4. THE MODEL

4.1 Model Philosophy: Data Versus Theory

The macroeconometric model of Spain presented here is comparatively small. It consists of 27 stochastic equations and 61 definitions. As a single-country model it reflects the main features of the Spanish economy. As the exchange rates, foreign demand and foreign prices are exogenous, the model cannot capture feedback from other regions - particularly the euro area, which is the destination of 60% of Spain's exports. Further exogenous variables are the short term interest rate, the oil price and government investment.

In the current model the focus is on aggregate demand, which is estimated in great detail. Gross fixed capital formation is disaggregated into construction on the one hand and machinery, equipment and other products on the other hand. Exports of goods are analysed for four different regional destinations: the euro area, the remaining EU-15 countries, the US and the rest of the world. Exports are thus much more disaggregated than in most prominent macro models[1].

The purpose of the model is threefold: economic analysis, policy simulations, short to medium term forecasts. The model is used for all three purposes at the same time. The focus of the analysis is on the short to medium term. Theories are usually formulated without taking the stochastic properties of the data into account. The econometrician who is estimating a model must then try to combine economic theory with the data. Often the two objectives are difficult to reconcile.

Therefore, priorities must be set. This is why many economists rely on theory as the only guideline. They use calibrated coefficients (often taken from micro studies),

[1] The Bank of Spain has recently published a re-estimation of its model of the Spanish economy with a division of goods exports into intra and extra euro area exports. Its estimation period ends in 1998, i.e. before the launch of EMU.

which are more in line with theory than the estimates they would obtain otherwise. Another approach is to focus on the data and their properties and to set a minimum of restrictions. In existing macro models we often find restrictions, which are supposed to ensure neoclassical properties in the long run. A prominent example is the assumption of long run (super)neutrality of money, an assumption, which is at least questionable.

In this model of the Spanish economy no calibration is applied. Although theoretical considerations play an important part in the *choice of variables* for the long-term relationships, data properties are given priority over theory. This means that no a-priori parameter restrictions derived from theory are imposed.

There is an additional reason why theory and the data are sometimes incompatible. Theory postulates a long-term relationship between a specific number of variables. This may exceed the number of variables, which are found to be cointegrated. If - for example - a cointegrating ("long-term") relationship has been found between only two variables no additional non-stationary series are needed for a stable long-term relationship *from a statistical point of view* - even if *from a theoretical point of view*, a long-term relationship between three variables would be expected. Data and theory are also difficult to reconcile due to the integration properties of the series, which are rarely given any attention in theoretical debates[2].

The model is backward-looking with adaptive expectations, i.e. expectations are modelled via lagged variables. All equations are estimated individually and subsequently combined into a macro model. The estimation strategy is described in detail in the next section.

4.2 Estimation Strategy

The equations of the model are generally estimated as error-correction equations[3]. In this approach the error-correction term reflects a stable long-term relationship, which is based on theoretical considerations, whereas (lagged) differences and other stationary variables model the short-term dynamics.

[2] For an interesting discussion based on money market analysis c.f. Juselius (1999).

[3] For a summary of cointegration and error-correction c.f. Appendix B.1

The appropriate error-correction equation is determined in a procedure involving three steps:

- The Engle/Granger (1987) approach is carried out for the set of variables defined by theory. This involves a static equation in levels with a specific to general approach. Starting from a bivariate estimation additional variables are added if necessary. The residuals are then tested for stationarity using the critical values of Mac Kinnon (1991). The Engle/Granger approach allows to decide on the appropriate deterministics. In contrast to the Johansen cointegration test it provides sensible results in the case of structural breaks such as level shifts, which are numerous in the Spanish data.
- In a second step the Johansen cointegration test is applied. The purpose of this test is to verify the results of the Engle/Granger approach in a dynamic setting. Unlike the Engle/Granger approach, the Johansen test allows to determine the number of cointegrating relationships. The deterministics are modelled as in the Engle/Granger approach. In the analysis a small sample adjustment of the Trace statistic is applied. The critical values of Pesaran et al. (2000), which allow for exogenous I(1) variables, like step dummies, are used as a proxy (instead of a simulation using software such as DISCO).
- The approach applied in the third step combines a test for cointegration with the actual estimation of the cointegrating vector. A non-linear estimator is applied (1987).

Although the direct estimate of error correction model (Stock approach) would be sufficient on its own, all three steps are carried out in order to assess the robustness of results. In the case of ambiguous results, however, priority is given to the estimated error-correction model.

Contrary to the specific to general approach followed in the tests for cointegration, a general to specific approach is applied in the estimation of the error correction equation once it is clear which variables form the long-term relationship. Usually a maximum lag length of four is chosen to allow for an adequate modelling of seasonal fluctuations. In cases, where considerable autocorrelation persists in the residuals the maximum lag-length is extended until the residuals are free of autocorrelation.

Extensive specification tests are carried out to ensure that the equations are appropriate. The tests include the Breusch-Godfrey serial correlation LM test, the ARCH LM test, White's heteroskedasticity test and the Jarque-Bera normality test. Stability is verified by the CUSUM and the CUSUM of squares tests. Ramsey's RESET test serves as a general specification test, as it allows to detect a number of misspecifications. The forecast performance of each equation is examined by dynamic out-of-sample forecasts. The specification tests are described in Appendix B.2. Their results are given along with the equations in this chapter. Results of the stability tests are given in Appendix B.3. The out-of-sample forecasts are documented in Appendix B.4.

The estimations, specification tests and simulations were carried out using the software package Eviews 4. PCGive 10 was applied for the Johansen cointegration tests. It offers a small sample adjustment of both the trace and the maximum eigenvalue statistic.

Throughout the model documentation lower case letters mean logarithms. In the model equations the t-statistic of each coefficient is given in parentheses. As the t-distribution is not valid in the case of the estimation of the long run relations in levels (Hassler 2004), the standard deviations are given in parentheses here.

4.3 Structure of the Model

The structure of the model follows the national accounts framework. We distinguish between aggregate demand (expenditure side of GDP), income and employment and prices. The latter can be interpreted as the supply side of the model. Due to insufficient data the government sector is not fully included. However, the main demand series are present in the model. All foreign aggregates, nominal exchange rates, commodity prices as well as monetary policy variables are exogenous. Public investment is to be used as a policy variable in the model simulations and is therefore also defined as exogenous. Of course, this approach is debatable, as the case of Germany has shown that public investment is often treated as a residual in economic policy. Figure 4.1 gives an overview of the structure of the model.

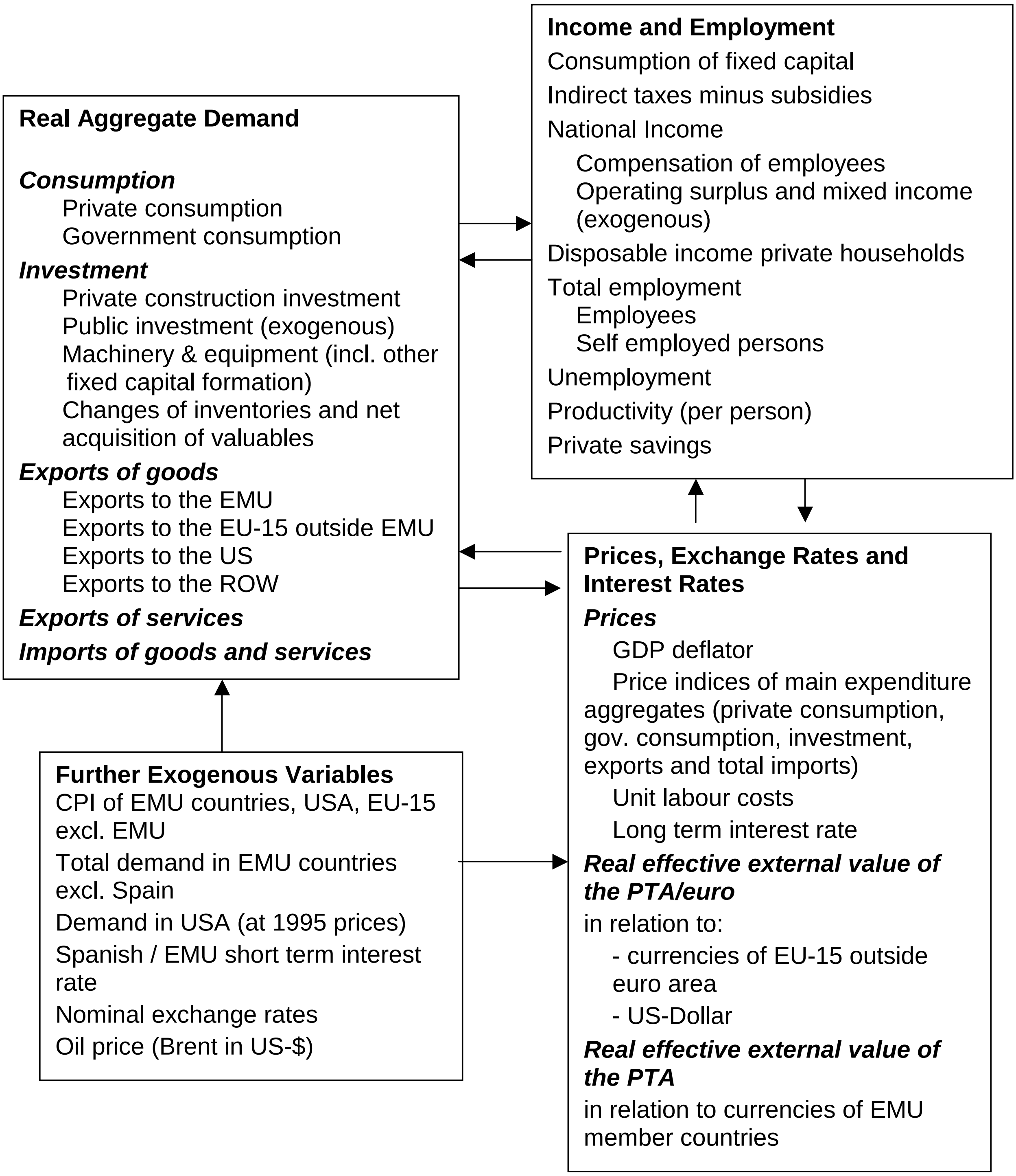


Figure 4.1: An illustration of the model structure

4.4 Aggregate Demand

4.4.1 Private Final Consumption Expenditure

Theoretical Considerations and Approaches Followed in Macroeconometric Models

Theoretical Considerations Different theoretical approaches offer a range of consumption hypotheses explaining the behaviour of private final consumption expenditure.

The so-called *absolute income hypothesis* was a popular explanation of private final consumption expenditure until the 1950s. It models private final consumption expenditure as a function of household disposable income. The marginal propensity to consume is assumed to lie between 0 and 1.

$$C_t = C_1(Y_t), 0 < \frac{dC_1}{dY_t} < 1 \tag{4.1}$$

This approach is often associated with John Maynard Keynes, although in his "General Theory of Employment, Interest and Money", Keynes offers several explanations of consumption behaviour ranging from income to revaluations of wealth or changes in the subjective discount factor of the individual as well as "expectations" (Keynes 1936, p. 91 ff.).

Brown (1952) published a new approach, which became known as the *habit persistence hypothesis*. It postulates that the individuals form habits which they keep, even if their income changes. Consumption today is thus a function of income and past consumption.

$$C_t = C_2(Y_t, C_{t-1}), 0 < \frac{\delta C_2}{\delta Y_t} < \frac{\delta C_2}{\delta C_{t-1}} < 1 \tag{4.2}$$

The *permanent income hypothesis* (Friedman 1957) and the *life cycle hypothesis* (Ando and Modigliani 1963) also offer an explanation of the relative smoothness of consumption. In these concepts the individuals anticipate their life-time income. This life-time income is the relevant income variable rather than the actual income of the respective period. Individuals are assumed to maximise their life-time utility, which is a function of consumption. The latter is subject to the budget constraint of the individual's life-time income. In this context the decision about today's consumption is an inter-temporal optimisation problem.

Under the assumption of no liquidity constraints all individuals can borrow at the prevailing real interest rate, which they are willing to do, if their subjective discount rate is higher than the prevailing real interest rate. Current disposable income is irrelevant in this framework. However, if we assume that - as we can observe in reality - at least some individuals cannot borrow freely i.e. face liquidity constraints, then current income plays a part in determining current consumption. The existence of liquidity constraints for some of the individuals is not denied and thus, besides wealth (in theory: assets + discounted future labour income) current income is a determinant of current consumption.

If the real interest rate is assumed to be constant and a significant part of the consumers is subject to liquidity constraints, the consumption function will be formulated as follows:

$$C_t = C_3(Y_t, W_{t-1}) \tag{4.3}$$

As wealth is a stock, it is included as lagged variable W_{t-1} referring to the value at the end of the previous period.

There is often a number of additional assumptions such as an infinite life-time. Models can discriminate between durable and non-durable goods. The share of individuals subject to liquidity constraints can be modelled explicitly. Precautionary saving can be taken into account. As the objective here is to derive a fairly simple macroeconometric consumption function based on highly aggregated data, these variations are not discussed in detail. For an extensive overview c.f. Muellbauer et al. (1999).

Hall (1978) criticised the inclusion of lagged values of income in regression estimations of consumption. He argued that under an expected life-time income today's consumption alone affects expected future utility (and consequently consumption). In other words this means that consumption follows a random walk. As the individuals form *rational* expectations, all relevant known information needed for the determination of future consumption is already included in today's consumption. No other lagged variables are needed. This implies that changes in consumption can be due only to unanticipated changes in permanent income. For forecasting the main conclusion from Hall's analysis is that C_{t-1} is the best forecast for C_t.

Overview of Consumption Functions As documented in Appendix B.6 finding an appropriate set of explaining variables for the consumption function proved difficult. Different sets of variables were tested for cointegration for a number of different periods. Besides the theoretical considerations, empirical results from other consumption functions served as a guideline. Church, Smith and Wallis (1996) have published an interesting survey of alternative consumption functions in seven models[4] of the British economy. All of them are based on the *life-cycle hypothesis/permanent income hypothesis of consumption* and therefore have disposable income and a wealth variable in their long-run relation. The same holds for the area wide model of the European Central Bank (Fagan, Henry, and Mestre 2001).

As Spain is not among the G7 countries there are only a few models of the Spanish economy[5]. From these the following four models based on quarterly data have been chosen for a comparative analysis of the consumption functions: the OECD's *Interlink*[6], the Spanish module of the ESCB-MCM (Estrada and Willman 2002)[7], its update (Estrada, Fernández, Moral, and Regil 2004) and the Spanish module of NIGEM. The consumption function of the latter has only recently been re-estimated within the framework of a study for the British treasury (Al-Eyd and Barrell 2004). It has recently been incorporated into the NIGEM. All consumption functions are estimated as error-correction models and thus follow a similar estimation methodology as the author.

Table 4.1 gives an overview of the variables explaining consumption in the four models and the estimated coefficients. [8]

[4] Of the eight equations analysed only one is not part of a macroeconomic model.

[5] c.f. Chapter 1

[6] There seems to be no detailed model documentation. The information provided in this paper is based on e-mail messages from Franck Sedillot and Hubert Strauß of August 1oth and July 15th, 2004

[7] Additional information was provided by Alpo Willman in his e-mail message of January 27th, 2004. He explained that the truncated step dummy used to model financial liberalisation equals 1 in the period from 1987Q1 until 1993Q1 and zero otherwise.

[8] As the focus is on the chosen variables and the interpretation of the estimated coefficients, standard

	Interlink	ESCB-MCM	ESCB-MCM (update)	NIGEM
Frequency	half-yearly	quarterly	quarterly	quarterly
Seasonal Adj.	SA	trend-cycle	SA	SA
Period	1980-2001	1981Q2 -1996Q4	1981Q1-1998Q4	1981Q2-2001Q2
Error correction term	-0.28	-0.30	-0.11	-0.08
$Constant$	0.009	0.37	0.24	0.23
ydr_{t-1}	-1.00		-0.96	-0.90
ydr_{t-3}		-1.00		
w_{t-1}			-0.04	-0.10
$(w/ydr)_{t-3}$		-0.02		
RLR_{t-1}			0.57	
DF_{t-3}		-0.03		
Δc_{t-2}			0.35	
Δydr	0.60		0.27	0.08
Δydr_{t-1}			0.11	
Δydr_{t-3}		0.26		
Δp	-0.001	-0.33		
$\sum_{i=0}^{5} \frac{1}{6} \Delta p_{t-i}$	-0.006			
Δrhw_{t-1}				0.06
Δrhw_{t-2}			0.14	
Δrfw_{t-2}			0.02	
RLR_t	-0.006			
$RLR_{t-1} + RLR_{t-4}$		-0.22		
ΔRLR				-0.0006
$\Delta RLR_{t-2} + \Delta RLR_{t-3}$			-0.14	
ΔDF_{t-3}		0.01		
ydr: real disposable income of households, w: real wealth, RLR: real long-term interest rate, DF: truncated step dummy, c: real private final consumption expenditure, p: consumer price index, rhw: real housing wealth, rfw: real financial wealth.				

Table 4.1: Selected consumption functions in the most important models of the Spanish economy

It is difficult to compare these consumption functions, because they apply to different time periods and employ different data sets. However, there are some similarities and some striking differences. The Spanish Module of the ESCB-MCM, its update and NIGEM include wealth in the consumption function. Whereas NIGEM and the updated version of the Spanish module of the ESCB-MCM employ real wealth in log levels, the

errors are not considered necessary for the overview below. In addition t-statistics/standard errors are available for only two of the three equations.

previous version of the Spanish Module of the ESCB-MCM uses the ratio of wealth over disposable income. The OECD, by contrast, models Spanish consumption according to the absolute income hypothesis. Whereas in Interlink as well as in the ESCB multi-country model the coefficient of real disposable income is restricted to -1, the *sum* of the coefficients of income and of wealth is restricted to -1 in NIGEM and the updated version of the Spanish module of the ESCB-MCM. For the Spanish modules in Interlink and the ESCB-MCM the adjustment coefficients in the consumption function are both near -0.3. The adjustment coefficients in the Spanish module of the updated version of the Spanish module of the ESCB-MCM as well as NIGEM are much lower in absolute terms (-0.11 and -0.08, respectively). The real long term interest rate appears in all four equations. Whereas it is included in levels in *Interlink* and the Spanish Module of the ESCB-MCM, it is included in differences in the consumption functions of the updated version of the Spanish module of the ESCB-MCM and in NIGEM. Obviously there are different assumptions about the stochastic properties of the real long term interest rate in different models. Another explanation may be different calculation methods.

From the analysis above we can conclude that, although estimation periods and available data do not differ too much, there is not *one* generally accepted specification of private final consumption in Spain. On the contrary, the consumption functions differ widely in terms of explanatory variables used and estimated coefficients. The differences are particularly striking for the estimated adjustment coefficients.

Cointegration Tests

As mentioned above finding a plausible cointegrating relationship for the consumption function proved extremely difficult. Following a "specific to general" approach beginning with a bivariate estimation a large number of variable combinations have been tested, starting with consumption and income as postulated by the absolute income hypothesis. It was assumed that income does have an influence on consumption in any case. Therefore, an attempt has been made to complement consumption and income by further variables until a cointegrating relationship is found. As in some of the consumption functions presented in the previous section, wealth variables were expected to fill the gap. Several series have been constructed and added to the bivariate system, but no cointegration has been found (for details c.f. Appendix B.6).

