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## PPP in the Medium Run Despite Oil Shocks: The Case of Norway

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PPP in the medium run despite oil shocks: The case of Norway

by

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# PPP in the medium run despite oil shocks: The case of Norway

Qaisar Farooq Akram\*

## Abstract

Existing studies generally reject purchasing power parity (PPP) on datasets from countries that have been affected by large real shocks, including Norway. However, we offer strong evidence of PPP between Norway and its trading partners during the post-Bretton Woods period, in which the Norwegian economy has experienced numerous real shocks such as discoveries of large petroleum reserves and oil price shocks. In particular, the behaviour of the Norwegian real and nominal exchange rates appears remarkably consistent with the PPP theory. Moreover, convergence towards PPP is relatively fast; the half-life of a deviation from parity is just about 1.5 years. We show that such deviations are eliminated by adjustments in the nominal exchange rate and we offer some explanations for the relatively fast convergence towards PPP.

**Key words:** *PPP, real exchange rate, cointegration analysis.*

**JEL Classification:** *C22, C32, C51, F31, F41.*

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## 1. Introduction

A number of recent empirical studies observe convergence towards purchasing power parity (PPP) in the long run, see e.g. Froot and Rogoff (1994), Rogoff (1996), Isard (1995) and MacDonald (1995). Accordingly, changes in nominal exchange rates outweigh changes in domestic prices relative to foreign prices in the long run, and real exchange rates exhibit reversion towards their constant equilibrium rates. However, the speed of reversion is reported to be relatively slow; estimates of the half-life of a deviation from an equilibrium level vary in the range of 3 to 6 years for industrial countries. Another common finding is that support for long-run PPP is stronger in data samples dominated by monetary shocks than in samples presumably dominated by real shocks, such as discoveries of natural resources, see e.g. Patel (1990) and Cheung and Lai (2000a).<sup>1</sup> In the latter type of samples, real exchange rate behaviour is often indistinguishable from a random walk. Also, support for PPP is often stronger in studies that employ wholesale prices, with a larger share of prices for tradables, rather than consumer prices. We present novel results against such a background.

We test for PPP between Norway and its trading partners by examining the behaviour of both real and nominal Norwegian exchange rates, and relative consumer prices. In addition, we investigate the extent to which deviations from parity are eliminated by adjustments in the nominal exchange rate and consumer prices. We apply standard time series techniques, i.e. the augmented Dickey Fuller (ADF) unit root test and the multivariate cointegration method of Johansen (1995), on quarterly data from the post-Bretton Woods period. In this period, the Norwegian economy has experienced numerous real shocks such as discoveries of huge oil and

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<sup>1</sup>Such shocks affect the national wealth and the foreign exchange earning potential of a country and thereby lead to substantial changes in the economic structure. Such changes are usually initiated and accompanied by changes in the relative price of traded goods to non-traded goods, which may lead to large deviations in aggregate prices across countries, see e.g. Corden (1984) and other references on the so called “Dutch disease”.

gas reserves and large fluctuations in (real) oil prices. We nevertheless find strong support for PPP, upheld by a remarkably stable equilibrium real exchange rate during the sample period. Moreover, deviations from PPP are eliminated relatively fast. The estimated half-life of a given deviation from the equilibrium rate is around 1.5 years, which is substantially below e.g., the median of half-life estimates for industrial countries reported by Cheung and Lai (2000a), which is 3.3 years.

Furthermore, our results seem to contradict the existing evidence on PPP between Norway and its trading partners, in particular evidence based on standard time series techniques. The existing studies report rejection of PPP between Norway and its trading partners, irrespective of whether they employ effective or bilateral exchange rates and whether they are based on annual, quarterly or monthly data, see e.g. Bahmani-Oskooee (1995), Jore et al. (1998), Chortareas and Driver (2001), Taylor (2001a) and Papell (1997).<sup>2</sup> Taylor (2001a), however, reports evidence of PPP between Norway and the USA, and a half-life of 2.7 years when he employs the generalised-least-square version of the Dickey Fuller (DF-GLS) test proposed by Elliot et al. (1996) on annual data. However, this estimate of half-life may be biased upward due to temporal aggregation bias. Taylor (2001b) shows this to be the case if e.g. annual data are used when the (true) adjustment horizon is of the order of quarters or months. The contrast between our results and those from the existing studies employing standard techniques may be largely due to differences in the time span of the data and to model formulations.

This paper is organised as follows: Section 2 examines the dataset in the light of the PPP theory and tests whether the Norwegian effective real exchange rate is an equilibrium-reverting process with a constant equilibrium rate.<sup>3</sup> In particular, Sub-

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<sup>2</sup>Papell (1997) does not reject the null hypotheses of a unit root in the bilateral real exchange rates between Norway and the USA, and between Norway and Germany on quarterly datasets at the 10% level of significance. However, this result for the bilateral real exchange rate between Norway and the USA is not supported on his monthly dataset.

<sup>3</sup>All empirical results and graphs are obtained using PcGive 9.10, PcFiml 9.10, GiveWin 1.24

section 2.2 carries out a sensitivity analysis of our findings by briefly examining the time series behaviour of bilateral real exchange rates for Norway's main trading partners the UK, Germany and the USA. Section 3 tests explicitly whether domestic and foreign prices have symmetrical and proportional effects on the Norwegian effective nominal exchange rate, as implied by the PPP theory. This section also examines the response of the nominal exchange rate and prices to deviations from parity. In particular, Subsection 3.2 reports the outcome of a comprehensive sensitivity analysis of our findings and points out their robustness in the face of extensions of our information set by additional variables and observations, and changes in model formulation. Section 4 endeavours to account for the relatively lower persistence of the Norwegian real exchange rate compared with that of other industrial countries. Section 5 concludes and the appendix presents precise definitions of variables, their sources and graphs of the bilateral nominal and real exchange rates, and consumer prices.

## **2. Data and tests of PPP in a univariate framework**

Norway aimed at fixed exchange rate arrangements with its trading partners (mainly western European countries and the USA) in the period 1972–1997, which covers our sample, see e.g. Alexander et al. (1997) for details. However, the nominal exchange rate fluctuated as a result of market pressure and official adjustments. In the period 1972–1986, Norway devalued a number of times to counteract deteriorating competitiveness and devaluation pressure; the last devaluation was in May 1986, see Norges Bank (1987). Since then, the nominal exchange rate has been relatively stable and fluctuations have been induced by (other) factors affecting the foreign exchange market.

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and 2.02, and PcGets 1.0, see Hendry and Doornik (1996), Doornik and Hendry (1996) and (1997) and Hendry and Krolzig (2001).

In the following we focus on the effective nominal and real exchange rates and foreign consumer prices, while Subsection 2.2 presents results for the bilateral real exchange rates. As Norwegian trade is not dominated by a single country, the effective exchange rate seems a better measure of Norway's competitive position against its trading partners.

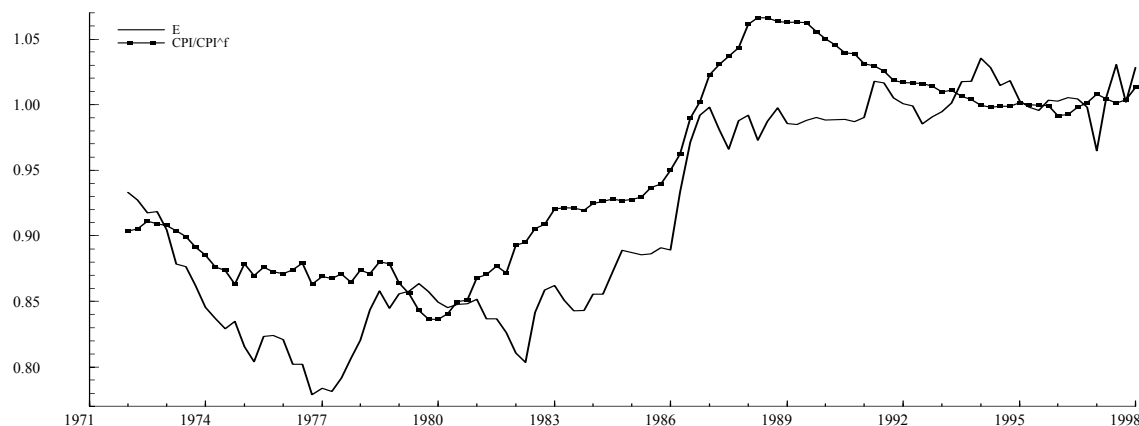


Figure 2.1: *The nominal exchange rate,  $E$ , (solid line) and relative consumer prices between Norway and trading partners,  $CPI/CPI^f$ , (boxed line).*

Figure 2.1 plots the Norwegian effective nominal exchange rate ( $E$ ) and the Norwegian consumer price index ( $CPI$ ) relative to a trade weighted index of foreign consumer prices ( $CPI^f$ ) over 1972:1–1997:4. The exchange rate reflects units of domestic currency per unit of foreign currency; the price data is not seasonally adjusted and the base year is 1995. The precise definitions of the variables and sources are presented in the appendix.

