Measurement issues in political economy

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MEASUREMENT ISSUES IN POLITICAL ECONOMY

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Chapter 1

Introduction

1.1 Aim of the study

A large number of the variables that are used in economic models are mental constructs. Providing numerical values for these variables is often difficult, since they cannot be directly observed. Other variables that *can* be observed are then needed in order to use as proxies of the unobservable ones. It is not always obvious which variables should be selected for this purpose, and in the selection and aggregation procedures subjective judgement plays an important role. Consequently, different researchers often create and / or use different proxies, possibly leading to different results and conclusions.

The first aim of this thesis is to provide valid measurement for a number of important variables in the field of political economy. More specifically, this thesis aims to provide measurements that do not depend on subjective judgement. Instead of constructing yet another indicator, the information that is contained in already existing indicators is used, and an objective measure is constructed using aggregation procedures that do not involve subjective judgement.

A second aim of this thesis is to apply these newly constructed measures in empirical models, in order to assess the influence of the variables under scrutiny on economic outcomes. It is particularly interesting to examine whether the relations that have been derived in previous studies, using already existing indicators, remain intact.

1.2 Measurement issues

The main property of the variables that we try to measure in this thesis is that they are not directly observable: they are mental constructs that are difficult to capture numerically. This type of variables is called *latent variables*, and our aim is to find related variables that *are* measurable. We can then use the latter variables as proxies or indicators of the construct we are interested in, and apply techniques that combine the information contained in the indicators.

In this thesis, the focus is on four important variables in the field of political economy: central bank independence, central bank conservativeness (i.e. inflation aversion), corporatism and economic freedom. We examine how they can be measured using a latent variables approach. For this purpose, factor analysis and principal components analysis procedures are applied. Before we discuss the specific properties of these procedures, the variables at hand are examined more closely.

The variables of interest are macroeconomic variables, i.e. they are measured on a national level. The first question is to which set of countries we limit our research. Since there are large institutional differences between developed and developing countries, we often cannot capture both groups using the same measurement procedures. If we want to consider them both, different latent variables models need to be constructed for developed and developing countries. However, the availability of empirical data is often problematic for the developing countries, especially in earlier time periods. Therefore, we have chosen to concentrate mainly on the developed world. Only in the measurement of economic freedom and the assessment of its relation to economic growth are developing countries also taken into account. In the other studies, we focus on countries that are members of the OECD.

However, restricting the set of countries to those for which data are readily available does not mean that all potential measurement problems are avoided. First, there is the problem of a small sample size. By focusing on industrial countries, we restrict ourselves to a small group. In 1980, the OECD consisted of 23 members. In econometric terminology, this is already considered a small sample. However, due to the fact that for certain countries data for the relevant indicators are not available, most of the studies in this thesis are performed using an even smaller group of countries. The measurement of corporatism, for instance, uses a sample of 16 countries, while the conservativeness of central banks is measured using a sample of only 14 countries. We will have to pay attention to the econometric consequences of estimating models of sizes this small, and correct for them if necessary.

A second, related, problem is the occurrence of missing data. In addition to complete unavailability of data, it often happens that for some countries data are only available for a part of the sample period. This problem can be solved by excluding countries and / or time periods from the sample, but that implies that a lot of valuable information is thrown away. Instead, we can look for ways to combine the information that *is* available. At several points in this thesis, procedures of this type are applied to ensure that the already small sample of countries is not decreased even further.

As mentioned before, a latent variables approach is used in this thesis to measure the concepts of central bank independence, central bank conservativeness, corporatism and economic freedom. The resulting variables are used in models that assess their impact on economic outcomes, such as inflation, economic growth and unemployment, and on monetary policy outcomes, such as interest rates. Specifically, we examine the impact of central bank independence on inflation and on the occurrence of political business cycles in short-term interest rates, the impact of central bank conservativeness on unemployment, the impact of (aspects of) corporatism on inflation and unemployment and the impact of economic freedom on economic growth. The results of these studies are compared to results that have been reported in the empirical literature.

A problem that arises when latent variables are used as regressors in a regression model is that the corresponding coefficient is underestimated. We acknowledge this problem here and describe a method to get rid of the underestimation.

1.3 Outline

The remainder of this thesis consists of seven chapters. Chapter 2 gives an overview of the methodology that is used throughout the studies in chapters three through seven. Chapter 3 describes how central bank independence can be measured, while chapter 4 answers the question whether central banks can be blamed for the occurrence of political business cycles. In chapter 5, the measurement of corporatism is the subject of research, and chapter 6 discusses the measurement of central bank conservativeness. In these studies, only industrial countries are taken into account. Chapter 7 discusses which economic freedoms contribute to growth, in developing as well as developed countries. Finally, in chapter 8 the conclusions of the research are summarized.

The largest part of chapter 2 deals with an overview of latent variables models. Models with one or multiple latent variables are presented in the form of factor analysis (FA / MFA) and principal components analysis (PCA) specifications. Further topics of discussion are the estimation of these models, model selection and model fit and the interpretation of the results. Also, the CALS estimator is discussed, which takes care of the underestimation of the coefficient of the latent variable in empirical applications. Apart from latent variables models, models exhibiting dynamic behaviour are used in this thesis. Chapter 2 concludes with a discussion of exact tests that can be used to test for dynamics and unit roots in dynamic panel data models.

Chapter 3 deals with the measurement of central bank independence (CBI). After disentangling the concepts of CBI and inflation aversion (conservativeness), a new indicator of CBI is constructed using the information contained in indicators taken from the literature. This is done using a latent variables approach. Some specific technical complications of the FA model of CBI are discussed, and the resulting indicator is used in several empirical models to assess the relation between CBI and inflation.

In chapter 4, the constructed CBI indicator is used in a model that tests whether central banks actively create political business cycles (PBCs). After a short discussion of PBC theory, cross-country and panel regressions are run in order to answer the question whether central banks should be blamed for the occurrence of PBCs. In the section on panel data, the exact tests of chapter 2 are applied to see whether the model exhibits dynamic behaviour and a unit root.

Chapter 5 describes how to measure the concept of corporatism using a number of indicators from the vast political and economic literature on the subject. After selecting the indicators suitable for use in a latent variables analysis, the resulting constructs are applied in an empirical model that examines the relation between (aspects of) corporatism and inflation and unemployment. The CBI indicator of chapter 3 is also included in this model.

In chapter 6, inflation aversion or conservativeness of central banks is discussed. The concept of conservativeness is quantified using factor analysis models for two different decades. Also, the relation between conservativeness and unemployment is examined.

Chapter 7 describes the measurement of economic freedom for a large sample of developed and developing countries. In this chapter, a self-constructed measure of economic freedom is compared to two alternatives. Extensive robustness checks are used to analyze whether economic freedom contributes to economic growth.

Chapter 8, finally, summarizes the results and offers some additional points of discussion.

Chapter 2

Methodology

This thesis deals with applied research in the field of political economy. Before we present the results of this research, its econometric context needs to be discussed. The majority of the models described in the following chapters deal with issues concerning latent variables and measurement error. The first section of this methodological chapter provides a general introduction to latent variables models and the techniques to estimate them, while section 2.2 describes how to deal with underestimation of the coefficient of a latent variable in a regression model. In some of the models that are used in the thesis, dynamics play a role. Section 2.3 describes two tests for panel data models: a test whether or not the model is in fact dynamic, and a test for the presence of a unit root. These tests do not rely on asymptotic results.

2.1 Latent variables

Economic theory tries to describe the relationships between variables using mathematical models. To be able to apply these models empirically, a quantitative measure of these variables has to be available. Some economic variables are clearly defined and straightforward to measure, such as the consumer price level, the total value of exported goods or the number of unemployed people in the labour force of a country. Others, however, are more difficult to capture numerically. For mental constructs such as the independence of central banks, the level of corporatism or the amount of economic freedom of a country, different definitions and quantifications exist. These variables are called *latent* variables, since they are not directly observable. In order to use them in empirical models, we need to find observable related variables, also called *proxies* or *indicators*. This process creates *measurement error*, since the proxies are only approximations of the true unobservable phenomenons. This section describes the basic concepts of latent variables models with measurement error and the techniques that can be used to estimate them. For a thorough discussion of the subject, see Wansbeek and Meijer (2000).

The standard linear multiple regression model is written as

$$y = \Xi\beta + \epsilon, \tag{2.1}$$

with y an observable $N-$vector, ϵ an unobservable $N-$vector of random disturbances, β a $k-$vector of unknown parameters and Ξ an $N \times k-$matrix containing the regressors. The disturbances are assumed to be independently identically distributed (i.i.d.) with expectation zero and variance σ_ϵ^2, and the regressors are assumed to be uncorrelated with the disturbances. Now, if there is measurement error in the regressors, the matrix X is observed instead of the unobservable Ξ:

$$X = \Xi + V. \tag{2.2}$$

Here, V is an $N \times k-$matrix of measurement errors. Its rows are assumed to be i.i.d. with zero expectation and covariance matrix Ω, and are assumed to be uncorrelated with Ξ and ϵ.

Factor analysis The observable regressors X in equation (2.2) can also be assumed to be generated by the following model, known as the *multiple factor analysis* (MFA) model:

$$x_{ni} = \tau_i + \lambda_i'\xi_n + \delta_{ni}. \tag{2.3}$$

In model (2.3), there are k latent variables. The subscript i corresponds to the different observable regressors, and n to the observational units. Since in this thesis the factor analysis model is applied in a macroeconomic context, the observational units n are from now on denoted as countries. Then, x_{ni} denotes indicator i for country n and ξ_n is a $k-$vector containing the aspects of the unobservable concept (the *factors*) that the indicators are supposed to measure, for country n. The parameter τ_i captures the mean of indicator i, while λ_i is a $k-$vector of parameters (the *factor loadings*) that capture both the scale of indicator i and the strength of its relation to the factors. Further, δ_{ni} is a random measurement error, with mean zero and variance ψ_{ii}, often called the *unique variance*, and δ_{ni} and δ_{nj} are assumed uncorrelated for $i \neq j$. Both are assumed uncorrelated with the factors ξ_n.

The factor analysis model was originally developed in psychology to model the dependencies among different measures of intelligence (Spearman, 1904). The model is illustrated graphically by way of a path diagram in figure 2.1, for the simple case of one latent variable and three indicators. Drawing a path diagram obeys certain conventions. Circled variables denote latent variables, i.e. hypothetical constructs. Variables in square boxes denote observed variables, such as the different indicators of the latent variable. Variables that are not circled and not in square boxes denote error terms. An arrow denotes a causal dependency in the order indicated.

From (2.3), it is clear that the mean and variance of ξ can be chosen arbitrarily, because a change in its mean or variance can be counteracted by changing the corresponding τ or λ accordingly, without changing the observed variables. Hence, it is customary to let ξ have mean zero and variance one. An introduction to this type of measurement model can be found in Bollen (1989, chapter 6) or Wansbeek and Meijer (2000, chapter 7).

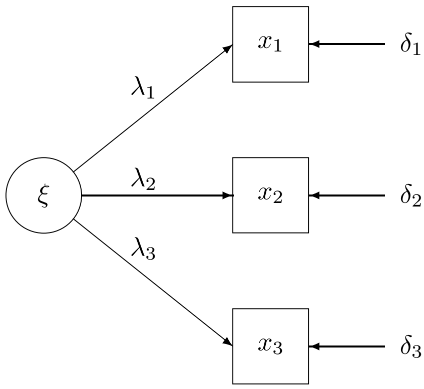


Figure 2.1: Path diagram of a factor analysis model with 1 factor and 3 indicators

Next, we discuss how the unknown parameters in model (2.3) can be estimated. For ease of exposition, we consider the simple factor analysis model with one factor, so $k = 1$. Estimation in MFA is in general a straightforward extension of this case. The parameters of the model, that is, the intercepts τ_i, the factor loadings λ_i, and the unique variances ψ_{ii}, are typically estimated from the means and covariance matrix of the indicators. From (2.3), the assumption that the errors are uncorrelated, and the imposed restriction that the mean of ξ is zero and its variance is one, it follows that the means of the indicators are

$\mu_i \equiv \mathrm{E}(x_i) = \tau_i$. Hence, a consistent estimator of τ_i is given by $\hat{\tau}_i \equiv \bar{x}_i$, the sample mean of the i-th indicator. The covariance of indicator i and indicator j is simply

$$\sigma_{ij} \equiv \mathrm{E}(x_i - \tau_i)(x_j - \tau_j) = \lambda_i \lambda_j, \tag{2.4}$$

where $i \neq j$. The variance of indicator i is

$$\sigma_{ii} \equiv \mathrm{E}(x_i - \tau_i)^2 = \lambda_i^2 + \psi_{ii}. \tag{2.5}$$

If there are at least three indicators, consistent estimators can be obtained by minimizing some sort of discrepancy between the sample variances and covariances and the theoretical variances and covariances as functions of the parameters, as given in (2.4) and (2.5). For example, (2.4) and (2.5) imply

$$\lambda_1 = \left(\frac{\sigma_{21}\sigma_{31}}{\sigma_{32}} \right)^{1/2}$$

and $\psi_{11} = \sigma_{11} - \lambda_1^2$. Hence, consistent estimators of λ_1 and ψ_{11} are given by

$$\hat{\lambda}_1 \equiv \left(\frac{s_{21}s_{31}}{s_{32}} \right)^{1/2}$$

and $\hat{\psi}_{11} = s_{11} - \hat{\lambda}_1^2$, where s_{ij} denotes the sample covariance between indicators i and j (provided that the expressions are nonnegative, cf. Dijkstra, 1992). However, when there are more than three indicators, generally more efficient estimators can be obtained that balance the discrepancies for different covariances optimally. Due to the small number of observational units in the empirical models, maximum likelihood procedures are used to estimate the unknown parameters throughout the rest of this thesis. Wansbeek and Meijer (2000) show that the estimation of the factor loadings and covariances in MFA comes down to solving a system of eigenequations. As will be shown later on, the corresponding eigenvalues are also used to determine the number of factors. The *reliability* of indicator i, denoted as r, is the squared correlation of the indicator and the unobservable concept ξ. As well as for the indicators, the reliability of the factor itself can also be estimated to assess the quality of the result.

In many cases, the scaling of the variables is arbitrary. For instance, there is no straightforward scale for central bank independence. In these cases, the model might be easier to interpret if the variables are rescaled such that they have variance 1. The corresponding solution of the FA model is called the *standardized solution*. The standardized solution is usually equivalent to the model estimated on (a reparametrization of) the correlation instead of the covariance matrix. In the one-factor model, the factor loadings of the standardized solution are simply the correlations of the indicators with the factor.

Now that the unknown parameters in (2.3) have been estimated, we would like to obtain values for the latent variable that we can use in empirical applications. These values are called *factor scores*, and the predictor $\hat{\xi}_n$ of ξ_n that has minimum mean squared error, under the restriction that the predictor is unbiased, is

$$\hat{\xi}_n = \lambda'\Sigma^{-1}(x - \tau)$$

for $k = 1$. This predictor is called the *Bartlett predictor*. Its expression for the MFA model follows analogously, see Wansbeek and Meijer (2000). If we rewrite the one-factor model as

$$\frac{x_{ni} - \tau_i}{\lambda_i} = \xi_n + \frac{1}{\lambda_i}\delta_{ni} = \xi_n + u_n,$$

we can write the Bartlett predictor as

$$\hat{\xi}_n = \lambda'\Sigma^{-1}\lambda(\xi_n + u_n) = \gamma\xi_n + v_n,$$

where $\gamma = \lambda'\Sigma^{-1}\lambda$ and $v_n = \lambda'\Sigma^{-1}\lambda u_n$.

Since the factor ξ is assumed to have expectation 0 and variance 1, and ξ and v_n are assumed to be uncorrelated, the variance of $\hat{\xi}_n$ is

$$\text{var}(\hat{\xi}_n) = \text{E}(\hat{\xi}_n^2) = \gamma^2 + \text{var}(v_n),$$

and consequently the estimated reliability $\hat{r}$ is

$$\hat{r} = \frac{\gamma^2}{\text{var}(\hat{\xi}_n)} = \frac{\gamma^2}{\gamma^2 + \text{var}(v_n)}. \tag{2.6}$$

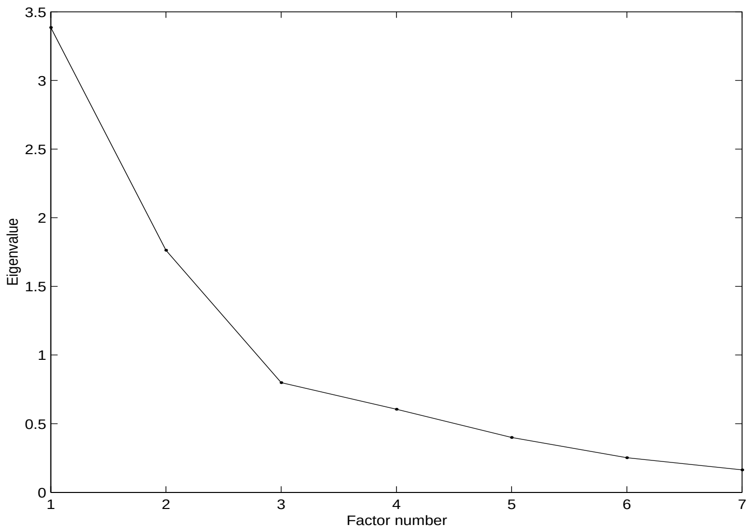
Finally, since the original factor ξ is assumed to have expectation 0 and variance 1, we would like the same to hold for our estimated counterpart. This means that $\hat{\xi}$ needs to be adjusted by a factor γ such that

$$\tilde{\xi}_n = \frac{1}{\gamma}\hat{\xi}_n = \xi_n + \frac{1}{\gamma}v_n = \xi_n + \tilde{v}_n. \tag{2.7}$$

The variance of $\tilde{v}$ in (2.7) is called the *measurement error variance*, and is denoted as ϕ.

Related issues The MFA model can be used in two different situations. In exploratory factor analysis (EFA), the analysis is purely exploratory and does not use subject matter theory to restrict the model parameters. Confirmatory factor analysis (CFA), however, takes subject matter theory as its point of departure and uses restrictions on the factor loadings and the covariance matrix of the factors. The EFA model, by contrast, is estimated on the basis of the correlation matrix. If we want to perform EFA, we have to choose the number of factors k to be used in the analysis. There are two commonly applied rules to do this, which are both based on the eigenvalues of the correlation matrix of the standardized variables. The first rule, which is also called the *Kaiser rule*, says that relevant factors correspond to eigenvalues larger than 1. The second rule uses the *scree plot*, which plots the number of factors against the eigenvalues. It states that the number of factors to select is the number of eigenvalues before a 'kink' that is often found in the scree plot. An example is shown in figure 2.2

Figure 2.2: Scree plot of a factor analysis



There are two eigenvalues larger than 1 in figure 2.2. Also, the number of factors before the kink in the plot is two, so using either rule leads to the same conclusion: we should select two factors in this case. If the two rules have different outcomes, the number of factors corresponding to the solution that is interpreted more easily is selected.

It may happen that the solution of a MFA is difficult to interpret. In that case, we can make use of the fact that the matrix of factor loadings is not identified: it can be multiplied with any orthonormal matrix without affecting the distribution of the indicators. Put differently, the factor loadings matrix is open to *rotation*, yielding a solution that may be easier to interpret because the matrix has a simpler structure. Ideally, each indicator is correlated with as few factors as possible. Several rotation methods can be applied: the one most frequently used in practice is varimax rotation. Another rotation method that is used in this thesis is direct oblimin rotation, which minimizes the correlation between columns of the factor loadings matrix. For a more detailed discussion of rotation and rotation methods see Wansbeek and Meijer (2000, pp. 167–169).

Finally, there are a number of criteria available to judge the fit of a factor analysis model. Of these criteria, we use the following two in this thesis: the χ^2-statistic, which compares the proposed restricted model to an unrestricted alternative (the *saturated model*), and the so-called *comparative fit index* (CFI) which considers the proposed model compared to a highly restrictive *null model*. In the null model, all factor loadings are restricted to zero. The CFI is an assessment of model fit that is especially valuable in small samples, which are encountered in most of the macroeconomic models in this thesis. An elaborate discussion of these and other model fit measures is found in chapter 10 of Wansbeek and Meijer.

Principal components analysis As we have mentioned above, the factor analysis model imposes a specific structure on the covariance matrix, implying assumptions that may not be satisfied in practical applications. As an alternative, one may drop the assumptions and try to find a Ξ and λ in (2.1) so that the resulting errors are small is some sense or other. This is the idea behind a data analysis method called *principal components analysis* (PCA). Here, the columns of the matrix Ξ are the principal components of y, which are uniquely obtained as a linear combination of the observed variables. If the number of indicators is not too small, the solutions of PCA and FA are quite similar. An application of PCA is found in chapter 7, where the measurement of economic freedom is discussed. More on PCA is found in Wansbeek and Meijer (2000).

2.2 Measurement error in a single regressor: the CALS estimator

When a latent variable is used as a regressor in a regression model, we are confronted with the problem that it can only be imperfectly measured. It is well known that a neglect of this problem leads to inconsistent estimation results. In particular, the coefficient of the latent variable will be underestimated. This phenomenon is quite widespread (as discussed earlier, many economic variables are theoretical ideals that allow no direct measurement) and is equally widely ignored. It may explain the experience, well known to applied researchers, that regression coefficient estimates are often disappointingly low.

If the variance of the differences between true and observed values of the regressor is known, it is possible to adapt the results and come up with consistent estimates. Meijer and Wansbeek (2000) and Wansbeek and Meijer (2000, section 5.2) describe how to do it, using the general approach due to Kapteyn and Wansbeek (1984) known as the consistent adjusted least squares (CALS) estimator.

Although the principle is easy to understand, the attention paid to it is modest. The main reason for this is that the condition of a known variance is usually not met in practice. However, if we know the reliability of the construct that has been derived from the unobservable phenomenon by factor analysis techniques, the adaptation to adjust the underestimation is carried out easily. Here, knowing the reliability should be interpreted in the sense of a consistent estimator being available. In the previous section, we have seen that the reliability of the factor can be estimated from the factor model. Consequently, CALS estimators can be computed.

Summarizing, our aim is to provide consistent estimation results and get rid of the underestimation of the coefficient of the latent variable. Then, the term we are interested in is the measurement error variance ϕ. From equation (2.7) in section 2.1 it follows that

$$\phi = \mathrm{var}(\tilde{v}_n) = \frac{\mathrm{var}(v_n)}{\gamma^2} = \frac{1}{r} - 1.$$

Since an estimate of r is available, the measurement error variance ϕ follows.

To correct for the downward bias, we define the matrix $A_N \equiv \frac{1}{N}X'X$ and scalar $\alpha \equiv e_1' A_N^{-1} e_1$. Here, X is the $N \times k$ matrix of observed variables of equation (2.2) and e_1 is an $N-$vector with first element equal to 1 and zeros otherwise. Further, let $\theta = \frac{1}{1-\phi\alpha}$. Then, if b_1 is the estimated coefficient of the latent variable using OLS, the CALS estimator that corrects for underestimation in case of measurement error is

$$\hat{\beta}_1 = \theta b_1.$$

Now, we can also derive the t-value that corresponds to a measurement error of size ϕ. The t-statistic in case of no measurement error is

$$t_0 = \frac{b_1\sqrt{N}}{\sqrt{s_\epsilon^2 \alpha}}.$$

Then, the t-statistic corresponding to a measurement error of size ϕ is given by

$$t_\phi = \frac{t_0}{\sqrt{1 + \frac{2}{N}(\theta - 1)^2 t_0^2}}.$$

Meijer and Wansbeek show that, while the coefficient estimate increases when the CALS estimator is used, the t-value decreases. Consequently, for large values of the measurement error variance ϕ, the coefficient estimate becomes insignificant.

2.3 An exact test for dynamic panel data models

In this section, we shortly step away from models containing latent variables to discuss econometric models that include dynamics. In particular, we are interested in panel data models that contain a lagged dependent variable among the regressors. We use the simple model specification

$$y = \gamma y_{-1} + X\beta + Z\alpha + u, \tag{2.8}$$

where y and y_{-1} are $NT \times 1$-vectors. For ease of exposition, we let the regressor matrix X be an $NT \times 1$-vector, with β its corresponding unknown parameter. The procedure outlined below can easily be generalized to the case of k regressors. The term $Z\alpha$ denotes the individual specific effects, with Z an $NT \times N$-matrix defined as $Z = \iota_T \otimes I_N$ and α an $N \times 1$-parameter vector. Here, ι_T is a vector consisting of T ones, and I_N is the identity matrix of order N. Finally, u is an $NT \times 1$-disturbance vector with variance matrix $\sigma^2 I$. Then, u/σ has expectation 0 and variance 1 and, under normality, does not depend on unknown parameters. Our aim is to derive an exact test for different values of γ. To do this, we use a test described in Van den Doel and Kiviet (1995) that does not rely on asymptotics.

Since the inclusion of a lagged dependent variable complicates the estimation procedure of a panel data model, it is useful to test first whether the model is dynamic. In this case, the null hypothesis is $H_0 : \gamma = 0$. Next, we adapt this test to test for unit roots, with null hypothesis $H_0 : \gamma = 1$. In chapter 4, these tests are applied in an empirical, macroeconomic context.

2.3.1 Testing for dynamics

To test for $\gamma = 0$ in model (2.8), we might use the fixed effects estimator of γ,

$$\hat{\gamma}_{FE} = \frac{y'_{-1} M_{XZ} y}{y'_{-1} M_{XZ} y_{-1}},$$

where the matrix $M_{XZ} = I_{NT} - (X, Z)\{(X, Z)'(X, Z)\}^{-1}(X, Z)'$ is the projector orthogonal to (X, Z). However, basing a test on $\hat{\gamma}_{FE}$ has the disadvantage that its distribution depends on unknown parameters. If we estimate γ from a different regression in an augmented model, this dependence can be avoided and the exact distribution of the resulting estimator of γ can be computed. In order to do so, define lag operators

$$B_0 = \begin{bmatrix} 0 & 0 \\ I_{T-1} & 0 \end{bmatrix},$$

and

$$B = B_0 \otimes I_N.$$

Our aim is to estimate γ in the augmented regression model

$$\begin{aligned} y &= \gamma y_{-1} + X\beta + BX\beta^* + Z\alpha + BZ\alpha^* + u \\ &= \gamma y_{-1} + W\theta + u, \end{aligned}$$

where $W = (X, BX, Z, BZ)$ and θ contains all parameters except γ. The estimator for γ in this model is

$$\hat{\gamma}^*_{FE} = \frac{y'_{-1} M_W y}{y'_{-1} M_W y_{-1}}, \tag{2.9}$$

which can also be written as

$$\hat{\gamma}^*_{FE} = \frac{u'B'M_W u}{u'B'M_W B u}. \tag{2.10}$$

Under $H_0 : \gamma = 0$, the distribution of this estimator does no longer depend on unknown parameters. Moreover, it is easily computed, since we can draw a large number of normally distributed vectors u and simulate the distribution of $\hat{\gamma}^*_{FE}$ using (2.10).