At last cointegration could be established between the following variables: private consumption, household disposable income, the nominal interest rate lagged by three quarters and a step dummy, which equals zero from 1980 until the fourth quarter of 1992 and 1 thereafter. The cointegrating relationship holds only for the period since 1986.

According to the approach of Engle and Granger the following equation in log levels is estimated for the period from 1986Q1 until 2002Q4.

$$
\begin{aligned}
cp95_t \;=\; & \underset{(0.271)}{0.332} - \underset{(0.006)}{0.025}\, Z_{1t} - \underset{(0.006)}{0.004}\, Z_{2t} - \underset{(0.006)}{0.008}\, Z_{3t} \\
& - \underset{(0.009)}{0.070}\, Z_{1t}SD - \underset{(0.009)}{0.046}\, Z_{2t}SD + \underset{(0.009)}{0.015}\, Z_{3t}SD \\
& - \underset{(0.005)}{0.029}\, SD9301_t + \underset{(0.024}{0.972}\, yd95_t - \underset{(0.006)}{0.058}\, NL_{t-3} + \hat{\nu}_t
\end{aligned}
\tag{4.4}
$$

Z_{it} are the centred seasonal dummies. $Z_{it}SD$ are centred seasonal dummies multiplied with a step dummy which is equal to one up to the fourth quarter of 1991 and zero afterwards. They model the change in the seasonal pattern of disposable income. $SD9301_t$ is the step dummy described above.

The residuals ($\hat{u}_t$) are analysed for their integration properties with an ADF-test. The test equation is estimated without constant beginning with four lags of differences eliminating insignificant lags step-by-step.

The t-value of the coefficient of the lagged residual series equals -4.44. Thus, the residuals are confirmed to be stationary according to the Mac Kinnon critical value of -4.12 cited in (Hassler 2004, Table 1, p. 111).

The Johansen procedure is carried out to test whether the results of the Engle/Granger approach hold in a multivariate context. Due to the difficulties in finding a plausible cointegrating relationship at all, the Johansen cointegration test has been repeated for a number of periods using the same set of variables. Two sets of centred seasonal dummies were included as in the Engle/Granger approach. Critical values corresponding to case III (cf. Appendix B.1) were used. As Table 4.2 shows the results are not very robust. Cointegration can only be confirmed for three of the 11 periods tested. However, for several other periods the trace statistic is close to the 5% critical value of 38.93 (Pesaran, Shin, and Smith 2000, p.339).

Sample	$cp95_t$	$yd95_t$	NL_{t-3}	$SD9301_t$	Lags (AIC)	Trace stat.
1982-2002	1.00	-0.96	0.01	0.06	1-4	34.91
1983-2002	1.00	-0.94	0.01	0.06	1-4	37.02
1984-2002	1.00	-0.91	0.01	0.06	1-4	39.50*
1985-2002	1.00	-0.86	0.01	0.05	1-4	48.14*
1986-2002	1.00	-0.90	0.01	0.05	1-4	46.04*
1982-2001	1.00	-0.97	0.01	0.06	1-4	35.93
1982-2000	1.00	-0.97	0.01	0.06	1-4	34.67
1982-1999	1.00	-0.99	0.006	0.04	1-4	35.12
1982-1998	1.00	-1.006	0.002	0.03	1-4	37.55
1982-1997	1.00	-1.02	-0.0005	0.02	1-4	37.76
1982-1996	1.00	-1.00	0.003	0.03	1-4	34.01

Table 4.2: Johansen tests: consumption, household disposable income, long-term nominal interest rate and step dummy

Both cointegration tests applied so far allow to reject the null hypothesis of no cointegration. Therefore we proceed with estimating an error-correction equation.

An Error-Correction Equation

The following equation is estimated for the period from 1986Q1 until 2002Q4.

$$\begin{aligned}
\Delta cp95_t = & -\underset{(-5.34)}{0.35}\,(-\underset{(-2.80)}{1.16} + \underset{(6.10)}{0.07}\,SD9301_t + c95_{t-1} - \underset{(-25.22)}{0.90}\,yd95_{t-1} + \underset{(6.60)}{0.01}\,NL_{t-4}) \\
& -\underset{(-5.08)}{0.07}\,Z_{1t} - \underset{(-3.15)}{0.02}\,Z_{2t} - \underset{(-3.19)}{0.05}\,Z_{3t} + \underset{(1.82)}{0.01}\,\Delta SD9301_{t-2} + \underset{(2.19)}{0.02}\,\Delta SD9301_{t-3} \\
& -\underset{(-4.21)}{0.31}\,\Delta yd95_{t-1} - \underset{(-4.08)}{0.20}\,\Delta yd95_{t-2} - \underset{(-1.87)}{0.11}\,\Delta yd95_{t-3} - \underset{(-2.60)}{0.25}\,\Delta cp95_{t-1} \\
& -\underset{(-4.99)}{0.41}\,\Delta cp95_{t-2} - \underset{(-2.03)}{0.20}\,\Delta cp95_{t-3} - \underset{(-6.98)}{0.01}\,\Delta UR_t - \underset{(-5.09)}{0.01}\,\Delta UR_{t-4} + \hat{\varepsilon}_t
\end{aligned} \tag{4.5}$$

According to the estimated error-correction equation the long-term elasticity of consumption with respect to disposable income equals 0.9.

With a t-value of -5.34 the adjustment coefficient is highly significant[9]. At 0.35 it is similar to those estimated in the consumption functions of *Interlink* and the Spanish module of the ESCB-Multi-Country Model.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.98
Residual tests	Probability
Normality test (Jarque-Bera)	[0.28]
Serial correlation LM test (lag 1)	[0.32]
Serial correlation LM test (lag 4)	[0.15]
Serial correlation LM test (lag 8)	[0.51]
ARCH LM test (lag 1)	[0.14]
ARCH LM test (lag 4)	[0.06]
White's heteroskedasticity test	[0.46]
RESET test (h=2)	[0.23]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.3: Diagnostic tests: private final consumption expenditure

The tests confirm that the residuals are well-behaved, i.e. the null hypotheses of no serial correlation, normal distribution and homoskedasticity cannot be rejected at the 5% level.

Both the CUSUM and the CUSUM of squares test suggest that the equation is stable over the estimation period (c.f. Appendix B.3).

The change in the seasonal pattern of real household disposable income is obviously reflected in the lagged disposable income. Therefore, the second set of seasonal dummies[10] is not needed. This is confirmed by the F-statistic of the redundant variables test. The null hypothesis of redundancy cannot be rejected. Consequently, the second set of seasonal dummies was removed from the equation.

[9] The 5 % critical value is: -3.91 (Hassler 2004, Table 4, p. 112).

[10] i.e. centred seasonal dummies multiplied by a step dummy equal to one up to the fourth quarter of 1991 and zero afterwards

The step dummy is required to model the level shift of the consumption series in the first quarter of 1993. As there is no similar level shift in any of the regressors or any other plausible economic explanation, the level shift in consumption is most probably the result of statistical problems in compiling national accounts according to the ESA 1995 for the period before 1995.

The change in the unemployment rate also affects private final consumption expenditure in the short run. The sign of the coefficient of the change in the unemployment rate is negative. This means - as one would expect - that consumption decreases as unemployment rises. If unemployment rises the risk of being unemployed rises as well, inducing households to increase their precautionary savings.

4.4.2 Government Consumption Expenditure

As in the case of private final consumption expenditure at constant prices of 1995, an income variable is the main explanatory variable of government consumption. As the government sector accounts are not available for the full period and revenue data are thus missing, GDP at constant prices, which is assumed to be closely correlated with government revenues, is chosen as a proxy.

GDP alone is not sufficient for modelling government consumption due to structural breaks in the government consumption series. The first is a change of the seasonal pattern in 1992, a feature we also see in other Spanish time series, the second is a change of the trend slope, also from 1992 onwards. This is why centered seasonal dummies multiplied with a step dummy (equal to one until 1991Q4 and zero afterwards) as well as an additional trend beginning in 1992 are included in the equation in log levels.

Cointegration Tests

As a first step we look at an equation in log levels.

$$\begin{aligned} cgov95_t \;=\; & \underset{(0.864)}{2.386} + \underset{(0.009)}{0.019}\; Z_{1t} + \underset{(0.008)}{0.004}\; Z_{2t} + \underset{(0.008)}{0.016}\; Z_{3t} \\ & - \underset{(0.011)}{0.021}\; Z_{1t}SD - \underset{(0.011)}{0.028}\; Z_{2t}SD - \underset{(0.011)}{0.003}\; Z_{3t}SD \\ & - \underset{(0.0003)}{0.005}\; KT9201_t + \underset{(0.001)}{0.008}\; t + \underset{(0.077)}{0.609}\; gdp95_t + \hat{\nu}_t \end{aligned} \tag{4.6}$$

The t-value of the lagged level of the residuals is -3.11 and thus not sufficient to confirm cointegration (5% critical value: -4.12). However, the two step approach in levels has a weak power and does not take into consideration the dynamic structure of the variables.

For the Johansen cointegration test a lag length of 4 is chosen according to both the AIC and the Schwarz criterion. The test finds at least two cointegrating relationships. Strictly speaking this would mean that one variable in the system is redundant. However, this would mean that either government consumption or GDP is stationary with respect to a broken trend - which, according to the unit root tests, is not the case - or GDP and government consumption are already cointegrated without adding the broken trend, which can also be denied. For this reason, the result of the Johansen cointegration test is ignored in the following.

Case	IV
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	57.97 (30.77)
Trace stat. $r \leq 1$	20.69 (15.44)

Table 4.4: Results of the Johansen cointegration test: government consumption

An Error-Correction Equation

The following error correction equation is estimated. The elasticity of government consumption with respect to GDP is roughly one. The short term dynamic consists mainly in the lagged dependent variable. A lag length of six is necessary to avoid autocorrelation in the residuals.

Contrary to the results of both the Engle/Granger approach and the Johansen test, cointegration is confirmed in the Stock approach. The t-value of the adjustment coefficient is below the 5% critical value of -4.12.

$$\begin{aligned}\Delta cgov95_t = & -\underset{(-5.08)}{0.28}\,(\underset{(1.01)}{2.15} + \underset{(8.49)}{0.005}\,KT9201_t - \underset{(-3.26)}{0.005}\,t + cgov95_{t-1} - \underset{(-5.35)}{1.02}\,gdp95_{t-1}) \\ & -\underset{(-3.16)}{0.02}\,Z_{1t} + \underset{(1.12)}{0.005}\,Z_{2t} - \underset{(-1.56)}{0.01}\,Z_{3t} - \underset{(-2.74)}{0.02}\,Z_{1t}SD + \underset{(1.55)}{0.01}\,Z_{2t}SD \\ & +\underset{(0.29)}{0.002}\,Z_{3t}SD - \underset{(-3.77)}{0.36}\,\Delta cgov95_{t-3} + \underset{(2.27)}{0.20}\,\Delta cgov95_{t-4} - \underset{(-1.82)}{0.17}\,\Delta cgov95_{t-6} \\ & -\underset{(-1.89)}{0.20}\,\Delta gdp95_{t-2} + \hat{\varepsilon}_t\end{aligned} \tag{4.7}$$

Except for some autoregressive conditional heteroskedasticity, particularly of first order, the residuals are well-behaved. The CUSUM and CUSUM of squares tests point to a problem of parameter instability.

Diagnostics	
Estimation period	1981Q4-2002Q4
Adjusted R^2	0.81
Residual tests	Probability
Normality test (Jarque-Bera)	[0.15]
Serial correlation LM test (lag 1)	[0.53]
Serial correlation LM test (lag 4)	[0.30]
Serial correlation LM test (lag 8)	[0.27]
ARCH LM test (lag 1)	[0.01]
ARCH LM test (lag 4)	[0.08]
White's heteroskedasticity test	[0.15]
RESET test (h=2)	[0.77]
Stability tests	outside error bands
CUSUM test	1993-1995
CUSUM of squares test	1994-2000

Table 4.5: Diagnostic tests: government consumption

4.4.3 Gross Capital Formation

Theoretical Considerations

There are basically four approaches for the explanation of investment behaviour: the accelerator approach, the neoclassical approach, Tobin's q, the financial accelerator.

The accelerator model (Chow 1967) postulates a stable relationship between the capital stock (K_t) and output (Y_t). Net investment (I_t^n), which is equivalent to a change in the capital stock, is initiated by a change in output. Formally, the core of the accelerator theory can be described by the following equations [11]:

$$K_t = \beta * Y_t \tag{4.8}$$

Net investment is then defined as:

$$\begin{aligned} I_t^n &= K_t - K_{t-1} \\ &= \beta * (Y_t - Y_{t-1}) \end{aligned} \tag{4.9}$$

As net investment is not usually available within the framework of national accounts statistics, the equations can be reformulated to apply to gross investment. For this purpose we have to recall that net investment equals gross investment less depreciation. This yields the following relationship for gross investment (I_t^g), where δ is the rate of depreciation:

$$\begin{aligned} I_t^g &= K_t - (1-\delta) * K_{t-1} \\ &= \beta * (Y_t - (1-\delta) * Y_{t-1}) \end{aligned} \tag{4.10}$$

The neoclassical approach (Jorgenson 1963) is the most widely applied one in macroeconometric models. It does not contradict the accelerator approach, but encompasses it. In this approach investment is seen as an adjustment of the capital stock to its profit-maximising level. The first order condition for maximum profits (under the constraint of the production function) is that the marginal product of capital equals its factor price. Thus, there is a long-run relationship between the capital stock, output

[11] Additional variables like relative prices can be added to explain an adjustment of the capital stock, but the emphasis is on the dynamics of output. For illustrative purposes, the presentation here is a simplified version of the model examined by Chow (1967).

and the user cost of capital. In the framework of a Cobb-Douglas production function the desired capital stock can be defined as follows:

$$K^* \;=\; \gamma * \frac{p{*}Y}{c}, \qquad (4.11)$$

where $p{*}Y$ is nominal output and c is the user cost of capital. The cost of capital includes a whole range of cost components: the real price of the capital good, interest rates, the depreciation rate, the corporate tax rate, which are partly offset by depreciation allowances and investment tax credits. γ is the elasticity of output with respect to capital.

The existence of adjustment costs explains, why investment is an ongoing process over several periods and not all additional capital is installed immediately. Thus, newly started investment projects (I_t^S) depend on investment projects started in past periods (I_t^P). Jorgenson (1963, p. 250) describes this relationship as a weighted average of past investment projects.

$$I_t^S \;=\; \sum_{i=0}^{\infty} w_i * I_{t-i}^P, \qquad (4.12)$$

where w is the share of projects completed in each period. Jorgenson goes into great depth describing lag structures. He differentiates between replacement investment and expansion investment. Net investment depends on the desired capital stock (determined by output, the price level, capital cost and parameters of the production function), past net investment and past values of the desired capital stock (or the variables defining it).

The neoclassical approach is often reproached with a lack of forward-looking elements. However, these can easily incorporated into the theory as Jorgenson shows.

Expectations are at the heart of the third explanation of investment: *Tobin's q approach.* Here investment is a function of "q", the ratio of a firm's market value to the replacement cost of its capital stock[12]. If q exceeds one, the firm has an incentive to invest, if it is below one, its capital stock will have to be reduced. For the decision

[12] Keynes' marginal efficiency of capital is a very similar concept: "The relation between the prospective yield of a capital asset and its supply price or replacement cost, i.e. the relation between the prospective yield of one more unit of that type of capital and the cost of producing that unit, furnishes us with the marginal efficiency of capital of that type." (Keynes 1936, p. 135)

marginal q rather than average q is relevant. This is the effect of one additional unit of capital on the present value of additional profits. As Hayashi (1982) points out this approach can easily be incorporated into the neoclassical paradigm.

Lately more and more attention is being paid to the transmission process of monetary policy and an additional explanation of cyclical fluctuations of investment has been found in the *financial accelerator*. This approach explains changes in investment by changes of the financing conditions of investment over the business cycle. At the beginning of an upswing investment can usually be financed internally by rising profits, but as the expansion continues internal funds are no longer sufficient and the importance of external financing increases. As long as the overall economy is in good shape, the value of the required collateral is usually high and external finance is thus easily accessible. The situation changes in a downturn or recession, particularly, when it is accompanied by an increase in interest rates. Then in addition to rising costs of finance, revenues usually decrease leading to liquidity problems and assets that could be used as collateral lose value. At the same time banks find themselves in a situation of asymmetric information. Problems of adverse selection and moral hazard increase in a downturn. This affects particularly small companies, for which information in the form of the share price is not available. These firms find it increasingly hard to obtain external finance. This change of the availability of external finance over the cycle acts as a pro-cyclical *financial accelerator* of the investment activity.

Theories of investment are usually concerned with private business investment. They do not differentiate between different types of investment such as construction or machinery and equipment. Sometimes housing is mentioned as a special case, because it is at the boundary of investment and purchase of a consumer durable. Chow (1967) points out that the interest effect on investment differs according to the durability of the respective investment goods. It can be assumed that the interest elasticity is smaller for less durable goods. Therefore, he suggests to classify investment according to the durability of the investment goods (Chow 1967, p.2).

All of the theories highlighted above are difficult to implement in a macroeconometric model. Both the accelerator and the neoclassical approach postulate a relationship between output and the capital stock, rather than output and investment. A permanent rise in output would thus lead to a temporary increase in investment. This is not what we observe in the data. The available data of the Spanish capital stock[13] are not stationary in differences, i.e. the capital stock is integrated of order 2.

In addition the capital stock is not included in the model described here, because it cannot be modelled consistently and the data are questionable. Nevertheless, an estimation of a capital stock proxy with output as an explaining variable has been tried out. The capital stock has been calculated as follows:

$$K_t = K_{t-1} + IFC95 - (CFC/(PIFC/100)) \tag{4.13}$$

The starting value was taken from the series used in the Spanish Module of the ESCB-Multi-Country Model (Estrada and Willman 2002). This series of the capital stock is also I(2). Nevertheless, the Johansen cointegration test finds a cointegrating relationship between real GDP and this capital stock proxy. However, an inclusion of the ensuing error correction model in the model resulted in an interruption of the baseline simulation due to negative values for national accounts aggregates, which do not normally turn negative. The approach is thus disqualified.

Tobin's q seems an elegant approach - in theory. However, there are just too many unknowns. Even if we assume that marginal q is equal to average q (which is questionable) and we could model the capital stock, we would still need a measure of the present value of future profits. Often the share price or share price index (on an aggregate level) is used as a proxy. In a macroeconometric model, however, we would have to estimate an equation for the share price index to make it an endogenous variable. It is a well-known fact, that econometrically the share price index is best characterised as a *random walk*, for which the best available forecast is yesterday's value. For longer forecast horizons, say, two to three years, however, such an approach would hardly make sense.