The overall impression obtained from Figure 2.1 is that the nominal exchange rate does not evolve independently of the relative consumer price ( $CPI/CPI^f$ ). A few exceptions to this tendency are the relatively strong exchange rate appreciation until 1976/77, which is not matched by a comparable fall in the relative consumer price, and the relative stability in the exchange rate during 1978–1981 and in particular 1990–1992. In these periods, the Norwegian krone was pegged to a currency



basket and to the European Currency Unit (ECU), respectively. Since 1992, the exchange rate has evolved around the relative consumer price, which may be interpreted as the equilibrium level of the nominal exchange rate in the PPP framework.

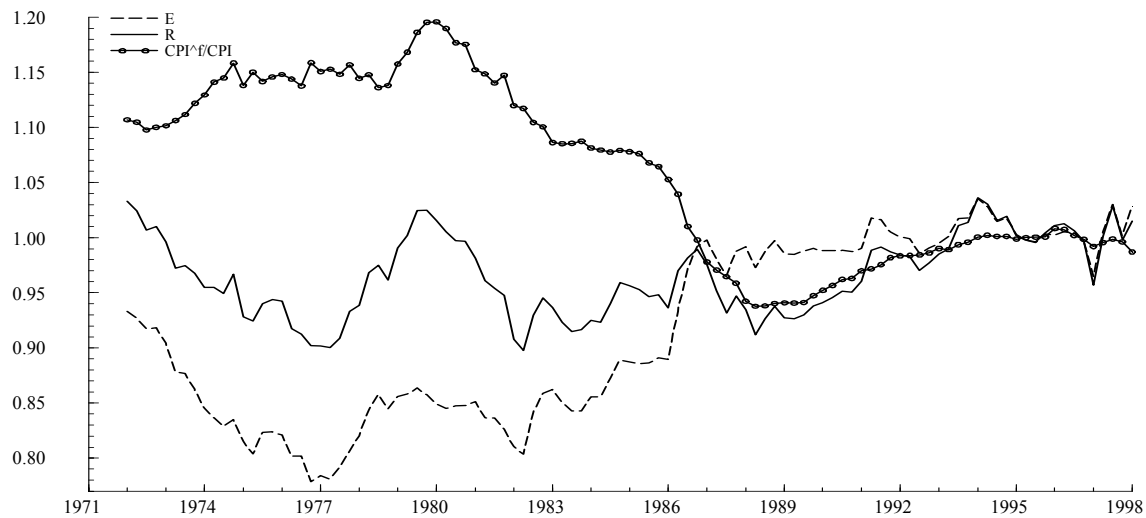


Figure 2.2: *Nominal exchange rate,  $E$ , (dashed line), real exchange rate,  $R$ , (solid line) and relative consumer prices between trading partners and Norway,  $CPI^f/CPI$  (circled line).*

Figure 2.2 suggests that changes in the nominal exchange rate and the relative consumer price (now defined as  $CPI^f/CPI$ ) tend to outweigh each other, especially before the 1990s. Thus fluctuations in the real exchange rate,  $R \equiv E(CPI^f/CPI)$ , have a smaller range than the fluctuations in the nominal exchange rate and the relative consumer price. In the late 1980s and the early 1990s, the nominal exchange rate is relatively stable and movements in the real exchange rate are mainly driven by the relative consumer price. The opposite is the case after 1992/93, when most of the fluctuations in the real exchange rate can be entirely ascribed to fluctuations in the nominal exchange rate, as the relative consumer price is quite stable. In particular, the sharp real appreciation in 1997:1 and the subsequent real depreciation are due to nominal appreciation and depreciation, respectively.

Figure 2.2 gives the impression that the real exchange rate is likely to have evolved around a constant level without an increasing fluctuation range over time. Such behaviour seems to be inconsistent with the random walk hypothesis for real exchange rates, cf. Mark (1990). The next subsection tests the random walk hypothesis for the real exchange rate against the alternative of a (weakly) stationary time series that fluctuates within a given range and reverts towards a constant equilibrium rate.

### 2.1. Random walk or equilibrium reverting process?

The PPP theory implies that the real exchange rate ( $R$ ) evolves around a constant equilibrium level,  $\gamma$ , over time. This can be formalised as follows:

$$R_t = \gamma + \sum_{i=1}^p \psi_i (R_{t-i} - \gamma) + \varepsilon_t. \quad (2.1)$$

Here  $\varepsilon_t$  is the error term, assumed to be identically, independently normally distributed with zero mean and constant variance,  $\sigma^2$ , i.e., IIDN(0,  $\sigma^2$ ).

The (absolute) value of  $\sum_{i=1}^p \psi_i$ , hereafter denoted  $\varrho$ , should be less than 1 for  $R$  to converge towards  $\gamma$  after a shock.  $R$  follows a random walk process if  $\varrho$  equals 1, in which case every shock has a permanent effect on  $R$ .

The augmented Dickey Fuller (ADF) test can be used to test the null hypothesis of  $\varrho$  equals 1, against the alternative hypothesis of  $\varrho$  equals less than 1, see e.g. Banerjee et al. (1993, ch. 4) for details. In order to separate out the unobservable  $\gamma$  from the actual rate  $R$ , we rephrase equation (2.1) in the ADF framework as follows, with the constant term  $\alpha$  defined as  $\gamma(1 - \varrho)$ :

$$\Delta R_t = \alpha - (1 - \varrho)R_{t-1} + \sum_{i=1}^{p-1} \psi_i \Delta R_{t-i} + \varepsilon_t. \quad (2.2)$$

This equation was initially formulated and estimated with a quite generous lag

length (p) equal to 9. However, given that an ADF test tends to loose power in the presence of redundant lags, we sequentially eliminated statistically insignificant terms of  $\Delta R_{t-i}$  from the model to minimise Akaike's information criteria (AIC). Accordingly, following terms were excluded  $\Delta R_{t-8}$ ,  $\Delta R_{t-6}$ ,  $\Delta R_{t-4}$  and  $\Delta R_{t-2}$ ; joint zero restrictions on their coefficients were accepted by an  $F$ -test at a  $p$ -value of 0.74. Table 2.1 sets out a parsimonious version of the general model, obtained by (sequential) omission of these terms and diagnostic test statistics. The parsimonious ADF model includes  $\Delta R_{t-7}$ ,  $\Delta R_{t-5}$  and  $\Delta R_{t-3}$ , which are statistically insignificant at the 5% level, since their exclusion increased the AIC-value.

The outcome of the ADF test is consistent with the PPP theory as the null hypothesis is rejected at the 1% level. Thus, the real exchange rate may be considered as an equilibrium reverting process. The derived estimate of the equilibrium level is  $0.161/0.167 \approx 0.96$ , see Table 2.1. The degree of equilibrium reversion is 0.167, which implies that the half-life of a disequilibrium is less than 4 quarters, when calculated by the commonly used formula:  $\ln(0.5)/\ln \hat{\varrho}$ , see e.g. Taylor (2001b, p. 474). However, impulse response analysis, which also takes into account all the dynamic terms in the ADF model, implies a half-life of about 7 quarters, see Figure 2.4.

The diagnostics of the estimated equation do not indicate systematic structure in the residuals, increasing the reliability of the coefficient estimates. However, due to the sharp real exchange rate fluctuations in 1997, the normality assumption is violated at the 1% level of significance. Inclusion of an impulse dummy that is 1 in 1997:1,  $-1$  in 1997:2 and zero elsewhere did not alter the conclusions.<sup>4</sup> Notably, the absolute value of  $t$ -ADF increased to 3.840, which strengthens our evidence against the null hypothesis a unit root in the real exchange rate.

The equilibrium real exchange rate  $\gamma$  seems to be remarkably stable over time. Figure 2.3 displays recursive OLS estimates of the equilibrium real exchange  $\hat{\gamma} \pm 2SE$

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<sup>4</sup>Limiting distribution of the DF-test statistics is not affected by allowance for such centered impulse dummies.

Table 2.1: A univariate model of the Norwegian real exchange rate

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$\Delta \hat{R}_t =$	0.161 (3.673)	$-$	0.167 (-3.681)	$R_{t-1} +$	0.212 (2.133)	$\Delta R_{t-1} +$	0.156 (1.491)	$\Delta R_{t-3}$
				$+ 0.156 \Delta R_{t-5} +$	0.202 (1.845)	$\Delta R_{t-7}$		
<i>Sample: 1972:2–1997:4, 103 Quarterly observations.</i>								
<i>t – ADF = –3.681, DF-critical values: 5% = –2.887, 1% = –3.489</i>								
<hr/>								
<b>Diagnostics</b>								
$R^2$				$=$	0.144			
Standard error of residuals: $\hat{\sigma}$				$=$	0.015			
Durbin Watson statistic: DW				$=$	1.98			
Autocorrelation 1-5: $F_{ar,1-5}(5, 92)$				$=$	0.92[0.48]			
ARCH 5: $F_{arch,1-5}(5, 87)$				$=$	1.10[0.36]			
Normality: $\chi^2_{nd}(2)$				$=$	7.1[0.03]*			
Heteroscedasticity: $F_{Xi^2}(10, 86)$				$=$	1.28[0.26]			
Heteroscedasticity: $F_{XiXj}, (20, 76)$				$=$	1.12[0.35]			
Model specification: RESET $F(1, 96)$				$=$	0.02[0.90]			

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Note: Ordinary t-values in parentheses below the coefficient estimates.  $F_{ar,1-5}$  (df1, df2) tests for autocorrelation in residuals up to 5 lags. df1 and df2 denote degrees of freedom.  $F_{arch,1-5}$  (df1, df2) tests for autoregressive conditional heteroscedasticity (ARCH) up to order 5, see Engle (1982). The normality test with chi-square distribution is that by Jarque and Bera (1980).  $F_{XiXj}$  (df1,df2) and  $F_{Xi^2}$  (df1, df2) tests for residual heteroscedasticity by omitting cross products of regressors and squares of regressors, respectively, see White (1980). RESET  $F(df1, df2)$  is a regression specification test. It tests the null hypothesis of correct model specification against the alternative hypothesis of misspecification, see Ramsey (1969). The results in this table are based on the implementation of these tests in PcGive 9.10, see Hendry and Doornik (1996). Here and elsewhere in this study, a raised star \* indicates rejection of the null hypothesis at the 5% level, while two stars \*\* indicate rejection of the null hypothesis at the 1% level. Furthermore, p-values are shown in square brackets.

over 1978:1–1997:4, where  $\hat{\gamma} = \hat{\alpha}/(1 - \hat{\varrho})$ . Estimates of  $\alpha$  and  $\varrho$  were derived by (forward) recursive estimation of the model in Table 2.1. The standard errors have been (recursively) estimated by following the procedure proposed in Bårdsen (1989).