The augmented regression comes down to adding BX, which contains the regressors X lagged by one period, to the regression, along with BZ. Since

$BZ = Z - e_1 \otimes I_N$, adding this term means an additional transformation to the original data: adding BZ is equivalent to ignoring the first N observations of X, BX and Z. For notational convenience, we still denote the transformed data as X, BX and Z, which are now two vectors of order $N(T - 1) \times 1$ and one matrix of order $N(T - 1) \times N$, respectively. The matrix W is redefined as $W = (X, BX, Z)$.

In order to compute $M_W u$ and $M_W Bu$ in (2.10), the matrix $W'W$ needs to be inverted. Since this matrix can become quite large, it is useful to make the computation more efficient by avoiding direct inversion of the matrix. This is done as follows. To compute $M_W u$, let $R = (X, BX)$ and consider

$$u = R\delta + Z\eta + \epsilon = W(\delta', \eta')' + \epsilon.$$

The unknown parameters δ and η are estimated by

$$(\hat{\delta}', \hat{\eta}')' = (W'W)^{-1}W'u,$$

hence the estimate of u is

$$\hat{u} = W(\hat{\delta}', \hat{\eta}')' = W(W'W)^{-1}W'u.$$

Then

$$M_W u = (I_{N(T-1)} - W(W'W)^{-1}W')u = u - \hat{u}.$$

Using the Frisch-Waugh theorem, see Wansbeek and Meijer (2000, p.352), the parameters δ and η can be estimated separately by

$$\hat{\delta} = (R'M_Z R)^{-1}R'M_Z u$$

and

$$\hat{\eta} = (Z'Z)^{-1}Z'(u - R\hat{\delta}).$$

To start with $\hat{\delta}$, note that

$$M_Z = (I_{N(T-1)} - Z(Z'Z)^{-1}Z') = I_{N(T-1)} - \frac{1}{T-1}(\iota_{T-1}\iota_{T-1}' \otimes I_N).$$

To write this in a more convenient form, use $\text{vec}(ABC) = (C' \otimes A)\text{vec}\, B$, see Wansbeek and Meijer (2000, p.350). If we let $u = \text{vec}\, U$, where the matrix U is of order $N \times (T - 1)$ with elements u_{it}, then

$$M_Z u = u - \frac{1}{T-1}\text{vec}(U\iota_{T-1}\iota'_{T-1}).$$

The computation of $M_Z R$ follows by analogy:

$$M_Z R = (X, BX) - \frac{1}{T-1}(\iota_{T-1}\iota'_{T-1} \otimes I_N).$$

Let $X = \text{vec}\,\tilde{X}$ and $BX = \text{vec}\,\widetilde{BX}$ with $\tilde{X}$ and $\widetilde{BX}$ matrices of order $N \times (T-1)$, then

$$M_Z R = (X, BX) - \frac{1}{T-1}\{\text{vec}\,(\tilde{X}\iota_{T-1}\iota'_{T-1}), \text{vec}\,(\widetilde{BX}\iota_{T-1}\iota'_{T-1})\},$$

where $\text{vec}\,(\tilde{X}\iota_{T-1}\iota'_{T-1})$ is simply a vector containing the column sums of $\tilde{X}$ stacked $T-1$ times. Then, $\hat{\delta} = (\hat{\delta}_1, \hat{\delta}_2)'$ is obtained by regressing $M_Z u$ on the columns of $M_Z R$, and

$$\begin{aligned} \hat{\eta} &= (Z'Z)^{-1}Z'(u - R\hat{\delta}) = \frac{1}{T-1}Z'(u - R\hat{\delta}) \\ &= \frac{1}{T-1}(\iota'_{T-1} \otimes I_N)(u - R\hat{\delta}). \end{aligned}$$

The expression for $\hat{\eta}$ can be computed more easily by noting that

$$u - R\hat{\delta} = u - \hat{\delta}_1 X - \hat{\delta}_2 BX = \text{vec}(U - \hat{\delta}_1\tilde{X} - \hat{\delta}_2\widetilde{BX})$$

and, using $\text{vec}(ABC) = (C' \otimes A)\text{vec}\,B$ again,

$$\hat{\eta} = \frac{1}{T-1}\text{vec}\{(U - \hat{\delta}_1\tilde{X} - \hat{\delta}_2\widetilde{BX})\iota_{T-1}\}.$$

Now, all that is left to obtain γ^*_{FE} in (2.10) is the term $M_W Bu$. This can be done by analogy to the computation of $M_W u$, using

$$Bu = W(\delta'_B, \eta'_B)' + \epsilon,$$

and estimating δ_B and η_B as above, yielding $\widehat{Bu} = W(\hat{\delta}'_B, \hat{\eta}'_B)'$. Then

$$\hat{\gamma}^*_{FE} = \frac{u'B'(u - \hat{u})}{u'B'(Bu - \widehat{Bu})},$$

which concludes the computation of the test statistic.

2.3.2 Testing for unit roots

The procedure described for testing dynamics can easily be adjusted to derive an exact test for a unit root. Let

$$\Gamma = \begin{bmatrix} 1 & & 0 \\ \vdots & \ddots & \\ 1 & \cdots & 1 \end{bmatrix} \otimes I_N.$$

Then, the aim is to estimate γ in the augmented regression model

$$\begin{aligned} y &= \gamma y_{-1} + X\beta + B\Gamma X\beta^* + Z\alpha + B\Gamma Z\alpha^* + u \\ &= \gamma y_{-1} + W\theta + u, \end{aligned}$$

where $W = (X, B\Gamma X, Z, B\Gamma Z)$ and θ contains all parameters except γ. The estimator for γ, under the null hypothesis $H_0 : \gamma = 1$, is

$$\hat{\gamma}^{**}_{FE} = \frac{y'_{-1}M_W y}{y'_{-1}M_W y_{-1}}, \tag{2.11}$$

which can also be written as

$$\hat{\gamma}^{**}_{FE} = 1 + \frac{u'\Gamma' B' M_W u}{u'\Gamma' B' M_W B\Gamma u}. \tag{2.12}$$

This estimator has a distribution that does not depend on unknown parameters. It is noted that $B\Gamma = \Gamma - I_{NT}$, so the augmented regression is computed using $W = (X, \Gamma X - X, Z, \Gamma Z - Z)$.

As in the previous subsection, the terms $M_W u$ and $M_W B\Gamma u$ in (2.12) can be computed more efficiently by avoiding direct inversion of the matrix $W'W$, using $R = (X, \Gamma X - X)$ and $Q = (Z, \Gamma Z - Z)$. Further, let

$$u = R\delta + Q\eta + \epsilon = W(\delta', \eta')' + \epsilon,$$

and

$$\Gamma u = R\delta_\Gamma + Q\eta_\Gamma + \epsilon = W(\delta'_\Gamma, \eta'_\Gamma)' + \epsilon,$$

to get $\hat{u} = W(\hat{\delta}', \hat{\eta}')'$ and $\hat{\Gamma u} = W(\hat{\delta}'_\Gamma, \hat{\eta}'_\Gamma)'$. Then equation (2.12) can be written as

$$\hat{\gamma}^{**}_{FE} = 1 + \frac{(u'\Gamma' - u')(u - \hat{u})}{(u'\Gamma' - u')(\Gamma u - \widehat{\Gamma u})(u - \hat{u})},$$

which concludes the computation of $\hat{\gamma}^{**}_{FE}$.

Chapter 3

Measuring central bank independence: a latent variables approach*

3.1 Introduction

Nowadays, it is widely believed that an inflation-averse ('conservative') and independent central bank may help to assure price stability. Indeed, many countries have recently adapted their central bank laws accordingly. There is quite some empirical evidence suggesting that central bank independence (CBI) helps to reduce inflation. For industrial countries, this evidence generally consists of cross-country regressions using proxies for CBI based on the statutes of the central bank. Still, this evidence has been challenged on various grounds; see Berger, de Haan and Eijffinger (2001) for an extensive review. For instance, in a widely cited paper, Campillo and Miron (1997) argue that once control variables are included, the CBI indicator they use plays no role in explaining cross-country inflation differentials in a sample of high-income countries.

One crucial question in this respect is how to measure independence, also making sure that independence is properly distinguished from what is commonly called *conservativeness* (i.e. the inflation aversion) of the central bank. Unfortunately, the concepts of independence and conservativeness are often not carefully disentangled in empirical indicators for CBI, as pointed out by Berger et al. (2001) and Romer (2001). Can central bank independence and conservativeness be measured on the basis of what the central bank law has to say on

*This chapter is an adapted version of De Haan, Leertouwer, Meijer and Wansbeek (2002).

the criteria deemed relevant? Forder (1999) argues "[W]hat are the criteria for a good measure? It cannot be 'one which gives a relation with inflation' since that makes the hypothesis unfalsifiable. [...] What we are lacking is an objective reason to prefer one [measure] over the other. Without this, there can be no test of the independence hypothesis."

Although Forder has a point, we argue that legal indicators are not useless. Independence and conservativeness are theoretical concepts defined on the basis of the unobservable loss functions of the government and the central bank. Hence they are latent variables that cannot be directly observed in practice. In order to include them in empirical models, indicators are needed to proxy independence and conservativeness. Usually this is done by setting up criteria that are thought to be related to the degree of independence, and by assigning scores to the central banks in the various countries under examination. Both steps require subjective judgement and substantive knowledge about central bank laws. It is, therefore, not surprising that different authors have come up with different indicators.

All available indicators try to measure the unobservable concepts of CBI and conservativeness and, no doubt, none of them is perfect. Following most of the literature, the focus in this chapter is on central bank independence. Conservativeness and its measurement are the topic of chapter 6. Here, we consider five CBI indicators. After disentangling the criteria related to conservativeness from those related to independence, we use a latent variables approach to develop an alternative measure of CBI. A latent variables approach to issues concerning the measurement of CBI has been used previously by Eijffinger and Schaling (1998). However, their context differs from ours: their aim is to test a model on the determinants of the optimal level of CBI, whereas in the present chapter we use a latent variable approach to develop an alternative indicator of CBI.

The remainder of this chapter is organized as follows. In section 3.2, we examine five CBI indicators that have been proposed in the literature around 1990, for 22 countries. First, we adjust some of these indicators to eliminate the aspect of conservativeness. Section 3.3 describes the factor analysis model as well as some technical issues that arise specifically in our case. In section 3.4 the results of the factor analysis are discussed. In section 3.5, we examine whether CBI is related to inflation differentials. Section 3.6 concludes.

3.2 Indicators of CBI

We consider five (legal) measures of CBI, as developed by Alesina (1988), Cukierman (1992), Cukierman, Webb and Neyapti (1992), Eijffinger and Schaling (1992) and Grilli, Masciandaro and Tabellini (1991). All indicators are based on central bank laws of the 1980s. At the end of this chapter, we consider other time periods. We have decided to use the aggregated CBI indicators in our analysis instead of the various elements on the basis of which they have been constructed, because it is the aggregate index that yields the degree of CBI according to the author(s) who constructed the index. This aggregate is basically the total score on a number of questions that, according to the author(s), are relevant in order to proxy CBI. These questions differ (sometimes) substantially across the various indicators. See Eijffinger and De Haan (1996) for a detailed comparison of the indicators used.

The indicators of Alesina (ALES) and Eijffinger and Schaling (ES) are constructed by quantifying the final responsibility for monetary policy, the number of government officials on the governing board and the appointment of board members by the government. Grilli, Masciandaro and Tabellini (GMT) present indices of political and economic independence, which are combined to serve as an indicator of legal independence. The indicators of Cukierman (CUK) and Cukierman, Webb and Neyapti (CWN) are aggregated from sixteen legal characteristics of central-bank charters, grouped into eight variables. The only difference between the two is that the CUK indicator is an unweighted average of the eight variables concerned while the CWN indicator consists of a weighted average. Because of this small difference, the CUK and CWN indicators are highly correlated. How we deal with this will be discussed later. Table 3.1 presents the values of the various indicators.

It appears that the indicators are scaled differently. More importantly, countries are ranked differently according to the various indicators. For instance, the central bank of Canada is seen as quite independent by GMT and the Cukierman indicators, but as quite dependent by ES. For an extensive review of the differences, see Eijffinger and De Haan (1996).

Before we describe how to handle the indicators we need to examine them a little closer. De Haan and Kooi (1997) show that CUK, CWN and GMT include issues related to the objective(s) of monetary policy. Although the concept of (goal) independence would suggest otherwise, the measures have a higher score if the law stipulates that stable prices are the primary objective of the central bank. The reason for this is that economic theory implies that a conservative and independent central bank are necessary to reduce inflation. In other words, CUK, CWN and GMT measure the conservativeness of central banks "as embodied in the law" (Cukierman 1992, p.377), in addition to the various aspects of legal independence proper. As we are only interested in independence proper, we adjust the indicators such that the element of conservativeness is removed.

Table 3.1: Indicators of CBI

Country	ALES	CUK	CWN	ES	GMT
Australia	1	.31	.36	1	9
Austria	–	.58	.63	3	9
Belgium	2	.19	.16	3	7
Canada	2	.46	.45	1	11
Denmark	2	.47	.50	4	8
Finland	2	.27	.28	3	–
France	2	.28	.29	2	7
Germany	4	.66	.69	5	13
Greece	–	.51	.55	–	4
Iceland	–	.36	.34	–	–
Ireland	–	.39	.44	–	7
Italy	1.5	.22	.25	2	5
Japan	3	.16	.18	3	6
Netherlands	2	.42	.42	4	10
New Zealand	1	.27	.24	3	3
Norway	2	.14	.16	2	–
Portugal	–	.41	.41	2	3
Spain	1	.21	.17	3	5
Sweden	2	.27	.29	2	–
Switzerland	4	.68	.64	5	12
UK	2	.31	.34	2	6
USA	3	.51	.49	3	12

This is done by subtracting the score on the question of whether the bank by law has to pursue monetary stability from the value of the indicator, after which the adjusted versions of CUK and CWN are rescaled. The adapted indicators are denoted by CUK-a, CWN-a and GMT-a and are shown, along with ALES and ES, in table 3.2. These are the idicators that are used in our latent variables approach.

3.3 The factor analysis model

In this section, we describe the latent variables approach we use in order to evaluate to which extent the CBI indicators describe the same unobservable phenomenon. The different indicators of CBI are correlated, as they should, because they are intended to measure the same concept. The correlation matrix of the indicators is given in table 3.3. Because of missing data (not all authors have considered the same set of countries, as is apparent from table 3.2), the correlations are computed by the expectation-maximization (EM) algorithm based on

Table 3.2: Indicators of CBI excluding conservativeness

Country	ALES	CUK-a	CWN-a	ES	GMT-a
Australia	1	.29	.35	1	8
Austria	–	.58	.62	3	8
Belgium	2	.22	.19	3	7
Canada	2	.49	.50	1	10
Denmark	2	.45	.48	4	7
Finland	2	.19	.18	3	–
France	2	.32	.28	2	7
Germany	4	.61	.64	5	12
Greece	–	.47	.51	–	4
Iceland	–	.35	.33	–	–
Ireland	–	.33	.37	–	6
Italy	1.5	.22	.26	2	5
Japan	3	.18	.21	3	5
Netherlands	2	.37	.35	4	9
New Zealand	1	.25	.22	3	3
Norway	2	.16	.20	2	–
Portugal	–	.47	.48	2	3
Spain	1	.15	.17	3	4
Sweden	2	.28	.31	2	–
Switzerland	4	.78	.75	5	11
UK	2	.32	.28	2	6
USA	3	.52	.49	3	11

the assumption of multivariate normality. The EM algorithm was suggested by Dempster, Laird and Rubin (1977) to solve maximum likelihood problems with missing data. It is an iterative method, in which the expectation step involves forming a log-likelihood function for the latent data as if they were observed and taking its expectation. In the maximization step, the resulting expected log-likelihood is maximized. More on the use of the EM algorithm in missing-data problems can be found in Ruud (1991).

Table 3.3: Correlations between the five CBI indicators

Indicator	ALES	CUK-a	CWN-a	ES	GMT-a
ALES	1.000				
CUK-a	.628	1.000			
CWN-a	.609	.980	1.000		
ES	.625	.439	.406	1.000	
GMT-a	.632	.624	.589	.349	1.000

As can be seen from table 3.3, the correlations between the different indicators are not perfect. Therefore, we consider the different indicators of CBI as imperfect measures of this concept, generated by the factor analysis model

$$x_{ni} = \tau_i + \lambda_i \xi_n + \delta_{ni}, \tag{3.1}$$

where x_{ni} is indicator i for country n, ξ_n is the CBI of country n, τ_i is the parameter that captures the mean of indicator i, λ_i is the factor loading that captures both the scale of indicator i and the strength of its relation to the factor ξ, and δ_{ni} is a random measurement error, with mean zero and variance ψ_{ii}. Further, δ_{ni} and δ_{nj} are assumed uncorrelated for $i \neq j$, and both are assumed uncorrelated with the factor ξ_n.

The covariance of indicator i and indicator j is

$$\sigma_{ij} \equiv \mathrm{E}(x_i - \tau_i)(x_j - \tau_j) = \lambda_i \lambda_j,$$

where $i \neq j$. The variance of indicator i is

$$\sigma_{ii} \equiv \mathrm{E}(x_i - \tau_i)^2 = \lambda_i^2 + \psi_{ii}.$$

We have to address a few further technical issues when applying the factor analysis model to the CBI data. The first one is that the indicators CUK-a and CWN-a are differently weighted combinations of the same components. Consequently, they are always expected to be highly correlated a priori, whether or not they measure CBI satisfactorily. This conflicts with the model assumptions, which state that two indicators should be approximately uncorrelated when one or both do not measure the concept well. There are two ways in which this problem can be solved. The first is to choose one of these indicators as the 'preferred' one and omit the other from the analysis. Then, we have a factor analysis model with four indicators that better satisfies the assumptions. The second solution is to augment the model with a parameter ψ_{23}, say, that reflects the covariance between the measurement errors of these two indicators. In this way, no information is lost and no arbitrary choice has to be made, whereas the new model is still reasonably consistent with our prior ideas. Therefore, we have chosen this solution. The proposed model is depicted in figure 3.1, where the assumed correlation is depicted using an arrow with arrowheads on both sides.

A second technical issue concerns the nonnormality of the data and the small sample size. The indicators are (weighted) combinations of yes/no dummies and are thus clearly nonnormally distributed. Explicit methods for nonnormal data, such as the asymptotically distribution-free method of Browne (1984), exist, but require large sample sizes (1000 or larger) to estimate the model well.

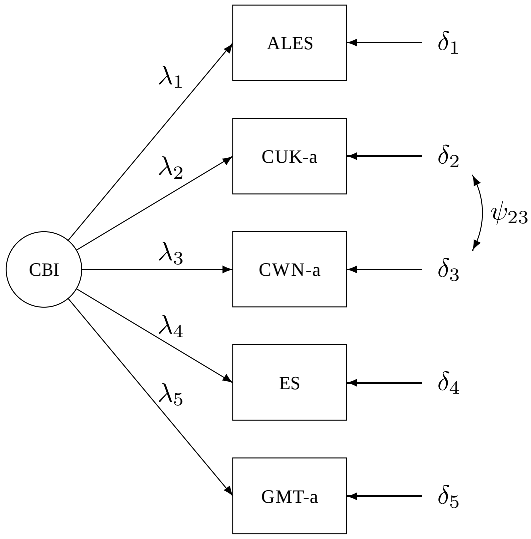


Figure 3.1: Path diagram of the proposed model

We have, however, a small sample of only 22 countries. The literature suggests that for small samples the maximum likelihood method based on the assumption of multivariate normality generally gives the most reliable estimates, even if the data are clearly nonnormally distributed (e.g. Chou and Bentler, 1995; West, Finch and Curran, 1995). Standard errors based on likelihood theory are, however, incorrect in this situation. Therefore, we use the method of White (1982) in order to get standard errors and t-statistics that are robust to this kind of misspecification.

The third technical issue is induced by the missing data. Not all indicators are available for all 22 countries. The two standard approaches to dealing with missing data are *listwise deletion*, which removes all countries that have at least one missing indicator, and *pairwise deletion*, which computes the covariance of two indicators on the basis of the countries that have no missing values on those two indicators, but may have other missing indicators. The drawback of listwise deletion is that many cases may be removed. In our case, 8 countries would have to be removed, leaving only 14 to analyze. This is, of course, a relatively large reduction and throws away a lot of information. Pairwise deletion uses almost all information in the data, but is not guaranteed to give a positive semidefinite covariance matrix. This is exactly what happens in our case. Fortunately, with maximum likelihood estimation, there is a solution that uses all information and that does not suffer from this problem. This full information maximum likelihood method is described in Arbuckle (1996) and implemented in the program AMOS (Arbuckle, 1997). Therefore, we use AMOS to estimate our FA model.

3.4 Empirical findings

In this section, we present the results of the analysis described above. The estimates of the parameters and corresponding White-corrected t-statistics are given in table 3.4. In the table, superindices * and ** denote significance at 5% and 1% levels, respectively. With a few exceptions, the parameters are statistically significantly different from zero at the conventional level of $\alpha = .05$. For the parameters that are not significantly different from zero, this is probably due to small sample size, and not because the parameter is actually zero.

Table 3.4: Factor analysis estimation results

	Estimate	t-stat	Reliability
Intercept (τ_i)			
ALES	2.1361^{**}	11.40	
CUK-a	0.3638^{**}	10.62	
CWN-a	0.3717^{**}	10.81	
ES	2.7710^{**}	11.29	
GMT-a	6.8743^{**}	11.58	
Loading (λ_i)			
ALES	0.7328^{*}	2.08	.7922
CUK-a	0.1192	1.78	.5503
CWN-a	0.1147	1.78	.5064
ES	0.6907	1.28	.4063
GMT-a	1.8622	1.86	.5059
Error variance (ψ_{ii})			
ALES	0.1409	1.20	
CUK-a	0.0116^{**}	3.41	
CWN-a	0.0128^{**}	3.69	
ES	0.6970^{**}	2.94	
GMT-a	3.3876^{**}	2.81	
Error covariance (ψ_{23})			
CUK-a, CWN-a	0.0117^{**}	3.44	

The χ^2-statistic has a value of 2.75 with 4 degrees of freedom, which lies well below the 5% critical value of 9.49. The comparative fit index CFI has a value of 1. According to these measures, the model fits very well.

In addition to the estimates and their t-values, the last column of table 3.4 gives the estimates of the reliability of each indicator, which is its squared correlation with the factor CBI. In terms of the model parameters as introduced above, the reliability of indicator i is given by λ_i^2/σ_{ii}.

According to this measure, ALES performs the best and ES the worst, with reliabilities of .7922 and .4063, respectively. The indicator of Alesina has the highest correlation with the latent variable CBI, but is still an imperfect reflection of it.[1] In other words, using the ALES indicator of CBI is not the best approach conceivable. The other four indicators, even though they correlate less with CBI than ALES, still convey some information on CBI that is not captured by ALES. Hence, a next logical step is to compute the factor scores for the latent variable CBI. As described in chapter 2, the linear combination of the indicators that has minimum mean squared error is given by $\lambda'\Sigma^{-1}(x-\tau)$, where λ, x and τ are vectors with typical elements λ_i, x_i and τ_i, respectively, and the matrix Σ has typical element σ_{ij}. Of course, in practical applications the parameters are replaced by the corresponding consistent estimators.

The computed factor scores are given in table 3.5. We order the countries by the values of their CBI (for the period to which the data apply), from high (Switzerland) to low (Spain). In fact, when the assumptions underlying the factor analysis model are maintained, these country scores are the best indicator of CBI in the 1980s given the input. In the remainder of this thesis, this indicator is dubbed the Factor Score (FS) indicator. Just as with the individual indicators, we can assess the quality of the result by its reliability. Its estimate is .8714, at least for those countries for which all five indicators are available. For countries for which not all indicators are available, reliabilities are lower, which reflects the according loss of information.

3.5 Is CBI related to inflation?

In order to see how the approach described in the previous section can be applied in practice, we present a simple model for (average) inflation during 1980–1989, where CBI is included as an explanatory variable, and compare the results for the different measures that have been discussed. We use the influential paper of Campillo and Miron (1997) as a benchmark. Campillo and Miron claim that CBI is a relatively unimportant determinant of inflation performance, using the CUK indicator as measure of CBI. Campillo and Miron do not examine how sensitive their results for high-income countries are with respect to the choice of the CBI indicator. As we will show below, this omission crucially affects their

[1]This finding is quite remarkable, as the Alesina index has been severely criticised in the literature; see Eijffinger and De Haan (1996).

Table 3.5: CBI scores of the FS indicator

Country	Score	Reliability
Switzerland	2.3923	.8714
Germany	2.1986	.8714
USA	1.1384	.8714
Austria	0.7298	.7477
Netherlands	0.2404	.8714
Japan	0.2141	.8714
Denmark	0.1775	.8714
Canada	0.0805	.8714
Iceland	0.0273	.5566
France	−0.2070	.8714
Greece	−0.2125	.6950
Belgium	−0.2262	.8714
UK	−0.2633	.8714
Finland	−0.3168	.8519
Ireland	−0.3222	.6950
Sweden	−0.3591	.8519
Portugal	−0.4774	.7477
Norway	−0.5660	.8519
Italy	−0.8686	.8714
Australia	−1.0340	.8714
New Zealand	−1.1252	.8714
Spain	−1.2206	.8714

conclusion.

We start with a very simple bivariate cross-sectional model that describes the relation between inflation and CBI for the 22 OECD countries mentioned earlier. Since the data for the various indicators of CBI refer to the 1980s, our sample period runs from 1980 until 1989. Inflation data, measured as the change in consumer prices, have been obtained from the IMF's International Financial Statistics, and have been averaged for the sample period of interest. A constant term is included in the specification, and the model coefficients are estimated using OLS. Because Greece and Iceland seemed to behave somewhat erratically during the sample period, we have added dummies to the model specifications that include the CBI indicators for these countries. Without these dummies, the explanatory power of the model is much lower, see also Eijffinger and De Haan (1996). The results, including the corresponding t-statistics and adjusted R^2 measures, are shown in table 3.6.

The results in the table indicate a negative relation between CBI and inflation, i.e. a greater independence of the central bank is related to a lower inflation

Table 3.6: Estimation results for a simple bivariate regression model of inflation

Measure	CBI-coeff	t-stat	$\bar{R}^2$
ALES	-0.0256^{**}	-5.356	0.610
CUK-a	-0.0681	-1.557	0.795
CWN-a	-0.0640	-1.446	0.792
ES	-0.0168^{*}	-2.527	0.211
GMT-a	-0.0105^{**}	-4.332	0.648
FS	-0.0251^{**}	-4.151	0.872

rate, which is consistent with theory. With the exception of the Cukierman indicators, the estimated coefficients are statistically significant. The model specification that uses the FS indicator has the highest $\bar{R}^2$-value.