The "financial accelerator" approach does not per se explain investment behaviour, but rather how its pro-cyclical effects exacerbate a downturn, which has already set in. For an estimate, again, we would need a whole set of financial market data, which

[13] The author has tested the series used in the Spanish Module of the ESCB-Multi-Country-Model (Estrada and Willman 2002) and a temporal disaggregation of the annual series given in the OECD Economic Outlook No. 74.

are hard to endogenise in a macroeconometric model. In single equation attempts to explain investment, however, combining this approach with the other two offers additional insights as a study of the NIESR for the G7 countries has shown (Ashworth and Davis 2001).

Thus, the inclusion of the long-term interest rate is the only element in the following, which is compatible with the theoretical considerations above. Output is not plausible as an explanatory variable for *investment* according to the theories above.[14].

Gross Fixed Capital Formation: Private Construction

As mentioned above investment is usually disaggregated into three sub-aggregates: public investment, private housing investment, private business investment. A closer look at the times series, however, confirmed that private construction other than housing is much more similar to private housing investment in its dynamics than to investment into machinery and equipment. This is why all private construction is estimated jointly in the following equation.

Housing accounts for almost half of Spanish construction investment. Therefore household disposable income has been chosen as demand variable here. An alternative with GDP95 has been tried out but rejected, because of autocorrelation in the residuals and because interest rates would no longer be significant in levels, which the author doubts. Contrary to any theoretical basis the long term nominal interest rate rather than the (long or short term) real interest rate helps to explain construction investment. This already holds for consumption and may in fact be a problem of a backward looking model with no adequate modelling of inflation expectations.

Cointegration Tests Again a changing seasonal pattern is captured by centred seasonal dummies multiplied with a step dummy. A level shift is modelled by a step dummy. The low Durbin-Watson statistic already indicates, that establishing a cointegrating relationship with an estimation in levels alone may prove difficult. Another striking fact is the long lag length of interest rates. However, given the necessity of obtaining a building

[14] Fair (2004) uses lagged values of investment, output and interest rates in the rest-of-the-world module of his well-known multi-county model.

permit prior to beginning the construction activity this is not totally implausible.

$$
\begin{aligned}
icon95pr_t = & \underset{(0.822)}{-11.95} - \underset{(0.015)}{0.09}\ Z_{1t} - \underset{(0.015)}{0.05}\ Z_{2t} - \underset{(0.015)}{0.03}\ Z_{3t} \\
& - \underset{(0.025)}{0.10}\ Z_{1t}SD - \underset{(0.025)}{0.05}\ Z_{2t}SD + \underset{(0.025)}{0.04}\ Z_{3t}SD \\
& - \underset{(0.014)}{0.10}\ SD9201_{t-1} + \underset{(0.073)}{1.90}\ yd95_t - \underset{(0.002)}{0.01}\ NL_{t-7} + \hat{\nu}_t
\end{aligned}
\tag{4.14}
$$

In the ADF-Test of the residuals three lags must be included to eliminate all remaining autocorrelation. According to this test the residuals cannot be considered stationary. The relevant t-statistic is -3.74 compared to a 5% critical value of -4.12.

In contrast the Johansen Cointegration test confirms the existence of a cointegrating relationship at the 5% level, if the lag length of five according to the AIC is chosen.

Case	III
exogenous variables	1
endogenous variables	3
Trace stat. $r = 0$	50.59 (38.93)
Trace stat. $r \leq 1$	13.78 (23.32)

Table 4.6: Results of the Johansen cointegration test: private construction investment

An Error-Correction Equation An error-correction equation with a significant adjustment coefficient is estimated. A cointegrating relationship between the log of private construction investment, real household disposable income, the nominal interest rate lagged by seven quarters and a step dummy (1992Q1) could thus be established. The coefficients look reasonable, but it has to be noted that they differ from those estimated in the Johansen test.

$$
\begin{aligned}
\Delta icon95pr_t \;=\; & -\underset{(-4.14)}{0.22}\,(\underset{(2.11)}{5.25} + \underset{(4.34)}{0.18}\; SD9201_{t-1} + icon95pr_{t-1} - \underset{(-6.18)}{1.33}\; yd95_{t-1} \\
& + \underset{(4.67)}{0.04}\; NL_{t-8}) - \underset{(-30.35)}{0.22}\; Z_{1t} - \underset{(-0.49)}{0.003}\; Z_{2t} - \underset{(-11.18)}{0.08}\; Z_{3t} \\
& + \underset{(3.78)}{0.07}\; ID9501_t - \underset{(-4.75)}{0.46}\; \Delta yd95_{t-2} + \underset{(2.30)}{0.007}\; \Delta NL_{t-2} \\
& - \underset{(-3.20)}{0.010}\; \Delta NL_{t-3} - \underset{(-3.16)}{0.010}\; \Delta NL_{t-7} + \underset{(2.69)}{0.009}\; \Delta NL_{t-8} + \hat{\varepsilon}_t
\end{aligned}
\tag{4.15}
$$

The residuals show no autocorrelation, they can be assumed to be homoskedastic and normally distributed. The RESET test indicates that the linear specification is suitable. However, according to the CUSUM of squares test the specification is not stable for the period from 1995 to 1997. The ARCH LM test suggests that there may be autoregressive heteroskedasticity of fourth order in the residuals (c.f. Table 4.7).

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.96
Residual tests	Probability
Normality test (Jarque-Bera)	[0.45]
Serial correlation LM test (lag 1)	[0.75]
Serial correlation LM test (lag 4)	[0.48]
Serial correlation LM test (lag 8)	[0.52]
ARCH LM test (lag 1)	[0.60]
ARCH LM test (lag 4)	[0.06]
White's heteroskedasticity test	[0.46]
RESET test (h=2)	[0.13]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1995-1997

Table 4.7: Diagnostic tests: construction investment

Gross Fixed Capital Formation: Machinery and Equipment

Investment in machinery and equipment was initially modelled as a function of GDP at constant prices and the nominal long-term interest rate. However, this specification introduced a destabilising feedback effect to the model. Any errors in either GDP or investment into machinery and equipment would build up into huge errors. Therefore, GDP has been replaced with the sum of goods exports and construction investment, which can both be expected to affect investment into machinery and equipment. The long-term interest rate has been kept as a regressor. As in the initial equation a step dummy is required to reflect a structural break.

Cointegration Test The estimation in levels yielded a coefficient with the wrong sign for the long-term interest rate. This contrasts with the results from the Johansen cointegration test and the direct estimation of the error-correction equation. Here, the long-term nominal interest rate showed a negative sign and it was not possible to confirm cointegration without it. For this reason the results of the Engle/Granger approach are regarded as spurious and we proceed directly to the Johansen test and the error correction equation.

The Johansen test suggests exactly one cointegrating relationship between the four variables: machinery and equipment at prices of 1995 in logs (imeq95), the sum of goods exports and construction investment at prices of 1995 in logs (xgicon95), the nominal long-term interest rate laged by two quarters (NL(-2)) and a step dummy for the second quarter of 1992(SD9202). The lag length of the VAR is 5 as recommended by the AIC. The results are reported in Table 4.8:

Case	III
exogenous variables	1
endogenous variables	3
Trace stat. $r = 0$	46.57 (38.93)
Trace stat. $r \leq 1$	9.27 (23.32)

Table 4.8: Results of the Johansen cointegration test: investment into machinery and equipment

An Error-Correction Equation The following error-correction equation is estimated with the variables mentioned above. Modelling investment into machinery and equipment is rather difficult, as no other time series in the model shows the same decline at the end of the sample as investment into machinery and equipment. To stabilise the model at the end of the sample period a step dummy beginning in the third quarter of 2001 has been added. For this reason no out-of-sample forecast is carried out.

$$\begin{aligned}\Delta imeq95_t = & -\underset{(-5.55)}{0.24}\,(-\underset{(-2.71)}{2.52}+\underset{(7.16)}{0.33}\,SD9201_{t-1}+imeq95_{t-1}-\underset{(-8.33)}{0.71}\,xgicon95_{t-1}\\ & +\underset{(3.30)}{0.03}\,NL_{t-3})-\underset{(-2.50)}{0.03}\,SD0103_t-\underset{(-3.01)}{0.10}\,Z_{1t}-\underset{(-2.25)}{0.03}\,Z_{2t}-\underset{(-4.14)}{0.14}\,Z_{3t}\\ & -\underset{(-1.85)}{0.19}\,\Delta imeq95_{t-1}+\underset{(2.22)}{0.22}\,\Delta imeq95_{t-4}-\underset{(-1.78)}{0.18}\,\Delta xgicon95_{t-4}+\hat{\varepsilon}_t\end{aligned} \tag{4.16}$$

The t-value of the adjustment coefficient is below the 5% critical value of -3.91, which provides some evidence of a cointegrating relationship. The residuals are uncorrelated and normally distributed. However, we have to reject the null hypothesis of the RESET test that the specification is appropriate. According to the CUSUM and the CUSUM of squares tests there are no stability problems.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.943
Residual tests	Probability
Normality test (Jarque-Bera)	[0.90]
Serial correlation LM test (lag 1)	[0.43]
Serial correlation LM test (lag 4)	[0.59]
Serial correlation LM test (lag 8)	[0.32]
ARCH LM test (lag 1)	[0.59]
ARCH LM test (lag 4)	[0.50]
White's heteroskedasticity test	[0.43]
RESET test (h=2)	[0.00]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.9: Diagnostic tests: investment into machinery and equipment

Change of Inventories and Net Acquisition of Valuables

The series to be explained is characterised by a significant seasonal pattern and a number of serious outliers. It can best be explained by its own lags and is thus exogenous to the model.

The series is stationary. As the changes of inventories and net acquisition of valuables can be both positive and negative, no transformation into logs is carried out.

$$\begin{aligned} IS95_t = \; & \underset{(3.01)}{254.10} + \underset{(0.59)}{135.56}\, Z_{1t} - \underset{(-2.14)}{641.00}\, Z_{2t} - \underset{(-3.11)}{876.95}\, Z_{3t} \\ & - \underset{(-1.49)}{0.12}\, IS95_{t-3} + \underset{(3.43)}{0.29}\, IS95_{t-4} + \underset{(2.31)}{0.19}\, IS95_{t-8} \\ & - \underset{(-3.24)}{2091.79}\, ID9003_t + \underset{(2.76)}{1762.24}\, ID9004_t + \underset{(2.84)}{1834.591}\, ID9701_t \\ & - \underset{(-2.86)}{1834.56}\, ID9703_t + \underset{(3.54)}{2277.76}\, ID0001_t - \underset{(-2.43)}{1628.56}\, ID0004_t + \hat{\varepsilon}_t \end{aligned} \tag{4.17}$$

Despite the lag length of eight, some higher order auto correlation (7,8,11,12) remains in the residuals. The null hypotheses of normal distribution and no heteroskedasticity in the residuals cannot be rejected. The estimation is stable over the estimation period from 1982Q1 until 2002Q4.

Diagnostics	
Estimation period	1982Q1-2002Q4
Adjusted R^2	0.78
Residual tests	Probability
Normality test (Jarque-Bera)	[0.07]
Serial correlation LM test (lag 1)	[0.64]
Serial correlation LM test (lag 4)	[0.40]
Serial correlation LM test (lag 8)	[0.05]
ARCH LM test (lag 1)	[0.56]
ARCH LM test (lag 4)	[0.17]
White's heteroskedasticity test	[0.98]
RESET test (h=2)	[0.07]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.10: Diagnostic tests: changes of inventories and net acquisition of valuables

4.4.4 Exports of Goods and Services

Theoretical considerations

From a theoretical point of view the modelling of export quantities is fairly straight forward. It is reflected by a standard demand function, where demand depends on income and the set of prices of the goods involved[15]. In the case of two countries and two goods this approach would yield:

$$\begin{aligned} X_E &= f(P_E, P_I, Y_I) \\ &= f(\tfrac{P_E}{P_I}, \tfrac{Y_I}{P_I}) \end{aligned} \tag{4.18}$$

Here X stands for the quantity exported. The subscript E refers to the exporting country, whereas the subscript I stands for the importing country. All prices and quantities are assumed to be denominated in the exporting country's currency. We have thus three explanatory variables[16]. The price level of the importing country P_I, the price level of the exporting country P_E and the income of the importing country Y_I. According to the theory we expect the following reactions of the export volume to changes of the variables:

- The export volume is affected negatively by an increase of the price level in the exporting country. In other words the own-price elasticity of exports is negative.
- The export volume reacts positively to an increase of the price level of the importing country or the cross-price elasticity is positive.
- The export volume increases, if the income of the importing country increases. This means that the income elasticity of the export volume is positive.

This model implies that imports and domestic goods are imperfect substitutes. Otherwise we would have to extend the model by a supply variable for the importing country (Leamer and Stern 1970, p.11). Persistent international price differences as well as differing commodity structures of foreign trade justify this assumption.

[15] A thorough discussion of the determinants of exports and imports can be found in Leamer and Stern (1970).

[16] With the pace of globalisation increasing rapidly, more recent literature mentions the international division of labour as an additional variable. Usually it is approximated by a deterministic time trend. Cf. for example Strauss (2002) or Stephan (2002).

The two-country-model is an extreme simplification that does not apply to the real world, where the typical industrial country exports to far more than 100 countries. Thus, we would actually have to expand our theoretical model to the n-country case. Whereas much of the above still holds for the multi-country case, an additional aspect must be taken into account: competition from third countries. Thus, in fact Spain's exports to the US depend not only on the bilateral relative prices and exchange rate, but also on the relative price of other competitors like France or Mexico.

Taking these effects into account would entail enormous additional data problems, determining the shares of the competitors in the respective market and their prices. It would also require additional disaggregation of data, because exporting to the same destination does not necessarily make a third country a competitor. Only exporting the same kind of commodities does. So we would have to look at disaggregated imports of the partner country with respect to regions *and* commodities. This seems an almost impossible task for the framework of a small, highly aggregated macro model.

Generally, it has been shown that a higher level of disaggregation improves the estimation quality - as long as the quality of the data does not deteriorate with the disaggregation, which is often the case (Goldstein and Khan 1985). The choice of the aggregation level of the export function for Spain is guided by two considerations:

- The increasing importance of the euro area is to be reflected in the model. The remaining goods exports are split into three aggregates, mainly because this makes modelling easier due to the choice of appropriate price and/or demand variables for the exports to the US and to the rest of the EU.

- The level of disaggregation should remain within manageable proportions and be appropriate for a small highly aggregated macro model (cf. section 4.1).

Exports of Goods to the Euro Area

Exports to the euro area account for about 60% of Spanish goods exports. The euro area is thus the most important destination of Spanish exports. However, Spain's market share in the euro area was only slightly less than 4% in the second half of the 1990s (Bravo and Gordo 2003). In the euro area France, Germany, Italy and Portugal are Spain's biggest export markets. The importance of Portugal (10%) for Spain's goods exports is one reason, why the cpi rather than unit labour cost is used in the calculation of the real effective exchange rate[17].

Real total demand (i.e. consumption + investment + exports = GDP + imports) in the euro area countries has been chosen as the relevant income variable. Whereas Germany exports mainly investment goods, more than 80% of Spanish goods exports are made up of consumption goods and intermediate goods. Investment goods account for slightly more than 10% of total goods exports. This justifies the use of a rather broad demand aggregate.

The need to model the increasing international division of labour has already been mentioned. In a regionally disaggregated framework one has to be careful with the use of such trends, as increasing integration in one area may actually come at the expense of trade with another. However, in the case of the euro area the continuation of the upward trend of the division of labour can safely be assumed. Therefore the inclusion of a trend in the equation of exports to the euro area has been tested, particularly because the high elasticity of exports with respect to the demand variable is worrisome. However, in the model with trend the elasticity still remained above 2 and the stability of the model weakened significantly as the CUSUM and CUSUM of squares tests showed. From an econometric point of view there is an additional argument against including a trend in the long-run relationship: a cointegrating relationship has already been established without the trend.

Cointegration Tests First a regression in log levels is carried out. If only total demand of the other euro area countries and the real effective exchange rate are included, the residuals are not stationary. They exhibit a level shift at the beginning of 1999, the year, when the euro was introduced. Obviously the move to the common currency caused a structural break in the cointegrating relationship. For this reason a step dummy, which is zero until 1998Q4 and 1 afterwards is added and proves highly significant. With the step dummy added, the null hypothesis of a unit root in the residuals can be rejected. The relevant t-statistic is -6.60 (5% critical value: -4.12).

[17] For Portugal unit labour cost data are not available for the whole estimation period.

$$xg95ewu_t = \underset{(0.714)}{-37.79} + \underset{(0.020)}{0.21}\ Z_{1t} + \underset{(0.020)}{0.10}\ Z_{2t} - \underset{(0.020)}{0.03}\ Z_{3t} + \underset{(0.047)}{3.65}\ ewuoes\_dtot_t - \underset{(0.116)}{1.08}\ rawewu_t - \underset{(0.029)}{0.15}\ SD9901_t + \hat{\nu}_t \tag{4.19}$$

The Johansen cointegration test confirms this result. The lag length of the VAR is 2 according to the AIC. Results are reported in Table 4.11:

Case	III
exogenous variables	1
endogenous variables	3
Trace stat. $r = 0$	41.94 (38.93)
Trace stat. $r \leq 1$	4.88 (23.32)

Table 4.11: Results of the Johansen cointegration test: exports of goods to the euro area

An Error-Correction Equation The elasticities estimated in the error-correction model are similar to those estimated in log levels and in the Johansen cointegration test.

$$\Delta xg95ewu_t = - \underset{(-6.60)}{0.60}\ (\underset{(40.04)}{36.01} + \underset{(5.35)}{0.16}\ SD9901_t + xg95ewu_{t-1} - \underset{(-57.38)}{3.51}\ ewuoes\_dtot_{t-1} + \underset{(7.02)}{1.02}\ rawewu_{t-1}) + \underset{(1.89)}{0.08}\ Z_{1t} - \underset{(-0.85)}{0.02}\ Z_{2t} - \underset{(-4.38)}{0.16}\ Z_{3t} + \underset{(4.25)}{2.30}\ \Delta ewuoes\_dtot_t - \underset{(-3.19)}{0.79}\ \Delta rawewu_{t-5} + \hat{\varepsilon}_t \tag{4.20}$$

Short-run fluctuations of goods exports to the euro area are determined by contemporaneous changes of euro area demand and the lagged change of the real effective exchange rate.

The residuals are well behaved (c.f. Table 4.12. The null hypothesis of the RESET test of no mis-specification cannot be rejected. The parameters are stable over the sample period. Despite this result the rather poor out-of-sample performance indicates some stability problems at the launch of EMU despite the inclusion of the step dummy (c.f. Appendix B.3).

Diagnostics	
Estimation period	1981Q3-2002Q4
Adjusted R^2	0.90
Residual tests	Probability
Normality test (Jarque-Bera)	[0.50]
Serial correlation LM test (lag 1)	[0.98]
Serial correlation LM test (lag 4)	[0.84]
Serial correlation LM test (lag 8)	[0.54]
ARCH LM test (lag 1)	[0.87]
ARCH LM test (lag 4)	[0.65]
White's heteroskedasticity test	[0.39]
RESET test (h=2)	[0.93]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.12: Diagnostic tests: exports of goods to the euro area

Exports of Goods to the Rest of the EU-15

The rest of the European Union[18] accounts for only slightly more than 10% of Spain's total goods exports. This share has changed only marginally over the past 20 years. Almost 90% of these exports go to the United Kingdom, Spain's fourth biggest export market (average 1980-2002). In 2002 more than half of the total were consumer goods, about one third intermediate goods and only about one tenth capital goods. This suggests the use of a broad income aggregate such as total demand. For the relative prices a real effective exchange rate of the Spanish peseta[19] with respect to the rest of the EU-15 has been calculated using the respective weights of each year in the sample.