The figure reveals that estimates of  $\gamma$  are remarkably stable around 0.95 and the constancy of  $\gamma$  cannot be rejected at the 5% level. This impression was also supported by backward recursive estimates of  $\gamma$  (not reported).

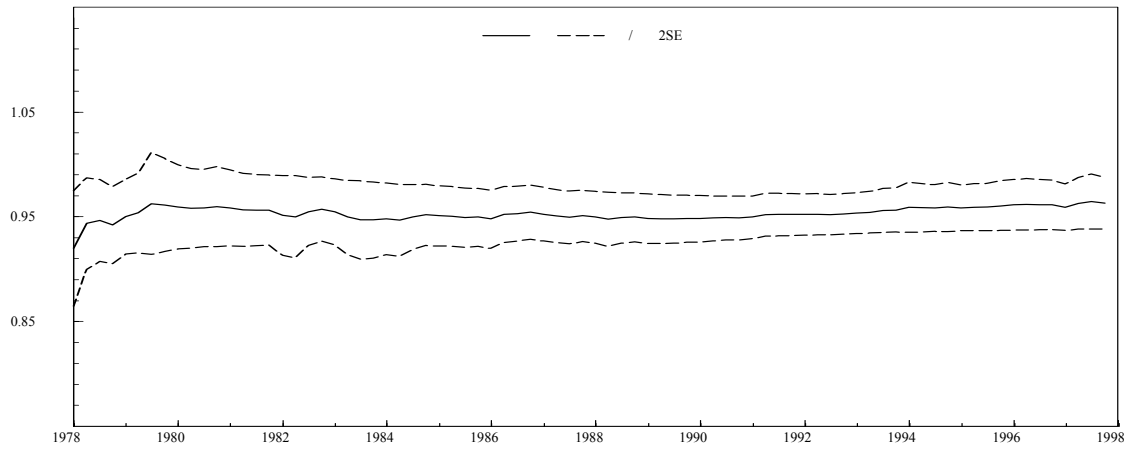


Figure 2.3: *Solid line shows recursive estimates of the equilibrium real exchange rate,  $\gamma$ . The recursive estimates have been derived for the period 1978:1-1997:4. The initial estimate is based on 22 observations for the period 1972:2-1977:4. The dashed lines represent the 95% confidence interval for  $\gamma$ .*

## 2.2. Sensitivity analysis: Evidence based on bilateral real exchange rates

This subsection presents evidence on the robustness of our findings using bilateral real exchange rates for Norway's main trading partners the UK, Germany and the USA.

Table 2.2 presents ADF-models of the Norwegian real exchange rates against the UK, Germany and the USA. These models were formulated by following the same specification strategy as in the case of the model of the effective real exchange rate. The table also reports estimates of half-lives for each of the real exchange rates and the diagnostics of each of the models. Graphs of the bilateral nominal exchange rates (indexed) and the relative consumer prices are presented in the appendix.

The table suggests that the null hypothesis of a unit root in each of the real exchange rates may be rejected at the 5% level. These results are supported by the visual impression from the graphs in the appendix. As in the case of the effective exchange rate, the evidence against the null hypothesis seemed to increase when impulse dummies were employed to filter out large residuals. For example, if we

Table 2.2: ADF models of the bilateral real exchange rates; Sample 1972:2-1997:4.

$$\begin{aligned}
\Delta R_t^{UK} &= -0.124 R_{t-1}^{UK} + \hat{\Gamma}'_{UK}(C, \Delta R_{t-1}^{UK}, \Delta R_{t-2}^{UK}, \Delta R_{t-4}^{UK}, \Delta R_{t-7}^{UK}) \\
&\quad (-2.936) \\
\Delta R_t^{Ger} &= -0.093 R_{t-1}^{Ger} + \hat{\Gamma}'_{Ger}(C, \Delta R_{t-1}^{Ger}, \Delta R_{t-3}^{Ger}, \Delta R_{t-5}^{Ger}) \\
&\quad (-2.970) \\
\Delta R_t^{USA} &= -0.100 R_{t-1}^{USA} + \hat{\Gamma}'_{USA}(C, \Delta R_{t-1}^{USA}, \Delta R_{t-2}^{USA}, \Delta R_{t-3}^{USA}, \Delta R_{t-7}^{USA}) \\
&\quad (-3.149) \\
\Delta R_t^{USA} &= -0.135 R_{t-1}^{USA} + \tilde{\Gamma}'_{USA}(C, ..., id84q3, id85q1) \\
&\quad (-4.325)
\end{aligned}$$

**Half-lives and Diagnostics**

Equation of:	$\Delta R_t^{UK}$	$\Delta R_t^{Ger}$	$\Delta R_t^{USA}$	$\Delta R_t^{USA}$
Half-life: $\ln(0.5)/\ln(\hat{\varrho})$	5.3	7.1	6.6	4.8
Half-life: Impulse resp.	8.8	10.9	—	6.0
Half-life: Impulse resp. corr.	7.5	7.0	—	4.1
$R^2$	0.11	0.16	0.24	0.35
DW	2.01	1.92	1.96	2.02
AR 1-5: $F_{ar,1-5}(5, df_2)$	0.82[0.54]	0.61[0.69]	0.17[0.97]	0.92[0.47]
ARCH 5: $F_{arch,1-5}(5, df_2)$	0.62[0.68]	0.92[0.47]	1.17[0.33]	0.17[0.95]
Normality: $\chi^2_{nd}(2)$	2.28[0.32]	11.49[0.00]	8.78[0.01]	8.81[0.00]
Hetero: $F_{Xi^2}(df_1, df_2)$	0.38[0.95]	0.92[0.50]	2.16[0.03]	2.16[0.76]
Hetero: $F_{XiXj}(df_1, df_2)$	1.03[0.44]	0.98[0.48]	1.31[0.20]	1.10[0.37]
RESET $F(1, df_2)$	0.18[0.68]	1.40[0.24]	0.49[0.48]	0.18[0.67]

Note: See the text and Table 2.1 for details. The DF critical t-values at the 5% and the 1% levels are -2.887 and -3.489, respectively. The right column reports results for the model of  $R^{USA}$  with impulse dummies.

extend the model of  $\Delta R^{USA}$  with two impulse dummies,  $id84q3$  and  $id85q1$ , the estimated adjustment coefficient and the corresponding t-ADF value become  $-0.135$  and  $-4.325$ , respectively, see Table 2.2.<sup>5</sup> The absolute value of this t-ADF value is larger than the DF-critical value, even at the 1% level.

The estimates of half-lives for the bilateral real exchange rates are also relatively low compared with the international evidence. The point estimates based on the half-life formula are around 6 quarters, see Table 2.2. Estimates based on impulse

<sup>5</sup>These impulse dummies probably reflect the large nominal and real depreciation of the Norwegian and other exchange rates relative to the US dollar in the period prior the Plaza Accord in September 1985, when finance ministers of the five major industrial countries agreed to undertake market interventions to drive down the value of the dollar.

