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The relation between unemployment and interest rate:
Some empirical evidence

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THE RELATION BETWEEN UNEMPLOYMENT AND INTEREST RATE:

SOME EMPIRICAL EVIDENCE

by

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Abstract

In Bierens (1987) a Granger causal relation was found between unemployment and the interest rate for the Netherlands. This phenomenon could be explained by the revenue maximization theory of Baumol (1959) augmented with a flexible labour effort rate. This implies that fixed costs are important in determining the employment decisions of firms. In the present paper we will investigate whether there exists a similar Granger causal relation between unemployment and interest rate for a number of other countries. It appears that, with our ARMAX modeling approach, this relationship is not confined to the Netherlands, but also holds for the USA, Canada, Japan, Germany, the UK and France. For these countries the interest rate is the main explanatory variable, together with industrial production (the latter with one exception), whereas for most countries the wage rate is of minor or no importance as a determinant of unemployment.

November, 1990

1. INTRODUCTION

Unemployment is one of the main economic and social problems of our time. Since Keynes' general theory numerous studies have been conducted on the causes of, and remedies for unemployment. However, after the first oil crisis in the early seventies economists and policy makers became increasingly aware that the old Keynesian theories and policy recipes don't work anymore. This resulted, for example, in studies like Hall (1975), Malinvaud (1977, 1980), Grubb, Jackman and Layard (1983), Layard and Symons (1984) and Blanchard and Summers (1987), among others. Also many country specific studies were conducted, like for example Sneesens and Drèze (1986) for Belgium, Nickell (1982), Nickell and Andrews (1983), Symons (1985), Layard and Nickell (1986) and Pissarides (1986) for the U.K., Kirkpatrick (1982), Franz (1983) and Franz and König (1986) for Germany, Hamada and Kurosaka (1984, 1986) and Sorrentino (1984) for Japan, Driehuis (1978, 1986), Kuipers and Buddenberg (1978), and Bierens (1987), for the Netherlands, and Clark (1985), Farmer (1985, 1989) and Evans (1989) for the U.S.

At the theoretical level the debate about the causes for unemployment has mostly been concentrated on the rivalry between the (neo-)Keynesian the neo-classical point of view. In neo-classical theory unemployment is caused by too high a real wage rate; if the wages are adjusted the labor market will return to its equilibrium value again. Thus, this theory assumes that the labor market will clear, eventually. The neo-classical theory augments the classical pre-Keynesian theory with the rational expectations (RE)-natural rate (NR) hypothesis. The RE-hypothesis asserts that the expectations of economic agents about the future state of the economy are basically conditional expectations, relative to all available information. Consequently, economic agents are able to anticipate policy measures such that they become impotent (see, e.g., Sargent and Wallace (1975, 1976)). Regarding unemployment, the RE hypothesis, together with the classical hypothesis of clearing markets, leads to the conclusion that the labor market will be permanently in equilibrium. The persistence of high unemployment is therefore attributed to a rise in the natural rate of unemployment, (cf. Friedman (1968), Phelps (1970)), hence the high unemployment in the seventies and eighties is to a large extent voluntary! On the other hand, the Keynesians assume that unemployment is caused by too low an effective demand. If firms are rationed on the goods market, this might have consequences for the labor market as well. Thus, this

theory assumes that in general markets do not clear.

In a more or less inductive investigation, Bierens (1987) used an ARMAX model to test the RE-NR hypothesis for the Netherlands. It appeared that the interest rate was significantly Granger causing unemployment, hence the RE-NR hypothesis could not be accepted as it asserts that unemployment only depends on its own past (cf. Hall (1975), Sargent (1976)). This empirical relation between interest rate and unemployment had to be given a theoretical justification, as it could not be explained by any of the established economic theories. It appeared that the managerial theory of the firm developed by Baumol (1959), augmented with a flexible labor effort rate as described in Bierens (1987) could explain the observed phenomenon.