Cointegration Tests Whereas the real effective exchange rate proves highly significant and shows the right sign, the demand variable does not seem to influence the evolution

[18] EU-15 excluding the euro area, i.e. the United Kingdom, Sweden and Denmark.

[19] The peseta exchange rates have been extended beyond the introduction of the euro, using the euro conversion rate of 166.386 pesetas per euro.

of exports to the rest of the EU-15 in levels. Instead a linear trend enters the equation in levels with a positive sign.

$$\begin{aligned} xg95reu_t = &+ \underset{(0.456)}{8.99} + \underset{(0.030)}{0.05}\ Z_{1t} - \underset{(0.030)}{0.01}\ Z_{2t} - \underset{(0.030)}{0.16}\ Z_{3t} \\ &+ \underset{(0.0004)}{0.02}\ t - \underset{(0.102)}{0.65}\ rawreu_t + \hat{\nu}_t \end{aligned} \qquad (4.21)$$

According to the ADF-Test the null hypothesis of a unit root in the residuals can be rejected at the 5% level. The respective t-value is -4.91 (-3.78).

The test result can be confirmed by the Johansen cointegration test only if the lag length of the VAR is set to one according to the Schwarz criterion. The Akaike criterion recommends a lag length of three, for which a cointegrating relationship cannot be confirmed.

Case	IV
exogenous variables	0
endogenous variables	2
Trace stat. $r = 0$	28.26 (25.77)
Trace stat. $r \leq 1$	3.87 (12.39)

Table 4.13: Results of the Johansen cointegration test: exports of goods to the rest of the EU-15

An Error-Correction Equation Whereas total demand of the three EU-15 countries outside the euro area does not play a significant part in the long-run relationship, it exerts a significant and strong influence in the short run. The elasticity of the real effective exchange rate is 0.4 and thus much lower than in the case of exports to the euro area. Several outliers have to be eliminated by impulse dummies.

$$\Delta xg95reu_t = \underset{(-5.07)}{-\ 0.38} \,(\underset{(-11.54)}{8.07} - \underset{(-35.62)}{0.02}\, t + xg95reu_{t-1} + \underset{(2.79)}{0.44}\, rawreu_{t-1})$$
$$+ \underset{(0.25)}{0.02}\, Z_{1t} + \underset{(2.94)}{0.15}\, Z_{2t} - \underset{(-3.44)}{0.11}\, Z_{3t} + \underset{(2.34)}{1.48}\, \Delta reu\_dtot_t$$
$$+ \underset{(3.49)}{2.31}\, \Delta reu\_dtot_{t-1} - \underset{(-2.94)}{0.24}\, \Delta xg95reu_{t-1} - \underset{(-2.85)}{0.21}\, \Delta xg95reu_{t-1}$$
$$- \underset{(-3.50)}{0.20}\, ID8201_t + \underset{(3.85)}{0.23}\, ID8401_t - \underset{(-3.38)}{0.19}\, ID9103_t$$
$$- \underset{(-3.61)}{0.21}\, ID9203_t + \hat{\varepsilon}_t \tag{4.22}$$

The residuals are well-behaved. Despite the result of the CUSUM of squares test, which points to some phases of parameter instability, the out-of-sample forecast (cf. Figure B.34 in Appendix B.3) shows an excellent forecast performance.

Diagnostics	
Estimation period	1980Q4-2002Q4
Adjusted R^2	0.84
Residual tests	Probability
Normality test (Jarque-Bera)	[0.41]
Serial correlation LM test (lag 1)	[0.71]
Serial correlation LM test (lag 4)	[0.78]
Serial correlation LM test (lag 8)	[0.73]
ARCH LM test (lag 1)	[0.56]
ARCH LM test (lag 4)	[0.25]
White's heteroskedasticity test	[0.24]
RESET test (h=2)	[0.10]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1993Q3, 1996Q4-1998Q1

Table 4.14: Diagnostic tests: exports of goods to the rest of the EU-15

Exports of Goods to the United States of America

Exports to the United States of America account for only slightly more than 4% of total Spanish goods exports. In the 1980s this share had exceeded 10% in some quarters. The US is thus a minor trading partner, but still ranks sixth among Spain's major export markets (average 1980-2002).

Cointegration Tests Spanish exports to the United States of America are determined by US total demand at prices of 1995 as well as the bilateral real exchange rate. The estimation period begins in 1980.

$$\begin{aligned} xg95us_t \;=\; & \underset{(0.912)}{-\,11.57} - \underset{(0.040)}{0.09}\; Z_{1t} - \underset{(0.040)}{0.03}\; Z_{2t} - \underset{(0.040)}{0.10}\; Z_{3t} \\ & + \underset{(0.061)}{1.52}\; us\_dtot95_{t-1} - \underset{(0.072)}{0.79}\; rawus_t + \hat{\nu}_t \end{aligned} \tag{4.23}$$

The null hypothesis of a unit root in the residual can be rejected only at the 10% level: -3.64 (5%:-3.78, 10%: -3.50).

In the Johansen cointegration test we can reject the first null hypothesis of no cointegration, but we cannot reject the second null hypothesis of at most one cointegrating relationship. For the VAR a lag length of 2 has been chosen in line with both information criteria.

Case	III
exogenous variables	0
endogenous variables	3
Trace stat. $r = 0$	36.64 (31.54)
Trace stat. $r \leq 1$	3.26 (17.86)

Table 4.15: Results of the Johansen cointegration test: exports of goods to the United States of America

An Error-Correction Equation The adjustment coefficient of the error-correction term is highly significant. Its t-value is -8.60 (5% critical value: -3.69) and thus corresponds to a confidence level of far less than 1%. The elasticity of exports with respect to the real exchange rate is -0.66. This means that an appreciation of the Peseta/Euro causes goods exports to the US to decline by 0.66%. An increase of demand by 1% raises Spanish goods exports to the US by 1.35%.

$$\begin{aligned}
\Delta xg95us_t = & - \underset{(-8.60)}{0.80} (\underset{(13.33)}{9.79} + xg95us_{t-1} - \underset{(-26.40)}{1.35} \; us_dtot95_{t-1} \\
& + \underset{(11.22)}{0.66} \; rawus_{t-1}) - \underset{(-3.05)}{0.08} \; Z_{1t} - \underset{(-0.51)}{0.01} \; Z_{2t} - \underset{(-3.14)}{0.08} \; Z_{3t} \\
& + \underset{(3.40)}{4.65} \; \Delta us_dtot95_{t-1} + \underset{(1.84)}{2.62} \; \Delta us_dtot95_{t-2} + \underset{(2.74)}{3.21} \; \Delta us_dtot95_{t-4} \\
& + \underset{(2.58)}{0.51} \; \Delta rawus_{t-1} + \hat{\varepsilon}_t
\end{aligned} \tag{4.24}$$

The residuals are well-behaved. There are no signs of autocorrelation, heteroskedasticity or outliers in the residuals. However, the RESET test indicates that the specification may not be ideal.

Diagnostics	
Estimation period	1981Q2-2002Q4
Adjusted R^2	0.62
Residual tests	Probability
Normality test (Jarque-Bera)	[0.43]
Serial correlation LM test (lag 1)	[0.83]
Serial correlation LM test (lag 4)	[0.33]
Serial correlation LM test (lag 8)	[0.66]
ARCH LM test (lag 1)	[0.15]
ARCH LM test (lag 4)	[0.38]
White's heteroskedasticity test	[0.32]
RESET test (h=2)	[0.06]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.16: Diagnostic tests: exports of goods to the United States of America

The parameters are stable during the whole estimation period and the out-of-sample forecast shows an excellent performance (cf. Figure B.35 in Appendix B.4).

Exports of Goods to the Rest of the World

The share of the rest of the world in Spain's goods exports declined sharply after the country entered the European Union in 1986. Whereas the share of countries outside the EU-15 and the US was still above 40% in the early 1980s, it is now only slightly higher than a quarter.

Modelling Spain's exports to the rest of the world is not an easy task as there is neither an appropriate demand variable nor a real effective exchange rate, which covers this particular group of countries. The share of most countries in this group is below one percent of Spain's total exports. Many of the countries involved offer only inadequate statistics.

Therefore, the OECD's index of the real effective exchange rate is used. It can be expressed as the product of the domestic price level, the reciprocal of the (weighted) foreign price level and the (weighted) nominal exchange rate. This relationship is used in the following equations. By dividing the overall real effective exchange rate by the Spanish CPI we obtain a new variable, which reflects the foreign price level and a weighted exchange rate. This variable is treated as exogenous. Changes of the Spanish price level will thus affect the overall real effective exchange rate in the model. This means that the real effective exchange rate is at least partly endogenous.

Of course, the OECD's real effective exchange rate can only be a rough proxy, as it also includes countries of the euro area or the EU-15 as well as the US, which are not part of the "rest of the world" as it is defined here.

The IMF offers an annual estimate of world GDP. The series is published with a considerable delay. In addition temporal disaggregation of World GDP and the subtraction of the GDP of the EU-15 (in US-Dollars) and of the US are time consuming. The result is not better than a simple linear time trend. Therefore, the deterministic trend rather than a complicated calculation of GDP is included in the following equation.

Cointegration Tests The introduction of the Euro in 1999 caused the second structural break for exports to the rest of the world after 1986. This effect is modelled with a step dummy. Due to the enormous structural change following EU accession, the estimation period begins in 1986.

$$xg95row_t = \underset{(0.565)}{14.03} - \underset{(0.021)}{0.03}\ Z_{1t} - \underset{(0.021)}{0.03}\ Z_{2t} - \underset{(0.021)}{0.11}\ Z_{3t} + \underset{(0.001)}{0.02}\ t - \underset{(0.123)}{1.56}\ raw_t - \underset{(0.026)}{0.12}\ SD9901_t + \hat{\nu}_t \quad (4.25)$$

An ADF-test of the residuals indicates cointegration. The t-value amounts to -7.02, which is way below the 5% critical value of -4.12.

The Johansen cointegration test confirms this result. For the VAR a lag length of 4 (AIC) is chosen:

Case	IV
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	53.94 (30.77)
Trace stat. $r \leq 1$	7.81(15.44)

Table 4.17: Results of the Johansen cointegration test: exports of goods to the rest of the world

An Error-Correction Equation The high adjustment coefficient of 0.8 points to a rapid reaction to deviations from the equilibrium. The t-value of the adjustment coefficient of -8.01 (5% critical value: -3.91) supports the results of the Engle/Granger approach and the Johansen test. Spanish exports to the rest of the world react strongly to changes in the real exchange rate. The respective elasticity is 1.32.

$$\Delta xg95row_t = -\underset{(-8.01)}{0.80}\ (-\underset{(-18.15)}{12.89} - \underset{(33.65)}{0.02}\ t + \underset{(3.12)}{0.11}\ SD9901_t + xg95row_{t-1} + \underset{(8.50)}{1.32}\ raw_{t-1}) - \underset{(-0.94)}{0.03}\ Z_{1t} - \underset{(-0.86)}{0.02}\ Z_{2t} - \underset{(-2.99)}{0.08}\ Z_{3t} + \underset{(2.42)}{0.20}\ \Delta xg95row_{t-4} - \underset{(-2.05)}{0.80}\ \Delta raw_{t-2} + \hat{\varepsilon}_t \quad (4.26)$$

The residuals of the estimation are normally distributed and free of autocorrelation. However, White's heteroskedasticity test allows to reject the null hypothesis of no heteroskedasticity in the residuals at confidence level of 7%. The CUSUM test shows that the estimated parameters are not stable over the whole sample. The relatively poor out-of-sample forecast performance also indicates some degree of instability. The diagnostics are presented in Table 4.18.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.74
Residual tests	Probability
Normality test (Jarque-Bera)	[0.48]
Serial correlation LM test (lag 1)	[0.11]
Serial correlation LM test (lag 4)	[0.26]
Serial correlation LM test (lag 8)	[0.23]
ARCH LM test (lag 1)	[0.53]
ARCH LM test (lag 4)	[0.41]
White's heteroskedasticity test	[0.07]
RESET test (h=2)	[0.97]
Stability tests	outside error bands
CUSUM test	1997Q1-1998Q4
CUSUM of squares test	none

Table 4.18: Diagnostic tests: exports of goods to the rest of the world

Exports of Services

Exports of services account for about one third of nominal exports. Most recently more than 50% of these exports originated from the tourist business, whereas the rest consisted of all kinds of services from transport to business services (Bravo and Gordo 2003, p.85). The heterogenous composition of service exports make an estimation difficult. Several approaches have been tested. The first concentrated on tourism and used income variables for the countries of origin. In 2002 28% of all tourists came from the United Kingdom and 20% were Germans (Banco de España, BDE 2003, pp. 58/59). However, various demand variables proved unsuccessful. Another approach was to concentrate on the other half of service exports. Often service exports are connected with goods exports, e.g. transport or financial services. This is why goods exports were tested as an explaining variable. The result is better than the alternatives, which were examined, but still not completely satisfactory.

Cointegration Tests A cointegrating relationship can be confirmed by both tests applied.

$$\begin{aligned} xs95_t = &\underset{(0.622)}{7.59} - \underset{(0.021)}{0.12}\ Z_{1t} + \underset{(0.021)}{0.09}\ Z_{2t} + \underset{(0.022)}{0.41}\ Z_{3t} - \underset{(0.138)}{1.47}\ raw_t \\ &+ \underset{(0.036)}{0.86}\ xg95_t - \underset{(0.037)}{0.33}\ SD9301_t + \underset{(0.070)}{0.23}\ ID9301_t + \hat{\nu}_t \end{aligned} \tag{4.27}$$

The residuals of the regression in levels can be shown to be stationary according to the ADF test. The relevant t-value is -4.26 (5% critical value: -4.12).

The Johansen cointegration test confirms the cointegrating relationship. The lag length of the VAR is 5, as recommended by the AIC.

An Error-Correction Equation The estimated error-correction model is a rather parsimonious equation. Except for the lagged dependent variable no lagged variables have a significant influence. The elasticity with respect to the real exchange rate is above one and thus relatively high.

Case	III
exogenous variables	1
endogenous variables	3
Trace stat. $r=0$	44.07 (38.93)
Trace stat. $r \leq 1$	11.17 (23.32)

Table 4.19: Results of the Johansen cointegration test: exports of services

$$\begin{aligned}\Delta xs95_t = & -\underset{(-4.41)}{0.18}\,(-\underset{(-5.26)}{6.48} + \underset{(2.97)}{0.20}\ SD9301_t + xs95_{t-1} - \underset{(-11.36)}{0.77}\ xg95_{t-1} \\ & + \underset{(4.14)}{1.04}\ raw_{t-1}) - \underset{(-2.74)}{0.11}\ Z_{1t} - \underset{(-0.30)}{0.01}\ Z_{2t} + \underset{(1.67)}{0.07}\ Z_{3t} \\ & - \underset{(-2.25)}{0.18}\ \Delta xs95_{t-1} + \underset{(8.10)}{0.71}\ \Delta xs95_{t-4} + \hat{\varepsilon}_t\end{aligned} \quad (4.28)$$

The residuals are not completely well-behaved. They are normally distributed and show no sign of autocorrelation, but they are not homoskedastic as both the ARCH LM test and White's heteroskedasticity test show. There is also a period of parameter instability according to the CUSUM of squares test.

Diagnostics	
Estimation period	1986Q6-2002Q4
Adjusted R^2	0.99
Residual tests	Probability
Normality test (Jarque-Bera)	[0.71]
Serial correlation LM test (lag 1)	[0.24]
Serial correlation LM test (lag 4)	[0.32]
Serial correlation LM test (lag 8)	[0.14]
ARCH LM test (lag 1)	[0.002]
ARCH LM test (lag 4)	[0.01]
White's heteroskedasticity test	[0.01]
RESET test (h=2)	[0.04]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1992Q1-1997Q1

Table 4.20: Diagnostic tests: exports of services

4.4.5 Imports of Goods and Services

Theoretical considerations

In the two-country case the theoretical basis for the import function is analogous to the export demand function described above. Import demand is characterised by a standard demand function with an income variable and all relevant prices as explanatory variables. Under the assumption of price homogeneity - i.e. if a change of all prices by the same proportion leaves imports unaffected - this relation can be rewritten as one of real income and *relative* prices.

$$\begin{aligned} M_I &= f(P_I, P_E, Y_I) \\ &= f(\tfrac{P_E}{P_I}, \tfrac{Y_I}{P_I}) \end{aligned} \tag{4.29}$$

M is the import volume. The subscript I stands for the importing country, whereas the subscript E refers to its partner i.e. the exporting country. Imports are imperfect substitutes for domestic products. Therefore no domestic supply variable has to be taken into account. Domestic supply affects imports in an indirect way via domestic prices (Leamer and Stern 1970). Taking into account the different commodity structure of imports and exports in Spain (Instituto Nacional de Estadística, INE 2004) as well as significant price differences between the goods produced in different countries, we can assume that domestic and imported goods are indeed imperfect substitutes.

In contrast to the export function the generalisation to an n-country case does not alter the main conclusions. On the one hand we have the importing country (i.e. Spain), on the other we have a large number of exporting countries (i.e. the rest of the world). The relevant market we look at is Spain and therefore any third market consideration need not worry us in the case of imports. The relevant relative price is that of imports to domestic prices. There is a number of potential income variables. The most general being either GDP or total demand (i.e. GDP+imports=total consumption+total investment + exports) in real terms. Depending on the structure of imports a combination of sub-aggregates is equally acceptable. In the current case the total of investment into machinery and equipment and exports shows the best fit. The Spanish foreign trade statistics show that the largest part of imports (almost 60% (Banco de España, BDE

2003)) consists of intermediate goods that can be put to a wide range of uses.

For the time being imports of goods and services are estimated as one aggregate. Additional steps could be taken towards a separate estimation of goods and services and the disaggregation of goods imports into price inelastic import demand (such as oil/energy imports) and price elastic import demand. This has been attempted by the author. However, import prices for the sub-aggregates proved difficult to estimate. Therefore, the approach offered no improvements to the current equation. As the oil price is not part of the long-term relationship in the import price equation and the share of energy imports has declined sharply since the early 1980s, there does not seem to be an urgent need for further disaggregation at the moment.

Cointegration Tests

Spanish imports are well explained by the total of exports and investment into machinery and equipment. At the same time the relative import price plays an important part.

$$\begin{aligned} m95_t \;=\; & -\underset{(0.184)}{3.91} + \underset{(0.016)}{0.01}\; Z_1 - \underset{(0.016)}{0.02}\; Z_2 - \underset{(0.016)}{0.07}\; Z_3 \\ & + \underset{(0.030)}{1.02}\; (x95_t + imeq95_t) + \underset{(0.038)}{0.74}\; pgdppm_t + \hat{\nu}_t \end{aligned} \tag{4.30}$$

The t-value of -3.44 (5% critical value: -3.78) is not sufficient to reject the null hypothesis of a unit root in the residuals. However, the Johansen cointegration test confirms cointegration, if - as recommended by the Schwarz criterion - the lag length of one is chosen.

