More precisely, the variables are defined as follows (* indicates trade weighted variables): $s = \log(\mathsf{TWI})$, $p = \log(\mathsf{CPI})$, $p^* = \log(\mathsf{CPI}^*)$, $r = \log[(1 + i/100)/(1 + \Delta p)]$, $r^* = \log[(1 + i^*/100)/(1 + \Delta p^*)]$, $op = \log(\mathsf{BRENT} \times \mathsf{USDNOK/CPI})$ and $\tilde{u} = \mathsf{LFS} - hp(\mathsf{LFS})$ where $hp(\mathsf{LFS})$ signifies the trend component in unemploy-

[11] The unemployment rate and price level variables are seasonally adjusted.

ment. We calculate the real interest rate by replacing expected inflation with realised inflation.[12]

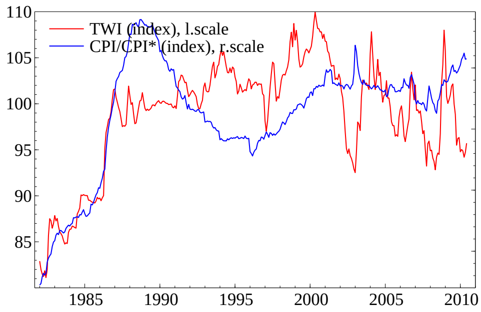

Figure 1. *TWI and price differential*

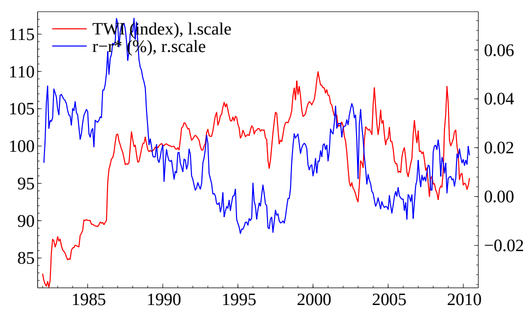

Figure 2. *TWI and interest rate differential*

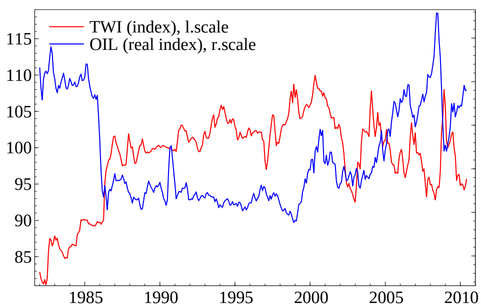

Figure 3. *TWI and oil price*

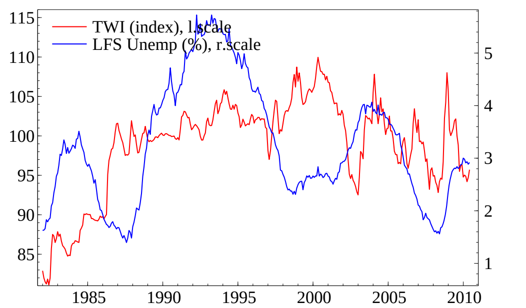

Figure 4. *TWI and LFS unemployment*

Figures 1–4 illustrate the time-evolution of the variables. Figure 1 indicates that higher price differential correlates with a weaker exchange rate. This is in line with the PPP-hypothesis and consistent with the results in Akram (2000a). However, concurrent with the introduction of inflation targeting the correlation appears to have weakened significantly. The current monetary policy regime is characterised by low and stable inflation and a freely floating exchange rate. Thus, the variance in the RER rate appears to be mainly attributed to the variance in the nominal exchange rate.[13]

Figure 2 indicates a structural break in the relation between the real interest

[12] Technically, I assume that the average forecasting error between actual and realised inflation is zero, i.e. $\mathbb{E}(\Delta_{12}p) - \Delta_{12}p \sim I(0)$ and $\frac{1}{12}\Delta_{12}p - \Delta p \sim I(0)$.

[13] See appendix A for a brief discussion of the evolution of the nominal exchange rate and inflation.

rate differential and the exchange rate after the inflation targeting regime was introduced. A higher interest rate differential seemed to correlate with a weaker currency during the fixed exchange rate regime. However, the converse is true after inflation targeting was introduced in 2001. The break is consistent with a freely floating, *credible* exchange rate regime. Specifically, in a regime with inflation targeting higher inflationary pressure and higher capacity utilisation typically leads to higher interest rates and an appreciation of the krone. Hence, a large interest rate differential is associated with stronger krone, as noted in Bjørnstad and Jansen (2007).