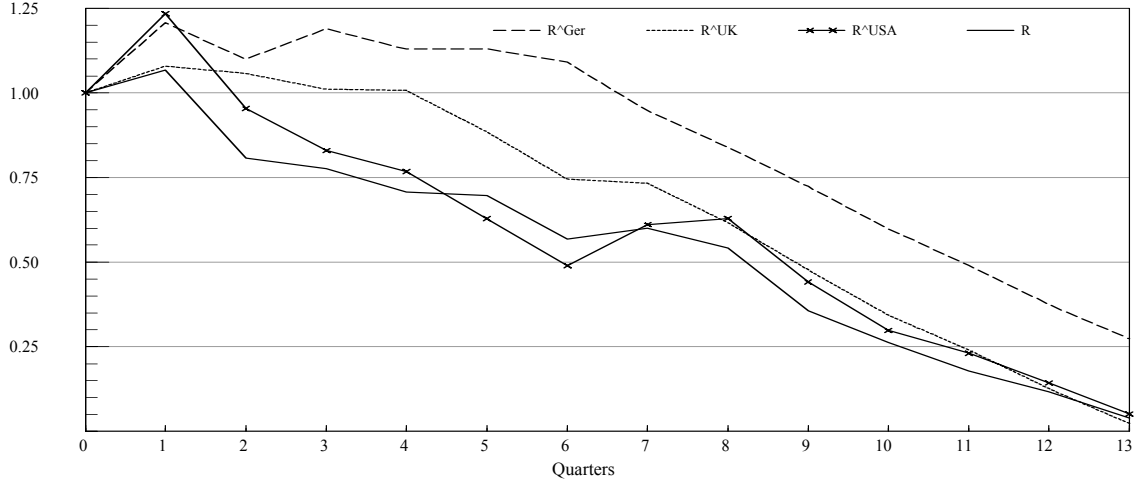


Figure 2.4: *Impulse response analyses: The figure depicts responses to a unit shock using the models of the Norwegian effective and the bilateral real exchange rates against UK, USA and Germany, in Tables 2.1 and 2.2, respectively.*

response analyses are higher, however, and vary in the range of 6 to 11 quarters. However, Figure 2.4 shows that the shock responses are initially amplified before dissipating. Cheung and Lai (2000b) points out that such non-monotonic dynamics can overstate half-life estimates. Row 4 of Table 2.2 presents the half-life estimates after the shock responses reach their maximum values and shows that the estimates become lower and closer to those based on the half-life formula, varying in the range of 4 to 8 quarters.

The difference between our estimate of half-life for  $R^{USA}$  based on (uncorrected) impulse response analysis, 6 quarters, and that of Taylor (2001a) (2.7 years) may be largely ascribed to differences in the data frequency; he applies a DF-GLS test on annual data. Taylor (2001b) shows that annual data may overestimate the half-life when the adjustment is of the order of quarters. To investigate this possibility we fitted an ADF model of  $R^{USA}$  with a constant and one lag (i.e.  $\Delta R_{t-1}^{USA}$ ) to annual data for the period 1972–1997. The t-ADF value from this model was  $-3.312$ , which rejected the null hypothesis of a unit root in the annual values of  $R^{USA}$  at the

5% critical value; the critical 5% DF-value is  $-2.966$ . The adjustment coefficient was  $-0.357$ , and (uncorrected) impulse response analysis suggested a half-life of 2.5 years, which is quite close to the estimate of Taylor (2001a).

### 3. Testing PPP in a system framework

The PPP theory implies symmetry and proportionality restrictions on domestic and foreign prices in a long-run nominal exchange rate equation. Specifically, these restrictions imply that  $\pi_1 = \pi_2 = 1$  in

$$e = \ln\gamma + \pi_1 cpi - \pi_2 cpi^f, \quad (3.1)$$

where the variables in small letters are the natural logs of the original variables.

In the previous section, these restrictions are imposed by definition of the real exchange rate. However, a number of studies have questioned the plausibility of these restrictions in the light of e.g. possible variation in the construction of aggregate prices across countries and measurement errors, see among others Froot and Rogoff (1994). This criticism finds support in numerous empirical studies that report considerable deviations from both restrictions, see e.g. MacDonald (1995). In addition, it is of interest to test whether deviations from parity are eliminated by adjustments in the nominal exchange rate, prices or both.

In order to test the above restrictions, we employ cointegration techniques, as ADF tests suggested that the (logs of) nominal exchange rate, domestic and foreign consumer prices are integrated of order one. The next subsection employs the multivariate cointegration procedure of Johansen (1995), in a small vector autoregressive (VAR) model of  $e$ ,  $cpi$  and  $cpi^f$ . In theory, cointegration between such a limited set of variables is robust to extension of the information set by new variables, see e.g.



Table 3.1: The VAR model

Trivariate VAR(5) model of: $e$ , $cpi$ and $cpi^f$					
<i>Unrestricted: C, CS, CS1 and CS2; restricted: trend (t)</i>					
Sample: Seasonally non-adjusted quarterly data, 1972:2–1997:4.					
Single equation and system diagnostics					
	$F_{ar,1-5}(5, 78)$	$\chi_{nd}^2(2)$	$F_{Xi^2}(32, 50)$	$F_{arch,1-4}(4, 75)$	$\hat{\sigma}$
$e$	1.41[0.23]	10.97[0.01]**	1.09[0.40]	1.06[0.38]	0.014
$cpi$	1.15[0.34]	7.84[0.02]*	0.70[0.86]	0.82[0.52]	0.006
$cpi^f$	1.63[0.16]	4.90[0.09]	1.02[0.46]	0.73[0.57]	0.003
	$F_{ar,1-5}(45, 196)$	$\chi_{nd}^2(6)$	$F_{Xi^2}(192, 274)$		
VAR	1.17[0.23]	24.84[0.00]**	0.81[0.94]		

Note: See Table 2.1 for an explanation of the tests.

Favero (2001, p. 70).

### 3.1. The VAR model

The VAR model contains five lags of the endogenous variables, a constant term ( $C$ ), three centered seasonal dummies ( $CS$ ,  $CS1$ ,  $CS2$ ) and a deterministic trend  $t$ . The time trend is restricted to the cointegration space, as recommended by Doornik et al. (1998) to safeguard against invalid inference on the cointegration rank,  $r$ : the number of cointegrating relations. The number of lags is motivated by the statistical significance of up to 5 lags of  $cpi^f$ , although fewer lags of  $cpi$  and  $e$  were significant at the 5% level.

The VAR model seems to be statistically well specified. Table 3.1 presents single equation and system diagnostics of the VAR model, and Figure 3.1 displays the residuals and their distributional properties. There do not seem to be violations of the standard residual assumptions, except for the normality assumptions. These violations may be due to a few outliers among the residuals from the exchange rate and the  $cpi$  equations in e.g. 1986 and 1997, see Figure 3.1. Thus the  $p$ -values in the squared brackets should be considered indicative, as they may deviate from the true significance levels.

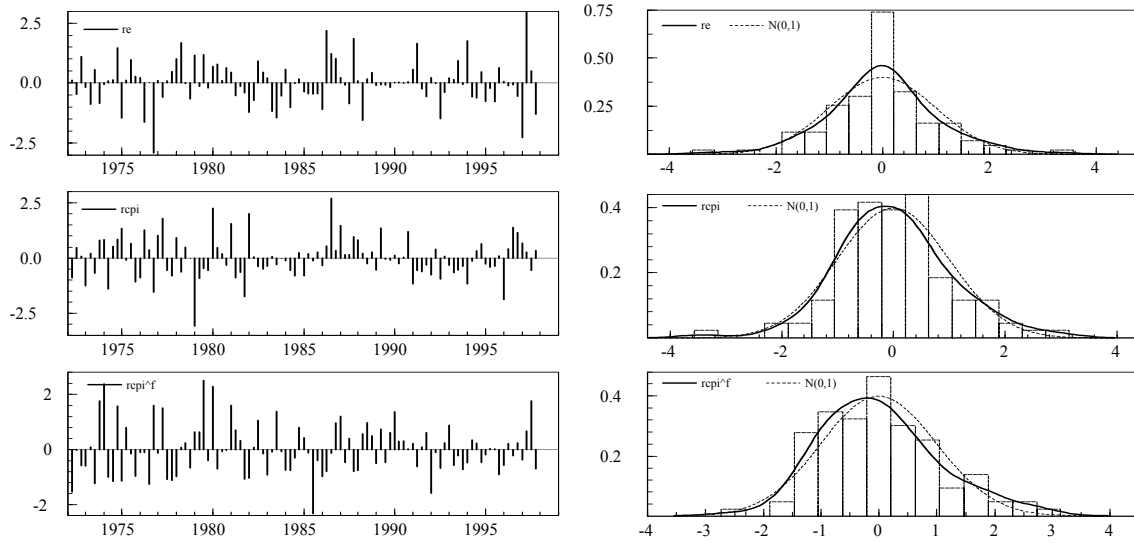


Figure 3.1: *Residual characteristics in the full VAR model. Scaled residuals in the first column and the distribution of residuals, plotted against the standard normal distribution, in the second column. The measures for skewness and excess kurtosis are also reported.*

The parameters of the VAR model appear to be constant over time, at least from 1981 onwards. Figure 3.2 displays the 1-step ahead recursively estimated residuals  $\pm 2SE$  and 1-step ahead forecast and breakpoint Chow tests for each of the three equations in the VAR model. The rejections of the Chow tests for the *cpi* and exchange rate equations in e.g. 1986 and 1997 may be ascribed to the outliers in these periods.