Of course, the model described above is extremely simple. In order to examine whether the results are robust, we look at the extended model in section 3.2 of Campillo and Miron (table 5), where the following additional variables are included:

- Average inflation in the previous period
- Political instability
- The degree of openness
- Log of GDP per capita
- Participation in a fixed exchange rate regime
- The need for tax revenue

As our sample periods runs from 1980 until 1989, we use averaged inflation over the period 1970–1979 as the first control variable (INFL7079). Political instability (POLINST) is measured as the total number of government changes during the sample period.[2]

Traditionally, the degree of openness is measured as the level of imports relative to GDP. Lane (1997) shows that the link between openness and inflation only holds for the most highly developed countries when country size is held constant. To correct for this effect, the log of real GDP in 1985 (LGDP85) is also included as a regressor. Romer (1993) argues that openness puts a check on the government's incentive to engage in unanticipated inflation, because of induced exchange rate depreciation. He has found that over a large sample of

[2]Alternatively, we have used the number of changes in the colour of the government (from left-wing to right-wing and vice versa) as measure of political instability. This produces similar results.

countries and over a broad time frame average inflation is significantly lower in smaller, more open countries. However, Terra (1998) argues that the negative link between inflation and openness is largely driven by the response of severely indebted countries to the debt crisis of the 1980s. The smaller a country's trade share, the more depreciated the real value of the currency must be in order to generate a given trade surplus. And the more depreciated the currency, the greater the pressure on the government's budget, and, hence, the pressure to inflate. See also Romer (1998) for a critique on this view.

A further control variable is the log of real GDP per capita in the year 1985 (LGDPPC85). A dummy variable is included to control for participation in an exchange rate system, the European Monetary System in this case. This variable has value 1 if the corresponding country is a member of the Exchange Rate Mechanism of the EMS and 0 otherwise. Finally, the need for tax revenue is measured by the level of debt relative to output at the start of the sample period (DEBT80). Due to data availability and the fact that we are only interested in OECD countries, we have excluded the Summers-Heston measure for quality of the data as a control variable.

Instead of including all control variables simultaneously, as Campillo and Miron suggest, we use a step-wise procedure. Each time, only one control variable is added to the bivariate model, in order to assess its specific effect on the relation between CBI and inflation. This avoids the possibility of overfitting due to a large number of regressors with a small sample size. In these regressions, the FS indicator of CBI is used. The results are given in table 3.7.

Table 3.7: Estimation results for regressions for inflation including control variables

Control	coeff	t-stat	CBI-coeff	t-stat	$\bar{R}^2$
INFL7079	0.9290^{**}	5.321	-0.0035	-0.614	0.941
POLINST	0.0004	0.116	-0.0249^{**}	-4.027	0.865
OPENNESS	-0.0552	-1.059	-0.0205^{**}	-3.172	0.870
LGDP85	-0.0095	-1.542			
LGDPPC85	-0.0724^{**}	-2.781	-0.0156^{*}	-2.511	0.900
EMS	-0.0113	-0.906	-0.0243^{**}	-4.053	0.870
DEBT80	-0.0258	-0.703	-0.0294^{**}	-4.005	0.868

From the table it is clear that the coefficient for CBI remains significant in all cases, except when inflation in the previous period is included in the regression. The conclusion by Campillo and Miron that CBI is a relatively unimportant determinant of inflation performance is largely based on this phenomenon. The model with inflation in the previous period included is, however, not very satisfactory. It merely states that countries with an above average inflation rate in the 1970s also had an above average inflation rate in the 1980s. It does not give an explanation of inflation differences among countries or why these differences are maintained, nor does it give directions based on which policy makers could attempt to reduce inflation rates if that is considered desirable.

The second reason why Campillo and Miron reach a different conclusion is because they only use the Cukierman indicator to measure CBI, which is insignificant in the bivariate model as well as when the control variables are included. In order to see to which extent the choice of indicator influences the results, we give adjusted R^2 values for the different model specifications using the various indicators of CBI in table 3.8.

Table 3.8: $\bar{R}^2$ values of the regressions for inflation including control variables, using different CBI indicators

	Indicator					
Control	ALES	CUK-a	CWN-a	ES	GMT-a	FS
INFL7079	0.657	0.943	0.943	0.742	0.831	0.941
POLINST	0.589	0.786	0.783	0.172	0.624	0.865
OPENNESS[a]	0.678	0.822	0.823	0.305	0.636	0.870
LGDPPC85	0.583	0.873	0.873	0.596	0.675	0.900
EMS	0.604	0.796	0.793	0.165	0.631	0.870
DEBT80	0.555	0.778	0.776	0.113	0.628	0.868

[a] Model specification includes LGDP85.

The results show that in all cases except (barely) the one including inflation in the previous period, the specification that uses the FS indicator has the highest explanatory power. We conclude that the negative, significant relation we have found between inflation and CBI is robust.

CALS estimators As discussed in section 2.2, the use of a latent variable as a regressor in a regression model leads to inconsistent estimation results due to measurement error, and underestimation of the coefficient of the latent variable. Since we have an estimate of the reliability of the construct of CBI here, we can get rid of the underestimation and get consistent estimates by calculating CALS estimators. The results for the specifications with and without control variables are presented in table 3.9. In the second and third column, estimated coefficients and t-values ignoring measurement error are given. In the last two columns, CALS estimates are shown.

It is clear from table 3.9 that, as predicted, the estimated coefficients increase while the t-values decrease when CALS estimators are used. The first effect is

Table 3.9: CALS estimates of the FS indicator

Control	coeff	t-stat	CALS coeff	CALS t
–	-0.0251^{**}	-4.151	-0.0305^{**}	-4.005
INFL7079	-0.0035	-0.614	-0.0055	-0.611
POLINST	-0.0249^{**}	-4.028	-0.0307^{**}	-3.878
OPENNESS[a]	-0.0205^{**}	-3.172	-0.0266^{**}	-3.053
LGDPPC85	-0.0156^{*}	-2.511	-0.0210^{*}	-2.430
EMS	-0.0243^{**}	-4.053	-0.0297^{**}	-3.910
DEBT80	-0.0294^{**}	-4.005	-0.0396^{**}	-3.684

[a] Model specification includes LGDP85.

much larger than the second, however, and in none of the cases the significance of the results is affected.

CBI in other decades The analysis in this chapter has been performed for the period 1980–1989. It can, however, be performed for other time periods as well. Data on the Cukierman indicators are also available for the periods 1950–1959, 1960–1971 and 1972–1979. An update for the period 1990–1997 has been constructed by van Lelyveld (2000).[3] The other indicators are only available for the 1980s. Since the level of CBI has not changed dramatically between 1960 and 1989, a factor analysis for 1960–1971 and 1972–1979 can be performed using the relevant values of the Cukierman indicators while assuming the values of the remaining indicators still valid for these periods. Due to the fact that there have been a number of changes in central bank laws since 1990, as van Lelyveld describes, this assumption is less straightforward for the 1990s. In spite of this, we have also constructed the FS indicator for the period 1990–1997, using the updated Cukierman indicators and the unchanged others. In empirical applications, however, we use both the FS indicator and the Cukierman indicator of CBI if the sample period includes the 1990s. The values of the FS indicator for the different decades are shown in table 3.10.

[3] Van Lelyveld's updated value for Spain contains an error. We have corrected this.

Table 3.10: Scores of the FS indicator for different decades

Country	1960–1971	1972–1979	1980–1989	1990–1997
Australia	−1.0664	−1.1892	−1.0340	−1.1623
Austria	0.8057	0.8263	0.7298	0.5872
Belgium	−0.1365	−0.2181	−0.2262	−0.0616
Canada	−0.0279	−0.0983	0.0805	−0.0901
Denmark	0.0590	0.0180	0.1775	0.0989
Finland	−0.1789	−0.2356	−0.3168	−0.2217
France	−0.1018	−0.2477	−0.2070	−0.1410
Germany	2.2598	2.2717	2.1986	2.3210
Greece	−0.6219	−0.0323	−0.2125	−0.3995
Iceland	−0.3610	−0.1974	0.0273	−0.3163
Ireland	−0.0232	−0.0645	−0.3222	−0.3280
Italy	−0.6896	−0.7800	−0.8686	−0.7318
Japan	0.7573	0.7144	0.2141	0.6032
Netherlands	0.0351	−0.0002	0.2404	0.1384
New Zealand	−1.2554	−1.3094	−1.1252	−1.2219
Norway	−0.2136	−0.2799	−0.5660	−0.3340
Portugal	−1.0824	−0.6282	−0.4774	−0.4407
Spain	−1.2265	−1.3146	−1.2206	−1.1518
Sweden	−0.1476	−0.2256	−0.3591	−0.2117
Switzerland	2.2638	2.2495	2.3923	2.3005
UK	−0.0881	−0.2873	−0.2633	−0.2927
USA	1.0401	1.0284	1.1384	1.0559

3.6 Conclusions

In this chapter, we have shown how the partial conflict between competing indicators of CBI can to a certain extent be resolved by using latent variables modeling. The resulting ordering of countries by the degree of CBI is of interest by itself, but the results also allow for more satisfactory statistical inference in regression models where CBI is an explanatory variable.

In contrast to the results of the influential study by Campillo and Miron (1997), we find that the CBI indicator we have constructed is significantly related to inflation, also when various control variables as suggested by Campillo and Miron are included.

Finally, the CALS estimator can be used to deal with problems concerning measurement error and underestimation of regression coefficients. This leads to a reasonable increase of the estimated coefficients of the CBI indicator, while the significance of the results remains intact.

Chapter 4

Who creates political business cycles: should central banks be blamed?*

4.1 Introduction

There are two principal reasons why central banks are independent. First, the inflationary bias is reduced, provided that the central bank is more conservative than the government. Many empirical studies provide evidence for that.[1] Second, "the most obvious advantage a fully independent central bank has is that of not being influenced by electoral deadlines" (Muscatelli, 1998). That the incumbent government may be inclined to stimulate the economy before elections to enhance re-election probabilities is well-known, see Alesina, Roubini and Cohen (1997) for an overview. Are central banks also influenced by electoral deadlines? Put differently, if we observe political business cycles (PBCs) in macroeconomic variables, such as unemployment and the growth rate, who is responsible for creating them – and who should or should not be blamed?

Surprisingly, the empirical literature has little to say about the exact role of governments and central banks when it comes to PBCs. Worse, in most previous studies different institutional features have largely been neglected. Often the scope for electorally-motivated monetary policies is reduced, since national or international restrictions bind central bankers. In a regime of fixed exchange

*This chapter is based on Leertouwer and Maier (2001) and Leertouwer and Maier (2002).

[1] See Eijffinger and de Haan (1996) and Berger et al. (2001) for overviews. Posen (1998), however, challenges this view by providing evidence that a higher degree of central bank independence does not reduce disinflation costs.

rates, for example, opportunistic policies are less likely to occur than in a flexible exchange rate system. Similarly, independent central banks are less likely to be involved in electorally motivated policies than central banks that are under the spell of the government. The restricting effects of these institutional features are recognized in economic theory, yet empirical papers on PBCs do not explicitly control for them.

Clark et al. (1998) argue that common cross-country studies of PBC models may be seriously flawed since they do not account for institutional differences that constrain national policymakers.[2] However, these authors only examine economic outcomes, namely output growth and unemployment. Although these variables are likely to be influenced by monetary policy, there are a number of other influences that may offset or reinforce the impact of monetary policy. Furthermore, the rational political business cycle model predicts that policymakers manipulate instruments while the effects on outcomes are less certain.

This chapter focuses on policy outcomes for which the central bank *can* be held responsible, namely the short-term interest rate. We can thereby answer the question of whether the central bank can be blamed for active opportunistic behavior. Our sample runs from the 1960s until 1997 and consists of monthly data for 14 OECD countries.

The results are simple and strikingly robust. The short-term interest rate shows hardly any sign of a PBC. We thus reject the hypothesis that central banks have actively been engaged in opportunistic behavior.

The outline of this chapter is as follows. In the next section, we explain the PBC models in more detail and show how national or international constraints can prevent politicians from using monetary policy for short-sighted purposes. Our estimation results are presented in sections 4.3 and 4.4. In section 4.5, we summarize our findings.

4.2 When do political business cycles occur?

4.2.1 Electoral pressure on the economy

A test for the existence of PBCs requires the following: first, we need a theoretical basis to explain why such short-sighted behavior might be pursued by the government or the central bank. Second, one has to account for restrictive institutional features that constrain implementation of such a policy. And finally, we need an appropriate measure for the central bank's policy stance.

[2] Clark and Hallerberg (2000) formulate this idea in terms of a theoretical model.

The theoretical framework

The first model on PBCs has been developed by Nordhaus (1975). It is based on the assumptions that politicians care only about their re-election and that voters judge the incumbent's performance by the state of the economy. The economy is characterized by an exploitable Phillips-curve and the incumbent can directly control the rate of inflation. Nordhaus assumes that the government is responsible for both monetary and fiscal policy.

Under the assumption of adaptive or non-rational expectations, the incumbent government has an incentive to pursue expansive economic policies before elections to enhance its probability of re-election by lowering the unemployment rate. After elections, the government has to reduce inflation with contractionary monetary policies, thereby raising the unemployment rate, before switching to expansionary policies again as the next election approaches. Due to the poor memory of the voters, this cycle might be repeated endlessly.

Such behavior is called 'opportunistic'. The testable prediction of Nordhaus' model is that before elections the unemployment rate drops due to expansionary policies, while after elections inflation is high and contractionary measures are taken. Similar patterns apply to economic instruments.

A common criticism has concerned the assumption of adaptive expectations. This has led to a reformulation of the model by Rogoff and Sibert (1988) who expand the framework to a 'rational political business cycle model' (RPBC). They assume that voters lack information about the competence of the politicians and in order to appear 'competent', policymakers manipulate policy instruments. The RPBC model predicts visible cycles in economic instruments, and short, possibly irregular cycles ('blips') in economic variables such as the inflation rate or the unemployment rate.

Restrictive institutional features

The above models make the simplifying assumptions that (a) the central bank and the government pursue similar policies, and (b) policymakers have sufficient national autonomy to implement their policies. Both assumptions need not hold in reality. The following two types of constraints may prevent governments from implementing an opportunistic policy:

National constraint If the central bank enjoys a low degree of statutory independence, then it is likely that pressure is applied such that monetary policy follows the opportunistic pattern set by fiscal policy. In this case, electoral cycles may be observed in monetary instruments. However, central banks may be independent. If we abstract from the idea that central banks have their own interests, due to which they prefer one government over another, then we should not expect them to behave opportunistically.[3] After all, one of the main arguments for making a central bank independent is that this enables decisions to be based on a longer time-horizon, which rules out short-sighted behavior. Still, even for central banks with a high degree of statutory independence, it might be rational to comply to the government's wishes, as independence might be threatened by a change in the central bank law. If this is the case, then independent as well as dependent central banks have an incentive to engage in PBC behavior. This argument has been put forward by Frey and Schneider (1981) and Berger and Schneider (2000).

International constraint Economic theory tells us that, under a regime of fixed exchange rates and high capital mobility, the scope for autonomous economic policies is reduced. Following the worldwide increase in capital mobility in the 1970s, we can thus expect more restricted opportunities for implementing a national monetary policy for countries that have either been member of a fixed exchange rate regime (such as the Bretton Woods system or the European Monetary System EMS), or that have pegged their currency unilaterally. Participation in a fixed exchange rate regime restricts national economic policies and restricts opportunities for PBCs.

Measuring monetary policy

Still unsolved is the question how monetary policy should be measured. Clearly unemployment or growth rates cannot be solely attributed to government or central bank policies, as has been suggested in Alesina and Roubini (1992), Soh (1986) and the original paper from Nordhaus (1975). Previous studies on PBCs in monetary policy have in general focused on monetary aggregates. Examples of this approach are Alesina, Roubini and Cohen (1997), Allen and McCrickard (1991), or Vaubel (1997), who all use M1. A survey of evidence for the US can be found in de Haan and Gormley (1997). Still, a PBC in, say, M1, does not necessarily imply active central bank behavior: if, for instance, the incumbent government uses expansionary fiscal policy before elections, and the central bank tolerates this behavior, then obviously a monetary aggregate must reflect preelectoral manipulation. Berger and Woitek (1997, 2001) have looked at the German case. They find cycles in M1, which could indicate opportunistic behaviour

[3] In Vaubel (1997) it is assumed that the central bank follows its own interests.

of the Deutsche Bundesbank. However, their findings indicate that the Bundesbank did not target a monetary aggregate, but rather economic variables such as inflation or output. Therefore, Berger and Woitek conclude that the cycle in M1 was demand-driven rather than supply-driven. However, it would be unfair to (fully) blame the central bank, as the PBC was created by the government. To answer the question of whether central banks regularly abuse monetary policy, evidence should be found in monetary instruments. This has the additional advantage that the distinction between PBC and RPBC no longer matters: both models predict cycles in policy instruments. Only under the assumption of non-rational expectations do these cycles translate into policy outcomes.

There is, however, one problem. It is nearly impossible to determine a 'key variable' that fully characterizes the current monetary policy stance. For most countries, focusing on one single instrument is not possible. In Germany, for example, different instruments have been used over time, and the relative weight of these instruments has changed considerably. Open market operations, for example, which were the most powerful monetary tool in the late 1980s and 1990s, were fully developed only in 1985. For most countries *the* monetary instrument does not exist.

Still, there is a possibility to circumvent these problems. The direction of monetary policy is reflected in the behavior of interest rates, as monetary instruments either directly or indirectly influence interest rates. Therefore, interest rates can be viewed as capturing the 'net effect' or the 'sum' of all monetary instruments. A similar view has been taken in Maier (2000).

There is a second argument why the choice of an interest rate might be appropriate. If politicians attempt to influence a central bank before elections, the political request will in most cases not be formulated in terms of a monetary aggregate ('increase the growth rate of M1!'), but in term of interest rates ('lower the interest rate!'). This point has first been made by Johnson and Siklos (1994), who use a short-term interest rate to estimate a VAR model.

What interest rate pattern can we expect before elections? If we assume a standard IS/LM model, three possibilities arise: first, only expansionary monetary policy is used and the LM-curve shifts to the right. Then we would expect the interest rate to fall before elections and (correctly) infer that the central bank caused the PBC.

Second, if only expansionary fiscal policy is used, the IS-curve shifts to the right and we would expect an increase in the interest rate before elections. Both cases can be easily distinguished. There is a third case that is more difficult: imagine a situation where monetary policy accommodates expansionary fiscal policy. Then, both curves shift to the right. If accommodation is perfect, the interest rate could remain unchanged. A similar pattern could be observed if both the government and the central bank try to create a PBC in such a way that the combination of both effects on the interest rate 'cancels out', or if the central bank follows an interest rate rule (horizontal LM-curve). In these cases we have an identification problem.

The above considerations show that the use of interest rates as an indicator of PBCs is not without caveats. In practice, however, these problems are less serious: all available evidence indicates that the third case is highly unlikely. In order to show up in a pooled regression, this case would require a PBC in fiscal policy in all countries in our sample. Such a cycle has for most countries been rejected by Alesina et al. (1997).

Therefore we have the following expectations: if interest rates remain unchanged (or go up) prior to elections, central banks have refrained from opportunistic manipulation. If the interest rate decreases before an election, we conclude that the central bank has *actively created* a PBC. This is reflected in a comment by Goodhart (1994, pp. 1426-27), who claims that interest rates are manipulated by politicians:

> "... those in charge of central banks generally regard monetary base control as a non-starter. The instrument which they can, and do, control is the short-term money market rate.
>
> Politicians ... suggest that an electorally inconvenient interest rate increase should be deferred, or a cut 'safely' accelerated. This political manipulation of interest rates .. leads to a loss of credibility..."

Few researchers have tested whether this claimed influence on interest rates does indeed exist. In this chapter we use short-term interest rates that are tightly controlled by the central banks and reflect their intentions.[4] If PBCs exist and if they have been actively created by central banks, they should be visible in the short-term interest rates.

4.2.2 Institutional constraints

National constraints

To account for national constraints, we need to classify the degree of statutory central bank independence in various countries. First, we use the index of CBI

[4] The German Bundesbank has considered the day-to-day rate as a 'key indicator' (Deutsche Bundesbank, 1995).

developed by Cukierman, Webb and Neyapti (1992), since it is the indicator most commonly used in the empirical literature. As an alternative, we apply the FS indicator that was developed in the previous chapter. These indicators are available for four subsets of the sample period, namely 1960–1971, 1972–1979, 1980–1989 and 1990–1997.

We divide the countries into two groups: those having scores above the median, which we consider as 'independent central banks', and those with scores below the median ('dependent central banks'). Countries ranked above the median value are marked with '+' (i.e. national constraints are present) and countries ranked below the median value with '-' (i.e. absence of national constraints). We expect this classification to give a reliable overall ranking of the degree of statutory independence: in the '+' countries PBCs are not likely to occur in monetary policy, as these countries have an independent central bank. Countries with a '-' have a central bank with a low degree of statutory independence and if PBCs are to be found, we expect them there. The classification, for the different time periods, is shown in table 4.1.

Table 4.1: Presence of national constraints

	1960–1971		1972–1979		1980–1989		1990–1997	
Country	CWN	FS	CWN	FS	CWN	FS	CWN	FS
Austria	+	+	+	+	+	+	+	+
Belgium	-	-	-	-	-	-	+	+
Canada	+	+	+	+	+	+	+	+
Denmark	+	+	+	+	+	+	+	+
Finland	-	-	-	-	-	-	-	-
France	+	+	-	-	-	+	+	+
Germany	+	+	+	+	+	+	+	+
Italy	-	-	-	-	-	-	-	-
Japan	-	+	-	+	-	+	-	+
Norway	-	-	-	-	-	-	-	-
Spain	-	-	-	-	-	-	-	-
Sweden	-	-	-	-	-	-	-	+
UK	+	+	-	-	-	-	-	-
US	+	+	+	+	+	+	+	+

From the table it is clear that the differences between the different indicators and the different time periods are small. The main difference is that CBI in Japan is classified as below-median by the CWN indicator and above-median by the FS indicator. Furthermore, if we assess the presence of a national constraint in terms of the CWN indicator, we see that the level of CBI in Belgium changes from below median until 1989 to above median in the 1990s. We also see that, in

1972, France and the UK experienced a change in CBI sufficiently large to change signs in table 4.1.[5] While CBI in the UK remains below the median value for the rest of the period, the value of France changes back to above median in the 1990s. If we use the FS indicator, we again see that central bank independence in France changes from above median to below median and back. For France, the level of CBI seems to hover around the median value. The level of CBI in Belgium and Sweden changes from below median until 1989 to above median in the 1990s. In the UK, both indicators show that the level of CBI is above the median until 1971 and below afterwards.

International constraints

Next, we examine when monetary policy could be controlled in an entirely autonomous way in the countries in our sample. Therefore, we check for participation in fixed exchange rate regimes.

Table 4.2: Participation in fixed exchange rate systems

Country	Snake	EMS	Unilateral peg
Austria	–	1995 : 01–1997 : 12	1973 : 04–1994 : 12
Belgium	1973 : 04–1978 : 12	1979 : 01–1997 : 12	–
Canada	–	–	–
Denmark	1973 : 04–1978 : 12	1979 : 01–1997 : 12	–
Finland	–	1996 : 01–1997 : 12	1977 : 01–1995 : 12
France	Intermittent	1979 : 01–1997 : 12	–
Germany	1973 : 04–1978 : 12	1979 : 01–1997 : 12	–
Italy	–	1979 : 01–1992 : 08	–
Japan	–	–	–
Norway	1973 : 04–1978 : 12	–	1979 : 01–1997 : 12
Spain	–	1989 : 01–1997 : 12	–
Sweden	1973 : 04–1976 : 12	–	1977 : 01–1997 : 12
UK	–	1990 : 10–1992 : 08	–
US	–	–	–

The 1970s mark a turning point in at least two relevant respects. First, international capital mobility has increased sharply since the 1970s. Secondly, the Bretton Woods system of fixed exchange rates collapsed in 1973. Clark et al. (1998) assume that during the Bretton Woods period capital mobility was sufficiently low for the fixed exchange rate system not to constrain monetary policy. While we generally follow their methodology in deriving the institutional constraints, we believe that capital mobility was sufficiently high before 1973 to

[5] In Clark et al. (1998), only the UK experiences a national constraint.

effectively constrain monetary policy. Therefore we interpret the Bretton Woods system as a constraint for all countries but the US, which dominated the system. Table 4.2 shows the time periods during which countries were part of a fixed exchange rate system.[6] The first two columns of table 4.2 show the participation in the Snake (the predecessor of the European Monetary System) and in the European Monetary System (EMS), respectively. The third column shows when countries have pegged their currencies unilaterally.

Italy and the UK left the EMS after the turmoil in 1992, see Deutsche Bundesbank (1993, p.93). As Germany was the anchor currency in the EMS, we do not count the EMS as a binding restraint for German monetary policy, see Lohmann (1998). Following Clark et al., we do not consider the Snake to be a binding constraint for France, as its participation was highly unstable. In table 4.3 we define the dummy variable FEX. The first column shows the periods during which countries have maintained fixed exchange rates. In other words, in these periods monetary policy autonomy is absent, and the dummy FEX has value $+1$. The second column shows when autonomous monetary policy could have been pursued ($FEX = 0$).

Table 4.3: Presence of international constraints

Country	Fixed Exchange Rates ($FEX = 1$)	No international constraint ($FEX = 0$)
Austria	1960 : 01−1997 : 12	−
Belgium	1960 : 01−1997 : 12	−
Canada	1960 : 01−1973 : 03	1973 : 04−1997 : 12
Denmark	1960 : 01−1997 : 12	−
Finland	1960 : 01−1973 : 03, 1977 : 01−1997 : 12	1973 : 04−1976 : 12
France	1960 : 01−1973 : 03, 1979 : 01−1997 : 12	1973 : 04−1978 : 12
Germany	1960 : 01−1978 : 12	1979 : 01−1997 : 12
Italy	1960 : 01−1973 : 03, 1979 : 01−1992 : 08	1973 : 04−1978 : 12, 1992 : 09−1997 : 12
Japan	1960 : 01−1973 : 03	1973 : 04−1997 : 12
Norway	1960 : 01−1997 : 12	−
Spain	1960 : 01−1973 : 03, 1989 : 01−1997 : 12	1973 : 04−1988 : 12
Sweden	1960 : 01−1997 : 12	−
UK	1960 : 01−1973 : 03, 1990 : 10−1992 : 08	1973 : 04−1990 : 09, 1992 : 09−1997 : 12
US	−	1960 : 01−1997 : 12

[6] Source: Clark et al. (1998).

Summary: National and international constraints

Summarizing the above, table 4.4 shows for each country the combination of both constraints, for both measures of CBI.

Table 4.4: National and international constraints

FEX	*CBI: CWN indicator*	
	above median	below median
No autonomy	Austria, Belgium (1990–1997), Denmark	Belgium (1960–1989), Norway, Sweden
For part of period	Canada, France (1960–1971 & 1990–1997), Germany, UK (1960–1971)	Finland, France (1972–1997), Italy, Japan, Spain, UK (1972–1997)
Always autonomy	US	
	CBI: FS indicator	
	above median	below median
No autonomy	Austria, Belgium (1990–1997), Denmark, Sweden (1990–1997)	Belgium (1960–1989), Norway, Sweden (1960–1989)
For part of period	Canada, France (1960–1971 & 1980–1997), Germany, Japan, UK (1960–1971)	Finland, France (1972–1979), Italy, Spain, UK (1972–1997)
Always autonomy	US	

Only the US experienced monetary policy autonomy during the entire sample period. However, this does not pose a problem, as we find substantial variation in the other parts of the table to give us reliable estimates of the influence of fixed exchange rate systems.

In a regression analysis, we would not expect to find PBCs in countries that are constrained in either way. Clark et al. (1998) have shown this hypothesis to hold for policy outcomes, such as inflation or unemployment rates. Their approach cannot, however, reveal the precise role of central banks, since these policy outcomes are influenced by many additional factors (e.g. supply and demand shocks). If we find cycles in policy outcomes, we cannot conclude that the central bank actively creates them. For cycles in the short-term interest rate, the central bank *can* be blamed. In the next section, we conduct country-specific tests of the PBC hypothesis. In order to examine the effects of cross-national differences in the constraints, we apply panel regressions in section 4.4.