Figure 3 shows that the correlation between the real oil price and the exchange rate varies considerably over time. In the 1990s the correlation was low and the oil price was relatively stable. During the 2000s, however, increases in the oil price seems largely to have coincided with the appreciation of the krone. Nevertheless, during some periods the krone has appreciated while the oil price was falling.

In economic theory, higher oil price should lead to crowding out of the manufacturing sector and a real appreciation of the exchange rate. However, the real appreciation of the krone is mitigated by the gradual phasing in of the wealth generated by the petroleum sector, and the sales of kroner on behalf of the Government Pension Fund Global.[14] Without this mechanism the effect of the oil price on the exchange rate would probably be greater. Nevertheless, the real oil price seems to coincide with stronger exchange rate during the last decade, suggesting that the variable should be included in the analysis.

Finally, figure 4 shows the evolution in the exchange rate and unemployment. A stable long-run relation does not appear to exist. Nevertheless, since 2001 there is some indication that the exchange rate appreciates when the unemployment gap is falling and vice versa. Furthermore, if Okun's law holds, we would also expect that

[14] This is the so-called "Handlingsregelen" described here (Norwegian only): `http://bit.ly/9SJjHE`. The sales of NOK on behalf of Government Pension Fund Global and its effect on the krone is described in Fidjestøl (2007).

the unemployment gap is stationary. The figure indicates that the unemployment rate is fluctuating around 3 percent over the sample and corroborates the finding of stationarity from table 1.

5 The long-run relation

5.1 The existence of a long-run relation

I focus on the sub-sample January 1983 – June 2010 omitting 12 lags for sequential lag reduction. Based on the PSS-test I find evidence of a stable long-run relationship for the krone exchange rate.[15] The results are summarised in table 2.

Table 2. *Summary of the bounds test*

Variable	F-statistic	t-statistic	p	Long-run
$F_s\ [s\vert(p-p^*), op, (r-r^*), \tilde{u}]$	4.56	-4.69	2	Exists
$F_{(p-p^*)}\ ((p-p^*)\vert s, op, (r-r^*), \tilde{u}]$	3.88	-2.66	3	Inconclusive
$F_{(r-r^*)}\ [(r-r^*)\vert s, (p-p^*), op, \tilde{u}]$	3.20	-3.71	3	Inconclusive
$F_{\tilde{u}}\ [\tilde{u}\vert s, (p-p^*), op, (r-r^*)]$	3.01	-2.32	6	Not Exist

Appendix B tabulates the relevant tables at the 5 and 10 percent level.

The reported F-statistics indicate that there exists a long-run relation between the nominal exchange rate, the price differential, the real oil price, the real interest differential and the unemployment gap. From Pesaran et al. (1999) table C1.iii with $k = 4$, the upper F-critical value is 4.01 at the 5 percent level. The corresponding 5 percent t-critical value is 3.99.[16] I reject the null of no long-run solution for both versions of the test.

The existence of a long-run relation between the variables allows us to test the statistical significance of the long-run regressors using the standard t-test. I find

[15]Following Pesaran et al. (1999), the testing procedure I follow in this paper is summarised as follows: starting from a general ADL(12), choose a suitable EqCM formulation. The model selection is based on the Akaike Information Criterion (AIC). From this parsimonious model I test for a single unique cointegrating relation amongst the variables.

[16]See Appendix B for tabulated critical values.

that the long-run coefficient on the unemployment gap is not statistically different from zero. Restricting the coefficient to zero and comparing the test regressions with a standard log-ratio test, I conclude that there is no long-run relation between these variables that includes the unemployment gap.[17] Table 2 summarises the results. This finding is similar to the results in Bjørnstad and Jansen (2007), Bjørnland and Hungnes (2008), and Bernhardsen (2008). However, I do not rely on implicitly assuming non-stationarity of the interest rate differential.