### 3.1.1. Cointegration analysis

Table 3.2 investigates the number of cointegrating relations between *e*, *cpi* and *cpi*<sup>f</sup>. The table reports the estimated eigenvalues ( $\hat{\mu}$ ), the log-likelihood values (*likl*) and the trace test statistics (*Trace*) under different null hypotheses about the cointegration rank *r*, the number of long-run relations. The trace test is fairly robust to violations of the normality assumption regarding the residuals, see e.g. Cheung

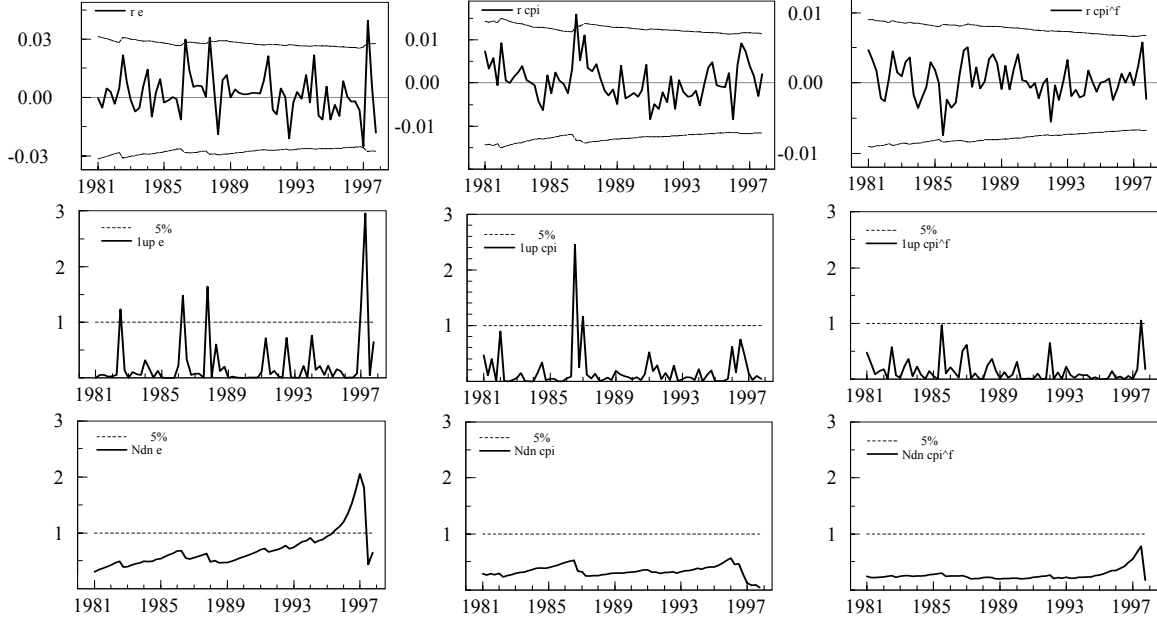


Figure 3.2: *Constancy test statistics for the VAR model, obtained by recursive estimation of the VAR model in the period 1981:1-1997:4. For each of the equations: one-step ahead residuals  $\pm 2SE_t$  in the top row; one-step ahead Chow statistics (1up Chows) in the middle row; and breakpoint Chows (Ndn Chows) in the bottom row. The Chow statistics are scaled by their critical values at the 5% level of significance.*

Table 3.2: Cointegration rank

$r$	0	1	2	3
$likl$	1569.6	1580.4	1587.7	1592.7
$\hat{\mu}$		0.19	0.13	0.09

$H_0 :$	$r = 0$	$r \leq 1$	$r \leq 2$
$Trace$	45.81[0.02]*	24.33[0.08]	9.73.[0.14]

Note: The  $p$ -values associated with the trace statistics ( $Trace$ ) have been provided by PcGive 10, see Doornik and Hendry (2001).

and Lai (1993). Testing the cointegration rank amounts to testing the number of eigenvalues different from zero. In the trace test, the null hypothesis is that the eigenvalues  $\mu_i = 0$ ,  $i = r + 1, r + 2$ , while the first  $r$  eigenvalues are non-zero. The table shows that only the null hypothesis of  $r = 0$  is rejected; the  $p$ -value is 2%.

This suggests one cointegrating relation between the three variables.

Figure 3.3 displays recursive estimates of the largest eigenvalue over 1985:1–1997:4. Their stability above zero supports the presence of one stable cointegrating relation in the sample period.

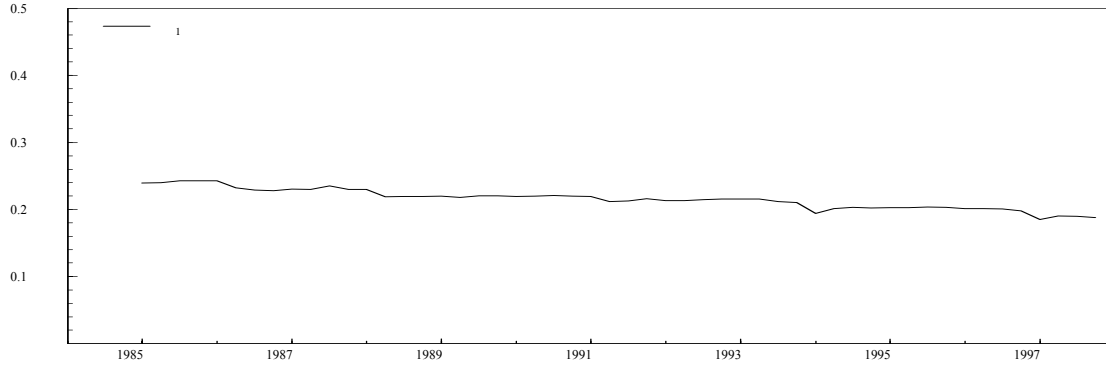


Figure 3.3: *Recursive estimates of the largest eigenvalue ( $\mu_1$ ). The initial estimation period is 1972:2–1984:4.*

### 3.1.2. Tests of PPP restrictions and the response of $e$ and $cpi$ to deviations from parity

Panel I of Table 3.3 tests whether the single long-run relation between  $e$ ,  $cpi$  and  $cpi^f$ , normalised on  $e$ , is consistent with the PPP theory. Row (a) shows that the deterministic trend is redundant since a zero restriction on the trend is accepted with a  $p$ -value of 0.85. Absolute values of the unrestricted coefficient estimates of  $cpi$  and  $cpi^f$  are fairly close to each other and row (b) shows that symmetry restriction on these coefficients is easily accepted at standard levels of significance. Moreover, the values of  $-1$  and  $1$ , implied by the proportionality restriction, fall within 95% confidence intervals for the coefficient estimates of  $cpi$  and  $cpi^f$ . Row (c) defines a long-run relation in strict accordance with the PPP theory by imposing both the symmetry and proportionality restrictions on  $cpi$  and  $cpi^f$  (in addition to the zero restriction on the trend). These restrictions are accepted with a  $p$ -value of 0.16.

Table 3.3: PPP and the response of  $e$ ,  $cpi$  and  $cpi^f$  to deviations from parity

<b>I. Testing PPP restrictions on the long-run relation</b>						
	$\beta'$	$e$	$cpi$	$cpi^f$	$t$	
(a)	<i>Unrestricted</i>	: 1	-0.66 (0.214)	0.57 (0.230)	0	, $\chi^2(1)$ : 0.035[0.85]
(b)	<i>Symmetry</i>	: 1	-0.77 (0.165)	0.77 (0.165)	0	, $\chi^2(2)$ : 3.880[0.14]
(c)	<i>PPP</i>	: 1	-1	1	0	, $\chi^2(3)$ : 5.202[0.16]
<b>II. Testing restrictions on adjustment coefficients</b>						
	$\alpha'$	$\Delta e$	$\Delta cpi$	$\Delta cpi^f$		
(d)	<i>Unrestricted:</i>	-0.148 (0.046)	0.035 (0.019)	-0.013 (0.012)		
(e)	<i>Unresponsive prices:</i>	-0.152 (0.046)	0	0	$\chi^2(5)$	9.769[0.082]
(f)	<i>Unresponsive exch. rate:</i>	0	-0.038 (0.020)	-0.014 (0.012)	, $\chi^2(4)$	16.738[0.000]

Note:  $\beta'$  is a  $1 \times 4$  vector of parameters defining the long-run relation and  $\alpha'$  is a  $1 \times 3$  vector of adjustment coefficients, see Johansen (1995). Standard errors in parentheses below the coefficient estimates and  $p$ -values in hard-brackets. Panel II tests restrictions on the adjustment coefficients when both PPP restrictions (c) are imposed.

Panel II examines whether deviations from PPP are eliminated through adjustments in the exchange rate or prices. Row (d) reports the unrestricted estimates of the adjustment coefficients (with standard errors in parentheses). These measure the response of the exchange rate and prices to deviations from PPP, defined by row (c) in Panel I. Numerically, both the exchange rate and domestic prices contribute to eliminate deviations from PPP. However, the response of domestic prices is much weaker than that of the exchange rate. As one would expect in the case of a small economy, the response of foreign prices is negligible. Statistically, the null hypothesis that foreign and domestic consumer prices are unresponsive to deviations from PPP is not rejected at the 5% level, see row (e). In contrast, the null hypothesis that the exchange rate is unresponsive to deviations from PPP is strongly rejected.