4.3 Country-specific tests

The country-specific tests we apply seek to examine whether a significant degree of covariation exists between elections and the short-term interest rate if we control for institutional restrictions. We start with a general model description incorporating changes in the international constraint. A model that incorporates changes in the national constraint is discussed later in this section. Following Alesina, Roubini and Cohen (1997), we start with the following model specification:

$$\begin{aligned} r_{it} &= \beta_{0i} + \beta_{1i} E_{it} + \beta_{2i} FEX_{it} + \beta_{3i} E_{it} * FEX_{it} \\ &\quad + \sum_{j=1}^{p} \beta_{j+3,i} r_{i,t-j} + \delta_{1i} IP_{it} + \delta_{2i} \pi_{it} + \delta_{3i} OC_t + u_{it}, \end{aligned} \tag{4.1}$$

for countries $i = 1, \cdots, N$ and time periods $t = 1, \cdots, T$. In model (4.1), r_{it} is the log of the nominal short-term interest rate. When estimating the model, using the log of the interest rate is preferable on econometric grounds: in our estimations, the disturbances u_{it} are assumed to be normally distributed, thus having positive and negative values. While the interest rate can only be positive, the log of the interest rate can take on negative as well as positive values and thereby fits the model assumptions better. It is noted, however, that the results are not affected qualitatively by the log transformation. E_{it} is the election variable, for which we use a *pre-election period* of 12 months. That is, the dummy variable E_{it} has value $+1$ in the month of a general election *and* in the eleven preceding months, and 0 otherwise. Different pre-election periods are discussed when we examine the robustness of our results. FEX_{it} is the coefficient for participation in fixed exchange rate systems as defined in table 4.3. The interaction term $E_{it} * FEX_{it}$ is also included, equaling $+1$ during electoral periods in countries that lack monetary policy autonomy. Finally, we have added three additional economic variables: industrial production IP_{it}, as a proxy for GDP growth, the inflation rate π_{it}, measured as the change in consumer prices, and a dummy variable that covers the impact of the oil crises in the 1970s (OC_t). To obtain comparable values for the different countries in our sample, we use monthly IFS data on the short-term interest rate for 14 OECD countries. The sample period starts for most countries in the 1960s and goes until 1997. Further details on the data can be found in appendix 4.A.

The interpretation of equation (4.1) is as follows: if a country does not participate in a fixed exchange rate regime and determines its monetary policy autonomously, then $FEX_{it} = 0$. In this case the model is reduced to a standard PBC model:

$$\begin{aligned} r_{it} &= \beta_{0i} + \beta_{1i}E_{it} + \sum_{j=1}^{p} \beta_{j+1,i} r_{i,t-j} + \delta_{1i} IP_{it} + \delta_{2i}\pi_{it} \\ &+ \delta_{3i} OC_t + u_{it}. \end{aligned} \tag{4.2}$$

If, however, $FEX_{it} = 1$ and an international constraint is present, then the sum $E_{it} + E_{it} * FEX_{it}$ becomes important: if our argument is correct that the absence of monetary policy autonomy decreases the probability that politicians will manipulate the macroeconomy for electoral purposes, we do not expect the sum of the coefficients of $E_{it} + E_{it} * FEX_{it}$ to be significantly different from zero. To examine this, we perform a Wald test in these cases to test for $\beta_1 + \beta_3 = 0$. Note, however, that this only makes sense if the election coefficient E_{it} is significantly different from zero, otherwise PBCs cannot be found in our sample.

For all country-specific tests, the models include lagged dependent variables, the order of which (p) is determined by examining the (partial) autocorrelation function. With respect to possible stationarity of the interest rate, we adopt the approach of Bierens (1997) and take the short-term interest rates to be nonlinear trend stationary processes. To see whether the inclusion of lagged disturbances is necessary, we have performed Breusch-Godfrey serial correlation LM tests and find no evidence of serial correlation. We use White's test to check for heteroscedasticity. If necessary, a heteroscedasticity-consistent covariance matrix is used to calculate standard errors. The model parameters are estimated using OLS techniques. Although OLS estimation in autoregressive models is known to give biased results, the very large number of time periods in our models validates the use of asymptotic results. More on this subject can be found in Hamilton (1994, chapter 5).

In all tables, the significance of the estimates is marked with superindices * and ** for significance at 5% and 1%-levels, respectively. In the last two rows of table 4.5, p shows the number of lags and $\bar{R}^2$ is the adjusted R^2 value. We do not report the estimates for the lagged interest rate coefficients in the tables for reasons of clarity. The estimation results of model (4.1) are shown in table 4.5.

The results are strikingly simple: for most countries we do not find any evidence of an electoral cycle. We observe that the coefficients for inflation and industrial production are not significant in most cases. The FEX_{it} coefficient is only significant for Germany and Italy, which means that interest rates are not significantly higher or lower under a fixed exchange rate regime for all other

Table 4.5: Country-specific tests

Country	E	FEX	$E*FEX$	IP	π	p	$\bar{R}^2$
Austria	-0.319^{**}	–	–	0.182	0.496	3	0.93
Canada	0.005	–	–	-0.023	0.370	2	0.92
Denmark	0.025	–	–	0.878	2.382^{*}	4	0.79
Finland	-0.004	0.003	0.014	-0.151	0.786^{**}	3	0.98
Germany	0.001	-0.046^{*}	-0.008	-0.428	2.310	3	0.85
Italy	0.016	0.022^{*}	-0.028	0.003	0.267	2	0.98
Japan	0.025^{**}	0.016	-0.028^{*}	0.552^{*}	0.142	4	0.99
Norway	-0.024	–	–	0.279	0.813	2	0.83
Spain	0.054	-0.037	-0.001	0.142	0.289	6	0.76
Sweden	0.012	–	–	0.080	0.628	3	0.84
US	-0.006	–	–	2.479^{**}	0.698^{*}	4	0.97

countries. The negative coefficient for Germany is probably due to the reunification, the positive coefficient for Italy shows that leaving the EMS after the turmoil in 1992 has led to lower interest rates.

The PBC coefficient is insignificant for most countries, which indicates that PBCs are not visible in the short-term interest rate in most OECD countries. There are, however, two exceptions: Austria and Japan.

Austria yields a negative, highly significant coefficient, which means that the interest rate decreases before elections. This would mean that Austria, despite its relatively independent central bank, experiences a PBC. This is rather puzzling, as Austria has closely followed the German interest rate policy since the mid 1970s. Only very small differences between the Austrian and the German interest rate exist, which we could assume to be too small to be used for systematic opportunistic monetary policy.[7] An explanation could be that the significant coefficient is a statistical coincidence rather than clear opportunistic evidence. This idea can easily be verified: if we find a PBC in Austria, this would mean that the *German* central bank creates it. That is, the German Bundesbank systematically misuses its monetary policy to influence Austrian election outcomes. This seems, at the very least, highly unlikely. Indeed, this speculation is not confirmed: as can be seen in table 4.6, the election coefficient for Austria is no longer significant if we use the German interest rate as the dependent variable.

Therefore, it seems that there *were* differences between the German and the Austrian interest rate, and that they were big enough to allow opportunistic

[7]For details on the Austrian exchange rate policy see, for example, Hochreiter and Winckler (1995).

Table 4.6: Results for Austria using the German interest rate as dependent variable

Country	E_{it}	IP_{it}	π_{it}
Austria	-0.011	0.068	0.160

policies. This could be explained by the fact that Austria is very corporatistic, which could result in a relatively close cooperation between the (on the surface relatively independent) central bank and the government. Following that reasoning, a PBC in Austria does not necessarily come as a surprise, but rather shows the weakness of CBI indicators. This has been formulated by Forder (1996): 'A central bank may be independent by statute, and it is nevertheless accepted – on all sides – that the government will have its wishes implemented.' The concept of corporatism is discussed extensively in chapter 5 of this thesis.

Japan also deserves a closer look, as both the election coefficient and the FEX-election interaction term are significant. Do we have a binding constraint here in the sense that the Bretton Woods system prevented electoral cycles in Japan's monetary policy? This would be the case if the sum of the two coefficients E_{it} and $E_{it} * FEX_{it}$ is not significantly different from zero. As mentioned earlier, this hypothesis can be tested using a Wald test, the outcome of which is shown in table 4.7.

Table 4.7: Wald-test for Japan

Country	$E + E * FEX$	χ^2-Wald	critical χ^2 (5%)
Japan	-0.003	0.051	0.103

As the results in table 4.7 show, the hypothesis that the sum of E_{it} and $E_{it} * FEX_{it}$ is zero is not rejected. In other words, the constraint is binding, and the Bretton Woods system has indeed been a restriction. This, however, does not mean that the Bank of Japan used monetary policy in an opportunistic sense.

First, it is noted that the sign of the election coefficient is positive: instead of lower (as the PBC model predicts), the short-term interest rate was higher before elections. Thus, the Bank of Japan used monetary brakes before elections instead of stimulating the economy. As has been documented in the literature, Japan is a special case, since its elections are endogenous. This means that the parliament can call elections when the ruling party experiences a favorable situation. There is a broad consensus that elections are more likely to be held when economic conditions are favorable for the incumbent (see Ito and Park, 1988), which is difficult to capture in a common PBC model, see also Cargill et al. (1997).

Second, the result is not robust. In fact, only for a pre-election period of 12 months (as we use in table 4.5), the election coefficient is significant. A pre-election dummy covering a period of 18 months is insignificant, as table 4.8 shows.

Table 4.8: Results for Japan using an 18 month pre-election dummy

Country	E_{it}	FEX_{it}	$E_{it} * FEX_{it}$	IP_{it}	π_{it}
Japan	0.008	0.016	-0.018	0.571^{**}	0.124

Moreover, if we use the plain nominal interest rate instead of its log, even the election coefficient using a 12-month pre-election period becomes insignificant. Further tests using different pre-election periods confirm that the result for Japan lacks robustness.

Finally, we take a look at the country-specific results for countries that have experienced a national constraint. For Belgium, France and the UK, and for Sweden only when the FS indicator is used, the score of the CBI variable changes over time. This is captured in our second country-specific regression model:

$$\begin{aligned} r_{it} &= \beta_{0i} + \beta_{1i}E_{it} + \beta_{2i}FEX_{it} + \beta_{3i}E_{it} * FEX_{it} + \beta_{4i}CBI_{it} \\ &+ \beta_{5i}E_{it} * CBI_{it} + \sum_{j=1}^{p} \beta_{j+5,i}r_{i,t-j} + \delta_{1i}IP_{it} + \delta_{2i}\pi_{it} \\ &+ \delta_{3i}OC_t + u_{it}. \end{aligned} \tag{4.3}$$

The dummy variable CBI_{it} has value $+1$ when the level of central bank independence is above-median, and 0 otherwise. We have added an interaction term for elections and CBI, as for elections and FEX before. For the periods in which the central banks have a low degree of independence, the model reduces to that in equation (4.1). The results are presented in table 4.9.

Table 4.9: Results in case of a national constraint

Country	E	FEX	$E * FEX$	CBI	$E * CBI$	IP	π
Belgium-both	-0.002	–	–	-0.002	0.036	0.182	0.529
France-CWN	< 0.001	0.013	0.003	-0.014	0.005	-0.023	0.467
France-FS	-0.001	0.032	-0.009	-0.030	0.021	-0.017	0.534^{*}
Sweden-FS	0.021	–	–	0.019	-0.036	0.079	0.745^{*}
UK-both	-0.043	0.011	0.056	-0.034	-0.017	0.959	-0.974^{*}

Again, we do not show the results for the lagged interest rate coefficients. The number of lags for Belgium and France is $p = 2$, for Sweden $p = 3$ and for the UK $p = 4$. The adjusted R^2 values are 0.90 for both models for Belgium, 0.97 for both models for France, 0.84 for Sweden and 0.87 for both models for the UK. None of the reported coefficients, apart from the one for inflation in some cases, is significant. This indicates that the central banks of Belgium, France, Sweden and the UK have not been engaged in creating PBCs. For the UK, similar findings have been reported by Clark et al. (1998, 2000).

Robustness checks We have performed a series of robustness checks. First, we have omitted the additional economic variables, and used the real interest rate (proxied by subtracting inflation from the nominal interest rate). This does not qualitatively change our results. Second, we have experimented with different lengths of the pre-election period. Economic theory gives no clear recommendation of how long we should expect the pre-election period to be. In the literature, periods of 12, 18 and 24 months are commonly used. We have found that our results in most cases do not depend on the length of the pre-election period. This does not hold for Austria and Japan. For Austria, the coefficient is highly significant for pre-election periods of 12 and 24 months, but not for 18 months. Also, the sign is reversed. The coefficient for Japan is only significant for a pre-election period of 12 months.

4.4 Panel data estimation

By pooling the data, we can examine the effects of cross-national differences in the national and international constraint. Since our focus is on a specific set of 14 countries instead of drawing the countries randomly from a large population, a fixed effects model is the appropriate specification here, see Baltagi (1995, p.10). By analogy to the country-specific case, it would seem natural to use an autoregressive panel data model. Before we do so, however, we examine whether this is indeed the correct specification, using the exact test for dynamics in panel data models that is described in chapter 2. Also, we test for unit roots.

4.4.1 Testing for dynamics and unit roots

The simplest dynamic panel data model with fixed effects to describe the influence of elections on the short-term interest rate is the following:

$$r_{it} = \gamma r_{i,t-1} + \alpha_i + \beta E_{it} + u_{it}, \tag{4.4}$$

where $i = 1, \cdots, N$, $t = 1, \cdots, T$, and the variables are defined as in the previous section. Data on the short-term interest rate are available for all 14 countries from January 1975 until December 1997, so $T = 252$ and $N = 14$. To examine whether a dynamic model specification is appropriate here, we test the hypothesis $H_0 : \gamma = 0$ versus $H_1 : \gamma > 0$. As shown in chapter 2, to test for $\gamma = 0$ we use the fixed effects estimator

$$\hat{\gamma}_{FE} = \frac{r'_{-1} M_W r}{r'_{-1} M_W r_{-1}}, \tag{4.5}$$

corresponding to the augmented regression model

$$\begin{aligned} r &= \gamma r_{-1} + E\beta + Z\alpha + BE\beta^* + BZ\alpha^* + u \\ &= \gamma r_{-1} + W\theta + u. \end{aligned}$$

Here, $Z = \iota_T \otimes I_N$ and, using

$$B_0 = \begin{bmatrix} 0 & 0 \\ I_{T-1} & 0 \end{bmatrix},$$

$B = B_0 \otimes I_N$. Also, $W = (E, Z, BE, BZ)$, and θ contains all parameters except γ. Equation (4.5) can also be written as

$$\hat{\gamma}^*_{FE} = \frac{u'B'M_W u}{u'B'M_W Bu}. \tag{4.6}$$

When u is a normally distributed random variable, the test statistic in (4.6) is a ratio of quadratic forms in standard normal variables, and its distribution can be assessed with great precision by simulation. Using 1000 replications, the distribution of $\hat{\gamma}^*_{FE}$ is shown in Figure 4.1.

From the figure, we find that the critical value of a one-sided test $H_0 : \gamma = 0$ versus $H_1 : \gamma > 0$ is 0.023 at the 95%-level. The value of the test statistic can be computed using (4.5) and is 0.968. The null hypothesis is therefore rejected, and a dynamic model specification is taken to be the appropriate one.

Next, we test for a unit root. The estimator for γ, under the unit root hypothesis $H_0 : \gamma = 1$, is the fixed effects estimator

$$\hat{\gamma}^{**}_{FE} = \frac{r'_{-1} M_W r}{r'_{-1} M_W r_{-1}}, \tag{4.7}$$

of the augmented regression model

$$\begin{aligned} r &= \gamma r_{-1} + X\beta + Z\alpha + B\Gamma X\beta^* + B\Gamma Z\alpha^* + u \\ &= \gamma r_{-1} + W\theta + u, \end{aligned}$$

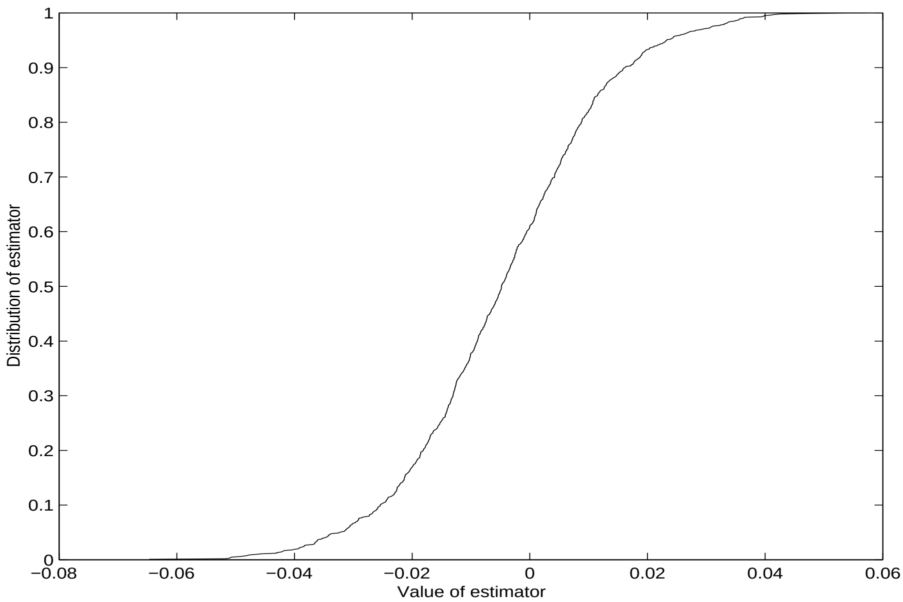


Figure 4.1: Distribution of $\hat{\gamma}^{*}_{FE}$ (y-axis) against the value of $\hat{\gamma}^{*}_{FE}$ (x-axis) under $H_0 : \gamma = 0$

where

$$\Gamma = \begin{bmatrix} 1 & & 0 \\ \vdots & \ddots & \\ 1 & \cdots & 1 \end{bmatrix} \otimes I_N.$$

The fixed effects estimator can also be written as

$$\hat{\gamma}^{**}_{FE} = 1 + \frac{u'\Gamma' B' M_W u}{u'\Gamma' B' M_W B \Gamma u}, \tag{4.8}$$

see chapter 2. Again using 1000 replications, the distribution of $\hat{\gamma}^{**}_{FE}$ is shown in Figure 4.2.

The critical value of a one-sided test $H_0 : \gamma = 1$ versus $H_1 : \gamma < 1$ is 0.978 at the 95%-level. The value of the test statistic can be computed using (4.7) and is 0.942. The null hypothesis of a unit root is therefore rejected.

4.4.2 Estimation results

Based on the test results from the exact tests, we use an autoregressive panel data model to examine the effects of cross-national differences in the national

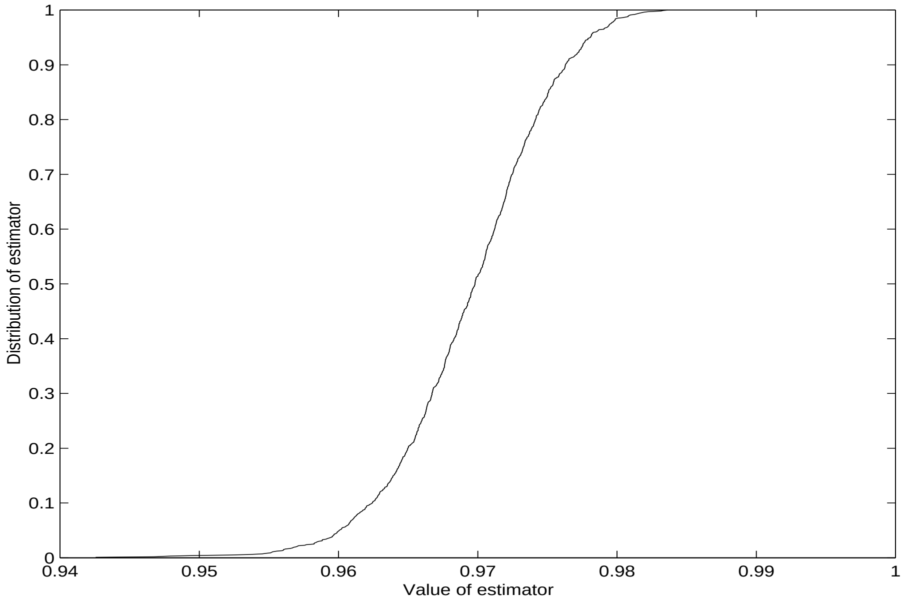


Figure 4.2: Distribution of $\hat{\gamma}^{**}_{FE}$ (y-axis) against the value of $\hat{\gamma}^{**}_{FE}$ (x-axis) under $H_0 : \gamma = 1$

and international constraint. We assume that the short-term interest rate does not exhibit a unit root. The constraint on PBC behaviour for the pooled sample can be modelled as follows:

$$\begin{aligned} r_{it} &= \beta_{1i} + \beta_2 E_{it} + \beta_3 FEX_{it} + \beta_4 CBI_{it} + \beta_5 E_{it} * FEX_{it} \\ &+ \beta_7 E_{it} * CBI_{it} + \beta_8 FEX_{it} * CBI_{it} + \beta_9 E_{it} * CBI_{it} * FEX_{it} \\ &+ \sum_{j=1}^{p} \beta_{j+9} r_{i,t-j} + \delta_1 IP_{it} + \delta_2 \pi_{it} + \delta_3 OC_t + u_{it}. \end{aligned} \tag{4.9}$$

In this model, we have included all possible combinations of central bank independence, fixed exchange rates and elections. The interpretation of the combination $E_{it} * FEX_{it}$ and $E_{it} * CBI_{it}$ has been given above. The interaction terms $FEX_{it} * CBI_{it}$ and $E_{it} * FEX_{it} * CBI_{it}$ are measures that tell us whether having an independent central bank influences a country's interest rate under a regime of fixed exchange rates. If such an effect is present, we would expect the coefficients to be negative.

Judson and Owen (1999) and Bun (2001) show that, for an unbalanced panel with a very large time dimension, the LSDV estimator is recommended. As in the country-specific case, White heteroscedasticity-consistent standard errors are computed. The lag length used in the panel data models is $p = 4$. Estimates of the lagged interest rate coefficients are not reported in the table. The estimation results of (4.9), for both measures of CBI, are reported in table 4.10.

Table 4.10: Results for the panel regression

	CWN indicator		FS indicator	
Variable	Coefficient	t-value	Coefficient	t-value
E	0.021	1.46	0.019	0.88
FEX	0.020**	2.99	0.007	0.81
CBI	0.032	1.66	−0.003	−0.25
$E * FEX$	−0.023	−1.40	−0.019	−0.81
$E * CBI$	−0.024	−1.48	−0.014	−0.64
$FEX * CBI$	−0.039*	−2.07	0.005	0.33
$E * FEX * CBI$	0.025	1.21	0.012	0.49
OC	0.007	0.72	0.007	0.73
IP	0.113*	2.00	0.114*	2.01
π	0.404**	2.59	0.376*	2.45

The adjusted R^2 value is 0.93 for both models. As in our previous regressions, the estimated coefficient for elections is insignificant. This confirms our findings of the country-specific model: we do not find any pattern compatible with the PBC model. The influence of the exchange rate systems is quite strong when the CWN indicator is used to measure CBI: the coefficient of FEX_{it} is significant at the 1% level. Its positive sign shows that countries that abandon their monetary policy autonomy by pegging their interest rate or participating in a fixed exchange rate regime have a tendency for higher interest rates. However, the significance of the coefficient disappears when the FS indicator is used. The interaction term $FEX_{it} * CBI_{it}$ is significant at the 5% level, which means that higher central bank independence lowers a country's interest rate under a fixed exchange rate regime. Again, this effect disappears when we use the FS indicator.

To summarize, the panel regressions confirm our previous findings. The main result is that the E_{it} dummy is insignificant, which implies that elections do not influence the short-term interest rate. Neither the degree of central bank independence nor participation in fixed exchange rates influences our findings. Given these results, we have to reject PBC theory as far as central banks are concerned. We do not find evidence that central banks have been actively engaged in short-sighted behavior before elections. Indeed, we have to conclude that, if cycles occur in monetary aggregates (as has been reported in previous studies), they are, in all likelihood, fiscally-induced. For this, central banks should not be held responsible. Moreover, we conclude that there is no evidence that national and international constraints effectively reduce the scope for electoral manipulations. Our estimation results suggest that central banks conduct monetary policy quite independent from upcoming elections. Overall, we conclude that central banks do not use their instruments as suggested by the PBC theory.

4.5 Conclusions

Making central banks independent is often justified by the concern that monetary policy might be opportunistically manipulated. According to Beck (1991), “the major argument for Fed independence is that monetary policy is politically neutral and technical. If the Fed is caught with its hand in the electoral cookie jar, then it can hardly claim to be apolitical in any sense of that word”. Using short-term interest rates, we have tested whether central banks in OECD countries indeed create PBCs and whether the degree of CBI is crucial to prevent that.

We have established evidence from two outcomes. First, we have conducted country-specific tests, based on the short-term interest rate for 14 OECD countries. The results of these tests show hardly any support for the PBC hypothesis. We conclude that central banks do not manipulate interest rates before elections. This suggests that either governments cannot impose their will on central banks, or that central banks have effectively resisted the wishes of the government. Our results do not suggest that the degree of statutory central bank independence matters in this respect. Alternatively, our results could be due to the fact that the short-term interest rate is not as tightly controlled by central banks as we have assumed. If financial markets have a strong impact on the short-term interest rate, under rational expectations manipulations are useless. This, however, would have the following implication. If, as the theory suggests, central banks use interest rates to manipulate monetary growth (and ultimately the inflation rate), and if their actions before elections have no effect on the short-term interest rate, then PBCs – if they exist in macroeconomic data, such as GNP growth or unemployment – cannot be due to central bank action, as their actions have no effect.

The second source of evidence stems from our panel data regressions. Their results provide more or less the same picture, namely that there is no evidence that central banks actively create PBCs. Overall, the implications are clear. If PBCs in macroeconomic variables such as unemployment show up, then the central banks should not be blamed.

Further research has to be done to reveal why cyclical behavior can be found in monetary aggregates. Still, if one believes that central banks have the means to control interest rates, then one has to reject the idea that central banks help governments to win elections. Electoral cycles that might appear in monetary aggregates could largely be demand-induced (perhaps due to fiscal behavior), but are not due to central bank action.

Appendix 4.A Data sources

Where available, we use monthly data from IMF's IFS statistics on industrial production, inflation and the short-term interest rate (line 60B). In addition, data have been provided directly by the central banks of Denmark, Sweden and the UK. Data for Germany have also been obtained from the CD-ROM "Deutsche Bundesbank: 50 Jahre Deutsche Mark". Data for the United States have been obtained from FRED (http://www.stls.frb.org/fred/).

Growth rates are computed as the change in the log of the raw series and have been detrended if necessary. All computed series are stationary. The election dummy is +1 eleven months before the election and during the election month, and 0 otherwise. The dummy for central bank independence, measured by the CWN indicator or the FS indicator, is +1 if the level of central bank independence is above-median, and 0 otherwise. The dummy for monetary policy autonomy (FEX) is +1 during participation in a fixed exchange rate regime (monetary policy autonomy is absent), and 0 if a country is not constrained.