5.2 Estimated long-run parameters

The estimated long-run exchange rate is as follows (standard errors underneath, estimated by Baardsens algorithm).[18]

$$s_t = \underset{(1.0755)}{4.5943} + \underset{(0.2383)}{0.6107} (p - p^*)_t - \underset{(0.4540)}{1.4292} (r - r^*)_t - \underset{(0.0201)}{0.0573} op_t \quad (8)$$

All coefficients have the anticipated signs.[19] Furthermore, the estimated size of the coefficients are approximately in line with economic theory. The 95 percent confidence interval on the price differential coefficient contains unity, in line with the PPP hypothesis. Also, the coefficient on the interest rate differential is -1 at the 5 percent level, in line with real UIP.[20] The results are similar to Akram (2004) using quarterly data. Bjørnland and Hungnes (2008) interpret the coefficient on the interest differential as representing the degree of international capital mobility relative to the elasticity of net exports. Such an interpretation implies a high degree of capital mobility for the Norwegian economy. This interpretation is consistent with Norway being a small open economy.

[17] The test statistic is $\chi^2_{(3)} = 2.90$.

[18] The model passes most specification tests, only showing signs of non-normality. Results are available from the author upon request. For an explanation of the Baardsen algorithm, see Doornik and Hendry (2006).

[19] The long-run standard error (σ_{LR}) was estimated as 0.16.

[20] Note, since my model is a reduced-form, interpretations of the coefficients must be done with care.

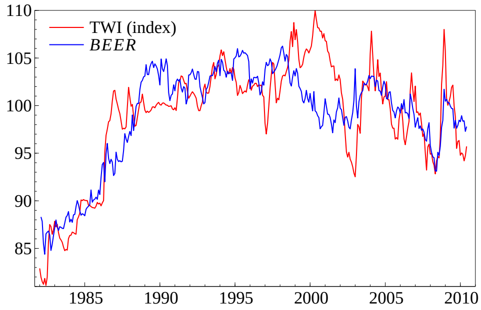

Figure 5. *TWI and estimated long-run TWI*

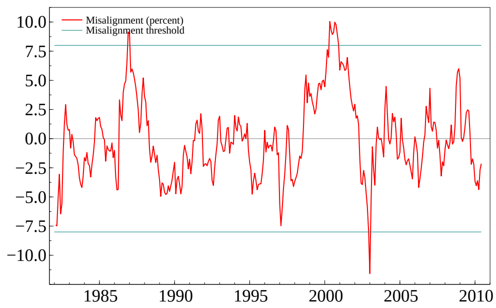

Figure 6. *Misalignment of the exchange rate (percent).*

Finally, the coefficient on the real oil price suggests that a 10 percent increase in the real oil price leads to approximately 0.6 percent appreciation in the TWI. Higher oil prices leads to higher natural wealth and raises demand, which in turn can only be met if the currency experiences real appreciation. Alternatively, an increase in the price of oil leads to an improvement of the current account causing a long-run appreciation of the exchange rate.

Figure 5 shows the estimated long-run exchange rate based on equation (8) and the TWI over the estimation period. The figure shows that the nominal exchange rate does not meander too far from the estimated long-run. Thus, shocks to the nominal exchange rate seems to dissipate quickly. A simple half-life calculation suggests that 50 percent of a shock to the nominal exchange rate is removed after approximately 7 months.[21] This short half-life is corroborated in several papers, but is often considered a stylised fact for developing countries with high and volatile rates of inflation rather than a highly developed industrialised country like Norway, see Habib and Kalamova (2007) and Akram (2000a) and the references therein. Furthermore, exchange rates that have short half-lives and relatively low volatility do not display the stylised characteristics of the PPP-puzzle.

Including the real oil price in the long-run relation helps explain the sustained appreciation of the krone from 2003 to late 2008. The financial turmoil which

[21] The estimated adjustment parameter is $\alpha = -0.092$.

began in late 2008 is captured well by the model. A general point from figure 5 is that the model appears to react slowly to sudden movements of the exchange rate, especially sudden depreciations, but captures the long term swings rather well.

Figure 6 shows that there are periods of large misalignments during the first half of the 2000s. The maximum deviation over the sample period is ± 10 percent. Focusing on the period from 2008, the volatility of the deviations does not seem excessive relative to the earlier period. As discussed in MacDonald and Stein (1999), a significant misalignment threshold is considered to be 1 standard deviation from zero. The implied long-run standard deviation is found to be 8 percent in the current exercise. In June 2010, the current misalignment was approximately -2 percent. Judging by figure 6, we see that the nominal exchange rate has only shown significant misalignment on three occasions over the sample. All instances resulted in sharp corrections of the nominal exchange rate.