Figure 3.4 substantiates the results in Table 3.3. This graphs the test statistics when we impose the restrictions defined in rows (b), (c), (f) and (e) recur-

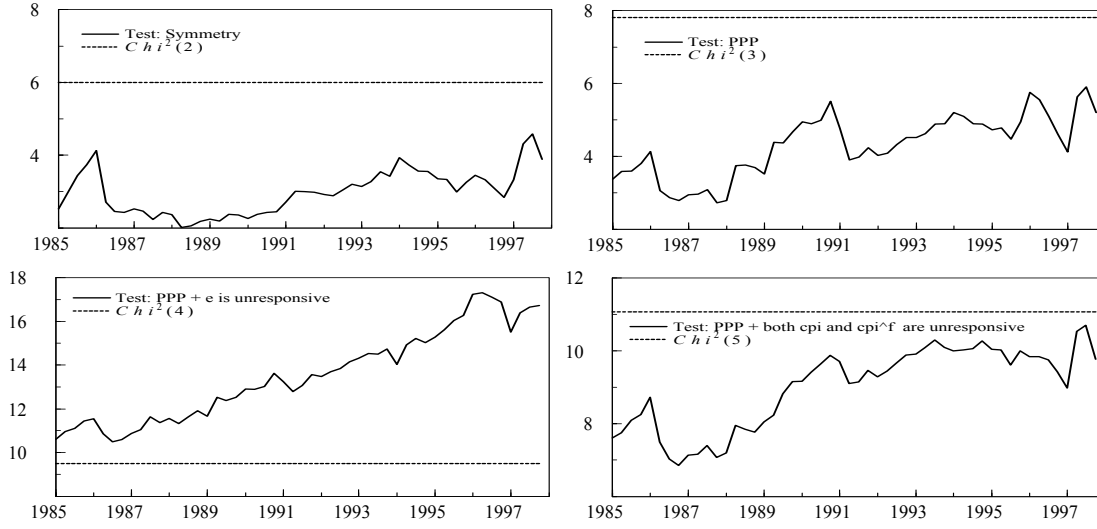


Figure 3.4: *Graphs of test statistics when we impose the restrictions defined in rows (b), (c), (f) and (e) recursively from 1985:1 to 1997:4. The associated one-off critical values at the 5% level are depicted as the dotted lines.*

sively from 1985:1, that is, on each of the sample periods 1972:2–1985:1, 1972:2–1985:2,...,1972:2–1997:4. The associated one-off critical values at the 5% level are depicted as dotted lines. Notably, we arrive at the same conclusions for each of the sample periods: PPP is supported and deviations from PPP are eliminated mainly through changes in the nominal exchange rate; the response of domestic and foreign prices is statistically insignificant.

### 3.2. Sensitivity analysis: Extensions of the information set

A robust cointegrating relation between a set of variables is characterised by invariance to the inclusion of new variables and observations in the information set. A number of sensitivity analyses demonstrate that our findings, including the PPP relation, are robust to such extensions of our information set.

First, our conclusions remain invariant to the addition of new variables to the VAR model. These include current account deficit, oil prices, domestic and foreign interest rates and a number of impulse dummies to control for the effects of oil price

shocks and the sharp exchange rate fluctuations in 1997. In addition, the results are supported in parsimonious versions of the extended VAR model in which we condition on oil prices, foreign consumer prices and interest rates. The details of this comprehensive sensitivity analysis are presented in Akram (2000a).

Second, PPP and the parity preserving response of the nominal exchange rate is strongly supported by a single equation non-linear equilibrium correcting model (EqCM) of the nominal exchange rate, with  $e-(cpi-cpi^f)$  as the equilibrium term. This model allows for non-linear oil price effects and controls for a number of short-run determinants of the exchange rate, see Akram (2000b) for details.

Table 3.4: EqCMs of the nominal exchange rate and domestic consumer prices

$$\begin{aligned}\Delta \widehat{e}_t &= -0.129[e-(cpi-cpi^f)]_{t-1} + \widehat{\Gamma}'_e(C, \Delta cpi_t^f, \Delta e_{t-1}, \Delta^2 cpi_{t-2}) \\ &\quad (0.033) \\ \Delta \widehat{cpi}_t &= 0.036 [e-(cpi-cpi^f)]_{t-1} + \widehat{\Gamma}'_{cpi}(C, \Delta cpi_t^f, \Delta cpi_{t-1}, \Delta cpi_{t-2}, \Delta cpi_{t-4}) \\ &\quad (0.016)\end{aligned}$$

*Extended sample: 1972:2–2001:3, 118 observations. Method: OLS*

Note: Both equations have been derived by applying PcGets 1.02 with default settings, except that the constant terms were imposed, see Hendry and Krolzig (2001). The effects of regressors in addition to the equilibrium terms, included in parentheses, have been suppressed to save space.  $\widehat{\Gamma}_e$  and  $\widehat{\Gamma}_{cpi}$  are vectors of coefficient estimates associated with the additional regressors in the exchange rate and the price equation, respectively.

Third, the conclusions do not seem to be a transient feature of the sample; The coefficient estimates are remarkably stable over time and are supported by out-of-sample observations. For the purpose of illustration, we have developed two single equation equilibrium correcting models of the nominal exchange rate and domestic consumer prices on an extended sample, 1972:2–2001:3, which contains 15 new quarterly observations. Table 3.4 presents specific versions of both models. These have been derived by applying the “general-to-specific” simplification strategy in the computer program PcGets, see Hendry and Krolzig (2001). The initial general model of  $\Delta e_t$  included  $\Delta cpi_t$ ,  $\Delta cpi_t^f$ ,  $[e-(cpi-cpi^f)]_{t-1}$ , 4 lags of  $\Delta e_t$ ,  $\Delta cpi_t$  and  $\Delta cpi_t^f$ , and a (fixed) constant term. The general model of  $\Delta cpi_t$  included 3 centered

seasonals and  $\Delta e_t$ , in addition to the other regressors in the model of  $\Delta e_t$ .

Table 3.4 shows that the equilibrium term  $[e-(cpi-cpi^f)]_{t-1}$  enters both equations and the coefficient estimates are close to those in the VAR model, see e.g. row (d) in Table 3.3. However, the coefficient estimate in the price equation is barely significant at the 5% level. It should be mentioned that it becomes insignificant if the sample ends in 1997:4, as in the case of the VAR model. We also note that the estimated adjustment coefficients in both equations are quite close to those derived within extended VAR models, see Akram (2000a). Furthermore, the estimated adjustment coefficient in the exchange rate equation is close to that in the non-linear EqCM mentioned above, see Akram (2000b).

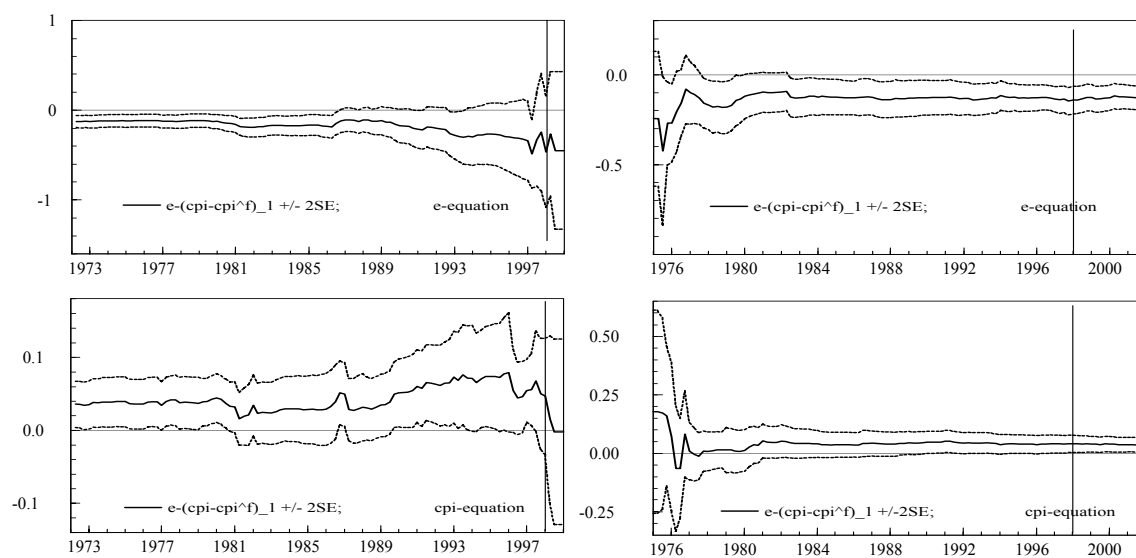


Figure 3.5: *Backward and forward recursive OLS estimates  $\pm 2SE$  of the adjustment coefficient in the exchange rate equation in the top row and those of the adjustment coefficient in the price equation in the bottom row. The equations are presented in Table 4.1. In all cases, the initial number of observations are 12. The dashed vertical line at 1998:1 marks the introduction of new observations. These covers the period 1998:1–2001:3.*

Figure 3.5 examines the stability of the adjustment coefficient in the exchange rate equation, in the top row. Both backward and forward recursive estimates



of the adjustment coefficient over  $t = 1975:1, \dots, 2001:3$  and  $t = 1998:4, \dots, 1972:2$ , respectively, point to substantial equilibrium reversion in the exchange rate; in the later as well as the early periods in the sample. Moreover, the coefficient estimates are remarkably stable over time and appear invariant to new observations.

In contrast, the contribution of domestic consumer prices to preserve PPP is relatively small and statistically insignificant over most of the sample period, see the bottom row of Figure 3.5.