The sample period differs for each country due to data availability. The period for which short-term interest rates are available for each country are shown in table 4.11, along with the corresponding sample size T. Due to lack of democratic elections until 1977, the sample period for Spain further reduces to 1977:01–1997:12.

Table 4.11: Availability of short-term interest rates

Country	Period	T
Austria	1967 : 01–1997 : 12	369
Belgium	1960 : 01–1997 : 12	444
Canada	1975 : 01–1997 : 12	274
Denmark	1972 : 01–1997 : 12	287
Finland	1972 : 10–1997 : 12	296
France	1964 : 01–1997 : 12	396
Germany	1960 : 01–1997 : 12	444
Italy	1971 : 01–1997 : 12	322
Japan	1960 : 01–1997 : 12	444
Norway	1971 : 08–1997 : 12	315
Spain	1974 : 01–1997 : 12	252
Sweden	1965 : 12–1997 : 12	382
UK	1960 : 01–1997 : 12	444
US	1960 : 01–1997 : 12	444

Chapter 5

How to use indicators of 'corporatism' in empirical applications*

5.1 Introduction

A large literature examines the relationship between the institutional settings of the labour market and the economic performance of the country concerned. In this context, the concept of corporatism plays an important role. Often corporatism refers here to the degree of centralization of wage bargaining. There is a debate whether a high degree of corporatism monotonically improves economic performance or whether the relation has an inverted U-shape. In line with the first view, Bruno and Sachs (1985) report a monotonic relation between their index of corporatism and the so-called misery index, i.e. the sum of the rise in inflation and the slowdown in real GDP growth. Calmfors and Driffill (1988), however, argue that employment is lowest in national single-industry bargaining, the argument being that both highly centralized and highly decentralized bargaining systems force unions to internalize the external effects of wage increases. Likewise, Alesina and Perotti (1997) find that the degree of shifting of labour taxation is a hump-shaped function of the degree of centralization of labour markets, peaking in countries with an intermediate degree of centralization.

Many subsequent studies have criticized this conclusion. For instance, Layard et al. (1991) argue that union coverage is a much more important factor af-

*This chapter is a revised version of Leertouwer and De Haan (2002).

fecting unemployment than centralization of wage bargaining. It has even been posed that corporatism may not be beneficial for employment growth. Proponents of what may be called the eurosclerosis view (Padovani and Gali, 2002) argue that corporatism distorts the allocative efficiency of markets, thereby reducing competitiveness and employment. The results of Woldendorp (1997) are in line with this hypothesis.[1]

From the cases discussed above it is already clear that, even though the concept of corporatism plays a crucial role in various studies, the concept is not very well-defined: different authors use diverging interpretations of corporatism. For instance, whereas Calmfors and Driffil argue that the degree of centralization of wage bargaining is crucial, Tarantelli (1986) posits that centralization is not the only relevant feature of industrial relations: the degree of consensus is also important. Here, consensus refers to a situation where the relevant groups broadly agree on the distribution of income. As Padovani and Gali conclude, corporatism clearly is a multi-faceted concept. Siaroff (1999), for instance, distinguishes six characteristics of corporatism, including unionization and the degree of centralization of wage bargaining. Unfortunately, most authors generally refer to a specific aspect or definition of corporatism, which reduces the comparability of the various studies on the subject.

In a recent strand of literature, the standard inflationary bias model of monetary policy is combined with the literature on labour market institutions, see Berger et al. (2001). One of the first contributions in this line of research is from Hall and Franzese (1998). They argue that the character of wage bargaining conditions the impact of central bank independence. Greater independence can reduce inflation without major employment effects where bargaining is coordinated, but it brings higher levels of unemployment where bargaining is uncoordinated.

In all of the studies referred to so far, some indicator for corporatism is used. There exists, however, a long list of empirical proxies for corporatism and it is often not obvious that a certain indicator has to be preferred. Several authors have therefore tried to use the information available in the existing indicators of corporatism to come up with some kind of ‘summary index’. For instance, Siaroff tries to disentangle the concept of corporatism using 23 indicators and suggests to replace it with an alternative indicator. However, there are shortcomings in his analysis. First, Siaroff does not take into account the fact that several indicators are based on each other. Second, he transforms the values of the different indicators of corporatism to a scale of 1 to 5. In doing so, a lot of

[1] See Teulings and Hartog (1998) for an extensive discussion of the literature on corporatism.

numerical information is thrown away.

In this chapter, we examine 29 indicators of labour market institutions that have been suggested in the literature, mainly referring to the 1970s and early 1980s. We examine the similarities and differences between these indicators, and try to extract common aspects using a latent variables approach. We consider the various indicators as imperfect measures of ‘corporatism’ and set out to find values for the parameters expressing the relation between the latent variable and the indicators. For this purpose, we apply factor analysis. This way, a set of variables is combined into a single variable that best reflects the original data, using all information that is available in the indicators. Our sample consists of 18 OECD countries.[2] Since the concept of corporatism is not uniformly defined by the different authors, it turns out to be impossible to capture it using one factor only. Instead, we find two factors that can be identified as the degree of coordination between employers and trade unions, and the organizational power of labour. Accordingly, we construct two new indicators of the institutional settings of the labour market of the countries in our sample. Finally, for illustrative purposes, we apply the resulting measures in the models of Hall and Franzese (1998) for inflation and unemployment. In contrast to various previous studies, we find that the organizational power of trade unions does not affect inflation and unemployment. We also do not find support for the hypothesis of Hall and Franzese that interaction effects between the levels of CBI and coordination play an important role, in particular with respect to unemployment.

The remainder of this chapter is organized as follows. Section 5.2 reviews the indicators of ‘corporatism’, while section 5.3 outlines our methodology. Section 5.4 presents the results of the factor analysis, which are used in section 5.5 in an empirical model for inflation and unemployment. The final section offers some concluding comments.

5.2 Indicators of ‘corporatism’

Many researchers have constructed indicators of the institutional setting of the labour markets in various countries, in order to test if and how labour market institutions matter for a country’s economic performance. Instead of creating another new indicator, we focus on the information contained in the existing ones. The indicators we use in this chapter are listed in table 5.5 in appendix 5.A,

[2]Greece, Iceland, Luxembourg, Spain and Portugal have been excluded due to lack of available data.

along with the studies in which they have been published, in chronological order of publication. Their numerical values can be found in table 5.6. Table 5.7 gives the correlations between the indicators, as well as the number of countries, N, for which they are available. Because of missing data, the correlations have been computed using the EM algorithm. It follows that many of the indicators are highly correlated. At the same time, it is also clear that some correlations are rather low. In other words, the various indicators do not necessarily refer to the same concept. Alternatively, errors in measurement can be responsible for these low correlations. After all, subjective judgement plays an important role in deciding upon indicator scores for the various countries. Since the latent variables approach is very well suited to deal with problems of this type, we use it in order to extract the common aspects in the indicators.

5.3 Latent variables approach

The main idea of this chapter is to consider the different indicators described in the previous section as imperfect measures of the unobservable concept of corporatism. It is possible that the unobservable concept cannot be captured completely in one factor but more factors are extracted, corresponding to different aspects of the phenomenon. Then, the indicators are assumed to be generated by the multiple factor analysis model that has been described in chapter 2:

$$x_{ni} = \tau_i + \lambda_i'\xi_n + \delta_{ni}. \tag{5.1}$$

If the number of factors is k, x_{ni} denotes indicator i for country n and ξ_n is a $k-$vector containing the aspects of the unobservable concept (the factors) that the indicators are supposed to measure, for country n. The parameter τ_i captures the mean of indicator i, while λ_i is a $k-$vector of parameters (the factor loadings) that capture both the scale of indicator i and the strength of its relation to the factors, and δ_{ni} is a random measurement error. Further, δ_{ni} and δ_{nj} are assumed uncorrelated for $i \neq j$, and both are assumed uncorrelated with the factors ξ_n.

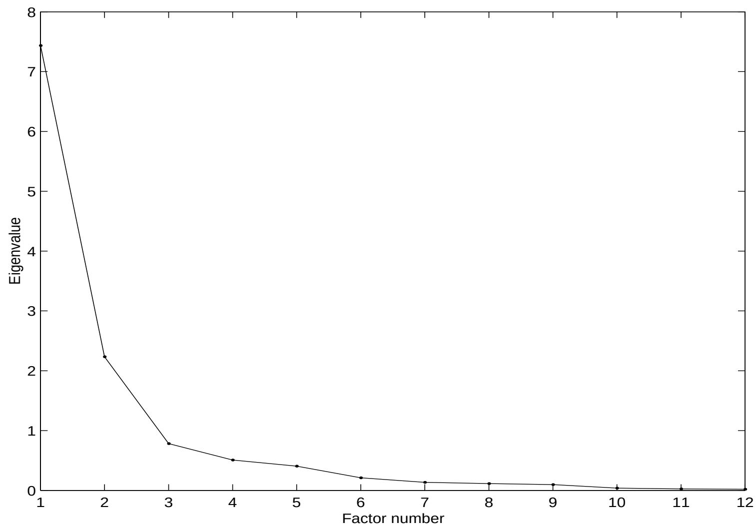
The problem with the 29 indicators of corporatism is that although different authors use different criteria to define and quantify corporatism, there are also a lot of similarities. This is apparent from the correlation matrix in table 5.7. Consequently, there may be a conflict with the model assumptions that the measurement errors are uncorrelated. There are two ways to solve this problem. The first is to augment the model with variables reflecting the covariance between the measurement errors. Since this would mean the inclusion of a lot of additional variables in our case, and our sample consists of only 18 countries, it would seriously reduce the number of degrees of freedom. Therefore, we have applied a second solution: to choose from clusters of correlated indicators one as the preferred indicator and omit the others from the analysis. As selection criterion we have mainly focused on the number of countries for which the indicators are available, as well as the fact that the selected indicators should more or less refer to the same time period. A detailed description of the selection procedure can be found in appendix 5.B.

Using the above selection procedure, we end up with a subset of 12 indicators: BLY, CAM, KEM, PAL, BRS, TAR, LNR, LNJ and the four OECD-indicators ODE, OBA, OCE and OCO. Unfortunately, not all selected indicators are available for Ireland and New Zealand. Therefore, we exclude these countries in our empirical applications, leaving a sample of 16 OECD countries for which we perform the factor analysis.

5.4 Results of the factor analysis

In this section we present the results of the factor analysis. In order to see how many factors are needed, we examine the scree plot in figure 5.1, which plots the number of factors against their eigenvalues.

Figure 5.1: Scree plot of the factor analysis



Using the Kaiser rule and examining the kink in figure 5.1, we conclude that we need to perform a factor analysis with two factors in order to capture the information contained in the indicators. Table 5.1 gives the corresponding standardized solution in the form of a matrix of factor loadings, the elements of which can be interpreted as correlations of the indicators with the factors. The

Table 5.1: Unrotated factor loadings

	Factor	
Indicator	1	2
BLY	.833	.213
CAM	.802	.470
KEM	.957	−.096
PAL	.779	−.554
BRS	.894	−.174
TAR	.747	−.344
LNR	.764	−.282
LNJ	.909	.122
ODE	.570	.749
OBA	.418	.442
OCE	.618	.505
OCO	.813	−.189

χ^2-statistic, which compares the proposed model to an unrestricted alternative, has a value of 48.0 with 43 degrees of freedom, which lies well below the 5% critical value of 59.3. According to this measure, the model fits very well. A graphical representation of the factor loadings is shown in figure 5.2.

However, it is not easy to give an interpretation to the two factors in table 5.1. One of the main problems is that the scores on the second factor are low for almost all indicators. In order to see whether the two factors can be polarized more, we rotate the solution in table 5.1 using varimax rotation. The results are shown in table 5.2 and figure 5.3.

Using the rotated factor matrix, we can try to interpret the two factors. The first factor correlates highly with the indicators of Keman, Paloheimo, Bruno and Sachs, Tarantelli, Lehner and the OECD coordination indicator. Since all these indicators mainly reflect coordination issues, we label this factor *coordination*. The second factor correlates highly with the indicator of Cameron and the OECD density indicator, which are both focusing on power resources of labour. Therefore, we label this factor *organizational power of labour*. Using the results of the factor analysis, we can combine the indicators in order to get constructs that have maximum correlation with the two aspects of corporatism as reflected by the factors. The resulting factor scores are given in table 5.3 for each country in the sample. It is shown in the table that, for instance, coordination is high in

Figure 5.2: Unrotated factor loadings

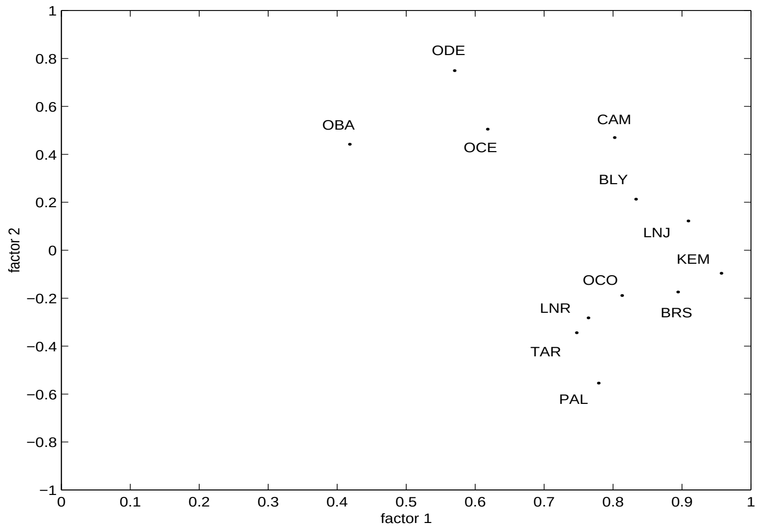


Table 5.2: Rotated factor loadings

	Factor	
Indicator	1	2
BLY	.569	.645
CAM	.399	.840
KEM	.845	.459
PAL	.956	−.020
BRS	.837	.359
TAR	.811	.136
LNR	.790	.196
LNJ	.683	.612
ODE	.050	.940
OBA	.097	.600
OCE	.227	.765
OCO	.779	.301

Austria and Switzerland, but low in the United Kingdom and the United States. Organizational power of labour is high in the Scandinavian countries, but low in France and Japan. A graphical representation is given in figure 5.4.

The pattern of the country scores in figure 5.4 has a very specific form: the centre of the figure is completely empty. This suggests that we may have a non-linear factor model here. To examine this, we have regressed the factor organizational power of labour on the squared values of the coordination factor.

Figure 5.3: Rotated factor loadings

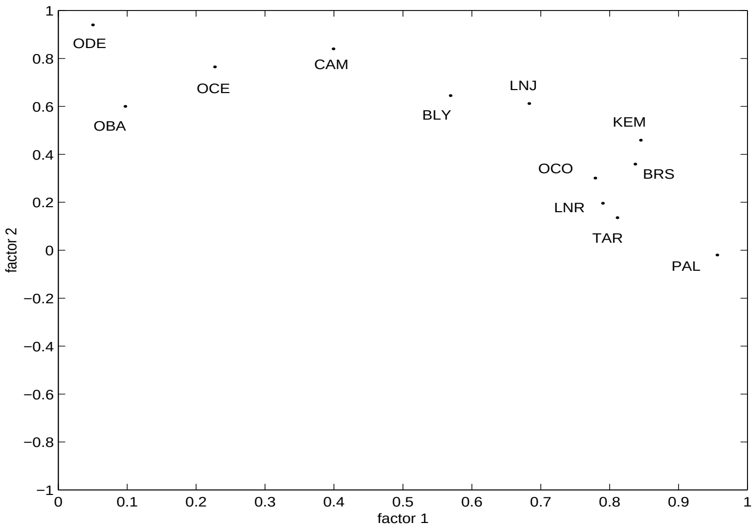
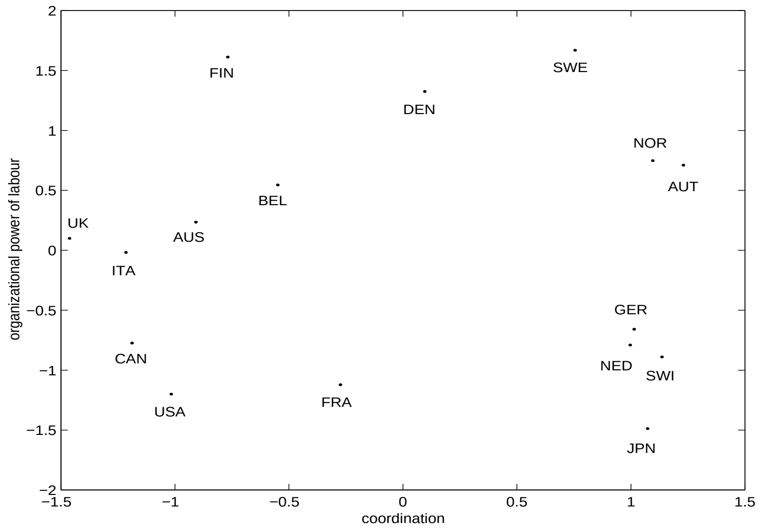


Table 5.3: Factor scores

Country	Coordination	Organizational power of labour
Australia	−0.909	0.235
Austria	1.229	0.710
Belgium	−0.550	0.545
Canada	−1.189	−0.774
Denmark	0.095	1.325
Finland	−0.769	1.612
France	−0.275	−1.122
Germany	1.013	−0.659
Italy	−1.216	−0.018
Japan	1.072	−1.488
Netherlands	0.996	−0.790
Norway	1.095	0.748
Sweden	0.754	1.669
Switzerland	1.135	−0.890
United Kingdom	−1.463	0.099
United States	−1.017	−1.200

This yields no significant coefficient, however, and we maintain our assumption that the model is linear.

Figure 5.4: Values of coordination and organizational power of labour



5.5 Using ‘corporatism’ in an empirical application

To show how the results of the approach described in the previous section can be applied in practice, we present a simple empirical model. As seen in table 5.8 in appendix 5.A, most studies analyze the relation between aspects of corporatism and inflation and / or unemployment. Of these studies, we use the influential paper by Hall and Franzese (1998) as a benchmark. We use their model as well as methodology mainly for illustrative purposes.

Instead of using Hall and Franzese’s regressors coordination of wage bargaining and union density, we examine whether the two aspects of corporatism extracted by the factor analysis, coordination and organizational power of labour, have a significant impact on inflation and unemployment. We start with a cross-sectional model for the period 1970–1985, since most indicators of corporatism have been constructed in the 1970s and early 1980s. We use rates of inflation (computed on the basis of the GDP-deflator) and internationally comparable unemployment rates, provided by Hall and Franzese.

The following control variables are included in the models for inflation and unemployment:

- Trade openness, defined as the sum of imports and exports divided by real GDP
- Left cabinet participation, defined as the percentage of cabinet seats held

by left wing parties

- The degree of central bank independence

Hall and Franzese expect interaction effects between the level of central bank independence and the level of wage coordination, especially with respect to the level of unemployment. Therefore, the cross product of coordination and CBI is included in the model. In the model for unemployment, the impact of international economic conditions on unemployment is analyzed by including the terms of trade, defined as export prices divided by import prices, as a control variable. Also, the cross product of the terms of trade and trade openness is included. Hall and Franzese include this cross product since domestic unemployment is expected to benefit from positive terms of trade shocks to the degree the economy is open to foreign trade. Terms of trade is not included in the inflation models since terms of trade movements are essentially defined as movements in domestic inflation relative to foreign inflation. Summarizing, the model specification is the following one:

$$\begin{aligned} Y_i &= \beta_{0i} + \beta_{1i} COORD_i + \beta_{2i} ORGPL_i + \beta_{3i} CBI_i \\ &+ \beta_{4i} COORD_i * CBI_i + \beta_{5i}' Z_i + u_i, \end{aligned} \tag{5.2}$$

where Y is inflation or unemployment, $COORD$ and $ORGPL$ are the degrees of coordination and organizational power of labour that have been constructed in the previous section, CBI is the degree of central bank independence and Z contains the additional control variables. Finally, u are the disturbances which are assumed to be normally distributed.

In addition to a cross-sectional analysis, Hall and Franzese examine two dynamic specifications: a pooled model that aggregates for different decades, and an annual model. We have also replicated these specifications, using a three period model instead of the decade model, for the periods 1970–1973, 1974–1979 and 1980–1985. We have chosen these periods instead of dividing the sample into the periods 1970–1974, 1975–1979 and 1980–1985 since values of most economic variables change dramatically from 1973 to 1974 due to the 1973 oil crisis. Following Hall and Franzese, an AR(1) process is incorporated in the residuals of the three-period model for inflation to allow for the temporal dependence in the observations. In the corresponding unemployment model, adding an autoregressive term does not improve the specification. Using the variables as defined in equation (5.2), the model specification is

$$
\begin{aligned}
Y_{it} &= \beta_{0i} + \beta_{1i}COORD_i + \beta_{2i}ORGPL_i + \beta_{3i}CBI_{it} \\
&+ \beta_{4i}COORD_i * CBI_{it} + \beta_{5i}' Z_{it} + u_{it},
\end{aligned}
$$

where $u_{it} = \rho u_{i,t-1} + \eta_{it}$ in the model for inflation. Note that the aspects of corporatism do not change over time.

The annual models are estimated as *pseudo-error correction* models with Beck-Katz panel-corrected covariance (PCSE) matrices. In short, this comes down to regressing the change in the dependent variable on the lagged change and lagged level of the dependent variable, changes in the independent variables and the first lag of the independent variables. See Beck (1991) and Beck and Katz (1995, 1996) for details. Thus, the model specification is

$$
\begin{aligned}
\Delta Y_{it} &= \beta_{0i} + \beta_{1i}COORD_i + \beta_{2i}ORGPL_i + \beta_{3i}CBI_{i,t-1} \\
&+ \beta_{4i}COORD_i * CBI_{i,t-1} + \beta_{5i}' \Delta Z_{it} + \beta_{6i}' Z_{i,t-1} \\
&+ \beta_{7i} \Delta Y_{i,t-1} + \beta_{8i} Y_{i,t-1} + u_{it}.
\end{aligned}
$$

Finally, Hall and Franzese also use logged real GDP per capita as a control variable in both static and dynamic models, on the premise that less-developed nations may be more tempted to rely on seignorage for revenue and more susceptible to high levels of unemployment. We do not think this is the appropriate control variable for a model examining the impact of aspects of corporatism on inflation and unemployment. Instead, we include the output gap in our dynamic specifications, which is defined as real GDP growth minus Hodrick-Prescott filtered GDP growth. This variable is supposed to control for the effects of the business cycle.

Data on central bank independence are taken from chapter 3, where an indicator for CBI (the FS indicator) is constructed using factor analysis. In this field, essentially the same measurement error problems arise as in the case of corporatism. The FS indicator combines the information contained in already existing indicators of CBI in an optimal way and does not rely on subjective judgement. Therefore, it is prefered to the measure that Hall and Franzese use, which is simply an unweighted average of existing indicators. Moreover, the FS indicator varies over time, something that Hall and Franzese do not take into account. Data on the other control variables have been obtained from the IMF’s International Financial Statistics, and from Hall and Franzese. The models are estimated using least squares and White’s heteroscedasticity-consistent covariance matrix, and the results are shown in table 5.4.

Table 5.4: Regression results for inflation and unemployment

Regressor	Inflation			Unemployment		
	average	period	annual	average	period	annual
Output gap	–	2.19	−0.19	–	−1.97**	−0.25**
		(1.90)	(−1.75)		(−3.16)	(−8.15)
Trade openness	−0.03	−0.03*	0.04	0.02**	0.18	0.06*
	(−1.70)	(−2.47)	(0.76)	(3.43)	(1.99)	(2.07)
Terms of trade	–	–	–	–	4.40	1.95
					(1.21)	(1.41)
Terms of trade	–	–	–	–	−0.15	−0.08**
× Openness					(−1.70)	(−2.76)
Left cabinet %	1.49	1.15	0.54	1.94	0.17	0.22
	(0.66)	(0.67)	(0.64)	(1.03)	(0.11)	(1.10)
CBI	−2.16**	−2.25**	−0.88*	−0.13	−0.10	−0.04
	(−4.51)	(−5.24)	(−2.07)	(−0.34)	(−0.26)	(−0.66)
Coordination	−0.99*	−1.30*	−0.51	−2.19**	−1.87**	−0.16**
	(−2.30)	(−2.31)	(−1.67)	(−5.00)	(−3.63)	(−2.90)
Org. power	0.33	0.53	0.11	−1.01	−0.82	−0.10*
	(0.97)	(1.16)	(0.44)	(−1.87)	(−1.69)	(−2.18)
Coordination	0.96	1.07**	0.43	−0.54	−0.55	−0.02
× CBI	(2.20)	(3.18)	(1.57)	(−1.55)	(−1.91)	(−0.40)
# observations	16	48	224	16	32	224
Adjusted R^2	0.79	0.65	0.30	0.72	0.59	0.59
DW-statistic	–	2.29	1.98	–	1.71	1.97

Since the models include interaction terms, interpretation of the t-statistics for the coefficients of the variables that are also included as cross-products should be done with caution. The estimated effect of a unit increase in coordination, for instance, is not given simply by its estimated coefficient: the coefficient of its cross-product with CBI times the level of CBI has to be added to this. In these cases, the significance of the estimated coefficient cannot be derived from the value of its t-statistic in a straightforward manner. However, the t-value gives an *indication* of the significance of a coefficient. For illustrative purposes, we report the conventional significance levels in table 5.4 using superindices * and ** for significance at a 5% and 1% level, respectively.

It follows from table 5.4 that organizational power of labour does not have a significant impact on inflation. The t-statistics corresponding to the coefficients of CBI and coordination, on the other hand, indicate that these variables influence inflation rates. Also, the interaction term of coordination and CBI is shown to have a significant impact on inflation. However, in the model for unemployment we do not find any significant interaction effects between the level of CBI and our variable for labour market coordination. This result contrasts with the findings of Hall and Franzese. The effect of CBI on unemployment does not seem very significant, either. Furthermore, the results show that coordination generally lowers unemployment, while organizational power of labour only has a significant impact in the annual unemployment model. Our general conclusions do not change when we add squared terms for our indicators.

5.6 Conclusions

In this chapter, two new indicators of the institutional settings of the labour market are constructed using the information provided by indicators of corporatism that have been suggested in the literature. By following a latent variables approach, two indicators that we label *coordination* and *organizational power of labour* are obtained. For illustrative purposes, these measures are applied in the models of Hall and Franzese for inflation and unemployment. In contrast to their results, we find no evidence that interaction effects between the level of central bank independence and coordination play an important role with respect to unemployment. Also, the impact of organizational power of labour on inflation and unemployment, which has been reported in various previous studies, is not found in our regressions.

5.A Indicators of labour market institutions and their properties

Table 5.5 lists all indicators used in this study, their sources and a short description. In table 5.6 the numerical values of the indicators are shown for the different countries in our sample, and table 5.7 displays the correlations between the indicators. Table 5.8 lists whether and how the original sources have used the indicators in empirical applications, as well as the time period to which they refer. Finally, table 5.9 gives some specific properties of the countries in the sample.

5.B Selection of the indicators

Based on the information in table 5.7, our selection procedure is as follows.