5.3 Stability of the long-run parameters

Ascertaining the stability of the long-run coefficients is important when checking for the robustness of the model. Figures 7–10 show recursive estimates for the long-run coefficients. Focusing on the latter part of the sample we observe that the financial crisis did have an impact on the coefficients but the effects were short lived. Furthermore, the long-run relationship appears relatively robust to the financial distress of 2008 and later.[22] Throughout the period of inflation targeting, the underlying cointegrating relationship appear rather robust.

In particular, the constant term has been relatively stable over the last decade, indicating no structural break. Clements and Hendry (2001) argue that breaks in the constant term are most harmful to the statistical performance of EqCM-type models.

[22] Flatner et al. (2010) describe a version of this model based on a sample ending in August 2008, inspection of the coefficients reveal that they are not statistically different from the model estimated here.

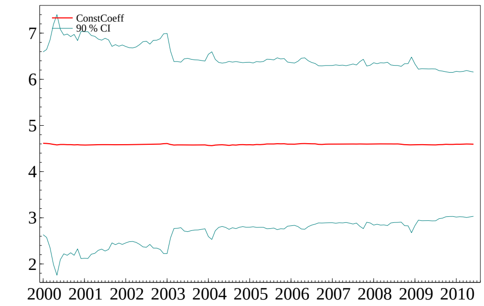

Figure 7. *Recursive estimates of the constant.*

Figure 8. *Recursive estimates on the price differential coeff.*

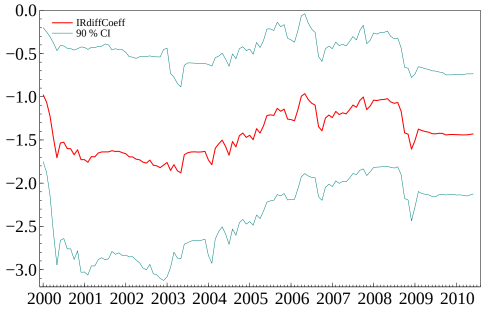

Figure 9. *Recursive estimates on the real interest differential coeff.*

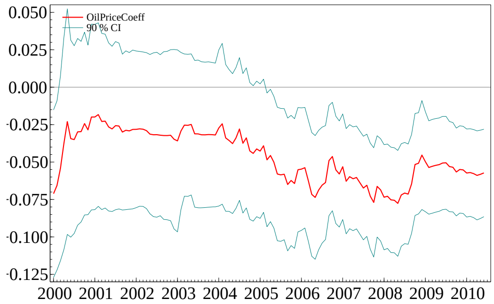

Figure 10. *Recursive estimates on the real oil price coeff.*

From figure 10 we see that the real oil price does not enter the long-run relation significantly from 2000 through 2003. Several authors, e.g. Bjørnland and Hungnes (2008) and Bjørnstad and Jansen (2007), find that the real price of oil is only marginally significant over the sample period. One explanation for this result could be attributed to the oil price fluctuating between 20 and 34 dollars per barrel from 2000 to 2004, whereas it has ranged between 31 and 135 USD during the last six years. Moreover, it is possible that the long-run effect of the (real) oil price on the Norwegian krone is non-linear.[23] Nevertheless, since 2004 the estimated long-run effect of oil on the krone exchange rate has been relatively stable and statistically significant.

[23] For an application of a non-linear effect of the oil price on the krone exchange rate, see Akram (2000b).

6 Conclusion

The analysis presented in this paper indicates that there exists a long-run relationship between the real exchange rate, the real oil price and the real interest rate differential. Moreover, the results corroborates the relatively common finding of a long-run benchmark for the trade-weighted NOK exchange rate. I document a finding of a single long-run relationship among several macroeconomic variables that nests long-run PPP as a special case.

Based on the estimated model we may analyse whether the exchange rate is expected to depreciate or depreciate going forward, based on current misalignment. The current misalignment can be joined with ad-hoc data in order to ascertain whether the exchange rate is under or overvalued at any given time.