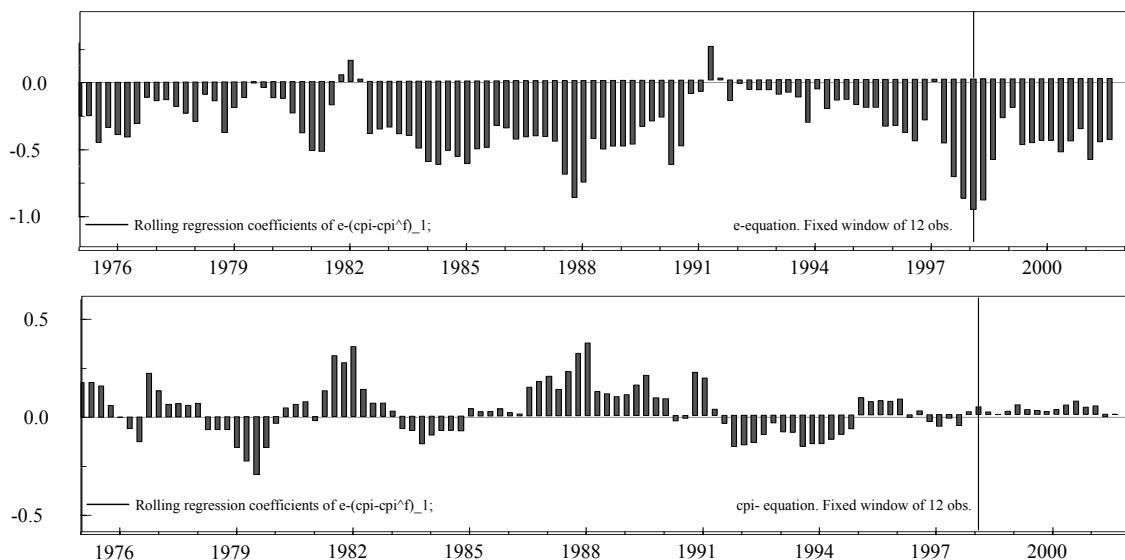


Figure 3.6: (a) *Forward rolling OLS regression estimates, represented by bars, of the adjustment coefficients in the exchange rate equation at the top, and (b) those of the consumer price equation at the bottom. Each of the estimates are based on a fixed window of 12 quarters. The dashed vertical line indicates the introduction of new observations.*

Figure 3.6 (a) displays rolling regression estimates of the adjustment coefficients in the exchange rate equation, based on a fixed window of just 12 overlapping observations. Even these reveal substantial PPP-preserving behaviour in the exchange rate. Almost all of the estimates are correctly signed, including the last one, which is based exclusively on new observations for the subperiod 1998:4–2001:3. There are, however, relatively large variations in the estimates. Their absolute values are rela-

tively small particularly around 1980 and the early 1990s. These may be explained by the peg to a currency basket in the period December 1978–February 1982 and to the ECU in the period October 1990–December 1992. It seems fair to say that the parity preserving response of the nominal exchange rate extends beyond the period of frequent devaluations which ended in 1986.

Figure 3.6 (b) reveals that the contribution of domestic consumer prices to bring about convergence towards PPP has been highly unstable; in many periods, they have actually contributed to divergence from parity, e.g. in the late 1970s and the early 1990s. Since the mid-1990s, the response of domestic prices to deviations from PPP has been negligible.

#### 4. Accounting for the fast convergence

Our evidence for Norway suggest that deviations from PPP are more short-lived than they are in other industrial countries. For example, the estimate of half-life (implied by impulse response analyses) of the Norwegian real exchange rate with the USA as a base country is 1.5 years (6 quarters), while the median half-life estimate of the real exchange rates of industrial countries vis-à-vis the USA reported by Cheung and Lai (2000a) is 3.3 years.<sup>6</sup> In this section we make an effort to account for this relatively low persistence of the Norwegian real exchange rate, or alternatively of deviations from PPP between Norway and its trading partners.

A number of studies including Cheung and Lai (2000a) ascribe cross country differences in the persistence of real exchange rates to differences in the nature of shocks, and countries' openness. Deviations from PPP initiated by nominal shocks

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<sup>6</sup>Their estimate is based on monthly data over 1973:4–1994:12. However, in this case, differences in the data frequency and the sample period cannot explain the difference between the half-lives. We fitted an ADF model of  $R^{USA}$  on monthly data for the periods 1972:1–1997:12 and 1973:4–1994:12, and in both cases obtained a half-estimate of about 16 months by means of impulse response analyses. The preferred ADF model included 3 lags of  $\Delta R^{USA}$  and 6 impulse dummies to control for outliers in the periods 1973:8, 1986:7, 1986:9, 1986:12, 1985:2 and 1985:4. The initial ADF model was formulated with up to 24 lags of  $\Delta R^{USA}$ .

are generally believed to be more short-lived than those initiated by real shocks. Nominal shocks, such as changes in money growth and devaluations, only contribute to deviations from PPP (or affect the real exchange rate) as a result of price stickiness. Hence, such deviations from PPP should not last more than one or perhaps two years. In contrast, real shocks, such as a productivity growth, government expenditures and discoveries of natural resources, are presumed to raise the price of non-tradables relative to tradables and thereby contribute to trend in real exchange rates. This is especially the case for relatively closed economies, which are exposed to weak arbitrage pressure.

Table 4.1: Explaining cross country differences in half-lives

	Growth	G-spending	Openness	Inflation	Std( $\Delta e$ )	C-index
Industrial	1.62	37.78	23.52	6.92	13	0.39
Norway	3.19 (2.21)	24.15 (32.6)	33.76	6.55	10	0.42

Note: The figures for the group of industrial countries are the estimated median values taken from (Cheung and Lai, 2000a, Table 4), except for the values of Std( $\Delta e$ ) and C-index. The values of Std( $\Delta e$ ), for both Norway and industrial countries, are based on (Taylor, 2001a, Table 6) and the values of the C-index are based on the subperiod-averages of the values reported in (Calmfors, 2001, Table 2). The remaining figures for Norway are sample averages over the period 1972-1997. Growth: average of GDP growth per capita; In parentheses: average of GDP growth in mainland Norway per capita. G-spending: average of government expenditures relative to GDP in %; in parentheses, relative to mainland GDP. Openness: average of the ratio in % of the average of exports and imports to the level of GDP; For Norway, the average of exports from mainland Norway and imports relative to the GDP in mainland Norway. Inflation: average of annual CPI-inflation in %. Std( $\Delta e$ ): standard deviation of the annual rate of nominal exchange rate depreciation relative to the US dollar. C-index: centralisation and coordination index used to characterise the system of wage bargaining.

Table 4.1 compares the estimates of productivity growth as measured by GDP growth per capita, government spending, openness and inflation (a proxy for nominal shocks), calculated by Cheung and Lai (2000a) for industrial countries with our estimates for Norway. The table shows that productivity growth in Norway has been almost twice as high as the median for industrial countries, 3.19% versus 1.62% per annum. About 1 percentage point of this is due directly to the Norwegian petroleum sector, given that the productivity growth exclusive of the petroleum sector is 2.21%,

which demonstrates the importance of oil- related shocks to the Norwegian economy. However, government spending in Norway relative to GDP (inclusive or exclusive the petroleum sector) has been relatively low and the Norwegian economy appears more open than that of other industrial countries. These factors may have contributed to the relatively lower persistence in deviations from PPP.

Nominal shocks, however, do not seem to account for the relatively low persistence. The table shows that the annual inflation rate has been in line with those of industrial countries. In addition, adjustment in the Norwegian nominal effective exchange rate in the face of deviations from PPP does not seem to become weak after the last devaluation in May 1986, see Figure 3.6 (a).

On the other hand, the Norwegian policy of managing the nominal exchange rate may have stabilised fluctuations in the real exchange rate over time. A number of studies, including Mussa (1986) and Taylor (2001a), record higher persistence in real exchange rates under a system of floating exchange rates than under a pegged or managed exchange rate regime.<sup>7</sup> Following Taylor (2001a), this “regime effect” is proxied by the standard deviation of annual changes in the bilateral nominal exchange rate against the USA for industrial countries and Norway, see the second last column of Table 4.1. The relatively lower variability in the Norwegian nominal exchange rate may have contributed to lower variability in the real exchange rate, and thereby stabilised deviations from PPP. However, this contribution seems to be modest, as the difference in the exchange rate variability is not that large.

Differences in the wage bargaining system is another factor that may explain the relatively low persistence in the real exchange rate. It has been pointed out that a system of centralised wage bargaining tends to take into account the effects of adverse shocks to the overall economy by wage moderation, see e.g. Layard et al. (1991, Ch.

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<sup>7</sup>This could be due to economic policies conducted in conjunction with, or in support of a pegged exchange rate regime. For example, a government may undertake devaluations, which is only possible under a pegged exchange rate system, and exercise fiscal restraint in order to improve competitiveness and neutralise the appreciating effects of real shocks to the real exchange rate.

2). Accordingly, shocks that may affect the viability of the sector for tradables, given its limited opportunities to offset higher factor costs by raising product prices, are countered by lower wage claims. Such restraints on wages are built into the Norwegian inflation model where the wage settlement for the sector for tradables is adopted by the sector for non-tradables, see e.g. Aukrust (1977). Potentially, this may contribute to stabilise the price of non-tradables relative to tradables and hence counteract trend behaviour in real exchange rates due to oil-related shocks.