Case	III
exogenous variables	0
endogenous variables	3
Trace stat. $r = 0$	33.34 (31.54)
Trace stat. $r \leq 1$	2.20 (17.86)

Table 4.21: Results of the Johansen cointegration test: imports of goods and services

An Error-Correction Equation

The elasticity of imports with respect to exports and investment into machinery and equipment is slightly above one. A rise in the relative domestic price of 1% results in an increase of 0.75%. There are two outliers in 1986 and 1992, which may have to do with EU accession and the introduction of the single market.

$$\begin{aligned}
\Delta m95_t \; = \; & -\underset{(0.066)}{0.37}\,(\underset{(0.254)}{3.89} + m95_{t-1} - \underset{(0.041)}{1.02}\, ximeq95_{t-1} \\
& - \underset{(0.056)}{0.75}\, pgdppm_{t-1}) - \underset{(0.016)}{0.04}\, Z_1 - \underset{(0.012)}{0.02}\, Z_2 - \underset{(0.02)}{0.10}\, Z_3 \\
& - \underset{(0.031)}{0.14}\, ID8601_t - \underset{(0.031)}{0.10}\, ID9204_t + \underset{(0.081)}{0.28}\, \Delta m95_{t-4} \\
& - \underset{(0.003)}{0.25}\, \Delta(ximeq95_{t-1}) + \hat{\varepsilon}_t
\end{aligned} \tag{4.31}$$

The specification tests suggest that the residuals are well-behaved, there is no indication of an incorrect functional form or missing variables. The specification is stable over the forecast period. This is also confirmed by the fairly good out-of-sample forecast performance.

4.5 Income and Employment

4.5.1 Employment

Theoretical considerations

As actual output is determined by aggregate demand, employment depends only on the production function.

This approach to modelling employment, which can be found in Layard, Nickell and Jackman (1991, p.364) is adopted in a number of recent macro models (Fagan, Henry, and Mestre 2001, Estrada and Willman 2002, Estrada, Fernández, Moral, and

Diagnostics	
Estimation period	1981Q2-2002Q4
Adjusted R^2	0.87
Residual tests	Probability
Normality test (Jarque-Bera)	[0.71]
Serial correlation LM test (lag 1)	[0.45]
Serial correlation LM test (lag 4)	[0.62]
Serial correlation LM test (lag 8)	[0.27]
ARCH LM test (lag 1)	[0.48]
ARCH LM test (lag 4)	[0.97]
White's heteroskedasticity test	[0.37]
RESET test (h=2)	[0.22]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.22: Diagnostic tests: imports of goods and services

Regil 2004). Employment is derived from the inverted production function. A Cobb-Douglas production function is assumed. It is defined as follows:

$$Y = AL^{\alpha}K^{1-\alpha},$$

where Y is output, A stands for technology[20], L is labour input, K is the capital stock, and α and $1-\alpha$ are the labour share and the capital share parameters, respectively. In logs the production function reads as follows (lower case letters mean logs):

$$y = a + \alpha l + (1-\alpha)k$$

then l is simply:

$$l = y/\alpha - a/\alpha - (1-\alpha)k/\alpha$$

This is indeed what we find e.g. in the ECB's Area Wide Model (AWM) (Fagan, Henry, and Mestre 2001, p.14). In the AWM this formulation is the cointegrating relationship used in the error correction equation.

In the case of Spain this would not be a balanced equation, since we would try to explain an I(1) variable (l) with other I(1) variables and an I(2) variable (k), i.e. the capital stock. In addition, as pointed out in Section 4.4.3, there is no reliable statistical basis of capital stock data.

[20] This includes trend total factor productivity. As the focus is on employment and the production function should be kept simple for illustrative purposes, it is not necessary to go into any details concerning the functional form of A.

This is not the only reason, why the capital stock plays no role in the employment function of this model. If we test for cointegration, we find exactly one cointegrating relationship between employment, a step dummy for the first quarter of 1992 (for statistical reasons) and output. It is not necessary to include further variables in the cointegrating relationship.

Employees

There are no statistics of hours worked, so the number of employees in thousands are estimated.

Cointegration Tests

$$\begin{aligned} ee_t = & - \underset{(0.252)}{0.91} + \underset{(0.006)}{0.04}\ Z_{1t} + \underset{(0.006)}{0.02}\ Z_{2t} + \underset{(0.006)}{0.03}\ Z_{3t} \\ & + \underset{(0.022)}{0.88}\ gdp95_t - \underset{(0.006)}{0.05}\ SD9201_t + \hat{\nu}_t \end{aligned} \tag{4.32}$$

The null hypothesis of a unit root in the residuals from the equation in levels can be rejected only at the 10% level.

The Johansen cointegration test allows to reject the null hypothesis of no cointegrating relationship, but does not permit to reject the hypothesis of one cointegrating relationship. A lag length of 4 has been chosen for the VAR as recommended both by the AIC and the Schwarz criterion.

Case	III
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	26.88 (23.32)
Trace stat. $r \leq 1$	4.29 (11.47)

Table 4.23: Results of the Johansen cointegration test: employees

An Error-Correction Equation In the error-correction equation the elasticity of employment with respect to output is higher than in the equation in levels (0.88) or in the long run relationship estimated in the Johansen cointegration analysis (0.91). In recent years we have indeed observed that employment grew almost as fast as output and productivity gains were low. In the short run changes of the real wage (deflated with the consumption deflator) affect employment negatively.

$$\begin{aligned}
\Delta ee_t = & - \underset{(-4.87)}{0.19} (\underset{(4.24)}{1.46} + \underset{(7.79)}{0.08} \, SD9201_t + ee_{t-1} - \underset{(-30.94)}{0.94} \, gdp95_{t-1}) \\
& - \underset{(-2.54)}{0.01} \, Z_{1t} + \underset{(2.76)}{0.01} \, Z_{2t} - \underset{(-1.33)}{0.01} \, Z_{3t} - \underset{(-3.23)}{0.10} \, \Delta rwee_{t-2} \\
& - \underset{(-2.61)}{0.08} \, \Delta rwee_{t-4} + \underset{(3.89)}{0.37} \, \Delta ee_{t-1} - \underset{(-2.44)}{0.13} \, \Delta gdp95_{t-1} + \hat{\varepsilon}_t
\end{aligned} \quad (4.33)$$

In the single-equation context the equation for the number of employees is well-specified. The residuals pass all diagnostic tests and the parameters are stable over time. The forecast performance is acceptable.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.77
Residual tests	Probability
Normality test (Jarque-Bera)	[0.88]
Serial correlation LM test (lag 1)	[0.99]
Serial correlation LM test (lag 4)	[0.84]
Serial correlation LM test (lag 8)	[0.96]
ARCH LM test (lag 1)	[0.22]
ARCH LM test (lag 4)	[0.61]
White's heteroskedasticity test	[0.18]
RESET test (h=2)	[0.92]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.24: Diagnostic tests: employees

Self-Employed Persons

The series of self-employed people is almost trend-stationary. At the same time it is difficult to find explanatory variables for the long run. Two trends seem to overlap:

while the number of self-employed persons in agriculture has been decreasing during the whole sample period, a rising number of self-employed persons has been registered in private services since the middle of the 1990s. For the time being a broken trend seems the best way of modelling the two tendencies. It has a negative slope from the beginning of the sample until the fourth quarter of 1994. In the short run the number of self employed people increases with economic activity.

$$\begin{aligned}\Delta es_t = & -\underset{(-3.40)}{0.26}\,(-\underset{(-485.70)}{7.93} - \underset{(-6.70)}{0.002}\; KT9404_t + es_{t-1}) - \underset{(-5.64)}{0.04}\; Z_{1t}SD \\ & - \underset{(-2.18)}{0.01}\; Z_{2t}SD - \underset{(-2.71)}{0.02}\; Z_{3t}SD + \underset{(4.26)}{0.38}\; \Delta gdp95_{t-3} \\ & + \underset{(2.84)}{0.28}\; \Delta gdp95_{t-4} + \underset{(2.08)}{0.21}\; \Delta gdp95_{t-5} + \underset{(1.98)}{0.17}\; \Delta gdp95_{t-6} \\ & - \underset{(-2.67)}{0.28}\; \Delta es_{t-6} - \underset{(-2.02)}{0.19}\; \Delta es_{t-12} + \hat{\varepsilon}_t\end{aligned} \tag{4.34}$$

The adjusted R^2 is rather low at 0.57. This is not a very good fit and there are some stability problems in the early 1990s. However, the residuals pass almost all diagnostic tests. Only the RESET test indicates some specification problems. Further, the parameters are not stable for the whole sample.

Diagnostics	
Estimation period	1983Q3-2002Q4
Adjusted R^2	0.57
Residual tests	Probability
Normality test (Jarque-Bera)	[0.54]
Serial correlation LM test (lag 1)	[0.26]
Serial correlation LM test (lag 4)	[0.50]
Serial correlation LM test (lag 8)	[0.70]
ARCH LM test (lag 1)	[0.51]
ARCH LM test (lag 4)	[0.45]
White's heteroskedasticity test	[0.54]
RESET test (h=2)	[0.05]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1992Q4-1995Q2

Table 4.25: Diagnostic tests: self-employed persons

Unemployment

Unemployment shows the disequilibrium in the labour market. It should actually be the difference between the labour supply and the labour demand, which is defined above. However, the total labour force, which could be seen as labour supply variable follows no clear pattern. An attempt has been made to estimate the total labour force on the basis of the real wage and a step dummy modelling the change in the unemployment statistics of 2001. However, the results proved very unsatisfactory in the model context. For this reason unemployment is estimated directly. Unlike employment, which is exclusively determined by output, unemployment seems to react to the *nominal* wage.

Cointegration Tests

$$\begin{aligned} u_t = & + \underset{(1.557)}{29.87} + \underset{(0.024)}{0.03}\ Z_{1t} + \underset{(0.024)}{0.09}\ Z_{2t} - \underset{(0.024)}{0.01}\ Z_{3t} \\ & - \underset{(0.181)}{2.91}\ gdp95_t + \underset{(0.068)}{1.43}\ wee_t - \underset{(0.038)}{0.30}\ SD0101_t + \hat{\nu}_t \end{aligned} \quad (4.35)$$

The residuals of the equation in levels are stationary according to the ADF-test. The respective t-value is -4.37 (5% critical value: -4.12).

The Johansen cointegration test finds exactly one cointegrating relationship if a lag length of three is chosen. However, in the case of the second hypothesis ($r \leq 1$) the trace statistic is close to the critical value, which indicates the possibility of two cointegrating relationships. Further, both the AIC and the Schwarz criterion recommended a lag length of five quarters, for which the second hypothesis can be rejected. This means that the Johansen cointegration provides only very weak support for the specification chosen.

Case	III
exogenous variables	1
endogenous variables	3
Trace stat. $r = 0$	58.22 (38.93)
Trace stat. $r \leq 1$	22.66 (23.32)

Table 4.26: Results of the Johansen cointegration test: unemployment

An Error-Correction Equation According to the equation the elasticity of unemployment with respect to GDP and the nominal wage is quite high.

$$\begin{aligned} \Delta u_t = & - \underset{(-6.89)}{0.15} \, (- \underset{(-12.56)}{38.16} + \underset{(5.75)}{1.46} \, SD0101_t + u_{t-1} + \underset{(10.67)}{3.82} \, gdp95_{t-1} \\ & - \underset{(-11.98)}{1.68} \, wee_{t-1}) - \underset{(-3.18)}{0.02} \, Z_{1t} - \underset{(-6.99)}{0.04} \, Z_{2t} + \underset{(1.56)}{0.01} \, Z_{3t} \\ & + \underset{(3.53)}{0.07} \, \Delta SD0101_{t-4} + \underset{(4.94)}{0.28} \, \Delta u_{t-1} + \underset{(3.28)}{0.22} \, \Delta u_{t-5} + \hat{\varepsilon}_t \end{aligned} \tag{4.36}$$

Except for some ARCH effects the residuals are well-behaved, the parameters are stable over the estimation period, there is no indication of a mis-specification.

4.5.2 The Income Side of GDP

The income side of GDP is made up of four components:

- compensation of employees ($COE = WEE * EE/1000$)

Diagnostics	
Estimation period	1981Q3-2002Q4
Adjusted R^2	0.81
Residual tests	Probability
Normality test (Jarque-Bera)	[0.53]
Serial correlation LM test (lag 1)	[0.37]
Serial correlation LM test (lag 4)	[0.18]
Serial correlation LM test (lag 8)	[0.46]
ARCH LM test (lag 1)	[0.07]
ARCH LM test (lag 4)	[0.45]
White's heteroskedasticity test	[0.38]
RESET test (h=2)	[0.67]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.27: Diagnostic tests: unemployed persons

- net[21] operating surplus and mixed income ($OSMIN$), which is equivalent with profits
- consumption of fixed capital (CFC, estimation described below)
- taxes less subsidies on production and income ($TIND$).

Compensation of employees does not have to be estimated, because it is by definition derived from two estimated variables.

Thus, three variables remain to be estimated: net operating surplus and mixed income, consumption of fixed capital and taxes less subsidies on production and imports.

Net Operating Surplus and Mixed Income

Initially, this aggregate was calculated as follows:

$OSMIN = GDP - COE - CFC - TIND$

However, in the context of the whole model this meant that all estimation errors in GDP were reflected in OSMIN, which is a sub-aggregate of household disposable income. Disposable income is an explaining variable both for private consumption and for construction investment. Thus, there would be a cumulating error in GDP, which

[21] "Net" means excluding consumption of fixed capital, which is part of gross operating surplus and mixed income.

would seriously destabilise the model. This was observed for the most recent observations in the initial model version.

For this reason OSMIN has to be estimated rather than calculated as a residual. This is difficult due to a lack of sensible explaining variables. The problem is largely a statistical one. Disposable income should include distributed profits (for which there are no data in the national accounts) rather than a series, which is above all a residual. The most reasonable explaining variable would probably be GDP, but it would suffer from the same drawbacks as calculating OSMIN as a residual. For this reason OSMIN remains exogenous for the time being.

Consumption of Fixed Capital

Consumption of fixed capital was initially estimated. However, it is not applied in the simulations in Chapter 5.

Cointegration Tests Consumption of fixed capital should be proportional to the capital stock. As the latter is not included in the model - due to the data problems described in Section 4.4.3 - nominal gross fixed capital formation is used as a proxy. A broken trend (KT9301i) has to be included.

$$\begin{aligned} cfc_t = \underset{(0.191)}{0.60} &+ \underset{(0.016)}{0.08}\ Z_{1t} + \underset{(0.016)}{0.13}\ Z_{2t} + \underset{(0.016)}{0.08}\ Z_{3t} + \underset{(0.019)}{0.91}\ ifc_t \\ &- \underset{(0.0004)}{0.006}\ KT9301i_t - \underset{(0.048)}{0.27}\ ID9301_t - \underset{(0.048)}{0.25}\ ID9301_{t-1} \\ &- \underset{(0.048)}{0.16}\ ID9301_{t-2} + \hat{\nu}_t \end{aligned} \tag{4.37}$$

The null hypothesis of a unit root in the residuals can be rejected only at the 10% level. The respective t-value in the equation of the ADF test is - 3.59 and thus not smaller than the 5% critical value of -3.78.

The Johansen cointegration test confirms a cointegrating relationship only at the 10% level (critical value: 20.75) The lag length of the VAR is five as recommended both by the AIC and the Schwarz criterion.

Case	III
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	22.51 (23.32)
Trace stat. $r \leq 1$	3.92 (11.47)

Table 4.28: Results of the Johansen cointegration test: consumption of fixed capital

An Error-Correction Equation Although the first two cointegration tests suggest at most a weak cointegrating relationship, an error-correction equation is estimated with the variables used in the Engle/Granger approach and the Johansen test. At -4.82 the t-value of the adjustment coefficient is sufficiently low to reject the null hypothesis of no cointegration at least by the standards of Banerjee et al. (critical value: -3.69).

The elasticity of consumption of fixed capital with respect to nominal investment seems a little on the high side, given the average share of consumption of fixed capital in nominal gross fixed capital formation of 63%.

$$\begin{aligned}\Delta cfc_t = & -\underset{(-4.82)}{0.36}\,(-\underset{(-2.82)}{1.15} + \underset{(9.22)}{0.01}\,KT9301I_t + cfc_{t-1} - \underset{(-22.28)}{0.86}\,ifc_{t-1}) \\ & -\underset{(-1.90)}{0.05}\,Z_{1t} + \underset{(1.55)}{0.02}\,Z_{2t} - \underset{(-0.16)}{0.01}\,Z_{3t} - \underset{(-4.56)}{0.16}\,ID9301_t - \underset{(-2.29)}{0.08}\,ID9301_{t-1} \\ & -\underset{(-3.13)}{0.30}\,\Delta cfc_{t-2} + \underset{(5.24)}{0.45}\,\Delta cfc_{t-4} + \underset{(1.75)}{0.32}\,\Delta ifc_{t-2} + \hat{\varepsilon}_t\end{aligned} \quad (4.38)$$

The residuals are well-behaved, there is no sign of mis-specification. The equation is stable over the forecast period. However, the relatively bad out-of-sample forecast performance indicates that there might be stability problems at the end of the sample. Consumption of fixed capital is strongly underestimated.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.90
Residual tests	Probability
Normality test (Jarque-Bera)	[0.29]
Serial correlation LM test (lag 1)	[0.23]
Serial correlation LM test (lag 4)	[0.63]
Serial correlation LM test (lag 8)	[0.67]
ARCH LM test (lag 1)	[0.50]
ARCH LM test (lag 4)	[0.92]
White's heteroskedasticity test	[0.77]
RESET test (h=2)	[0.35]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.29: Diagnostic tests: consumption of fixed capital

Taxes Less Subsidies on Production and Imports

Taxes less subsidies on production an imports were initially estimated. However, the equation is not applied in the simulations in Chapter 5.

Cointegration tests One would naturally expect the taxes less subsidies on production and imports to be well-explained by gross domestic product, but here, again we run into data problems, as nominal GDP is I(2) and thus not suitable for the explanation of an I(1) variable. Thus, this aggregate is modelled by real GDP and deterministics. A linear trend serves as a proxy for the changing price level. Thus, we can estimate a balanced equation with only I(1) variables and deterministics.

$$\begin{aligned} tind_t \;=\; & -\underset{(3.378)}{12.26} + \underset{(0.029)}{0.28}\, Z_{1t} + \underset{(0.027)}{0.06}\, Z_{2t} + \underset{(0.027)}{0.11}\, Z_{3t} - \underset{(0.045)}{0.22}\, Z_{1t}SD \\ & - \underset{(0.044)}{0.21}\, Z_{2t}SD - \underset{(0.046}{0.10}\, Z_{3t}SD - \underset{(0.067)}{0.32}\, ID9003_t + \underset{(0.064)}{0.31}\, ID9204_t \\ & + \underset{(0.302)}{1.81}\, gdp95_t + \underset{(0.002)}{0.01}\, TREND_t + \hat{\nu}_t \end{aligned} \tag{4.39}$$

The null hypothesis of a unit root in the residuals of the estimation in levels can be rejected with a wide margin. The respective t-statistic is -6.97 and thus much smaller than the critical value of -3.78.