- The indicators by Cameron and Schmitter are constructed in a very similar way. We prefer Cameron’s indicator since it is available for a langer number of countries.
- The indicators by Lehmbruch and Lehner are based on the one by Czada, hence the high correlation between the three. Since the scaling of Lehner’s indicator is the most detailed, and since the indicator is available for all 18 countries, we prefer this one.
- The indicator by Schott is largely based on the ones by Schmitter and Lehmbruch and is therefore excluded.
- The indicator of Lijphart and Crepaz is an average of standardized values of a number of other indicators, so it does not provide additional information.
- The indicators by Wiarda and Schmidt (1986) are largely based on the one by Lehmbruch and are therefore excluded.
- The indicator by Hall and Franzese is based on Soskice, while missing values have been extrapolated using Layard, Nickell and Jackman and Crouch.
- The indicators by Braun, Soskice, Compston and Wiarda are available for only a small part of our set of countries.
- The indicator by Bruno and Sachs is largely based on the one by Crouch. Bruno and Sachs use a more detailed scaling, but their indicator is not

available for Ireland whereas Crouch's is. We have no clear preference in this case.

Based on these arguments, we have first decided to exclude the indicators of Schmitter, Braun, Czada, Lehmbruch, Schott, Schmidt (1986), Soskice, Lijphart and Crepaz, Compston, Wiarda and Hall and Franzese, leaving a set of 18 indicators.

However, table 5.7 shows that there are two more clusters of highly correlated variables:

- The indicator of Keman is highly correlated with the one by McCallum. In addition, the indicator of McCallum is highly correlated with Crouch's indicator. Moreover, all three are available for the 18 countries in our sample. If we select McCallum's indicator, we would lose both the indicators by Keman and Crouch, while a selection of the Keman indicator only results in the loss of the one by McCallum. Therefore, the Keman indicator is the preferred one.

- The indicators by Cameron, Garrett and Lange and Calmfors and Driffill are highly correlated. Garrett and Lange has the lowest data availability, but a choice between Cameron and Calmfors and Driffill is not straightforward. However, a closer inspection of the correlation table shows that the indicator by Calmfors and Driffill is, in general, correlated more strongly with the remaining indicators. Therefore, we select the indicator by Cameron.

With the 14 remaining indicators we perform a factor analysis, but we do not obtain a feasible solution. Hence, we exclude the indicator by Marks, since this is the one that has the largest amount of missing values. Now, we get a feasible solution if we apply factor analysis, but the model fit is not very good. Therefore, we exclude another indicator from the set. Of the different specifications we attempt, the one excluding Schmidt's 1982 indicator has the best fit. Thus, our final set of indicators consists of the following twelve: Blyth, Cameron, Keman, Paloheimo, Bruno and Sachs, Tarantelli, Lehner, Layard, Nickell and Jackman, OECD-density, OECD-bargaining, OECD-centralization and OECD-coordination.

Table 5.5: Indicators of labour market institutions

Name	Source	Focus
BLY	Blyth (1979)	Ranking of centralization of structure: extent to which trade union and employer organizations are federated or joined into strong central bodies at national level with substantial executive powers.
SMT	Schmitter (1981)	Ranking of social corporatism.
BRA	Braun (1983)[a]	Summary scale of neo-corporatism.
CZA	Czada (1983)	Classifications of neo-corporatism.
SD1	Schmidt (1983)	Ranking of corporatism.
CAM	Cameron (1984)	Ranking of power resources concerning organizational power of labour, defined as the sum of confederation power in collective bargaining and organizational unity of labour, divided by unionized labour force %.
KEM	Keman (1984)[a]	Ranking of neo-corporatism, using degree of social partnership and state intervention.
LHM	Lehmbruch (1984)	Cumulative scale of corporatism.
PAL	Paloheimo (1984)	Ranking of economic consensus.
SCT	Schott (1984)	Ranking of corporatism.
BRS	Bruno and Sachs (1985)	Ranking of corporatism using degree of union centralization, extent of shop-floor union power, employer coordination and presence of work councils.
CRO	Crouch (1985)	Dichotomous scale of corporatism.
GAR	Garrett and Lange (1986)	Ranking of corporatism using leftist participation in cabinet and labour organization.
MKS	Marks (1986)	Ranking of neo-corporatist incomes policy.
MCL	McCallum (1986)	Classification of economic consensus based on strike levels and corporatist institutions.
SD2	Schmidt (1986)	Classification of policy coordination.
TAR	Tarantelli (1986)	Ranking of neo-corporatism which reflects a key role of consensus.
CFD	Calmfors and Driffill (1988)	Index of centralization of wage bargaining.
LNR	Lehner (1988)	Summary-scale of public-private interaction.
SSK	Soskice (1990)	Ranking of economy-wide coordination.
LNJ	Layard, Nickell and Jackman (1991)	Scores of union and employer coordination in collective bargaining.
LYP	Lijphart and Crepaz (1991)	Combination of twelve neo-corporatist rankings.
CMP	Compston (1997)	Ranking of union influence over economic policy.
ODE	OECD (1997)	Trade union density, measured as % of workers belonging to trade unions.
OBA	OECD (1997)	Bargaining coverage, measured as % of workers covered by collective agreements.
OCE	OECD (1997)	Ranking of centralization of wage bargaining.
OCO	OECD (1997)	Ranking of coordination of wage bargaining.
WRD	Wiarda (1997)	Ranking of corporatism.
HFR	Hall and Franzese (1998)	Ranking of coordination of wage bargaining.

[a]The indicators of Braun (1983) and Keman (1984) are taken from the appendix of Keman et al. (1985).

Table 5.6: Values of the indicators

	AUS	AUT	BEL	CAN	DEN	FIN	FRA	GER	IRE	ITA	JPN	NED	NZL	NOR	SWE	SUI	UK	USA
BLY	2	3	2	1	3	3	1	2	1	1	2	1	3	3	3	3[a]	1	1
SMT	.	26	18	8	20.5	20.5	6	15	8	3	.	19	.	23.5	20.5	9	5	8
SD1	2	3	2	1	2	2	1	2	1	1	3	2	2	3	3	3	1	1
BRA	.	3	2	.	3	2	1	2	2	1	.	3	.	3	3	2	1	.
CZA	1	3	2	1	2	2	2	2	2	2	1	3	1	3	3	2	2	1
CAM	28	90	66	10.8	64.8	65.8	4.8	32	25.6	16.4	4.8	33.6	.	97.5	105	24	31.5	8.4
KEM	2	5	2	1	4	3	2	3	1	2	3	4	3	5	5	4	1	1
LHM	1	4	3	1	3	3	.	3	3	2	.	4	1	4	4	3	2	1
PAL	1	3	2	1	2	1	2	3	1	1	3	3	.	3	3	3	1	1
SCT	1	3	2	1	2	2	1	2	1	1	3	2	.	3	3	3	1	1
BRS	0	4	0.5	0	3	1.5	0	4	.	0.5	1.5	4	0.5	4	4	2	0	0
CRO	1	2	1	1	2	2	1	2	1	1	1	2	1	2	2	2	1	1
GAR	2	4	2	1	4	3	1	2	.	1	1	2	.	4	3	.	2	1
MKS	.	4	4	1	2	3	1	2	2	1	.	3	.	4	4	1	2	1
MCL	0	4	0.5	0	3	1.5	0	4	0	0.5	1.5	4	0.5	4	4	4	0	0
SD2	1	3	2	1	2	2	2	3	2	1	3	2	1	3	2	3	1	1
TAR[b]	7	12	6	6	9	7	4	11	2	1	10	7	5	9	9	10	2	6
CFD	4	6	4	2	4.75	4.75	3.25	4.5	.	3.25	3	4.25	4	5	5	3	3.25	2
LNR	3	4	3	1	3	3	1	3	3	2	5	4	.	4	4	5	2	1
SSK	.	5	.	.	.	.	1.5	3.5	.	2	5	3	.	4	4	4	0	0
LNJ	3	6	4	2	6	6	4	5	3	2	4	4	3	6	6	4	2	2
LYP	-1.02	1.6	0.26	-1.34	0.52	0.43	-0.73	0.48	-0.53	-0.85	0.05	1.01	-1.11	1.53	1.4	0.51	-0.86	-1.34
CMP	.	9.7	5.3	.	6.9	7.1	3	4.4	6.1	6.6	.	5.9	.	6.9	9.2	8	3.8	.
ODE	48	56	56	36	76	70	18	36	57	49	31	35	56	57	80	31	50	22
OBA	88	98	90	37	69	95	85	91	.	85	28	76	67	75	86	53	70	26
OCE	2.25	2.25	2.25	1	2.25	2.5	2	2	.	1.75	1	2	2	2	3	2	2	1
OCO	2.25	3	2	1	2.5	2.25	1.75	3	.	1.5	3	2	1.5	2.5	2.5	2.25	1.5	1
WRD	.	3	.	.	2	.	1	2	.	1	.	3	.	.	3	.	2	.
HFR	0.25	1	0.5	0	0.75	0.75	0.25	0.75	0	0.25	0.75	0.5	0.25	1	1	0.75	0	0

Note: a dot (.) denotes that for that country the value of the indicator is missing.

[a] Although Blyth does not include Switzerland explicitly in his ranking tables, it is classified as strongly corporatistic in his paper.

[b] The original scaling of the indicator of Tarantelli is opposite to the scaling of the other indicators. For ease of exposition, we have reversed its scaling here.

Table 5.7: Correlations between measures of corporatism * 100%

	BLY	SMT	SD1	BRA	CZA	CAM	KEM	LHM	PAL	SCT	BRS	CRO	GAR	MKS	MCL	SD2	TAR	CFD	LNR	SSK	LNJ	LYP	CMP	ODE	OBA	OCE	OCO	WRD	HFR
BLY	100																												
SMT	82	100																											
SD1	81	84	100																										
BRA	62	88	75	100																									
CZA	26	64	39	76	100																								
CAM	65	85	53	77	74	100																							
KEM	78	82	87	84	64	64	100																						
LHM	41	82	68	91	96	81	77	100																					
PAL	52	63	81	66	57	36	82	82	100																				
SCT	81	84	95	75	52	54	88	84	86	100																			
BRS	59	77	75	86	74	68	91	91	83	81	100																		
CRO	63	81	64	77	70	67	84	79	65	72	88	100																	
GAR	84	89	64	80	67	89	75	74	40	62	66	81	100																
MKS	56	79	66	69	76	95	61	81	48	66	66	50	79	100															
MCL	65	74	81	75	67	61	92	82	86	87	98	89	66	55	100														
SD2	48	66	72	51	51	31	67	82	85	83	75	64	47	46	74	100													
TAR	65	79	84	66	28	38	75	60	79	83	75	73	58	43	78	77	100												
CFD	65	86	54	78	72	88	75	78	43	52	74	71	93	84	65	42	46	100											
LNR	66	71	91	77	45	42	78	84	78	89	73	60	55	61	78	76	73	41	100										
SSK	84	82	94	79	34	35	81	79	84	94	68	56	50	63	72	84	81	42	86	100									
LNJ	75	90	73	75	63	71	84	80	66	78	79	82	82	69	77	75	72	83	59	74	100								
LYP	58	86	73	94	86	74	87	96	81	85	91	86	79	72	88	80	68	78	76	78	88	100							
CMP	80	65	75	61	47	53	73	51	34	75	55	63	76	38	59	39	55	56	71	90	61	69	100						
ODE	55	57	25	51	41	79	35	41	-5	26	41	36	82	67	25	04	05	75	19	35	48	41	51	100					
OBA	24	42	17	04	52	57	27	48	10	13	34	36	44	60	29	24	19	69	09	24	50	45	15	42	100				
OCE	62	69	40	50	49	77	48	50	13	32	37	50	76	74	34	19	27	77	28	32	67	53	53	71	72	100			
OCO	62	78	81	71	42	46	74	72	73	79	75	64	59	63	73	83	88	64	75	88	81	76	56	30	41	44	100		
WRD	52	77	85	84	89	87	80	94	76	85	79	78	71	98	79	45	62	75	95	70	58	82	60	44	27	67	56	100	
HFR	80	85	89	75	60	65	93	79	79	92	88	81	74	65	89	78	82	75	76	89	91	88	74	37	41	53	89	68	100
N	18	15	18	13	18	17	18	16	17	17	17	18	15	15	18	18	18	17	17	11	18	18	13	17	17	17	17	8	18

Table 5.8: Empirical application of the indicators

	Period	Empirical application of corporatism measure
BLY	1970s	No empirical application.
SMT	1958–1975	No empirical application.
SD1	1974–1980	Correlation between corporatism, unemployment and inflation.
BRA	1970s	No empirical application.
CZA	1970s	Correlation of aspects of corporatism with unemployment and income policy.
CAM	1965–1982	Correlation of organizational power of labour with inflation and unemployment.
KEM	1965–1982	Impact of corporatism on inflation and unemployment.
LHM	1950–1980	No empirical application.
PAL	1960–1981	Price inflation, wage inflation and economic growth for different levels of corporatism.
SCT	1970s	Ranking of countries by economic performance criteria and political arrangements.
BRS	1973–1979	Impact of corporatism on inflation.
CRO	mid–1970s	Qualitative analysis of impact of neocorporatism on inflation and unemployment.
GAR	1974–1982	Impact of level of domestic political structures index on changes in economic growth.
MKS	1950–1980	No empirical application.
MCL	1980–1984	Impact of corporatism on unemployment.
SD2	1960–1984	Relation between corporatism and Okun's misery index.
TAR	1968–1983	Impact of neocorporatism on the misery index.
CFD	1960–1979	Correlation of aspects of corporatism with inflation and unemployment.
LNR	1960–1979	Phillips curve analysis for different levels of corporatism.
SSK	1985–1989	Impact of economy-wide coordination on unemployment.
LNJ	1956–1988	Impact of different aspects of corporatism on unemployment.
LYP	1950–1980	Impact of consensus democracy and dominant tendency in government on corporatism.
CMP	1972–1993	Impact of union participation on unemployment.
ODE		
OBA	1978–1982	Correlation of aspects of corporatism with measures of economic performance, such as
OCE		unemployment, price inflation and wage inflation. Also regression results.
OCO		
WRD	1950–1995	No empirical application.
HFR	1955–1990	Impact of level of wage-bargaining coordination on inflation and unemployment.

Table 5.9: Specific properties of the countries

Country	Specific properties
Australia	Weak economic consensus.
Austria	Very powerful unions, centralized coordinating role. Medium strong employer organizations.
Belgium	Medium economic consensus, highly centralized collective bargaining.
Canada	Weak economic consensus.
Denmark	Medium economic consensus.
Finland	Weak economic consensus, highly centralized collective bargaining.
France	Tacit government coordination, some sectoral corporatism. Essentially pluralist.
Germany	Medium strong union coordination, strong employer organizations.
Italy	Informal employer coordination. Some help from union confederations.
Japan	Very powerful tacit employer coordination, more or less centralized. Weak unions. Properties also described as paternalistic-liberal capitalism.
Netherlands	Medium union coordination, strong employer organizations.
Norway	Powerful centralized employer organizations, centralized union confederations. Additional coordinating role by the government.
Sweden	Powerful centralized employer organizations, centralized union confederations.
Switzerland	Very powerful employer organizations, weak decentralized unions. Weak centralization of government. Properties also described as liberal capitalism.
United Kingdom	Zero employer and union coordination.
United States	Zero employer and union coordination.

Chapter 6

How inflation-averse are central banks?

6.1 Introduction

In the theoretical literature on central bank independence, the degree of inflation aversion plays a central role. However, as already mentioned in chapter 3, in the empirical literature not much attention has been paid to the distinction between the degree of central bank conservativeness and central bank independence proper. In the context of central banks, conservativeness is defined as whether the central bank officially pursues monetary stability among its goals - thus, it reflects the degree of inflation aversion. Central bank independence proper reflects the extent to which the independence of a central bank is formally defined in the legislation of a country. For a thorough discussion of these concepts, see De Haan and Kooi (1997). They show that 'conservativeness as embodied in the central bank law' is not related to the inflation performance of OECD countries. The measurement of central bank independence has been discussed extensively in chapter 3 of this thesis. In this chapter, the focus is on central bank conservativeness, i.e. we examine how inflation-averse central banks are. In particular, we are interested in how conservativeness can be measured.

One of the problems in dealing with the degree of conservativeness of central banks is that it is a variable that cannot be observed. Therefore, a latent variable analysis is applied in order to include conservativeness as a variable in macroeconomic models. In section 6.2, the economic variables that may serve as indicators for conservativeness of central banks are discussed, while in section 6.3 the factor analysis model is presented for the 1980s. Section 6.4 extends the model to include the degree of conservativeness in the 1990s. In section 6.5, the constructed measure of conservativeness is applied empirically. Section 6.6 concludes.

6.2 Possible indicators of conservativeness

The setup of the research is to let macroeconomic theory imply a structure on a model of conservativeness of central banks, in the 15 countries that were a member of the European Union in the late 1990s. This model is consequently tested by the latent variables analysis that is performed, using confirmatory factor analysis. In order to do this, we need variables that may serve as possible indicators of conservativeness, and their relations. One of the most influential papers in the modern theoretical literature on central bank independence and conservativeness is Rogoff (1985). From his work and that of Eijffinger and Hoeberichts (1998), we expect to find that the rate and variability of inflation and the variability of output are related of the degree of conservativeness of a central bank. To a large extent, these variables behave very similarly for all developed countries. In order to obtain a clearer view of their country-specific behaviour, the variables are corrected using their averages vis-à-vis the average of the OECD countries:

$$\pi^* = \pi - \pi_{OECD},$$

where π^* denotes the corrected inflation rate used in this chapter, π is the common inflation rate and π_{OECD} is the average rate of inflation for all OECD countries. The variability of inflation is defined by its standard deviation, denoted as std π^*.

Whereas the inclusion of inflation as a variable expressing policy can be justified, output variability is a variable that clearly reflects economic shocks as well as economic policy. Therefore, we need to take a closer look at how this variable can be defined more precisely, in order to isolate the policy component. First, we specify the standard deviation of real GDP growth as a measure of output variability. To correct for economic shocks, we then subtract the standard deviation of the terms of trade. Terms of trade is defined as the ratio between export and import price indices. The resulting variable, named std Y^*, now mainly reflects the impact of economic policies. Unfortunately, since terms of trade data are largely unavailable for Portugal, we have to exclude this country from our sample. The resulting sample set therefore contains 14 EU countries.

A negative relation is expected between conservativeness and the rate and variability of inflation, whereas a positive relation is expected between conservativeness and output variability. Since we assume that conservativeness adjusts slowly to changes in macro-economic variables, we also include the rate and variability of inflation and the variability of output in the previous time period as possible indicators. Data on inflation, defined as changes in consumer prices, and real GDP have been obtained from the IMF's International Financial Statistics.

A final indicator for conservativeness is obtained from chapter 3. There, we have decomposed the CBI indicators of Cukierman (CUK), Cukierman et al. (CWN) and Grilli, Masciandaro and Tabellini (GMT) into indicators for central bank independence proper and conservativeness. Due to lack of variation in the GMT indicator of conservativeness and its availability for only a limited number of countries, we do not use it here. Since the component of conservativess is the same for CUK and CWN, we end up with one additional indicator for conservativeness, which for the rest of the chapter is referred to as CI. This indicator is available for the periods 1960–1971, 1972–1979, 1980–1989 and 1990–1997. A positive relation is expected between the degree of conservativeness and the CI indicator.

Since measurement of CBI has been described for the 1980s, we start our analysis of conservativeness in the same period. Later, we extend our model to include the 1990s as well, in order to detect changes in conservativeness. To be able to perform a static factor analysis, the values of π^* and Y^* have been averaged. Summarizing, the indicators that are used in our initial latent variables model are: π^*, std π^*, std Y^*, averaged for the period of interest (1980–1989) and the previous period (1970–1979), and CI in 1980–1989.

6.3 Conservativeness in the 1980s

In this section, we sketch the latent variables approach that is used to evaluate to which extent the indicators that were mentioned in the previous section describe the same unobservable phenomenon. Here, the unobservable phenomenon is the conservativeness of central banks. Since the indicators are intended to measure the same concept, they should be correlated. Table 6.1 gives the correlation matrix of the variables used for the time period under consideration.

It is clear from the table that the correlations between the different indicators are not perfect. Therefore, the different indicators are considered to be imperfect measures of the unobservable concept of conservativeness, generated by the

Table 6.1: Correlations between the indicators * 100%

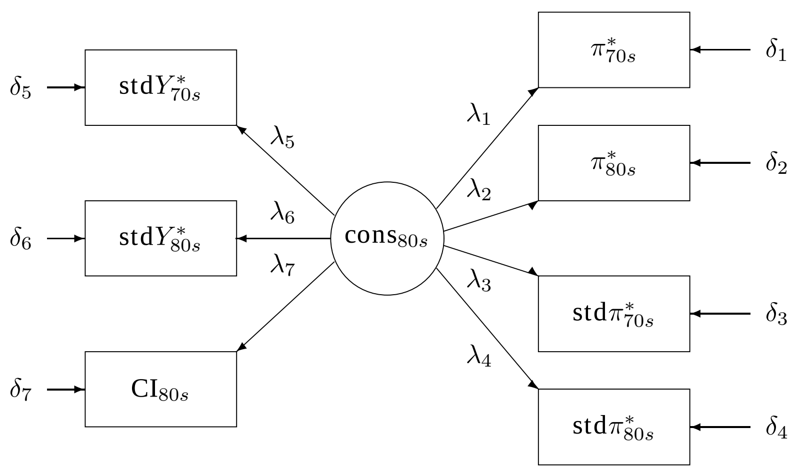
	π^*_{80s}	π^*_{70s}	std π^*_{80s}	std π^*_{70s}	std Y^*_{80s}	std Y^*_{70s}	CI_{80s}
π^*_{80s}	100						
π^*_{70s}	71	100					
std π^*_{80s}	53	48	100				
std π^*_{70s}	64	64	37	100			
std Y^*_{80s}	-9	-3	-6	-11	100		
std Y^*_{70s}	-8	-74	-66	-52	3	100	
CI_{80s}	1	-2	-4	33	25	-6	100

following factor analysis model:

$$x_{ci} = \tau_i + \lambda_i \xi_c + \delta_{ci}, \tag{6.1}$$

where x_{ci} is indicator i for country c and ξ_c is the unobservable concept of conservativeness, denoted as cons80s. Further, τ_i and λ_i are parameters that allow for the differences in mean and scaling between the indicators, and δ_{ci} is a random measurement error. We assume ξ to have mean zero and variance one, and the errors δ_{ci} and δ_{di} to be uncorrelated for $c \neq d$ and independent of ξ_c. A graphical representation of the model is shown in figure 6.1

Figure 6.1: Path diagram of the factor analysis for the 1980s.



As in chapter 3 for the case of central bank independence, we have estimated model (6.1) using maximum likelihood procedures incorporated in the software package AMOS because of the small sample size. The resulting estimates are shown in table 6.2. The values in the second column of the table are the estimates of the coefficients in model (6.1) for the different indicators. Corresponding t-statistics, calculated using the method of White, are given in the second column.

Table 6.2: Estimation results for the factor analysis of the period 1980-1989

	Estimate	t-stat
Intercept (τ_i)		
π^*_{80s}	-0.0198	-1.8279
π^*_{70s}	-0.0127	-1.7169
std π^*_{80s}	-0.0198^{**}	-8.1979
std π^*_{70s}	-0.0281^{**}	-8.7711
std Y^*_{80s}	0.0483^{**}	4.1190
std Y^*_{70s}	0.0315^{**}	4.2146
CI_{80s}	-0.4714^{**}	-5.1403
Loading (λ_i)		
π^*_{80s}	-0.0388	-1.5544
π^*_{70s}	-0.0245^{**}	-3.4208
std π^*_{80s}	-0.0057	-1.5119
std π^*_{70s}	-0.0086	-1.6751
std Y^*_{80s}	0.0030	0.1611
std Y^*_{70s}	0.0228^{**}	2.7103
CI_{80s}	-0.0199	-0.1100
Error variance (ψ_{ii})		
π^*_{80s}	0.0006	1.8145
π^*_{70s}	0.0002^{*}	2.3724
std π^*_{80s}	0.0001^{**}	3.2499
std π^*_{70s}	0.0001^{**}	2.9372
std Y^*_{80s}	0.0019	1.8615
std Y^*_{70s}	0.0003^{**}	2.6792
CI_{80s}	0.1174^{**}	5.2559

The χ^2-statistic, which compares the proposed model to the saturated model, has a value of 4.83 with 14 degrees of freedom, lying way below the 5% critical value of 23.7. The comparative fit index CFI has value 1. According to these measures, the model fits very well. The estimated reliability of the indicators is 0.90.

From table 6.2 it follows that the estimated coefficients all have the expected sign. It is also clear that the results are dominated by the 'lagged' variables: the loadings of π^*_{70s} and std Y^*_{70s} are the only ones that are significant.[1] The Cukierman indicator is not significant.

[1]Other measures of output, such as the variability of real GDP and the output gap have been experimented with. However, they all turn out to be insignificant in a factor analysis.

Our model contains observations on three variables in the current as well as previous time period, which are expected to be correlated *a priori*. As seen in table 6.1, especially the correlations among the rates and variabilities of inflation are high. This casts some doubt on the assumed independence of the corresponding measurement errors. Therefore, we have also performed a factor analysis on an augmented model. In this augmented model, a parameter is added that reflects the correlation between the measurement errors of the variables that have been included for two different time periods. In figure 6.1, this comes down to adding two-sided arrows between the error terms δ_1 and δ_2, δ_3 and δ_4 and between δ_5 and δ_6. Consequently, the results of the factor analysis include an estimate for these correlations. Since these estimated correlations all turn out to be small and insignificant, however, we stick to our original model specification.

Using the results of the factor analysis, the concept of conservativeness can be quantified. Using the estimated parameters, we can calculate the predicted factor scores, which are shown in table 6.3 in descending order of inflation aversion.

Table 6.3: Predicted factor scores

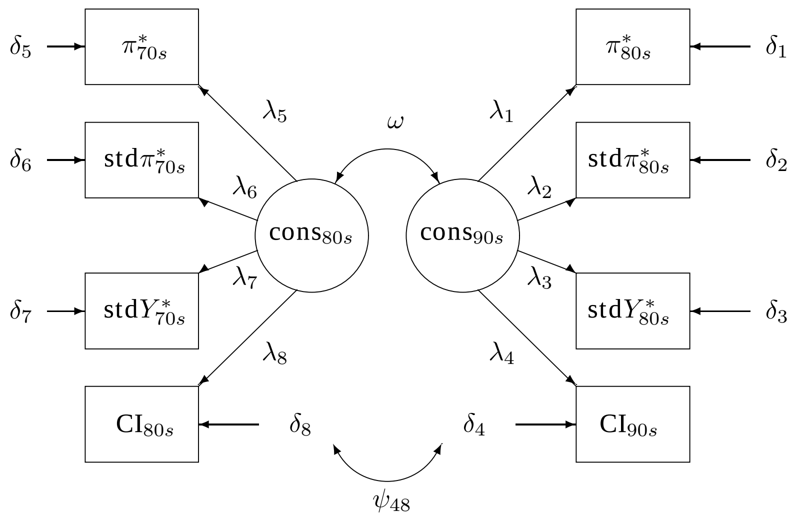
Country	Score
Austria	1.1328
Germany	1.0585
Belgium	0.9370
Netherlands	0.9098
Sweden	0.4951
France	0.4875
Denmark	0.4840
Finland	0.2974
Norway	0.2064
UK	−0.7247
Ireland	−1.1252
Italy	−1.2353
Spain	−1.2968
Greece	−1.6265

The table shows that for the period 1980–1989, Austria can be classified as the most inflation-averse country in the EU, closely followed by Germany, Belgium and the Netherlands. The southern European countries are the ones that are most inflation-prone. Since we are also interested in inflation-aversion in the EU for the period 1990–1999, we would like to extend the model (6.1). This is done in the next section.