A Inflation regimes in Norway

The data, can be divided into (at least) two distinct periods. The first, January 1982 – December 1996, is characterised by high (albeit falling) inflation. The second period, January 1997 – June 2010, is characterised by low and stable inflation. Figure 11 illustrates. The figure also shows that there is a pronounced increase in the 12-month volatility of the NOK post January 1997 and that the persistence in the inflation rate is much higher than for the exchange rate. Figures 1 and 11 suggest that the volatility in the RER is caused by the volatility in the nominal exchange rate over the sample.

From table 12 and 13 we observe that the volatility of the interest rates has halved post January 1997. This could be a result of inflation stabilising after the introduction of a floating exchange rate.

Figure 11. *Annual CPI-ATE inflation and the exchange rate*

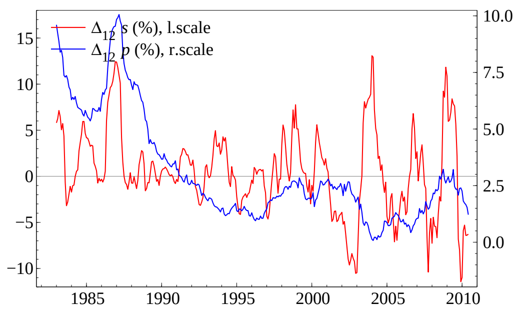

Figure 12. *Correlation matrix February 1982 – December 1996*

	Δs	Δp	Δp*	Δr	Δr*	Δop	$\Delta\tilde{u}$	σ
Δs	100.00 %	—	—	—	—	—	—	0.74 %
Δp	9.50 %	100.00 %	—	—	—	—	—	11.60 %
Δp*	-4.38 %	20.72 %	100.00 %	—	—	—	—	14.41 %
Δr	2.77 %	12.78 %	-11.44 %	100.00 %	—	—	—	-2.94 %
Δr*	-18.15 %	-5.41 %	7.33 %	27.29 %	100.00 %	—	—	2.09 %
Δop	12.04 %	-10.73 %	18.57 %	-4.99 %	9.83 %	100.00 %	—	-1.95 %
$\Delta\tilde{u}$	0.74 %	11.60 %	14.41 %	-2.94 %	2.09 %	-2.68 %	-1.95 %	100.00 %

Figure 13. *Correlation matrix January 1997 – June 2010*

	Δs	Δp	Δp*	Δr	Δr*	Δop	$\Delta\tilde{u}$	σ
Δs	100.00 %	—	—	—	—	—	—	1.48 %
Δp	-2.86 %	100.00 %	—	—	—	—	—	0.37 %
Δp*	-24.22 %	35.56 %	100.00 %	—	—	—	—	0.15 %
Δr	-18.22 %	3.36 %	31.40 %	100.00 %	—	—	—	0.28 %
Δr*	-3.17 %	7.05 %	37.34 %	65.60 %	100.00 %	—	—	0.17 %
Δop	-15.81 %	20.70 %	49.83 %	19.70 %	33.09 %	100.00 %	—	8.72 %
$\Delta\tilde{u}$	-6.99 %	-0.95 %	-11.20 %	-28.98 %	-32.01 %	-2.68 %	100.00 %	0.09 %

B Tabulated critical values for the PSS test

For ease of reference, table 14 and 15 tabulates a truncated version of Pesaran et al. (1999) table C1.iii: Case III and C2.iii: Case III. The row highlighted in bold font represents the critical values for the test used in the paper. For the complete tables, please see the reference.

Figure 14. *Critical values for the PSS* F-*test*

	90 %		95 %	
k	*I*(0)	*I*(1)	*I*(0)	*I*(1)
1	4.04	4.78	4.94	5.73
2	3.17	4.14	3.79	4.85
3	2.72	3.77	3.23	4.35
4	2.45	3.52	2.86	4.01
5	**2.26**	**3.35**	**2.62**	**3.79**

Figure 15. *Critical values for the PSS* t-*test*

	90 %		95 %	
k	*I*(0)	*I*(1)	*I*(0)	*I*(1)
1	-2.57	-2.91	-2.86	-3.22
2	-2.57	-3.21	-2.86	-3.53
3	-2.57	-3.46	-2.86	-3.78
4	-2.57	-3.66	-2.86	-3.99
5	**-2.57**	**-3.86**	**-2.87**	**-4.19**

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