The empirical evidence acknowledges some contribution from the Norwegian inflation model in accounting for the relatively low real exchange rate persistence. The last column of Table 4.1 reports the median estimates of an index for the centralisation and coordination of the wage setting (C-index) for industrial countries and Norway. These estimates are based on the values of the C-index defined and presented in Calmfors (2001, Table 2). The relatively higher value of the C-index for Norway is favourable to relatively low persistence in the real exchange rate, but the difference in the C-index is fairly small, 0.42 versus 0.39. Actually, this is consistent with the weak and unstable response of consumer prices to deviations from PPP displayed in Figure 3.6 (b). This suggests that the Norwegian wage and price process has played a modest role, if any, in correcting for deviations from PPP and preserving competitiveness.

## 5. Conclusions

Existing empirical studies generally reject or present weak support for PPP for countries that have been predominantly exposed to real shocks. This is especially the case for existing studies of Norway, which has experienced large oil-related shocks in the post-Bretton Woods period. Moreover, deviations from PPP are found to be quite persistent for industrial countries in general, with half-life estimates in the range of 3 to 6 years. This paper presents novel results against this background.

In a number of standard tests for PPP between Norway and its trading partners, we find clear evidence of convergence towards PPP in the medium run; the half-life of a given deviation from parity is just about 1 1/2 years. In addition, the Norwegian equilibrium real exchange rate appears to have been constant over the sample period 1972–1997. We also find that deviations from parity are mostly corrected by adjustment of the nominal exchange rate; the contribution of domestic prices appears to be weak and ambiguous. This suggests that in the long run, the direction of causation is from domestic prices to the nominal exchange rate.

We have undertaken a number of sensitivity analyses to demonstrate the robustness of our findings, which are primarily based on quarterly observations of the Norwegian effective real and nominal exchange rates and consumer prices over the period 1972:2–1997:4. In particular, we show that the evidence of PPP in the medium run is supported by analyses of the Norwegian bilateral real exchange rates vis-à-vis Norway’s main trading partners the UK, Germany and the USA. We have also unveiled PPP between Norway and the USA by examining annual and monthly observations of the bilateral real exchange rate. In addition, we report that the support for PPP based on empirical analyses of the effective nominal exchange rate is robust to extension of the information set by additional variables and changes in model formulations. We also demonstrate that the evidence of PPP, and the response of the nominal exchange rate to deviations from PPP, are robust to extensions of the information set by post-sample observations for the period 1998:1–2001:3.

Finally, we have made an effort to account for the relatively low persistence of the Norwegian real exchange rate, despite oil shocks, compared with those of other industrial countries. Arguably, lower government spending relative to GDP, higher openness to international trade, a stable exchange rate regime, and the Norwegian system of centralised and coordinated wage bargaining may have contributed to outweigh the real appreciation effects of the oil shocks, and preserved the international

competitiveness of the Norwegian economy over time. Our account of the relatively low persistence is merely indicative, however, and more research on this issue is warranted.

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## Appendix: Data definitions

Unless otherwise stated, the data source is OECD-Main Economic Indicators. Other sources include RIMINI and TROLL8. The former refers to the database of the macroeconometric model of Norges Bank, RIMINI. TROLL8 database is also maintained by Norges Bank.

$CPI$  : Consumer price index for Norway, 1995 =1. Quarterly observations: CPI.

Source: RIMINI. Annual observations: NOR.CPALTT01.IXOB.A; 1995 = 100. Monthly observations: NOR.CPALTT01.IXOB.M; 1995 = 100.

$CPI^f$ : Trade weighted average of consumer price indices for Norway's trading partners; 1995 =1. Quarterly observations: PCKONK. Source: RIMINI.

$CPI^{Ger}$  : Consumer price index for Germany; 1995 = 100. Quarterly observations: DEU.CPALTT01.IXOB.Q.

$CPI^{UK}$  : Consumer price index for United Kingdom; 1995 = 100. Quarterly observations: GBR.CPALTT01.IXOB.Q.

$CPI^{USA}$ : Consumer price index for United States; 1995 = 100. Quarterly observations: USA.CPALTT01.IXOB.Q. Annual observations: USA.CPALTT01.IXOB.A and monthly observations: USA.CPALTT01.IXOB.M.

$CS$  : Centered seasonal dummy variable (mean zero) for the first quarter in each year. It is 0.75 in the first quarter and -0.25 in each of the three other quarters, for every year.

$E$  : Trade weighted nominal value of NOK, 1995 = 1. The quarterly observations are averages of daily observations: PBVAL. Source: RIMINI.

$E^{Ger}$  : The nominal NOK/DEM spot exchange rate; value of 100 DEM. Quarterly observations: K9301212. Source: TROLL8.

$E^{UK}$  : The nominal NOK/GBP spot exchange rate. Quarterly observations: K9300012.

Source: TROLL8.

$E^{USA}$  : The nominal USD/GBP spot exchange rate. Annual, quarterly and monthly observations are: A9300412, K9300412 and M9300412. Source: TROLL8.

Export for Norway: Total exports less exports of capital, oil, gas and shipping services, fixed baseyear prices. Quarterly observations: AF. Source: RIMINI.

GDP for Norway: GDP volume for Norway. Indexed, 1995=100. Annual observations: A.14299BVPZF.... Source: IMF\_IFS.

GDP for mainland Norway: GDP volume for mainland Norway. Annual observations. Source: Statistics Norway. Quarterly observations: YF. Source: RIMINI.

Government expenditures for Norway: Sum of public consumption expenditures (CO) and gross investment expenditures in fixed capital (JO). Fixed baseyear prices. Quarterly observations: CO and JO. Source: RIMINI.

$id19AAq_i$  : Denotes an impulse dummy that takes on a value of 1 in quarter  $i$  of the year 19AA, and zero elsewhere.

Imports for Norway: Total imports, fixed baseyear prices. Quarterly observations: B. Source: RIMINI.

OPENNESS for Norway: Average of  $(AF + B)/YF$  over the period 1972:1–1997:4.

Population in Norway: Annual observations. Source: Statistics Norway.

$R$  : Trade weighted real exchange rate of Norway;  $R \equiv (E \times CPI^f)/CPI$ .

$R^{Ger}$  : Bilateral real exchange rate of Norway vis-à-vis Germany.

$R^{UK}$  : Bilateral real exchange rate of Norway vis-à-vis UK.

$R^{USA}$  : Bilateral real exchange rate of Norway vis-à-vis USA.

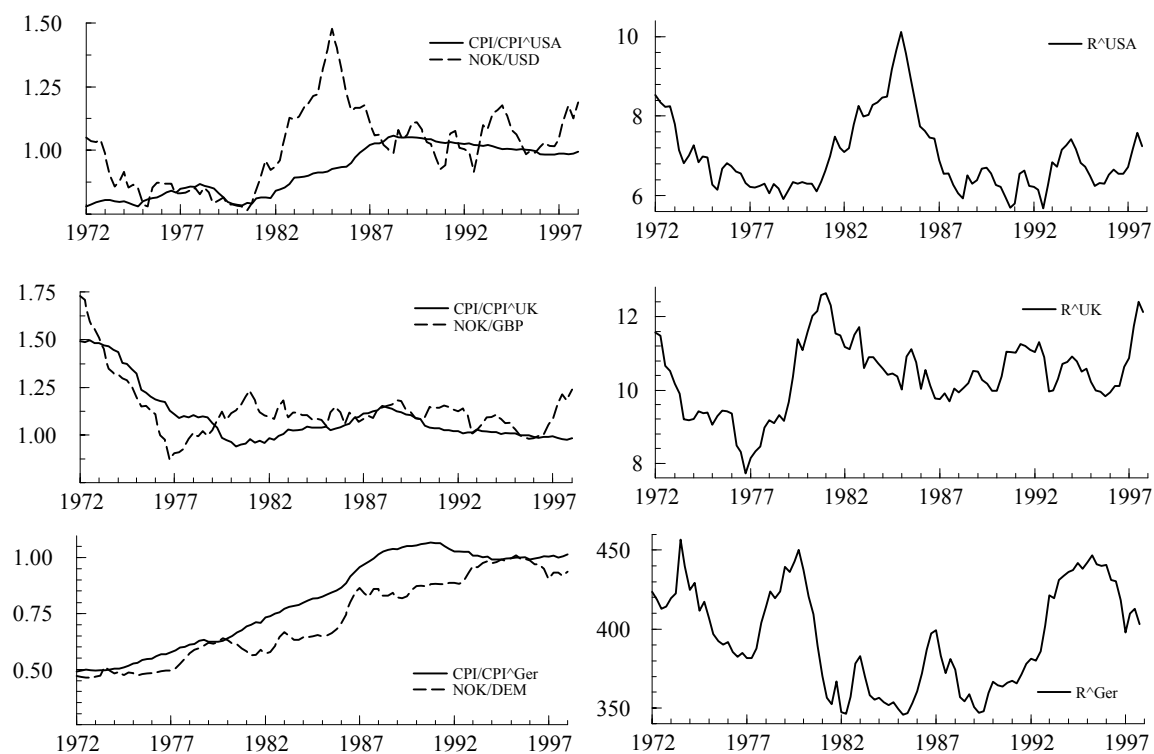


Figure 5.1: *Left column: Relative consumer prices and spot exchange rates for Norway vis-à-vis the USA, the UK and Germany. The spot exchange rates have been indexed with 1995 =1, to facilitate comparison. Right column: The bilateral real exchange rates of Norway vis-à-vis the USA, the UK and Germany.*



**KEYWORDS:**

PPP

Real exchange rate

Cointegration analysis