6.4 Conservativeness in the 1990s

In order to get a measure for the level of conservativeness in the 1990s, we add a second latent variable to our model, which is denoted as cons90s. Since table 6.2 shows the 'lagged' variables to have the largest weight in constructing the measure of conservativeness in the 1980s, we use the variables averaged for the 1970s as indicators for cons80s, together with the CI indicator for the 1980s. The variables averaged for the 1980s, as well as the CI indicator for the 1990s, are taken as indicators for cons90s.[2] The correlation between CI_{80s} and CI_{90s}, which is not shown in table 6.1, is 0.95. Therefore, we include the assumption that these indicators are correlated in the factor model. The correlations between CI_{90s} and the rest of the model variables are below 0.30 and insignificant. We also assume that the latent variables cons80s and cons90s are correlated. The resulting two-factor model is shown graphically in figure 6.2. The assumed correlations are depicted using arrows with arrowheads on both sides.

Figure 6.2: Path diagram of the factor analysis for the 1980s and 1990s.



We have again estimated the model using maximum likelihood procedures incorporated in AMOS. The estimated results are shown in table 6.4. The χ^2-

[2]We have also experimented with specifications that include the economic variables averaged for the 1990s. However, due to the small number of observations, these models turn out to be too complex to be estimated correctly.

statistic has a value of 13.5 with 18 degrees of freedom, lying well below the 5% critical value of 28.9. The comparative fit index CFI has value 1. According to these measures, the model fits very well.

Table 6.4: Estimation results of the two-factor model

	Estimate	t-stat
Intercept (τ_i)		
π^*_{80s}	-0.020	-1.76
π^*_{70s}	-0.013	-1.66
std π^*_{80s}	-0.020^{**}	-7.90
std π^*_{70s}	-0.028^{**}	-8.46
std Y^*_{80s}	0.048^{**}	3.97
std Y^*_{70s}	0.032^{**}	4.06
CI_{90s}	-0.500^{**}	-5.62
CI_{80s}	-0.471^{**}	-4.91
Loading (λ_i)		
π^*_{80s}	-0.035^{**}	-3.31
π^*_{70s}	-0.025^{**}	-3.92
std π^*_{80s}	-0.006^{*}	-2.41
std π^*_{70s}	-0.009^{**}	-2.89
std Y^*_{80s}	0.003	0.26
std Y^*_{70s}	0.023^{**}	3.35
CI_{90s}	-0.031	-0.36
CI_{80s}	-0.047	-0.52
Error variance (ψ_{ii})		
π^*_{80s}	0.0006	1.13
π^*_{70s}	0.0002	1.35
std π^*_{80s}	0.0001^{*}	2.21
std π^*_{70s}	0.0001	2.09
std Y^*_{80s}	0.002^{*}	2.55
std Y^*_{70s}	0.0003	1.90
CI_{90s}	0.102^{*}	2.54
CI_{80s}	0.118^{*}	2.53
Error covariance (CI_{90s}, CI_{80s})	0.105^{*}	2.48
Covariance between factors	0.924^{**}	6.35

Again, we have also performed a factor analysis on an augmented model, in which a parameter is added that reflects the correlation between the measurement errors of the variables that have been included for two different time periods. Since the estimated correlations all turn out to be small and insignificant, we stick to the model specification in figure 6.2. Using the estimates in table 6.4, we can calculate the predicted factor scores for both the 1980s and the 1990s. These are shown in table 6.5.

Table 6.5: Predicted factor scores and corresponding country rankings

Country	1980–1989		1990–1999	
	Score	Rank	Score	Rank
Austria	1.0252	1	0.8909	4
Germany	0.9403	2	0.9433	2
Belgium	0.8436	3	0.5496	6
Netherlands	0.7999	4	0.9543	1
Denmark	0.7585	5	0.9042	3
France	0.7381	6	0.4922	7
Sweden	0.6201	7	0.7201	5
Finland	0.2933	8	0.4902	8
Norway	0.1728	9	0.0283	9
UK	−0.7851	10	−0.4103	10
Ireland	−1.1538	11	−1.4177	13
Italy	−1.1778	12	−1.2476	12
Spain	−1.4154	13	−1.0295	11
Greece	−1.6598	14	−1.8681	14

First of all, the scores for the 1980s do not differ much from the scores in table 6.3, as should be the case. The values are slightly different, due to the fact that the model includes additional parameters and restrictions, and the factor is constructed such that it has expectation zero and variance one. The main difference is that the rank order of Sweden, France and Denmark, the scores of which are very close in both analyses, is reversed in table 6.5. Due to the restrictions on the mean and variance of the factors, a comparison of absolute values between both decades is not very informative. When we look at their relative positions, however, we see that Austria is no longer the country in the EU with the highest degree of inflation aversion. The Netherlands, Germany and Denmark have overtaken Austria in the 1990s. The southern European countries still rank among the most inflation-prone, although Ireland has moved to thirteenth place in the 1990s.

6.5 Conservativeness in an empirical model

Now that we have quantified the concept of conservativeness, we would like to use the resulting construct in an empirical model. In the literature, the impact of conservativeness on inflation is often the topic of research. Examining this relation here would lead to circular reasoning, however, since the rate and variation of inflation are used to construct our measure of conservativeness. Consequently, regressing conservativeness on inflation would almost certainly lead to a significant result. Instead, we examine the impact of conservativeness on unemployment in the period 1980–1989. For this purpose, we use the factor scores of table 6.3. Using the factor scores given in table 6.5 for the 1980s does not qualitatively change the results. We follow the approach used in the empirical model for CBI in section 5 of chapter 3, first using a bivariate model and then adding control variables one-by-one. As control variables, we include CBI and trade openness. As in chapter 3, we also include the log of GDP (value for 1985) in the model containing openness. The estimates, along with their t-values, are shown in table 6.6. The superindex * denotes significance of the coefficient at a 5% level.

Table 6.6: The impact of conservativeness on unemployment

Control	coeff	t-stat	cons-coeff	t-stat	R^2
–	–	–	-1.9750	-1.8456	0.129
CBI	-0.5045	-0.3066	-1.7224	-1.2778	0.056
Trade openness	19.1842^*	2.5665	-3.0218^*	-3.1339	0.301
log GDP (1985)	1.7098	1.8815			

From the table, it follows that the effect of conservativeness on unemployment has a negative sign, and is only significant if trade openness is included in the model. To correct for the underestimation of the coefficient of conservativeness, we have also estimated the models using the CALS estimator. The resulting coefficients of the conservativeness variable are shown in table 6.7.

Table 6.7: CALS estimates of the indicator of conservativeness

Control	cons-coeff	t-stat	CALS coeff	CALS t
–	-1.9750	-1.8456	-2.2532	-1.8367
CBI	-1.7224	-1.2778	-2.1453	-1.2690
Trade openness[a]	-3.0218^*	-3.1339	-3.5551^*	-3.0675

[a] Model specification includes log GDP in 1985.

It is clear from table 6.7 that, using the CALS estimator, there is a reasonable increase in the absolute value of the coefficients. The t-values decrease, as they should, but this effect is small. The significance in the model that includes trade openness is not affected.

6.6 Conclusions

This chapter is an attempt to measure the degree of conservativeness of central banks using a latent variables approach. For a sample of 14 European countries it is shown that, unsurprisingly, the rate of inflation is the most important indicator of conservativeness. The indicator of Cukierman, which has been designed specifically to measure conservativeness, does not seem to be of much use. Also, when results for the 1980s and 1990s are compared, the analysis shows that there have not been major changes in the degree of conservativeness. Finally, the impact of conservativeness on unemployment is examined. Only when trade openness is also included in the model, the relation is significant.

Chapter 7

Does economic freedom contribute to growth?*

7.1 Introduction

In this chapter, we examine the relation between economic freedom and economic growth. What connection, if any, is there between economic development and economic liberties? Economic theory indicates that economic freedom affects incentives, productive effort, and the effectiveness of resource use. Indeed, since the time of Adam Smith, if not before, economists and economic historians have argued that the freedom to choose and supply resources, competition in business, trade with others and secure property rights are central ingredients for economic progress. Using various proxies for economic freedom, an increasing number of recent empirical studies suggest that economic freedom may be important in explaining cross-country differences in economic performance (e.g. De Vanssay and Spindler, 1994, De Haan and Siermann, 1998, Dawson, 1998, Nelson and Singh, 1998 and De Haan and Sturm, 2000).[1]

In most of the studies referred to above, a single overall measure of economic freedom based on an aggregation of various underlying components is used. Recently, Heckelman and Stroup (2000) have rightly criticized these procedures as being ad hoc. They argue that aggregated measures of economic freedom also include some components that are not, or even negatively, related to growth and that, therefore, it should be possible to find better aggregated measures if an adequate weighting procedure for the various components can be found.

*This chapter is a revised version of Sturm, Leertouwer and De Haan (2002).

[1]De Haan and Sturm (2000) conclude that the level of economic freedom is not related to economic growth. In contrast, changes in economic freedom are robustly related to growth.

Heckelman and Stroup come up with an alternative aggregation methodology in which aggregation is directly based upon the relevance of each component for growth, as determined by a multivariate regression analysis. On the basis of cross country regressions for 49 countries over the period 1980–1990, they conclude that "given the proper framework, differences in economic freedoms between nations can explain almost half of the variation in growth".

In this chapter, we argue that Heckelman and Stroup criticize the existing aggregation procedures for good reasons. However, their aggregation procedure has serious shortcomings. We propose an alternative indicator using a latent variables approach. We also criticize Heckelman and Stroup's conclusion on the importance of economic freedom, as it is not based on the by now standard robustness analyses as proposed by Levine and Renelt (1992) and Sala-i-Martin (1997a,b). Using these analyses, we find that our index of economic freedom is not robustly related to economic growth, while the index of Heckelman and Stroup is only weakly related to economic growth.

The remainder of this chapter is organized as follows. The next section discusses the aggregation procedure as proposed by Heckelman and Stroup and our suggested alternative. Section 7.3 summarizes the robustness analysis that we employ, while in section 7.4 our results are presented. The final section offers some concluding comments.

7.2 Aggregation of components of economic freedom

Heckelman and Stroup employ the (components of the) economic freedom indicators as constructed by Gwartney et al. (1996). According to the latter, an index of economic freedom should measure the extent to which rightly acquired property is protected and individuals are free to engage in voluntary transactions. In an economically free society, the fundamental function of the government is the protection of private property and the enforcement of contracts. When a government fails to protect private property, takes property itself without full compensation or establishes restrictions that limit voluntary exchange, it violates the economic freedom of its citizens. Institutional arrangements that restrain trade, increase transaction costs, weaken property rights, and create uncertainty will reduce the realization of gains from trade and also the incentive of individuals to engage in productive activities.

Gwartney et al. choose various measures and rate a large number of countries on each of these measures on a scale of 0-10, in which zero means that a country is completely unfree and ten means it is completely free. The measures are in four broad areas: money and inflation (I), government operations and regulations (II), 'takings' and discriminatory taxation (III) and international exchange (IV). The components in the monetary area reflect the availability of sound money to the citizenry. The components in the economic structure are indicators of reliance on markets rather than the political process to allocate resources. In the takings area, the index is designed to measure the degree to which governments treat individuals equally rather than engage in transfer activities and impose taxes. Finally, the components in the international area are designed to measure the presence of policies consistent with free trade, see also Gwartney et al. (1999).

These components are combined in aggregated rankings in three ways. In the first index (**Ie**) each component is assigned a weight equal to the inverse of its standard deviation, while in the index **Is1** the importance of the components is based on a survey among experts in the field of economic freedom. Finally, in the index **Is2** the weighting is also based on a survey, but this time the survey is held among a number of country experts.

Heckelman and Stroup rightly criticize these aggregation procedures, arguing that they are ad hoc and may lead to misspecification problems. Alternatively, they first run a multivariate regression of growth on all 14 components of economic freedom for which data are available over 1980–1990. Gwartney et al. (1996) distinguish 17 components, but Heckelman and Stroup do not take the following three components into account in their analysis: Price controls, Freedom of private business to compete in markets and Equality of citizens under the law and Access of citizens to a non-discriminatory judiciary, as they are not available for 1980. The weights in Heckelman and Stroup's aggregation procedure are determined by calculating the contribution of each t-statistic to the sum of the absolute values of all the t-statistics. They claim that "in this way, we can construct an overall index for each country based on the components of economic freedom which are weighted stronger toward those freedoms which were found to have a higher significance level for impacting on growth (either positively or negatively)." (p. 537). Consequently, using bivariate regressions this time, they find that, in contrast to the aggregate indices of Gwartney et al. (1996), their index (constructed for 1980) is significantly related to economic growth measured over the period 1980–1990 in a sample of 49 countries.

The aggregation procedure of Heckelman and Stroup can be criticized for a number of reasons. The basic problem is that their procedure is based on circular reasoning. First, the relationship of the various components to economic growth is estimated using a multivariate regression. Then, the weights are determined on the basis of this relationship. Consequently, it is not surprising that, in a bivariate regression, their index is related to growth. The same effect would take place if we decided to regress the measure of conservativeness that has been constructed in the previous chapter using, among others, inflation and the standard deviation of inflation, on inflation itself. The weighting scheme as suggested by Heckelman and Stroup will almost always result in a significant index explaining economic growth. To illustrate this, we have simulated their procedure using random numbers instead of the 14 components provided by Gwartney et al. (1996). We create 14 series of 49 observations with draws from a uniform distribution ranging from 0 to 10. After running an OLS regression using these 14 series and a constant 'explaining' economic growth, we use the resulting t-statistics to construct an aggregated index. The bivariate relationship between this aggregated index and economic growth is then estimated, and the entire procedure is repeated 10,000 times. Only in 0.2% of these regressions the aggregated index based on randomly constructed series turns out to be *insignificant* at a 5% level.

Moreover, the use of t-statistics as weighting factors is rather suspect as t-statistics do not measure the impact of a variable on growth. The index of Heckelman and Stroup leads to some very counterintuitive rankings of countries. For instance, according to their ranking Italy (rank 3) has a substantial higher level of freedom than e.g. the UK (rank 16) and the US (rank 26), see table 7.5 for further details.

A final criticism that can be raised is that Heckelman and Stroup do not examine how sensitive their index is to the selection of countries in the sample. We have analyzed this issue as follows. From the sample of 49 countries, we randomly draw 40 countries and run bivariate regressions of growth on all 14 components of economic freedom to determine the weights of these components using the t-statistics. This is repeated 10,000 times, giving a distribution of the weights of the various components. The results are shown in table 7.1.

The numbering of the indicators in the first column corresponds to the numbering in table 4 of Heckelman and Stroup. It follows that most of the weights are extremely sensitive to the selection of countries. For instance, the standard deviations of the weights are often higher than the absolute value of their averages.

Still, Heckelman and Stroup are right to criticize the existing aggregation procedures as being ad hoc. The basic problem is that various components may provide some information on an unobservable variable called economic freedom. From this perspective, a latent variable approach is therefore preferable.

Table 7.1: Sensitivity of the weighting scheme to the selection of countries

Indicator	Description	mean	stdev	min	max
IA	Money Expansion	−0.034	0.037	−0.137	0.112
IB	Inflation Variability	0.125	0.029	−0.056	0.221
IC	Foreign Currency Accounts	0.018	0.028	−0.133	0.105
ID	Deposits Abroad	−0.032	0.038	−0.168	0.111
IIA	Government Consumption	0.058	0.042	−0.067	0.238
IIB	Government Enterprises	0.047	0.046	−0.146	0.190
IIF	Credit Market	−0.085	0.043	−0.227	0.108
IIIA	Transfers & Subsidies	0.088	0.042	−0.084	0.207
IIIB	Marginal Tax Rates	−0.176	0.037	−0.319	0.008
IIIC	Conscription	−0.024	0.038	−0.123	0.136
IVA	Trade Taxes	0.032	0.049	−0.148	0.205
IVB	Exchange Rate Controls	0.139	0.053	−0.036	0.418
IVC	Exp. Size of Trade Sector	0.009	0.043	−0.147	0.171
IVD	Capital Restraints	−0.011	0.041	−0.179	0.121

We consider the fourteen components of economic freedom as imperfect measures (indicators) of economic freedom. In order to see how the indicators are related to each other, their correlation matrix is shown in table 7.2.

As in the previous chapters, we first perform a factor analysis to construct measures of (the components of) economic freedom. However, this does not result in a specification that is satisfactory. Obviously, the assumptions regarding the structure of the covariance matrix that the FA model imposes are not satisfied in this case. Therefore, we employ principal components analysis instead. As described in chapter 2, this is a method of combining a set of variables into a single variable that best reflects the original data, using all information that is available in the indicators without imposing a specific structure on the covariance matrix. It is neither based on subjective judgements nor on circular reasoning.[2]

The procedure partitions the variance of a set of variables and uses it to determine the linear combination of these variables that maximizes the variation of the newly constructed principal component. In order to see how many components we should include in the analysis, we examine the scree plot in figure 7.1, which plots the number of components against their eigenvalues.

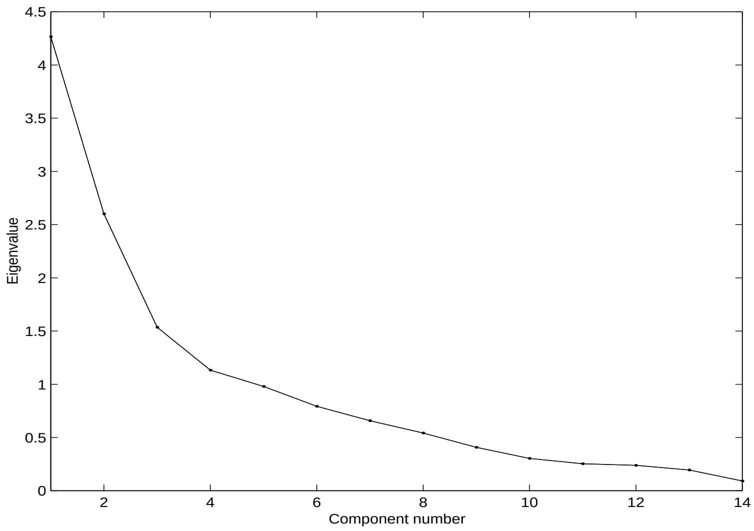
Using the Kaiser rule and examining the kink in the scree plot, we end up

[2]After we had finished the first draft of this chapter, we found out that in the 2001 edition of *Economic Freedom of the World*, PCA is also used to construct an aggregated measure of economic freedom.

Table 7.2: Correlations between the indicators * 100%

	IA	IB	IC	ID	IIA	IIB	IIF	IIIA	IIIB	IIIC	IVA	IVB	IVC	IVD
IA	100													
IB	62	100												
IC	-8	19	100											
ID	-7	4	71	100										
IIA	-31	-25	2	14	100									
IIB	27	29	29	30	12	100								
IIF	58	58	33	20	-11	54	100							
IIIA	-20	-56	-22	6	52	10	-18	100						
IIIB	-14	-20	30	35	23	21	6	39	100					
IIIC	23	-11	-24	-8	-10	5	8	16	-6	100				
IVA	51	70	13	-7	-40	27	46	-60	-27	-25	100			
IVB	28	45	32	28	-7	36	43	-45	-1	4	52	100		
IVC	30	19	-6	6	-11	-5	15	-7	-14	-5	26	18	100	
IVD	45	44	24	35	-11	44	47	-32	11	10	57	65	6	100

Figure 7.1: Scree plot of the components and eigenvalues



with three or four principal components. Since the solution using three components is extremely difficult to interpret, we use four principal components that are labeled PC1 to PC4. The (unrotated) standardized solution of the PCA is given in table 7.3.

Interpretation of the components shown in table 7.3 is still not very straight-

Table 7.3: Unrotated standardized solution of PCA

Indicator	PC1	PC2	PC3	PC4
IA	0.661	−0.282	0.451	−0.002
IB	0.775	−0.250	0.023	0.169
IC	0.373	0.620	−0.468	−0.039
ID	0.661	0.755	−0.195	−0.082
IIA	−0.351	0.511	0.267	0.430
IIB	0.539	0.476	0.295	0.120
IIF	0.756	0.095	0.326	0.161
IIIA	−0.522	0.361	0.578	0.330
IIIB	−0.062	0.767	0.101	−0.101
IIIC	−0.023	−0.101	0.632	−0.661
IVA	0.766	−0.302	−0.166	0.192
IVB	0.721	0.176	−0.105	−0.121
IVC	0.324	−0.267	0.106	0.450
IVD	0.735	0.295	0.125	−0.248

forward. To see if we can obtain a solution that is easier to interpret, we rotate the solution using the rotation methods described in chapter 2. Direct oblimin rotation gives the most satisfying results, and the rotated standardized solution is given in table 7.4.

The loadings of the standardized solution are also the correlations between the indicators and the components, and we use a superindex * to denote correlations that are significant at the 5% level. It follows from the table that the first component displays the highest number of significant correlations with the indicators of economic freedom. We will therefore use this component in our empirical research. The empirical analyses that follow have, however, also been done using each of the other three principal components. This yields the same general conclusion.

Using the standardized loadings in table 7.4, we can calculate scores of the latent variable for the different countries. Table 7.5 shows the scores and ranking of the 49 countries according to our PC1 index and compares it to those of Heckelman and Stroup and Gwartney et al. (1996).[3] The ordering in table 7.5 is based on the index of Heckelman and Stroup. In parentheses the ranking of the other indicators is shown. The indicator of Gwartney et al. is their equal impact

[3]We have corrected the Gwartney et al. coding of Canada as reported in Heckelman and Stroup's table 5 (7.5 instead of 4.7). Although Gwartney et al. do not provide information for all 14 indicators for Malta, Heckelman and Stroup came up with data for all the indicators for this country. As Heckelman kindly provided their data set to us, we are able to report estimates for the same 49 countries as in their study.

Table 7.4: Standardized solution of PCA using Oblimin rotation

Indicator	PC1	PC2	PC3	PC4
IA	0.773^*	-0.079	-0.256	-0.287^*
IB	0.780^*	0.061	-0.452^*	0.118
IC	0.090	0.801^*	-0.133	0.225
ID	0.062	0.825^*	0.115	0.004
IIA	-0.173	0.161	0.772^*	0.112
IIB	0.550^*	0.542^*	0.173	-0.149
IIF	0.815^*	0.285^*	-0.106	-0.114
IIIA	-0.232	-0.091	0.904^*	-0.152
IIIB	-0.149	0.635^*	0.445^*	-0.201
IIIC	0.014	-0.119	0.015	-0.903^*
IVA	0.723^*	0.051	-0.565^*	0.264
IVB	0.553^*	0.490^*	-0.391	-0.047
IVC	0.485^*	-0.207	-0.065	0.282^*
IVD	0.590^*	0.571^*	-0.268	-0.303^*

(Ie) index.

It is clear from the table that our ranking resembles the one by Gwartney et al. (1996) and is at odds with the ranking of Heckelman and Stroup. The correlation coefficients of our index and the indices of Heckelman and Stroup and Gwartney et al. are 0.37 and 0.79, respectively. The correlation coefficient between the latter two indices is 0.13.

Now, we can examine whether our alternative index of economic freedom is related to economic growth by applying it in an empirical model. Before we do so, however, we briefly discuss our modeling strategy.

7.3 Robustness analysis

The empirical analysis of Heckelman and Stroup (2000) is based on simple regressions of the average growth rate of GDP per capita over the period 1980–1990 on (components of) various aggregated indicators of economic freedom. In other words, no control variables are included. Consequently, their conclusion that differences in economic freedoms between nations can explain almost half of the variation in growth is based on biased estimates. Which control variables should be included? A serious problem in this regard is that economic theory does not provide enough guidance for the proper specification of empirical growth models. Sala-i-Martin (1997a,b) identifies, for instance, around 60

Table 7.5: Rankings of economic freedom (1980)

	Country	H-Stroup	PC1	rank	G'ney	rank
1	South Korea	333.4	0.07	(24)	4.0	(31)
2	Philippines	325.9	0.56	(13)	4.7	(19)
3	Italy	313.5	0.00	(26)	3.8	(36)
4	Singapore	301.2	2.04	(1)	7.0	(3)
5	Cyprus	286.7	-0.22	(31)	3.6	(41)
6	Indonesia	284.9	-0.75	(41)	4.9	(18)
7	South Africa	283.3	0.18	(20)	4.4	(25)
8	Spain	263.8	-0.05	(29)	4.0	(32)
9	Japan	262.4	1.33	(5)	6.4	(8)
10	Taiwan	259.2	0.00	(27)	5.4	(13)
11	Finland	250.8	0.78	(11)	5.0	(16)
12	Portugal	247.7	-0.35	(33)	3.3	(45)
13	Austria	229.3	0.88	(10)	5.2	(14)
14	Netherlands	227.9	1.45	(4)	6.4	(9)
15	Malaysia	223.9	0.16	(21)	6.0	(11)
16	United Kingdom	222.1	0.29	(18)	4.7	(20)
17	Sri Lanka	214.8	-0.49	(36)	3.7	(39)
18	Belgium	213.8	1.79	(2)	6.8	(5)
19	Ireland	203.4	0.13	(22)	4.5	(22)
20	Turkey	201.2	-1.85	(47)	2.0	(48)
21	Egypt	199.2	-1.04	(43)	3.5	(42)
22	Switzerland	197.6	1.06	(8)	7.3	(2)
23	Sweden	195.7	0.23	(19)	4.0	(33)
24	Germany	193.9	1.32	(6)	6.6	(6)
25	Denmark	182.8	0.91	(9)	4.3	(26)
26	United States	182.7	1.27	(7)	6.9	(4)
27	Greece	179.5	-0.61	(39)	3.8	(34)
28	New Zealand	177.7	0.69	(12)	5.1	(15)
29	Israel	176.9	-1.75	(46)	2.4	(47)
30	Tunisia	170.0	-0.51	(37)	3.2	(46)
31	Australia	166.9	0.55	(14)	6.0	(10)
32	Bolivia	166.6	-0.95	(42)	4.4	(24)
33	Norway	166.2	0.52	(15)	3.8	(37)
34	Chile	165.2	-0.42	(35)	4.1	(27)
35	France	162.6	0.33	(17)	4.6	(21)
36	India	126.2	-0.59	(38)	3.8	(35)
37	Canada	125.5	1.46	(3)	7.5	(1)
38	Guatemala	125.1	0.10	(23)	6.4	(7)
39	Kenya	118.2	-0.02	(28)	4.0	(30)
40	Malta	117.8	0.05	(25)	4.5	(23)
41	Argentina	116.8	-2.29	(48)	3.3	(44)
42	Mexico	101.8	-1.61	(45)	3.7	(38)
43	Ivory Coast	90.8	-0.67	(40)	4.0	(29)
44	Fiji	89.7	0.34	(16)	4.9	(17)
45	Zambia	86.8	-0.31	(32)	3.4	(43)
46	Ghana	80.3	-2.39	(49)	1.8	(49)
47	Zimbabwe	69.0	-0.15	(30)	3.7	(40)
48	Malawi	34.7	-0.38	(34)	4.1	(28)
49	Uruguay	1.5	-1.11	(44)	6.0	(12)

variables that have been suggested to be correlated with economic growth. The so-called *extreme bounds analysis* of Levine and Renelt (1992) is therefore often used to examine how 'robust' the relation between a certain variable of interest and economic growth is. In this approach, equations of the following general form are estimated:

$$\Delta Y_i = \alpha M_i + \beta F_i + \gamma Z_i + u_i, \tag{7.1}$$

where the subscript refers to country i, ΔY_i is the average growth of per capita GDP of country i, and M_i is a vector of 'standard' economic explanatory variables. F_i is the variable of interest (in this case: an indicator of economic freedom), Z_i is a vector of up to three possible additional economic explanatory variables, which according to the literature may be related to economic growth, and u_i is an error term. The extreme bounds test for the variable F says that if the lower extreme bound for β - i.e. the lowest value for $\hat{\beta}$ minus two standard deviations - is negative, while the upper extreme bound for β - i.e. the highest value for $\hat{\beta}$ plus two standard deviations - is positive, the variable F is not robust.