Subsequently the Johansen cointegration test is carried out. If the lag length is five as recommended by the AIC, no cointegration is found. This changes if we shorten the lag length to one as recommended by the Schwarz criterion. Then the trace statistic allows to reject the null hypothesis of at most zero cointegrating relationships. At the same time we cannot reject the second hypothesis of at most one cointegrating relationship.

Case	IV
exogenous variables	0
endogenous variables	2
Trace stat. $r = 0$	64.11 (25.77)
Trace stat. $r \leq 1$	3.49 (12.39)

Table 4.30: Results of the Johansen cointegration test: taxes less subsidies on production and imports

An Error-Correction Equation The t-value of the estimated adjustment coefficient is -11.73 (5% critical value: -3.69) confirms the existence of a cointegrating relationship. The centred seasonal dummies are jointly insignificant and therefore excluded. Thus, only the centred dummies multiplied with a step dummy remain. The strict general to specific approach did not yield any sensible results in the case of this equation. There was either autocorrelation in the residuals or the residuals were not normally distributed. Therefore lags were reintroduced until an acceptable specification was found.

$$\begin{aligned}\Delta tind_t = & - \underset{(-11.73)}{1.11} (\underset{(2.42)}{8.16} - \underset{(-3.91)}{0.01}\ TREND + tind_{t-1} - \underset{(-4.80)}{1.45}\ gdp95_{t-1}) \\ & - \underset{(-2.42)}{0.11}\ Z_{1t}SD - \underset{(-2.50)}{0.11}\ Z_{2t}SD - \underset{(-0.43)}{0.02}\ Z_3SD - \underset{(-4.24)}{0.25}\ \Delta tind_{t-3} \\ & + \underset{(2.09)}{0.13}\ \Delta tind_{t-8} - \underset{(-4.67)}{0.38}\ ID9003_t + \hat{\varepsilon}_t\end{aligned} \tag{4.40}$$

The White test result rejects the null hypothesis of no heteroskedasticity in the residuals at the 5% level. Further, there is a period of instability from 1995Q2 until 1996Q2. In all other respects the specification seems well-chosen.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.81
Residual tests	Probability
Normality test (Jarque-Bera)	[0.20]
Serial correlation LM test (lag 1)	[0.35]
Serial correlation LM test (lag 4)	[0.18]
Serial correlation LM test (lag 8)	[0.36]
ARCH LM test (lag 1)	[0.71]
ARCH LM test (lag 4)	[0.70]
White's heteroskedasticity test	[0.04]
RESET test (h=2)	[0.12]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1995Q2-1996Q2

Table 4.31: Diagnostic tests: taxes less subsidies on production and imports

4.5.3 Disposable Income

Disposable income (YD) is defined as follows (for details cf. Appendix A.3.1).:

$YD = COE + OSMIN - TD - SC + TRR$

It is thus the total of labour income (COE, compensation of employes) and income from profits (OSMIN, operating surplus and mixed income - net of consumption of fixed capital) minus direct taxes (TD) and social contributions (SC) plus transfers (TRR). Of course, taxes are exclusively paid to the government sector. However, social contributions (SC) also include private pension schemes, insurance premiums etc. Equally net transfers to households are paid by different sectors. Compensation of employees is given by a definition based on other estimated variables, only direct taxes and social contributions are estimated directly. Net operating surplus and mixed income and transfers remain exogenous.

Direct Taxes Paid by Households

Direct taxes paid by households depend on households' gross income. Therefore we would expect a stable relationship between direct taxes and the total of COE and OSMIN, which serves as a proxy for gross income. Both direct taxes and the income variable are characterised by a change in the trend slope in the early 90s. Nevertheless, a broken trend with a negative slope from the beginning of the sample until the first quarter of 1992 is needed to ensure cointegration.

Cointegration Tests

$$\begin{aligned} td_t = & \underset{(0.437)}{0.62} - \underset{(0.018)}{0.06}\ Z_{1t} - \underset{(0.018)}{0.07}\ Z_{2t} - \underset{(0.018)}{0.04}\ Z_{3t} \\ & + \underset{(0.014)}{0.75}\ coeosmin_t - \underset{(0.002)}{0.02}\ KT9201i_t + \hat{\nu}_t \end{aligned} \qquad (4.41)$$

The null hypothesis of a unit root in the residuals of the estimation in levels cannot be rejected. The t-statistic of -2.61 is larger than the critical value of -3.78. However, it is well-known that the approach of Engle and Granger (Engle and Granger 1987) often fails to reject the null hypothesis, because the dynamics cannot be modelled adequately.

The Johansen cointegration test overcomes this drawback. Here one cointegrating relationship is found for the system with a lag length of 5 (according to both AIC and the Schwarz criterion).

Case	III
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	28.63 (23.32)
Trace stat. $r \leq 1$	5.74 (11.47)

Table 4.32: Results of the Johansen cointegration test: direct taxes paid by households

An Error-Correction Equation The following error-correction equation is estimated:

$$\begin{aligned}
\Delta td_t = & \underset{(-6.76)}{-\ 0.27} \ (\underset{(0.69)}{0.26} + \underset{(7.80)}{0.02}\ KT9201i_t + td_{t-1} - \underset{(-25.10)}{0.82}\ coeosmin_{t-1}) \\
& \underset{(-8.24)}{-\ 0.08}\ Z_{1t} + \underset{(0.16)}{0.001}\ Z_{2t} - \underset{(-8.66)}{0.05}\ Z_{3t} + \underset{(0.30)}{0.004}\ Z_{1t}SD + \underset{(1.70)}{0.03}\ Z_{2t}SD \\
& + \underset{(4.45)}{0.05}\ Z_{3t}SD + \underset{(9.99)}{0.57}\ \Delta td_{t-1} + \underset{(5.09)}{0.29}\ \Delta td_{t-4} - \underset{(-3.92)}{0.20}\ \Delta td_{t-5} \\
& + \underset{(5.60)}{0.25}\ \Delta td_{t-8} - \underset{(-7.37)}{0.11}\ ID8601_t - \underset{(-4.46)}{0.06}\ ID9204_t + \underset{(2.63)}{0.03}\ ID9301_t + \hat{\varepsilon}_t
\end{aligned} \tag{4.42}$$

The residuals can be assumed to be normally distributed, homoskedastic and uncorrelated. The RESET test gives no indication of a misspecification. The parameters are stable over the estimation period. However, the poor forecast performance implies that there might be some stability problems at the end of the sample.

Social Contributions of Households

The total of compensation of employees and (net) operating surplus and mixed income also serve as an explaining variable for social contributions, although not all contributions are linked to income.

Diagnostics	
Estimation period	1986Q1-2002Q4
Adjusted R^2	0.98
Residual tests	Probability
Normality test (Jarque-Bera)	[0.63]
Serial correlation LM test (lag 1)	[0.50]
Serial correlation LM test (lag 4)	[0.59]
Serial correlation LM test (lag 8)	[0.71]
ARCH LM test (lag 1)	[0.92]
ARCH LM test (lag 4)	[0.23]
White's heteroskedasticity test	[0.58]
RESET test (h=2)	[0.31]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.33: Diagnostic tests: direct taxes on households' income

Cointegration Tests

$$\begin{aligned} sc_t = & - \underset{(0.082)}{2.22} - \underset{(0.012)}{0.06}\, Z_{1t} - \underset{(0.012)}{0.09}\, Z_{2t} - \underset{(0.012)}{0.04}\, Z_{3t} \\ & + \underset{(0.008)}{1.08}\, coeosmin_t + \hat{\nu}_t \end{aligned} \tag{4.43}$$

Both variables had been found to be stationary (with respect to a broken trend). Both the estimation in levels and the Johansen cointegration test confirm this result.

Case	III
exogenous variables	0
endogenous variables	2
Trace stat. $r = 0$	23.83 (17.86)
Trace stat. $r \leq 1$	8.73 (8.07)

Table 4.34: Results of the Johansen cointegration test: social contributions paid by households

An Error-Correction Equation The income elasticity of the social contributions is close to one, as we would expect.

$$\begin{aligned}\Delta sc_t = & - \underset{(-6.70)}{0.31} (\underset{(20.22)}{2.10} - \underset{(-114.88)}{1.07} \; coeosmin_{t-1} + sc_{t-1}) - \underset{(-4.18)}{0.03} \; Z_{1t} \\ & + \underset{(0.05)}{0.0003} \; Z_{2t} - \underset{(-4.85)}{0.02} \; Z_3 + \underset{(4.99)}{0.07} \; ID9301_t + \underset{(3.05)}{0.10} \; \Delta sc_{t-1} \\ & + \underset{(6.26)}{0.46} \; \Delta sc_{t-4} + \underset{(4.36)}{0.28} \; \Delta sc_{t-8} + \hat{\varepsilon}_t \end{aligned} \tag{4.44}$$

Except for some instability according to the CUSUM of squares test, the equation seems well-specified and the residuals are well-behaved. However, the out-of-sample forecast performance (cf. Appendix B.45) is rather poor.

Diagnostics	
Estimation period	1982Q2-2002Q4
Adjusted R^2	0.98
Residual tests	Probability
Normality test (Jarque-Bera)	[0.79]
Serial correlation LM test (lag 1)	[0.41]
Serial correlation LM test (lag 4)	[0.27]
Serial correlation LM test (lag 8)	[0.56]
ARCH LM test (lag 1)	[0.59]
ARCH LM test (lag 4)	[0.49]
White's heteroskedasticity test	[0.42]
RESET test (h=2)	[0.84]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1986Q1-1989Q1

Table 4.35: Diagnostic tests: social contributions of households

4.6 Prices, Wages and the Long-Term Interest Rate

4.6.1 The GDP Deflator

A Philips Curve Approach

The Spanish GDP-Deflator is an I(2) series. One way of dealing with I(2) price series is estimating inflation instead of the price level. This actually means that we estimate some kind of a Phillips-Curve relationship, i.e. a relationship between the inflation rate and the unemployment rate. For this, again, we have to take the time series properties of the variables involved into account. Schreiber and Wolters (2003) distinguish five different combinations of integration properties of inflation and unemployment and their implications for the estimation of a Phillips-Curve relationship:

1. Both the inflation rate and unemploment are I(0): this implies a steady state defined by the means of unemployment and inflation. There is no long-run Phillips-Curve relationship.

2. The inflation rate is I(0), but the unemployment rate is I(1). This actually implies a *horizontal* "Phillips-Curve".

3. The inflation rate is I(1) and the unemployment rate is I(0). This is the case of a vertical Phillips-Curve, where the unemployment rate fluctuates around a constant NAIRU (not accelerating inflation rate of unemployment), while inflation follows a stochastic trend.

4. Both the inflation rate and the unemployment rate are I(1) and there is no cointegrating relationship. In this case we may either have the case of a time-varying NAIRU or further variables are missing in the estimation.

5. Both inflation and the unemployment rate are I(1) and there is a cointegrating relationship between both series. This is the case of a Phillips-Curve, where there is a long-term relationship between inflation and unemployment.

For the estimation period from 1980 until 2002, the properties of the data correspond to case 4. This leaves us with two possibilities:

1. We can estimate a time-varying NAIRU.

2. We can look for further variables which together with inflation and the unemployment rate lead to a cointegrating relation.

Deciding for approach 1 implies two serious drawbacks. The first is that we would actually have to model a relationship between the *change* of the inflation rate and the departure of unemployment from the NAIRU of the respective period. As we are interested in the price *level* rather than the inflation rate or even its change, the approach would lead rather far away from the variable of interest. By differencing twice we would lose even more information. Another difficulty is the measurement of the NAIRU. As Estrada, Hernando and López-Salido (2000) have found out testing a wide range of methods for the estimation of the Spanish NAIRU, the different approaches yield very different results. They sum their findings as follows: "Thus, our main conclusion is that the usefulness of the NAIRU concept as a general guideline for discussing and analyzing macroeconomic policy is very limited, given the current state of economic research on this area" (Estrada, Hernando, and López-Salido 2000, p.23).

For this reasons, the second approach is chosen for the price equation. We would find cointegration, if we added a trend to the bivariate regression. However, this trend represents a missing variable. If we interpret the unemployment rate as an indicator of excess supply or demand in the Spanish economy, it seems most likely to find a third variable for the equation among variables reflecting competitiveness or other external influences. Several combinations of variables have been tried out for this purpose: the real exchange rate, the ratio of import prices over export prices, the ratio of import prices over the GDP deflator, the ratio of the weighted CPI of the EMU over the Spanish CPI. Due to strong exchange rate fluctuations in the 1990s the real effective exchange rate cannot be used to explain the falling trend. The relationship between the EMU price level and the Spanish price level is not a suitable variable either as it excludes an exchange rate. Although it is not completely satisfactory from a theoretical point of view, the ratio of the import deflator over the GDP deflator is chosen, because it is found that a cointegrating relation with unemployment and the inflation rate exists.

Estimation with a Phillips-Curve Approach Complemented by a Third Variable (Version 1) For the first estimation approach the unemployment rate and the ratio of import prices over the GDP deflator have been chosen as explaining variables. The result is described in the next paragraphs.

Cointegration Tests

$$\Delta pgdp_t = \underset{(0.003)}{0.02} + \underset{(0.002)}{0.02}\ Z_{1t} + \underset{(0.002)}{0.003}\ Z_{2t} + \underset{(0.002)}{0.01}\ Z_{3t} - \underset{(0.003)}{0.02}\ Z_{1t}SD$$
$$- \underset{(0.003)}{0.02}\ Z_{2t}SD - \underset{(0.003)}{0.03}\ Z_{3t}SD - \underset{(0.015)}{0.04}\ UR1_t + \underset{(0.002)}{0.02}\ (pm_t - pgdp_t)$$
$$+ \underset{(0.006)}{0.02}\ ID8601_t - \underset{(0.005)}{0.02}\ ID8701_t + \hat{\nu}_t \tag{4.45}$$

We can reject the null hypothesis of a unit root in the residuals of the equation above. The respective t-value from the ADF test equation is -7.29 (5% critical value: -3.78).

The Johansen cointegration test confirms this result, if the lag length of 2 is chosen for the VAR as recommended by the Schwarz criterion (the AIC suggests a lag length of 5).

Case	III
exogenous variables	0
endogenous variables	3
Trace stat. $r = 0$	44.16 (31.54)
Trace stat. $r \leq 1$	4.73 (17.86)

Table 4.36: Results of the Johansen cointegration test: Inflation (log difference of the GDP deflator, Version 1)

An Error-Correction Equation In the estimated error correction equation the same long-run coefficients have been estimated as in the estimation above. However, the coefficients are quite small and the effect that the unemployment rate and the relative import prices have on inflation is quite weak.

$$\Delta^2 pgdp_t = \underset{(-11.24)}{-0.84} \Big(\underset{(-4.55)}{-0.01} + \underset{(2.20)}{0.03}\ UR1_{t-1} + \Delta pgdp_{t-1} - \underset{(-7.78)}{0.02} (pm_{t-1} - pgdp_{t-1}) \Big)$$
$$+ \underset{(6.27)}{0.02}\ Z_{1t} + \underset{(0.89)}{0.002}\ Z_{2t} + \underset{(5.90)}{0.02}\ Z_{3t} - \underset{(-3.37)}{0.01}\ Z_{1t}SD - \underset{(-0.86)}{0.004}\ Z_{2t}SD$$
$$- \underset{(-2.17)}{0.01}\ Z_{3t}SD + \underset{(4.65)}{0.02}\ ID8601_t - \underset{(-4.95)}{0.02}\ ID8701_t + \underset{(2.42)}{0.19}\ \Delta^2 pgdp_{t-4}$$
$$+ \underset{(1.69)}{0.11}\ \Delta^2 pgdp_{t-5} + \underset{(2.81)}{0.5}\ \Delta wee_{t-1} + \underset{(5.16)}{0.11}\ \Delta wee_{t-2} + \underset{(2.76)}{0.05}\ \Delta wee_{t-4}$$
$$- \underset{(-2.53)}{0.04}\ \Delta(pm_{t-5} - pgdp_{t-5}) + \hat{\varepsilon}_t \qquad (4.46)$$

The residuals of the error-correction equation are well-behaved. The RESET test confirms that the linear specification is appropriate and there are no missing variables. The CUSUM of squares test detects some parameter instability.

Diagnostics	
Estimation period	1981Q4-2002Q4
Adjusted R^2	0.95
Residual tests	Probability
Normality test (Jarque-Bera)	[0.21]
Serial correlation LM test (lag 1)	[0.27]
Serial correlation LM test (lag 4)	[0.70]
Serial correlation LM test (lag 8)	[0.25]
ARCH LM test (lag 1)	[0.32]
ARCH LM test (lag 4)	[0.89]
White's heteroskedasticity test	[0.87]
RESET test (h=2)	[0.56]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1995Q1-Q4, 1997Q2, 1998Q1

Table 4.37: Diagnostic tests: first difference of the GDP deflator (Version 1)

Although in total the specification of the Phillips-Curve are acceptable, some drawbacks remain. There are strong signs that the Spanish wage bargaining process can best be explained by an insider-outsider model. In this case unemployment would be almost irrelevant for wage setting. Therefore, the traditional Phillips-Curve approach is not unproblematic, because it postulates that unemployment affects the change of the price level via changes in the wage level.

An Approach Based on Labour Cost (Version 2)

The biggest drawback of the equations presented above is that they ignore the direct effect of wage cost on prices. If one compares inflation rates and the changes of unit labour cost across Europe, it seems quite obvious that countries with higher unit labour cost increases also have higher inflation rates. Therefore it would be desirable to reflect such a relationship of prices and wage cost in the model. As we have seen above this is not easy due to the integration properties of prices on the one hand and wages and unit labour cost on the other.

In this context the idea of the New Keynesian Phillips Curve[22] could be helpful. It differs from the traditional Phillips Curve approach in two respects. It includes a proxy for average marginal cost as the cyclical variable and it is forward looking[23] and thus incorporates inflation expectations rather than past inflation. The Phillips-Curve is specified as follows:

$$\pi_t = \beta E_t(\pi_{t+1}) + \lambda \overline{mc}_t$$

$E_t(\pi_{t+1})$ is the expected value of the inflation rate in period $t+1$, $\overline{mc}_t$ is average real marginal cost, measured as a %-deviation from its steady state.

In the context of a Cobb-Douglas production function real marginal cost is defined as the ratio of the real wage to the marginal product of labour.

For the estimation we can approximate this by the labour income share, which is equivalent to real unit labour cost (RULC)[24]:

$$RULC_t = (W_t * N_t)/(P_t * Y_t) = (W_t/P_t)/(Y_t/N_t)$$

It has to be stressed that the approach has not been adopted as described in the

[22] For details and empirical applications see (Galí and Gertler 1999) and (Galí, Gertler, and López-Salido 2001).
[23] Hybrid models with both backward-looking and forward-looking entrepreneurs were also estimated by Galí and his co-authors.
[24] The two are not quite identical, if we allow for self-employment.

two articles mentioned above. Forward-looking elements are not to be incorporated into the model at this time and the falling trend in the change of the price level suggests that real unit labour cost rather than its deviation from the trend makes sense [25]. The objective of this experiment is to include a wage cost variable in the estimation rather than copy the concept of the New Phillips Curve one to one.

The result is given below.

Cointegration Tests

$$\begin{aligned}\Delta pgdp_t = & -\underset{(0.025)}{0.11} + \underset{(0.002)}{0.02}\ Z_{1t} + \underset{(0.002)}{0.01}\ Z_{2t} + \underset{(0.002)}{0.01}\ Z_{3t} - \underset{(0.004)}{0.01}\ Z_{1t}SD \\ & - \underset{(0.004)}{0.01}\ Z_{2t}SD - \underset{(0.003)}{0.03}\ Z_{3t}SD + \underset{(0.006)}{0.02}\ ID8601_t - \underset{(0.006)}{0.02}\ ID8701_t \\ & + \underset{(0.013)}{0.07}\ ((wee_t - pgdp_t) - (gdp95_t - et_t)) - \underset{(0.001)}{0.01}\ SD9201_t + \hat{\nu}_t\end{aligned} \tag{4.47}$$

The null hypothesis of a unit root in the residuals can be rejected at a confidence level way below one %. The respective t-value is -8.28, whereas the 5% critical value is only -3.78.

In the Johansen cointegration test we cannot reject the null hypothesis of no cointegrating relationship. The lag length is five as suggested both by the AIC and the Schwarz criterion.

Case	III
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	18.58 (23.32)
Trace stat. $r \leq 1$	5.71 (11.47)

Table 4.38: Results of the Johansen cointegration test: Inflation (log difference of the GDP deflator, Version 2)

[25] Obviously Galí et al. manage to incorporate falling inflationary expectations into their model.

An Error-Correction Equation As in the previous version the adjustment coefficient is high, but the elasticities of the inflation rate with respect to the explaining variables is relatively low. We need two sets of seasonal dummies to model the change of the seasonal pattern. Four impulse dummies are required to eliminate outliers in the GDP deflator.

$$\begin{aligned}
\Delta^2 pgdp_t = & - \underset{(-11.75)}{0.84} \; (\underset{(6.81)}{0.16} + \underset{(9.71)}{0.01} \; SD9201_t + \Delta pgdp_{t-1} \\
& - \underset{(-7.80)}{0.10} \; ((wee_{t-1} - pgdp_{t-1}) - (gdp95_{t-1} - et_{t-1}))) \\
& + \underset{(10.09)}{0.03} \; Z_{1t} + \underset{(2.00)}{0.004} \; Z_{2t} + \underset{(6.97)}{0.02} \; Z_{3t} - \underset{(-5.48)}{0.02} \; Z_{1t}SD - \underset{(-3.17)}{0.01} \; Z_{2t}SD \\
& - \underset{(-5.51)}{0.02} \; Z_{3t}SD + \underset{(5.81)}{0.02} \; ID8601_t - \underset{(-5.06)}{0.02} \; ID8701_t - \underset{(-3.40)}{0.01} \; ID8801_t \\
& - \underset{(-3.23)}{0.01} \; ID9702_t - \underset{(-4.52)}{0.09} \; \Delta rweepgdp_{t-1} - \underset{(-5.33)}{0.09} \; \Delta rweepgdp_{t-3} + \hat{\varepsilon}_t
\end{aligned} \tag{4.48}$$

The residuals are well-behaved and there is no sign of a mis-specification. However, the parameters are not stable over the estimation period. The poor in-sample and out-of-sample forecast performance confirm the existence of stability problems.

A more severe problem than the bad out-of-sample forecast performance occurs, when the equation is included in the model. The dangerous property of this specification is that inflation reacts negatively to productivity increases. If we model e.g. a euro area demand shock, Spanish GDP surges. As employment rises by less than GDP, there is a productivity gain, which has a negative impact on the inflation rate. Thus, the price reaction to a demand shock would be negative in the model context - a highly implausible result, that we cannot accept. Maybe the inclusion of inflation expectations could offset this effect.

From the two approaches tested version 1 is inserted into the model as it produces more plausible results. Nevertheless, there is a strong need for further research in this

Diagnostics	
Estimation period	1981Q1-2002Q4
Adjusted R^2	0.95
Residual tests	Probability
Normality test (Jarque-Bera)	[0.53]
Serial correlation LM test (lag 1)	[0.54]
Serial correlation LM test (lag 4)	[0.42]
Serial correlation LM test (lag 8)	[0.55]
ARCH LM test (lag 1)	[0.71]
ARCH LM test (lag 4)	[0.76]
White's heteroskedasticity test	[0.63]
RESET test (h=2)	[0.92]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1993Q2-1998Q4

Table 4.39: Diagnostic tests: first difference of the GDP-Deflator (Version 2)

area, as the weak effect of wages on prices does not seem plausible.

4.6.2 The Private Consumption Deflator

Like the GDP-Deflator the private consumption deflator is an I(2) series, i.e. it is not stationary in differences as most other series in the model. Thus, a Phillips-Curve might as well have been estimated with the private consumption deflators's first difference as the dependent variable. The GDP-Deflator could then be simply calculated. However, the decision to place the GDP-Deflator at the heart of the price estimations also has to do with the properties of the private consumption deflator. Besides being an I(2) series, the private consumption deflator shows an extreme change of the seasonal pattern in 1992 and is therefore even more difficult to estimate. As a consequence, the focus is on the GDP-deflator and the other deflators, which are also I(2), are simply linked to the GDP-deflator via an autoregressive distributed lag (ADL) model in levels. A general to specific approach is followed. With enough lags, there is no longer any spurious regression problem.

As we would expect the private consumption deflator is explained by a combination of the domestic price level (GDP-deflator) and the foreign price level (import deflator). Seasonal dummies multiplied with a step dummy model the abrupt change of the seasonal pattern. The usual set of centred seasonal dummies is not significant.

$$
\begin{aligned}
pc_t \;=\; & -\underset{(4.22)}{0.11} - \underset{(-2.80)}{0.02}\; Z_{1t}SD + \underset{(6.23)}{0.06}\; Z_{2t}SD + \underset{(8.66)}{0.04}\; Z_{3t}SD + \underset{(4.16)}{0.23}\; pgdp_{t-1} \\
& + \underset{(8.41)}{0.92}\; pc_{t-1} - \underset{(-2.75)}{0.27}\; pc_{t-2} + \underset{(2.05)}{0.09}\; pc_{t-4} + \underset{(2.41)}{0.03}\; pm_{t-2} - \underset{(-1.74)}{0.04}\; pm_{t-5} \\
& + \underset{(2.45)}{0.06}\; pm_{t-6} - \underset{(-3.62)}{0.04}\; pm_{t-8} + \hat{\varepsilon}_t
\end{aligned}
\tag{4.49}
$$

At 0.88 the long-run impact of the GDP deflator is very strong, whereas import prices hardly affect the consumption deflator in the long run (long-run coefficient: 0.04). As Table 4.40 shows the residuals "pass" all diagnostic tests. There are no hints of a mis-specification or instability.

Diagnostics	
Estimation period	1982Q1-2002Q4
Adjusted R^2	1.00 (rounded)
Residual tests	Probability
Normality test (Jarque-Bera)	[0.92]
Serial correlation LM test (lag 1)	[0.92]
Serial correlation LM test (lag 4)	[0.80]
Serial correlation LM test (lag 8)	[0.36]
ARCH LM test (lag 1)	[0.93]
ARCH LM test (lag 4)	[0.72]
White's heteroskedasticity test	[0.11]
RESET test (h=2)	[0.16]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.40: Diagnostic tests: consumption deflator

4.6.3 The Government Consumption Deflator

The largest part of government consumption are the services provided by the public administration. Its costs consist mainly of the compensation of government employees. Therefore it is not surprising that wages best explain the evolution of the government consumption deflator. The use of compensation of *government* employees would certainly be more appropriate. Given the high degree of aggregation of the model, compensation of employees in the whole economy is used as a proxy. Further, wages in the government sector showed a similar development as those in the total economy. In 1993, compensation of employees in the total economy "jumps" to a higher level. This level shift indeed represents payroll developments of the private industries. This is why the step dummy SD9301 is required in the estimation. However, the trend slope of compensation of employees is more or less identical for the total economy and the government sector.

Cointegration Tests

$$\begin{aligned} pcgov_t = & - \underset{(0.068)}{2.69} + \underset{(0.010)}{0.07}\ Z_{1t} + \underset{(0.010)}{0.05}\ Z_{2t} + \underset{(0.010)}{0.05}\ Z_{3t} \\ & + \underset{(0.014)}{0.14}\ Z_{1t}SD + \underset{(0.014)}{0.16}\ Z_{2t}SD + \underset{(0.014)}{0.05}\ Z_{3t}SD \\ & + \underset{(0.009)}{0.86}\ wee_t - \underset{(0.008)}{0.04}\ SD9301_t + \hat{\nu}_t \end{aligned} \tag{4.50}$$

Cointegration is confirmed by the estimation in levels. The null hypothesis of a unit root in the residuals can be rejected. The respective t-value is -6.44 (5% critical value: -3.78).

The Johansen cointegration test equally allows to reject the null hypothesis of no cointegrating relationship at the 5% level. The lag length of the VAR is 5 according to both information criteria.

Case	III
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	48.11 (23.32)
Trace stat. $r \leq 1$	1.95 (11.47)

Table 4.41: Results of the Johansen cointegration test: government consumption deflator

An Error-Correction Equation

The following error-correction equation has been estimated.

$$\begin{aligned}
\Delta pcgov_t = & -\underset{(-6.05)}{0.14}\,(\underset{(14.57)}{2.14} + \underset{(5.87)}{0.08}\,SD9301_t + pcgov_{t-1} - \underset{(-49.94)}{0.81}\,wee_{t-1}) \\
& - \underset{(-0.89)}{0.002}\,Z_{1t} - \underset{(-2.57)}{0.005}\,Z_{2t} - \underset{(-2.92)}{0.01}\,Z_{3t} + \underset{(4.18)}{0.02}\,Z_{1t}SD - \underset{(-0.87)}{0.003}\,Z_{2t}SD \\
& + \underset{(4.44)}{0.01}\,Z_{3t}SD + \underset{(4.94)}{0.02}\,ID8301_t - \underset{(-5.18)}{0.02}\,ID8701_t - \underset{(-4.85)}{0.02}\,ID8801_t \\
& + \underset{(4.01)}{0.02}\,ID9201_t + \underset{(4.14)}{0.02}\,ID9501_t - \underset{(-4.50)}{0.02}\,ID9701_t + \underset{(2.63)}{0.14}\,\Delta pcgov_{t-2} \\
& + \underset{(1.56)}{0.03}\,\Delta wee_{t-2} - \underset{(-7.44)}{0.13}\,\Delta wee_{t-4} + \hat{\varepsilon}_t
\end{aligned} \tag{4.51}$$

The residuals show no autocorrelation. They can be assumed to be normally distributed and homoskedastic. The low p-value of the RESET test however points to some specification problems, such as the wrong functional form. This is not very surprising, when you look at the series in differences (cf. Appendix A.1.2).

4.6.4 The Fixed Investment Deflator

The fixed investment deflator proved extremely difficult to estimate. It is characterised by an abrupt change of the seasonal pattern in 1992 as well as a change of the trend slope in 1999 that is not paralled in any of the other potential explaining variables. These are the GDP deflator, the import deflator, the US dollar exchange rate, interest rates (long term, short term, real and nominal). The series is a borderline case. It can be classified as either I(1) or I(2). There is no cointegrating relationship with the GDP deflator or other variables.

Thus, the estimation approach is an ADL model in levels. All the variables mentioned above have been tried as well as investment demand.

Diagnostics	
Estimation period	1981Q2-2002Q4
Adjusted R^2	0.98
Residual tests	Probability
Normality test (Jarque-Bera)	[0.90]
Serial correlation LM test (lag 1)	[0.73]
Serial correlation LM test (lag 4)	[0.27]
Serial correlation LM test (lag 8)	[0.45]
ARCH LM test (lag 1)	[0.27]
ARCH LM test (lag 4)	[0.31]
White's heteroskedasticity test	[0.40]
RESET test (h=2)	[0.00]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.42: Diagnostic tests: government consumption deflator

The variables for demand in combination with exchange rates or unit labour cost showed significant t-values, but in both cases, which were estimated, the forecast performance proved desastrous[26]. This is why, for the time being, an autoregressive equation with a step dummy for the first quarter of 1999 has been chosen as the specification with the best fit and stability. In fact the series thus remains exogenous to the model.

$$\begin{aligned} pifc_t = \; & \underset{(0.029)}{0.12} + \underset{(0.003)}{0.004}\, Z_1SD + \underset{(0.003)}{0.01}\, Z_2SD - \underset{(0.003)}{0.01}\, Z_3SD \\ & + \underset{(0.064)}{0.90}\, pifc_{t-1} + \underset{(0.112)}{0.37}\, pifc_{t-4} - \underset{(0.096)}{0.29}\, pifc_{t-5} + \underset{(0.003)}{0.01}\, SD9901 + \hat{\varepsilon}_t \end{aligned} \tag{4.52}$$

Except for a certain degree of heteroskedasticity the residuals are well-behaved. There is no indication of a mis-specification and the stability of the parameters is confirmed both by the CUSUM and CUSUM of squares test and the out-of-sample forecast performance.

[26] Various specifications have also been tested for the deflators of construction investment and investment into machinery and equipment. The use of demand variables in the price equation actually contradicts the interpretation of the price equation as the supply side of the model, which is another reason besides the poor forecast performance, why these specifications were not chosen.

Diagnostics	
Estimation period	1981Q2-2002Q4
Adjusted R^2	1.00 (rounded)
Residual tests	Probability
Normality test (Jarque-Bera)	[0.44]
Serial correlation LM test (lag 1)	[0.73]
Serial correlation LM test (lag 4)	[0.65]
Serial correlation LM test (lag 8)	[0.22]
ARCH LM test (lag 1)	[0.38]
ARCH LM test (lag 4)	[0.19]
White's heteroskedasticity test	[0.04]
RESET test (h=2)	[0.11]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.43: Diagnostic tests: fixed investment deflator

4.6.5 The Export Price Function

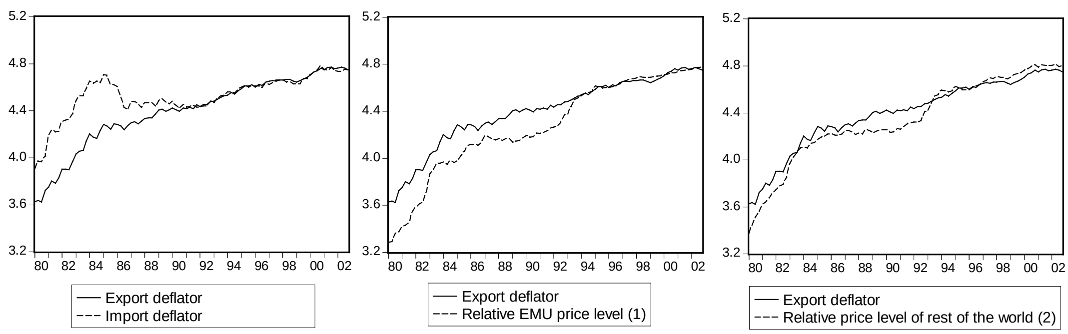
The Choice of Variables

In a market with imperfect competition prices are set as cost plus a mark-up. In the context of international trade, the focus of research has recently shifted to the question to what extent changes in exchange rates are reflected in import and export prices. Researchers noted that often, movements of the exchange rate are not fully reflected in foreign trade prices, i.e. the exchange rate pass-through is incomplete or we have an instance of pricing to market.

Pricing to market is a long term pricing strategy. In order to secure market shares in export markets, exporters find it rational not always to pass through exchange rate movements completely. Rather, exporters vary the mark up according to the relative cost of competitors, the exchange rate as well as the demand pressure in the export market. Generally, this concept can best be studied at industry level, but numerous studies have found evidence for pricing to market at the macro level, e.g. (Warmedinger 2004). Stephan (2005a) stresses that pricing to market is typical of markets with high product differentiation and argues that this is one reason, why pricing to market is the appropriate modelling strategy for Germany because of the high share of finished products in its exports and imports (74% and 88% in 2003, respectively). However, she finds that pricing to market plays almost no part in German export prices, which are mainly influenced by the domestic cost situation.

In contrast Spain's exporters seem to set prices in accordance with the price level of their export markets as Estrada, Fernández, Moral and Regil (2004, p.37) have found in their estimation of a quarterly macroeconometric model of the Spanish economy. With a coefficient of 0.73 the effect of pricing to market is particularly high for the prices of exports to the euro area.

In contrast, the author has found evidence that besides the domestic price level import prices may be one determining factor of export prices. As Figure 4.2 shows, the import price level has moved parallel to the export price level since the early 1990s and may explain a large part of the export price developments in recent years. This seems plausible if we take into account that almost 60% of Spain's imports were intermediate goods in 2002 (Banco de España, BDE 2003). An argument against this interpretation is the relatively low degree of openness of the Spanish economy. The share of (nominal) exports in GDP is below 30%.



(1) multiplied by the ecu exchange rate as index divided by 100
(2) calculated as inverse of the OECD real effective exchange rate for Spain multiplied by the Spanish CPI multiplied by 100

Figure 4.2: Export deflator and potential explaining variables (in logs)

As the euro area is Spain's main export market the weighted average CPI of the euro area multiplied by the exchange rate has equally been tested. A third variable which could be relevant besides the domestic price level might be a wider relative price index, derived from the OECD's real effective exchange rate[27].

[27] This relative price series has been obtained by multiplying the reciprocal of the real effective exchange rate with the Spanish CPI as well as 100.

As the graph shows the import deflator seems an appropriate variable to explain export prices. However, the price levels of the euro area countries play a vital part in all three series. They are one of the main explaining variables in the import deflator (c.f. section 4.6.6) and they account for a large share in the OECD's real effective exchange rate. Thus, not the actual outcome for the model is very different, but its interpretation.

The Export Deflator

The export deflator is an I(2) variable. The equation is estimated in log levels. From the original ADL model with eight lags a very parsimonious specification remains after a general to specific approach is applied.

$$\begin{aligned} px_t = \; & \underset{(5.84)}{0.31} - \underset{(-1.48)}{0.01}\; Z_{1t} + \underset{(2.14)}{0.01}\; Z_{2t} + \underset{(0.12)}{0.001}\; Z_{3t} + \underset{(1.48)}{0.01}\; Z_{1t}SD \\ & - \underset{(-1.95)}{0.02}\; Z_{2t}SD - \underset{(-4.19)}{0.04}\; Z_{3t}SD - \underset{(-3.07)}{0.53}\; pgdp_{t-1} \\ & + \underset{(4.08)}{0.68}\; pgdp_{t-3} + \underset{(10.41)}{0.67}\; px_{t-1} + \underset{(4.08)}{0.11}\; pm_{t-1} + \hat{\varepsilon}_t \end{aligned} \tag{4.53}$$

The long-run impact of the GDP deflator is thus 0.45. At 0.33 the long-run coefficient of the import price level is rather high - especially compared to the low impact of import prices on the private consumption deflator (c.f. Equation 4.49).