Sala-i-Martin (1997a,b) argues that the test applied in the extreme bounds analysis is too strong for any variable to really pass it. If the distribution of β has some positive and some negative support, then one is bound to find one regression for which the estimated coefficient changes sign if enough regressions are run. Instead of analyzing the extreme bounds of the estimates of the coefficient of a particular variable, Sala-i-Martin suggests to analyze the entire distribution. He considers the distance of the point estimates for β from zero, averaged over a large set of regression models. Broadly speaking, if the averaged 90 per cent confidence interval of a regression coefficient does not include zero, Sala-i-Martin classifies the corresponding regressor as a variable that is strongly correlated with economic growth. He concludes that a substantial number of variables are strongly related to growth. We use this approach in our empirical analysis.

7.4 Results

First, we present the results of the simple regression model (7.1). Table 7.6 reports the estimated coefficients and corresponding t-values in parentheses. The $M-$vector consists of initial income (Y_{1980}), average investment as share of GDP (I) and the average population growth. The indicators of economic freedom are our PC1 index, the index of Heckelman and Stroup and the **Ie** index of Gwartney et al.

Table 7.6: Estimated coefficients (t-values) for the regression model

Variable	PC1 index		H-Stroup		Gwartney et al	
Constant	16.27	(4.94)	14.56	(5.16)	16.11	(5.37)
Y_{1980}	−2.13	(−5.36)	−1.97	(−5.61)	−2.08	(−5.18)
I	0.27	(6.87)	0.22	(5.54)	0.27	(6.89)
Pop. growth	−1.46	(−6.68)	−1.39	(−6.71)	−1.44	(−6.29)
Econ. freedom	−0.02	(−0.08)	0.01	(1.92)	−0.08	(−0.49)
Adj. R^2	0.68		0.71		0.68	

It follows that in the basic regression, only the indicator by Heckelman and Stroup is significantly related to growth. As explained in the previous section, this does not imply that there is a robust relationship. To examine whether the results are robust, figures 7.2, 7.3 and 7.4 show the distribution of t-statistics of the three indicators of economic freedom if various combinations of up to three additional explanatory variables are included in the model. For a description of the additional variables, see table 7.7 in the appendix. The source for the first 13 variables in the table is the Penn World Table (PWT 5.6a), as discussed most recently in Heston and Summers (1996). The other variables are obtained from the Barro-Lee dataset described in Barro and Sala-i-Martin (1995). The histograms are based on 12,383 regressions.

Figure 7.2: Histogram using the PC1 index

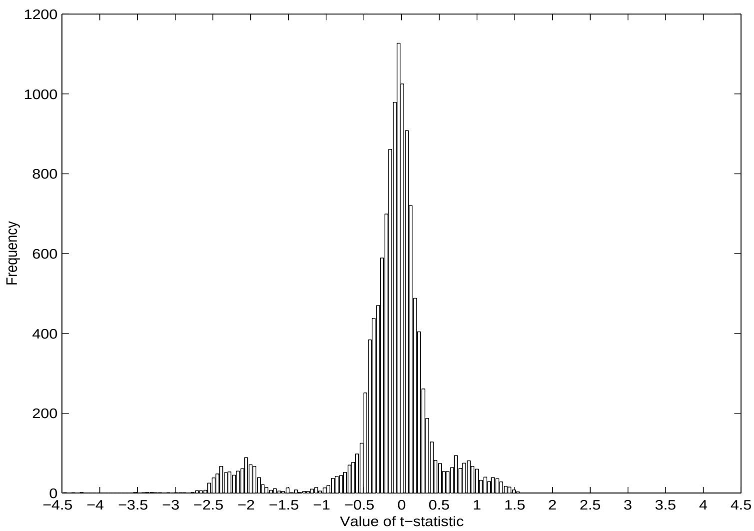


Figure 7.3: Histogram using the Heckelman-Stroup index

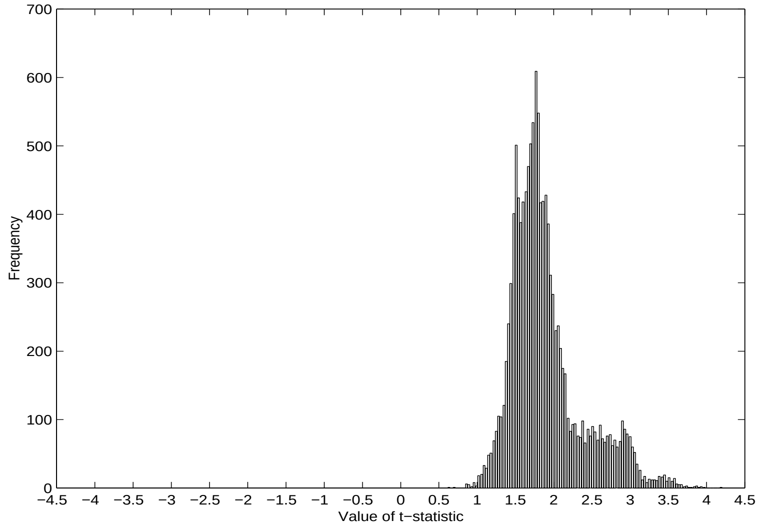
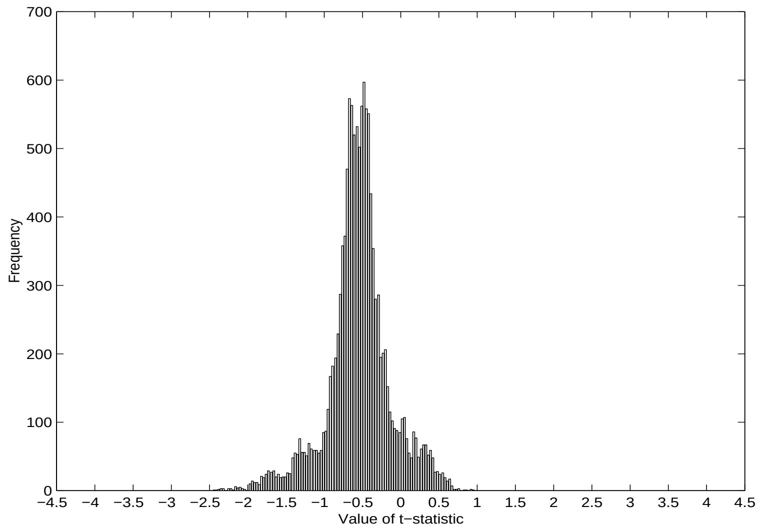


Figure 7.4: Histogram using the equal impact indicator of Gwartney et al.



The figures clearly lead to two conclusions. First, our index of economic freedom and the Gwartney-index are not robustly related to economic growth. This conclusion is fully in line with the results reported in De Haan and Sturm (2000), where the indicators of Gwartney et al. (1996) are used. Second, the indicator of Heckelman and Stroup is weakly related to economic growth. Its coefficient is always positive. However, in only 9% of all regressions the coefficient of their indicator is significantly different from zero at the 1% level. Applying a 5% significance level, the percentage of significant coefficients increases to 26. Based on these results, we conclude that economic freedom does not significantly influence economic growth.

7.5 Conclusions

Most studies on the relationship between economic freedom and growth employ a measure of economic freedom based on an (ad hoc) aggregation of various underlying components. We argue that the alternative aggregation procedure as recently suggested by Heckelman and Stroup (2000) - in which aggregation is directly based upon the relevance of each component for growth, as determined by multivariate regression analysis - has serious shortcomings. We present an alternative index based on latent variable estimation techniques. Using standard robustness analyses we find that the level of economic freedom is not robustly related to economic growth.

7.A Control variables

In table 7.7, the control variables that are used for Z in model (7.1) are described.

Table 7.7: Description of the control variables

Variable	Description
GRGDPTT	Growth rate Real GDP per capita (Terms of Trade)
GDPTT	Real GDP per capita (Terms of Trade)
INVEST	Real Investment (1985 prices)
GRPOP	growth rate Population
C	Real consumption (1985 prices)
G	Real govt (1985 prices)
Y	CGDP relative to US
OPEN	Openness
IPRI	Gross Domestic Private Investment
STLIV	Standard of Living index
INFL	Growth rate Price level GDP
DCPI	Growth rate Price level Consumption
GRXR	Growth rate Exchange rate with US$
HUMAN	Average schooling years in the total population over the age of 25
PYR	Average years of primary schooling in the total population over the age of 25
SYR	Average years of secondary schooling in the total population over the age of 25
HYR	Average years of higher schooling in the total population over the age of 25
PRI	% 'primary school attained' in the total population
PRIC	% 'primary school complete' in the total population
SECC	% 'secondary school complete' in the total population
HIGH	% 'higher school attained' in the total population
HIGHC	% 'higher school complete' in the total population
S	Total gross enrollment ratio for secondary education
H	Total gross enrollment ratio for higher education
POP15	Population Proportion under 15
FERT	Total fertility rate (children per woman)
MORT	Infant Mortality Rate (ages 0-1)
LIFEE0	Life expectancy at age 0
GOVWB	Nominal govt 'consumption' expenditure / Nominal GDP
GEXP	Nominal govt current expenditure (incl. interest payments & transfers) / Nominal GDP
GDE	Nominal govt expenditure on defense / Nominal GDP
GEEREC	Recurring nominal govt expenditure on education / Nominal GDP
GEETOT	Total nominal govt expenditure on education / Nominal GDP
INVPUB	Nominal public domestic investment (fixed capital formation) / Nominal GDP
GGCFD	Real public domestic investment (using HS deflator for investment) / Real GDP (deflated)
GVXDXE5	Real govt 'consumption' expenditure net of spending on defense & education / Real GDP
ASSASSP	Number of assassinations per million population per year
COUP	Number of coups per year
REVOL	Number of revolutions per year
PINSTAB	Measure of political instability
POLRIGHT	Index of political rights (from 1 to 7; 1=most freedom)
CIVLIB	Index of civil liberties, (from 1 to 7; 1=most freedom)
BMP	Black market premium
BMPL	Log (1+BMP)
TOT	Terms of trade shock (export price growth rate minus import price growth rate
LLY	Liquid liabilities / GDP

Chapter 8

Concluding remarks

The aim of this thesis is to provide valid measurements for a number of important variables in the field of political economy that do not depend on subjective judgement. Subsequently, these newly constructed measures have been applied in empirical models in order to assess the influence of the variables under scrutiny on economic and monetary policy outcomes. In this chapter, the results are summarized.

In chapter 2, we have discussed the methodology that is used throughout the chapters 3 to 7. An overview of latent variables models is given, as well as the procedures to estimate them. Corresponding issues such as correcting the underestimation of the coefficient and criteria concerning model selection and model fit are discussed. Also, an exact test that can be used to examine the presence of dynamics and unit root behaviour in panel data models is described.

In chapter 3, we have shown how the partial conflict between competing indicators of central bank independence can to a certain extent be resolved using latent variables modeling. We have disentangled the concepts of CBI and central bank conservativeness and constructed a measure of CBI proper in 22 OECD countries, for the period 1980–1989. The resulting ordering of countries by the degree of CBI is of interest by itself, but the results also allow for more satisfactory statistical inference in regression models in which CBI is an explanatory variable. In contrast to the results of the influential study by Campillo and Miron (1997), we find that the CBI indicator we have constructed is significantly related to inflation, also when various control variables as suggested by Campillo and Miron are included. Finally, the CALS estimator has been used to deal with problems concerning measurement error and underestimation of the coefficient of CBI. Use of this estimator leads to a reasonable increase of the estimated coefficient, while the significance of the results remains intact. We have also created measures of CBI for other time periods.

In chapter 4, we have examined whether central banks can be held responsible for political business cycles in monetary policy outcomes. Using short-term interest rates to measure the monetary policy stance, we have tested whether central banks in 14 OECD countries have indeed created PBCs and whether the degree of CBI is crucial to prevent that. We have established evidence from two outcomes. First, we have conducted country-specific tests in which short-term interest rates are regressed on election dummies, variables describing national and international constraints and a number of control variables. The results of these tests show hardly any support for the PBC hypothesis. The second source of evidence stems from panel data regressions. After determining the correct model specification using exact tests to test for dynamics and unit root behaviour, we find that the estimation results provide more or less the same picture as the country-specific results. There is no evidence that central banks actively create PBCs. Overall, the implications are clear. If PBCs in macroeconomic variables such as unemployment show up, then the central banks should not be blamed. Either their actions have no effect, or they simply do not engage in short-sighted behavior.

In chapter 5, two new indicators for the institutional settings of the labour market in the 1970s and early 1980s have been constructed using the information provided by indicators for corporatism that have been suggested in the literature. Following a latent variables approach, two indicators that we label *coordination* and *organizational power of labour* have been obtained. For illustrative purposes, these measures are applied in the models of Hall and Franzese for inflation and unemployment. In contrast to their results, we find no evidence that interaction effects between the level of central bank independence and coordination play an important role with respect to unemployment. Also, the impact of organizational power of labour on inflation and unemployment, which has been reported in various previous studies, is not found in our regressions.

Chapter 6 returns to the conservativeness of central banks. The degree of conservativeness is measured using the rate and variability of inflation, variability of output and the conservativeness component that was isolated from one of the CBI indicators in chapter 3. For a sample of 14 European countries it is shown that, unsurprisingly, the rate of inflation is the most important indicator of conservativeness. The indicator of Cukierman, which has been designed specifically to measure conservativeness, does not seem to be of much use. Also, when results for the 1980s and 1990s are compared, the analysis shows that there have not been major changes in the degree of conservativeness. Further, the impact of conservativeness on unemployment has been examined using the newly constructed measure. Only if trade openness is included as a control variable in the model, the coefficient of conservativeness is signifcant. Finally, CALS estimators that correct for the underestimation of the conservativeness coefficient have been used. This leads to a reasonable increase of the (absolute) coefficient values, while the (in)significance of the results is unaffected.

In chapter 7, finally, the relationship between economic freedom and economic growth is assessed. Instead of using a measure of economic freedom based on an (ad hoc) aggregation of various underlying components, as is often the case in the literature, we construct a measure of economic freedom using principal components analysis. We argue that the alternative aggregation procedure as recently suggested by Heckelman and Stroup (2000) - in which aggregation is directly based upon the relevance of each component for growth, as determined by multivariate regression analysis - has serious shortcomings. Using standard robustness analyses, we find that the level of economic freedom is not robustly related to economic growth.

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Samenvatting (Summary in Dutch)

Een groot deel van de variabelen die worden gebruikt in economische modellen zijn theoretische constructies. Het is vaak moeilijk om hieraan getalswaarden toe te kennen, aangezien ze niet direct waarneembaar zijn. Dit type variabelen wordt latente variabelen genoemd. Andere variabelen, die wel kunnen worden waargenomen, zijn dan nodig om te dienen als indicatoren voor de niet waarneembare variabelen. Het is niet altijd duidelijk welke variabelen geselecteerd zouden moeten worden voor dit doel, en bovendien spelen subjectieve oordelen een belangrijke rol in de selectie- en aggregatieprocedures. Als gevolg hiervan creëren en / of gebruiken verschillende onderzoekers verschillende indicatoren, wat mogelijk leidt tot verschillende resultaten en conclusies.

Het eerste doel van dit proefschrift is de ontwikkeling van een valide meetinstrument voor een aantal belangrijke variabelen op het gebied van de politieke economie. In plaats van nog meer indicatoren te ontwikkelen, wordt de informatie gebruikt die is opgeslagen in reeds bestaande indicatoren, en wordt een objectieve maatstaf ontwikkeld met behulp van aggregatieprocedures die geen subjectieve oordelen behelzen. Een tweede doel is de toepassing van deze nieuw geconstrueerde maatstaven in empirische modellen om de invloed van de betreffende variabelen op economische uitkomsten te bepalen. In het bijzonder is het interessant te onderzoeken of verbanden die in eerdere studies zijn afgeleid, gebruikmakend van bestaande indicatoren, overeind blijven.

Onze focus ligt op vier belangrijke variabelen op het gebied van de politieke economie: centrale bank onafhankelijkheid, centrale bank conservatisme (dat wil zeggen de mate waarin een centrale bank een afkeer heeft van inflatie), corporatisme en economische vrijheid. Onderzocht wordt hoe deze variabelen kunnen worden gemeten met behulp van latente variabelen modellen. Hierbij worden methoden van factoranalyse (FA) en principale componenten analyse (PCA) toegepast. Alvorens de specifieke eigenschappen van deze methoden te bespreken, worden de variabelen nader bekeken.

De vier genoemde variabelen zijn macro-economische variabelen, dat wil zeggen dat ze worden gemeten op nationaal niveau. Ten eerste moet worden bepaald tot welke verzameling landen we het onderzoek beperken. Aangezien er grote institutionele verschillen bestaan tussen ontwikkelde en ontwikkelingslanden kunnen meetmethoden vaak niet worden toegepast op beide groepen landen. Als we toch beide groepen willen bekijken, moeten dus verschillende latente variabelen modellen worden gecreëerd voor ontwikkelde en ontwikkelingslanden. De beschikbaarheid van empirische gegevens voor ontwikkelingslanden is echter vaak problematisch, zeker in eerdere tijdsperioden. Derhalve hebben we besloten ons voornamelijk te richten op de groep ontwikkelde landen. Slechts bij het meten van economische vrijheid worden ontwikkelingslanden ook in beschouwing genomen. De overige studies worden uitgevoerd voor landen zijn aangesloten bij de OESO.

Beperking van de verzameling landen tot de groep waarvoor gegevens goed beschikbaar zijn, betekent echter niet dat alle potentiële problemen met betrekking tot meetbaarheid zijn ondervangen. Ten eerste is de grootte van de steekproef een probleem. Door ons te richten op geïndustrialiseerde landen beperken we ons tot een kleine groep. In 1980 telde de OESO 23 leden. In econometrische termen is dit al een kleine steekproef. Gezien het feit echter dat voor een aantal landen gegevens van de relevante indicatoren niet beschikbaar zijn, worden de meeste studies in dit proefschrift uitgevoerd voor een nog kleinere groep landen. Voor het meten van corporatisme wordt bijvoorbeeld een steekproef van 16 landen gebruikt, terwijl conservatisme van de centrale bank wordt gemeten met behulp van een steekproef van slechts 14 landen. We zullen aandacht moeten besteden aan de economische consequenties van het schatten van modellen met dergelijke kleine steekproeven, en hiervoor corrigeren indien nodig. Een tweede, gerelateerd, probleem is het gedeeltelijk ontbreken van gegevens. Het gebeurt dikwijls dat, naast het volledig niet beschikbaar zijn van gegevens, voor bepaalde landen gegevens slechts beschikbaar zijn voor een gedeelte van de periode waarop de studie betrekking heeft. Dit probleem zou kunnen worden opgelost door landen en / of tijdsperioden volledig uit te sluiten, wat echter tot gevolg heeft dat veel waardevolle informatie wordt weggegooid. In plaats hiervan kan ook worden gezocht naar manieren om de informatie die wel beschikbaar is volledig te benutten. Op verschillende plaatsen worden dergelijke procedures toegepast om ervoor te zorgen dat de toch al kleine steekproeven niet nog verder worden ingeperkt.

Behalve het meten van een aantal belangrijke variabelen op het gebied van de politieke economie worden in dit proefschrift de resulterende maatstaven toegepast in empirische modellen, waarin hun invloed wordt onderzocht op economische uitkomsten en monetaire instrumenten. Zo onderzoeken we de invloed van centrale bank onafhankelijkheid op inflatie en het optreden van *political business cycles* in korte-termijnrentes, de invloed van centrale bank conservatisme op werkloosheid, de invloed van (aspecten van) corporatisme op inflatie en werkloosheid en de invloed van economische vrijheid op economische groei. De resultaten van deze studies worden vergeleken met resultaten die zijn gemeld in de empirische literatuur. Een probleem dat ontstaat wanneer latente variabelen worden gebruikt als regressor in een regressiemodel is dat de betreffende coëfficiënt wordt onderschat. We erkennen dit probleem en beschrijven een methode om deze onderschatting kwijt te raken.

Hoofdstuk 2 geeft een overzicht van de methodologie die wordt gebruikt in de rest van het proefschrift. Hoofdstuk 3 beschrijft hoe centrale bank onafhankelijkheid kan worden gemeten, terwijl in hoofdstuk 4 de vraag wordt beantwoord of centrale banken verantwoordelijk kunnen worden gehouden voor het optreden van *political business cycles*. Dit zijn golfbewegingen die het resultaat zijn van manipulatie van economische politiek door politici met het oog op het vergroten van de kans op herverkiezing. In hoofdstuk 5 staat het meten van corporatisme centraal, en in hoofdstuk 6 wordt het meten van centrale bank conservatisme besproken. Hoofdstuk 7 analyseert welke economische vrijheden een bijdrage leveren aan economische groei. In hoofdstuk 8 worden tenslotte de resultaten samengevat.

Overzicht en resultaten

Het grootste deel van hoofdstuk 2 betreft een bespreking van latente variabelen modellen. Modellen met één of meer latente variabelen worden gepresenteerd in de vorm van FA- en PCA-specificaties. Verdere onderwerpen van discussie zijn welk model dient te worden geselecteerd, hoe dit kan worden geschat, hoe goed de modelspecificatie voldoet en hoe de resultaten kunnen worden geïnterpreteerd. Tevens wordt de CALS-schatter besproken, die de onderschatting van de coëfficiënt van de latente variabele in empirische toepassingen tenietdoet. Naast latente variabelen modellen worden in dit proefschrift ook modellen gebruikt waarin dynamiek een rol speelt. Hoofdstuk 2 besluit met een bespreking van een toets die niet gebaseerd is op asymptotische overwegingen om te toetsen op het voorkomen van dynamiek en/of *unit roots* in panel data modellen.

Hoofdstuk 3 beschrijft het meten van centrale bank onafhankelijkheid. Nadat de concepten onafhankelijkheid en inflatie-aversie (conservatisme) van elkaar gescheiden zijn, wordt een indicator voor centrale bank onafhankelijkheid geconstrueerd met behulp van een latente variabelen model dat vijf indicatoren voor onafhankelijkheid uit de literatuur gebruikt. Enkele technische complicaties van het FA-model voor centrale bank onafhankelijkheid worden besproken, en de resulterende maatstaf wordt toegepast in enkele empirische modellen om de relatie tussen centrale bank onafhankelijkheid en inflatie te onderzoeken. In tegenstelling tot de invloedrijke studie van Campillo en Miron (1997) vinden we een significant verband tussen onze maatstaf voor centrale bank onafhankelijkheid en inflatie, ook wanneer controlevariabelen worden opgenomen. Toepassing van de CALS-schatter om onderschatting aan te pakken, leidt tot een aanzienlijke toename van de geschatte coëfficiënt van centrale bank onafhankelijkheid, terwijl de significantie van het verband overeind blijft.

In hoofdstuk 4 wordt de invloed van de zojuist geconstrueerde maatstaf voor centrale bank onafhankelijkheid gebruikt in een model dat onderzoekt of centrale banken op actieve wijze *political business cycles* (PBCs) in monetaire beleidsuitkomsten creëren. Na een korte bespreking van de PBC-theorie worden verschillende cross-sectie modellen geformuleerd waarin korte-termijnrentes worden geregresseerd op verkiezingsgegevens, variabelen met betrekking tot nationale en internationale beperkingen en een aantal controlevariabelen. De resultaten van deze regressies geven weinig aanleiding te geloven dat centrale banken actief PBCs creëren. Tevens bekijken we een panel data model, waarbij de exacte toetsen uit hoofdstuk 2 worden gebruikt om te toetsen op dynamiek en *unit roots*. De resultaten van de panelregressies schetsen hetzelfde beeld. Wanneer PBCs in economische uitkomsten zoals werkloosheid voorkomen, moet de schuld daarvoor derhalve niet bij de centrale banken worden gezocht.

In hoofdstuk 5 worden twee maatstaven voor de institutionele organisatie van de arbeidsmarkt geconstrueerd, gebruikmakend van een groot aantal indicatoren uit de politieke en economische literatuur over corporatisme. Na selectie van de indicatoren die geschikt zijn voor gebruik in een latente variabelen analyse, en de uitvoer hiervan, worden de twee resulterende maatstaven toegepast in de empirische modellen van Hall en Franzese (1998) die de relatie onderzoeken tussen (aspecten van) corporatisme en inflatie en werkloosheid. De maatstaf voor centrale bank onafhankelijkheid uit hoofdstuk 3 wordt tevens gebruikt in deze modellen. In tegenstelling tot Hall en Franzese, en een aantal andere eerdere studies, vinden we geen enkel verband tussen institutionele organisatie van de arbeidsmarkt en werkloosheid.

Hoofdstuk 6 keert terug naar de mate van conservatisme van centrale banken. Conservatisme wordt gekwantificeerd met behulp van FA-modellen voor twee verschillende decennia, waarbij inflatie de belangrijkste indicator voor conservatisme blijkt. Tevens wordt de invloed van conservatisme op werkloosheid onderzocht, waarbij slechts een significant verband wordt gevonden wanneer *trade openness* als controlevariabele in het model wordt opgenomen. Toepassing van de CALS-schatter leidt wederom tot een toename van de geschatte coëfficiënt maar beïnvloedt de significantie niet.

Hoofdstuk 7 beschrijft het meten van economische vrijheid voor een groot aantal ontwikkelde en ontwikkelingslanden. Economische vrijheid is hierbij gedefinieerd als de mate waarin op legale wijze verkregen bezit wordt beschermd en mensen vrij zijn om vrijwillige transacties met elkaar aan te gaan. Met behulp van PCA-methoden wordt een maatstaf voor economische vrijheid afgeleid en vergeleken met twee alternatieven uit de literatuur. We laten zien dat de alternatieve aggregatieprocedure van Heckelman en Stroup (2000) ernstige tekortkomingen vertoont, en vinden, in tegenstelling tot hun resultaten, geen robuust verband tussen de mate van economische vrijheid en economische groei.

Hoofdstuk 8 besluit met een samenvatting van de resultaten en conclusies.

Theses on Systems,
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