The equation yields fairly well-behaved residuals. Except for some indication of heteroskedasticity the residuals pass all diagnostic tests. There is a short spell of parameter instability according to the CUSUM test.

Diagnostics	
Estimation period	1980Q4-2002Q4
Adjusted R^2	1.00 (rounded)
Residual tests	Probability
Normality test (Jarque-Bera)	[0.22]
Serial correlation LM test (lag 1)	[0.35]
Serial correlation LM test (lag 4)	[0.55]
Serial correlation LM test (lag 8)	[0.72]
ARCH LM test (lag 1)	[0.31]
ARCH LM test (lag 4)	[0.61]
White's heteroskedasticity test	[0.04]
RESET test (h=2)	[0.29]
Stability tests	outside error bands
CUSUM test	1998Q4-1999Q4
CUSUM of squares test	none

Table 4.44: Diagnostic tests: export deflator

4.6.6 The Import Price Function

Theoretical Considerations and the Choice of Variables

In the two country case imports are just the mirror image of exports and the same holds for import prices. In a multi-country context this is no longer true. The importing country always faces a perfectly elastic supply. This is why the domestic demand situation should not affect the import price level (Stephan 2005a, p.79). If pricing to market is the strategy of the foreign exporters in the Spanish market then we would expect the Spanish price level to explain the import price dynamics to some extent. The import price function would thus include the foreign price level (as proxy for the cost of the exporters) and the domestic (i.e. Spanish) price level. Besides these variables energy prices usually have a special impact on import prices. As the demand for energy is inelastic in the medium term, energy prices can be fully passed through.

Warmedinger (2004) has recently examined to what extent pricing to market effects play a rôle in the import functions of the five biggest euro area economies. For this purpose he analysed the estimated long-run relationships of the respective equations in recent versions of the ESCB-Multi-Country Model. In almost all cases competitors' prices, the domestic price level and energy prices appeared as explaining variables. In addition all equations required step dummies in addition. According to the estimations the Netherlands, a small and highly open economy, is the only country, where the domestic price level is irrelevant for import prices. In the case of Spain the coefficient of the domestic price level is 0.18, which indicates a weak but significant pricing to market effect.

This result differs markedly from the specifications published by the Bank of Spain (cf. (Estrada and Willman 2002, Estrada, Fernández, Moral, and Regil 2004)). The import price function in the original version of the Spanish Block of the ESCB-Multi-Country model relies only on the exporting countries' price level and energy prices. Cointegration is not confirmed according to the statistics given. In the updated version of this model import prices are disaggregated further. Estrada et al. (2004) differentiate between imports of goods from the euro area, imports of goods from the rest of the world and imports of services. Here, the foreign price level is the main explaining variable for all import prices. In addition import prices from the rest of the world are also driven by energy and raw material prices. Pricing to market is only relevant in the case of services, where the coefficient of 0.45 is relatively high.

Unfortunately, Warmedinger does not inform the reader, which domestic price index he uses for the Spanish economy. As he refers to the national accounts data in his section on data, he may have used the GDP deflator. However, the use of an I(2) variable to explain an I(1) variable as import prices is problematic. So is the use of two step dummies in one equation without any explanation of the particular effect the dummy is to capture - particularly in 1997.

The author has chosen to model import prices as a function of foreign prices[28] and the ECU and US-Dollar exchange rates. The oil price exerts only a short term influence.

The Import Deflator

Although Spain's imports mainly come from the euro area the US-Dollar exchange rate has a much bigger impact on import prices than the price level of the EU-countries converted at the ECU/Euro exchange rate. The coefficients change slightly in the error correction model below, but their relative importance remains largely unchanged.

[28] Here the CPIs of the euro area countries were used, as they account for the largest share of Spain's imports.

Cointegration Tests

$$\begin{aligned} pm_t = \underset{(0.114)}{0.46} + \underset{(0.019)}{0.03}\ Z_{1t}SD + \underset{(0.019)}{0.02}\ Z_{2t}SD - \underset{(0.025)}{0.01}\ Z_{3t}SD \\ + \underset{(0.027)}{0.50}\ usd_t + \underset{(0.016)}{0.17}\ (cpiewu_t + ecu_t) + \hat{\nu}_t \end{aligned} \tag{4.54}$$

The null hypothesis of a unit root in the residuals from the estimation in levels can be rejected at the 5% level. The respective t-value is -3.93 (critical value: -3.78). The series can therefore be assumed to be cointegrated. However, the Johansen Cointegration test finds no cointegrating relationship at lag length 2 as recommended by the AIC.

Case	III
exogenous variables	0
endogenous variables	3
Trace stat. $r = 0$	24.00 (31.54)
Trace stat. $r \leq 1$	11.06 (17.86)

Table 4.45: Results of the Johansen cointegration test: import prices

Nevertheless an error-correction equation is estimated.

An Error-Correction Equation The highly significant adjustment coefficient is an indicator for the existence of a cointegrating relationship between the variables in the error-correction term. Its t-value of -4.11 is far below the critical value of -3.69. Contemporaneous changes of the US-Dollar have a strong influence in the short run.

$$\begin{aligned} \Delta pm_t = & - \underset{(-4.11)}{0.22} (- \underset{(-4.69)}{1.42} + pm_{t-1} - \underset{(-3.29)}{0.11}\ (cpiewu_{t-1} + ecu_{t-1}) - \underset{(-7.95)}{0.43}\ usd_{t-1}) \\ & + \underset{(6.84)}{0.06}\ Z_{1t}SD + \underset{(2.05)}{0.02}\ Z_{2t}SD - \underset{(-1.58)}{0.01}\ Z_{3t}SD + \underset{(7.46)}{0.32}\ \Delta usd_t \\ & + \underset{(3.77)}{0.06}\ \Delta oil\$_t + \underset{(4.19)}{0.06}\ \Delta oil\$_{t-1} + \underset{(3.16)}{0.22}\ \Delta pm_{t-3} + \hat{\varepsilon}_t \end{aligned} \tag{4.55}$$

The estimated equation passes all specification tests (c.f. Table 4.46). The residuals are well-behaved. There are no signs of non-linearities or instability. The equation shows an excellent out-of-sample forecast performance.

Diagnostics	
Estimation period	1981Q1-2002Q4
Adjusted R^2	0.76
Residual tests	Probability
Normality test (Jarque-Bera)	[0.49]
Serial correlation LM test (lag 1)	[0.69]
Serial correlation LM test (lag 4)	[0.76]
Serial correlation LM test (lag 8)	[0.46]
ARCH LM test (lag 1)	[0.82]
ARCH LM test (lag 4)	[0.28]
White's heteroskedasticity test	[0.67]
RESET test (h=2)	[0.30]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.46: Diagnostic tests: import deflator

4.6.7 The Consumer Price Index

The consumer price index is used in the calculation of four real effective exchange rates (cf. Appendix B.5). Like the GDP-deflator, the private consumption deflator and the export deflator it is not stationary in differences. An ADL model is estimated with a general to specific approach starting with eight lags.

$$\begin{aligned} cpi_t = & \underset{(3.24)}{+0.10} + \underset{(2.09)}{0.01}\ Z_{1t}SD + \underset{(2.75)}{0.01}\ Z_{2t}SD + \underset{(3.66)}{0.01}\ Z_{3t}SD \\ & + \underset{(7.45)}{0.71}\ cpi_{t-1} + \underset{(2.70)}{0.31}\ cpi_{t-3} - \underset{(-2.56)}{0.20}\ cpi_{t-7} + \underset{(3.79)}{0.26}\ pc_{t-1} \\ & - \underset{(-3.33)}{0.24}\ pc_{t-4} + \underset{(2.31)}{0.14}\ pc_{t-8} + \hat{\varepsilon}_t \end{aligned} \tag{4.56}$$

As Table 4.47 shows there may be some higher-order (> 4) auto-correlation in the residuals. However, the test statistics do not permit to reject the null hypotheses of no auto-correlation at the 5% level.

Diagnostics	
Estimation period	1982Q1-2002Q4
Adjusted R^2	1.00 (rounded)
Residual tests	Probability
Normality test (Jarque-Bera)	[0.69]
Serial correlation LM test (lag 1)	[0.64]
Serial correlation LM test (lag 4)	[0.11]
Serial correlation LM test (lag 8)	[0.08]
ARCH LM test (lag 1)	[0.59]
ARCH LM test (lag 4)	[0.53]
White's heteroskedasticity test	[0.30]
RESET test (h=2)	[0.64]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	none

Table 4.47: Diagnostic tests: consumer price index

4.6.8 Real Compensation per Employee (Wage Equation)

Theoretical Considerations and Choice of Variables

The most prominent approach to the estimation of wages is the bargaining approach described in Layard et al. (1991). It describes wage formation as a negotiation process between employers and unions. Whereas the employers seek to maximise profits, the employees want to maximise their utility, which depends on the real wage and the reservation wage (i.e. the transfer they would receive, if they remained unemployed). High unemployment weakens the bargaining position of the unions/employees and thus exerts a downward pressure on wages. Positive price surprises equally have a negative effect on the real wage. In addition, there is a number of institutional factors, which affect wage pressure, such as the duration of benefit payments, the tax wedge or the replacement ratio. Trend productivity (reflected by the capital intensity of production) also causes the wage level to rise. In the context of an open economy competitiveness has a positive impact on wages.

We thus obtain the following wage equation (Layard, Nickell, and Jackman 1991, p. 389):

$$w - p = \gamma_0 - \gamma_1 u - \gamma_2 \Delta^2 p + \gamma_3 c + z_w + \beta_4 (k - l) \qquad (4.57)$$

p is the log of the price level, w is the log of the wage level, u stands for the unemployment rate, c is external competitiveness, z_w is a variable that reflects institutional factors, k is the log of the capital stock and l is the labour supply. Thus, $k - l$ stands for the log of the capital intensity, which can be interpreted as trend productivity.

If we estimate the real wage (which has been identified as trend stationary) rather than the nominal wage, we can avoid an I(2) price variable in the equation.

Due to the structure of the model we are actually interested in the real wage from a supply perspective[29]. That means we should take the tax wedge into account and choose a price index based on consumption.

From a theoretical point of view this would be desirable. However, the existing national accounts data base does not offer any data of the *net* compensation of employees in Spain for the estimation period - neither on a quarterly nor on an annual basis. Similar difficulties apply to the other institutional variables such as the replacement ratio. The Bank of Spain has produced its own series for the ESCB-Multi-Country Model and kindly supplied it to the author. However, the series has not contributed significantly to explaining the real wage in Spain. When you look at Figure 4.3 you will understand why. Obviously, the actual benefit payments were somehow "weighted" with the reciprocal of the unemployment rate[30]. This means that a large part of the information in the series is already included in the unemployment rate. Thus, we might at best run into problems of multi-collinearity using the series.

As the availability of institutional variables is limited, we are left with the unemployment rate and productivity to explain the evolution of real wages. The consumption deflator has been chosen as price variable to deflate nominal wages.

Cointegration Tests

The real wage depends solely on productivity. The unemployment rate is irrelevant both in the short and in the long run. This is quite plausible in the case of Spain,

[29] For a detailed discussion of the determinants and macroeconomic rôle of wages both from a consumer's and a producers perspective cf. (Deutsche Bundesbank 2000)

[30] Unfortunately, there is no exact documentation of how the replacement ratio has been obtained.

Figure 4.3: The replacement ratio and unemployment rate

where a minority of union members negotiates wage agreements which are automatically extended to all employees.

$$\begin{aligned} rwee_t \;=\; & \underset{(0.419)}{4.66} - \underset{(0.010)}{0.04}\; Z_{1t} - \underset{(0.010)}{0.04}\; Z_{2t} - \underset{(0.010)}{0.04}\; Z_{3t} - \underset{(0.014)}{0.06}\; Z_{1t}SD \\ & - \underset{(0.014)}{0.14}\; Z_{2t}SD - \underset{(0.014)}{0.07}\; Z_{3t}SD + \underset{(0.047)}{0.42}\; prodet_t + \underset{(0.009)}{0.09}\; SD9101_t + \hat{\nu}_t \end{aligned} \tag{4.58}$$

The real wage is treated like an I(1) variable, although - according to the Perron tests (cf. Chapter 3) the series can be considered stationary around a broken trend. However, there is never any absolute certainty and it would not make sense to try and explain the real wage by nothing more than a broken trend (What explains the broken trend?) or to choose an explaining variable merely on the basis of its deterministics. For this reason the Perron Test result is not taken as a dogma.

The null hypothesis of a unit root in the residuals of the estimation equation in levels can be rejected. The respective t-value is -5.66 (5% critical value: -3.78).

The Johansen cointegration test finds only one cointegrating relationship only for lag length three, which is not in line with the information criteria. Both the AIC and the Schwarz criterion recommend a lag length of 5, for which two cointegrating relationships would be found.

Case	III
exogenous variables	1
endogenous variables	2
Trace stat. $r = 0$	36.00 (23.32)
Trace stat. $r \leq 1$	9.91 (11.47)

Table 4.48: Results of the Johansen cointegration test: real wage

An Error-Correction Equation

Real wages react relatively slowly to a change in productivity. The elasticity of 0.43 corresponds to the estimation in levels.

$$\begin{aligned}
\Delta rwee_t = & -\underset{(-4.92)}{0.26}\,(-\underset{(-5.50)}{4.52} + rwee_{t-1} - \underset{(-4.71)}{0.43}\,prodet_{t-1} - \underset{(-4.48)}{0.07}\,SD9101_t) \\
& - \underset{(-3.41)}{0.04}\,Z_{1t} - \underset{(4.36)}{0.02}\,Z_{2t} - \underset{(-4.36)}{0.03}\,Z_{3t} - \underset{(-0.66)}{0.01}\,Z_{1t}SD - \underset{(-5.01)}{0.04}\,Z_{2t}SD \\
& - \underset{(-0.44)}{0.003}\,Z_{3t}SD + \underset{(4.56)}{0.37}\,\Delta rwee_{t-4} + \underset{(3.75)}{0.26}\,\Delta rwee_{t-8} - \underset{(-2.81)}{0.31}\,\Delta prodet_{t-1} \\
& - \underset{(-2.65)}{0.28}\,\Delta prodet_{t-4} + \underset{(3.42)}{0.03}\,\Delta SD9101_{t-6} + \underset{(2.10)}{0.02}\,\Delta SD9101_{t-8} + \hat{\varepsilon}_t
\end{aligned} \tag{4.59}$$

The results of the specification tests are given in Table 4.49. According to the tests applied the equation seems well-specified. The White test indicates some heteroskedasticity, but the null hypothesis of no heteroskedasticity cannot be rejected at the 5% level. With one value of the CUSUM of squares statistic outside the confidence lines, stability is not really an issue. The out-of-sample forecast performance is also quite good. However, it has to be kept in mind that the real wage is almost constant in the forecast period. Therefore a good forecast performance is not surprising.

Diagnostics	
Estimation period	1982Q2-2002Q4
Adjusted R^2	0.98
Residual tests	Probability
Normality test (Jarque-Bera)	[0.98]
Serial correlation LM test (lag 1)	[0.70]
Serial correlation LM test (lag 4)	[0.66]
Serial correlation LM test (lag 8)	[0.63]
ARCH LM test (lag 1)	[0.87]
ARCH LM test (lag 4)	[0.16]
White's heteroskedasticity test	[0.07]
RESET test (h=2)	[0.15]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1998Q2

Table 4.49: Diagnostic tests: real compensation per employee

4.6.9 The Long-Term Interest Rate

Cointegration Tests

From the expectations theory of the term structure of interest rates we would expect the long term interest rate to be cointegrated with the short term interest rate, which is equivalent to a stationary interest rate spread (Shiller 1979). The long-term interest rate is indeed cointegrated with the three-month interbank offered rate. Thus, no further variables are needed for modelling the long-run behaviour of interest rates from an econometric point of view. Seen from an economic perspective though, the results are not quite satisfactory. We know that inflation expectations have a strong impact on the long-term interest rate. In a backward-looking model like this, various lagged price variables can be used as proxies for inflation expectations. For modelling the short-term dynamics in the error-correction equation the author tried the second differences GDP-deflator, the private consumption deflator and the CPI as well as the first difference of unit labour cost (all in logs). However, none proved significant. Thus, we are left with the unsatisfactory result, that the short-term interest rate is the only variable that has an effect on the long-term interest rate.

$$NL_t = + \underset{(0.304)}{2.09} + \underset{(0.025)}{0.81} \; NS_t + \hat{\nu}_t \tag{4.60}$$

The null hypothesis of a unit root in the residuals of the estimation in levels can be rejected at the one percent level. The respective t-value in the ADF test equation is -4.87 (5% critical value: -3.41). Contrary to theoretical considerations the estimated coefficient clearly differs from one.

However the Johansen cointegration test does not confirm cointegration.

Case	II
exogenous variables	0
endogenous variables	2
Trace stat. $r = 0$	16.88 (20.18)
Trace stat. $r \leq 1$	0.28 (9.16)

Table 4.50: Results of the Johansen cointegration test: long term interest rate

An Error-Correction Equation

The adjustment coefficient of the the error-correction term points to a relatively slow adjustment. However, the elasticity of the longterm interest rate with respect to the three months rate is close to one.

$$\begin{aligned}\Delta NL_t = & - \underset{(-4.78)}{0.26} \; (- \underset{(-1.98)}{1.10} - \underset{(-19.04)}{0.87} \; NS_{t-1} + NL_{t-1}) + \underset{(4.69)}{0.44} \; \Delta NL_{t-1} \\ & + \underset{(7.80)}{0.40} \; \Delta NS_t - \underset{(-2.79)}{0.19} \; \Delta NS_{t-1} - \underset{(-2.34)}{0.12} \; \Delta NS_{t-3} - \underset{(-2.24)}{0.10} \; \Delta NS_{t-5} + \hat{\varepsilon}_t\end{aligned} \tag{4.61}$$

The residuals are well-behaved. According to the RESET test there is no indication of either missing variables, incorrect functional form or a correlation of the residuals with the regressors. The CUSUM of squares test detects parameter instability during the period from 1994Q2 until 1997Q1. The fact that the long-term interest rate was underestimated in all out-of-sample forecasts may in fact be the consequence of missing inflation expectations.

Diagnostics	
Estimation period	1981Q3-2002Q4
Adjusted R^2	0.57
Residual tests	Probability
Normality test (Jarque-Bera)	[0.49]
Serial correlation LM test (lag 1)	[1.00]
Serial correlation LM test (lag 4)	[0.84]
Serial correlation LM test (lag 8)	[0.49]
ARCH LM test (lag 1)	[0.32]
ARCH LM test (lag 4)	[0.30]
White's heteroskedasticity test	[0.10]
RESET test (h=2)	[0.27]
Stability tests	outside error bands
CUSUM test	none
CUSUM of squares test	1994Q2-1997Q1

Table 4.51: Diagnostic tests: long-term interest rate