Baumol's (1959) theory asserts that managerial firms are revenue maximizers rather than profit maximizers. Profits only play a role as a constraint: a minimum profit level is required to safeguard the firms' continuity. A high interest rate implies more interest payments and hence more fixed costs. Therefore, if the interest rate is high enough it might happen that profit falls below its minimum level. Bierens (1987) augments this theory with the hypothesis of a flexible labor effort rate: it is not the actual amount of labor that determines the production level, but the labor effort. The idea involved is actually extremely simple: if workers work harder, they can produce more. The labor effort rate, however, is not fixed but may vary between zero and an upper bound. Now if the fixed costs of a firm increase such that profit falls below the profit constraint, the firm has to reorganize. Since by nature fixed costs are difficult to cut down, the firm will seek cost reduction in laying off workers, in particular those workers that can easily be replaced, and increase the labor effort rate of the remaining workers towards its upper bound. In case the labor effort rate is already on its maximum level, employment as well as production have to decrease until the minimum profit level is retained, and if this is not possible the firm has to close down. Moreover, an increase of the interest rate will likely reduce the demand for expensive durables that have to be financed by loans, and via the fixed costs of households like mortgage payments, rent, et cetera, also the demand for nondurables. An increase in the wage rate has a similar effect on the firm as an increase of the interest rate, but an opposite effect on demand, especially if the wage increase is due to collective bargaining. The increase of demand caused by the increase of the wage rate may therefore offset the negative effect on the profits of firms. Indeed, Bierens (1987)

found that only the interest rate has a significant (positive) effect on unemployment in the Netherlands, and that other variables such as wage and price inflation, and even the growth of industrial production had no significant impact.

The purpose of this paper is to test this unemployment-interest relationship for a number of other industrial countries as well, along the approach of Bierens (1987), in order to investigate whether this phenomenon is typically Dutch or of general nature. The countries to be considered are the United States, Canada, Japan, the Federal Republic of Germany, Great Britain, France and again the Netherlands. As in Bierens (1987), we shall work with monthly data.

As an appetizer, consider the figures 1 and 2 below. These figures suggest that this unemployment-interest rate relation also applies to the US. Comparing the unemployment series in figure 1 with the interest rate series in figure 2, we see that unemployment and the interest rate (lagged about $1\frac{1}{2}$ year) have a remarkably common pattern, particularly in the early seventies and eighties.

In section 2 we review the ARMAX modeling approach. Section 3 is devoted to a description and an analysis of the data. In section 4 the estimation and test results will be discussed and in section 5 we make some concluding remarks.

2. THE ARMAX MODELING APPROACH

2.1. Introduction

An ARMAX model is a model that explains a dependent variable out of lagged dependent variable and other (X-) variables plus a MA disturbance. In essence it is an ARMA model with additional explanatory variables. In this way it is a combination of an econometric model and a time series model. This implies that it represents the dynamic properties of the time series and on the other hand includes a role for economic theory. Regarding y_t as the dependent variable and $x_t \in \mathbb{R}^k$ as a vector of explanatory variables (possibly including seasonal dummy variables), our ARMAX model has the general form

$$\phi(L)y_t = \mu + \alpha(L)x_t + \theta(L)\theta(L)\varepsilon_t, \quad (2.1.1)$$

where μ is the constant and

$$\phi(L) = 1 - \sum_{j=1}^p \phi_j L^j, \quad \alpha(L) = \sum_{j=1}^r \alpha_j L^j,$$

Moreover, we assume a multiplicative MA error structure with

$$\theta(L) = 1 + \sum_{j=1}^q \theta_j L^j, \quad \Theta(L) = 1 + \sum_{j=1}^Q \Theta_j L^{sj},$$

where s indicates the seasonality, i.e., $s=4$ in case of quarterly data, $s=12$ in case of monthly data. L is the usual lag operator, i.e., $L^j z_t = z_{t-j}$, and the ε_t 's are the residuals, which are assumed to be martingale differences. If the ARMAX model is to be stationary, $\phi(L)$ should have all its roots outside the unit circle and the process generating x_t has to be stationary. Similarly the ARMAX model is invertible if $\theta(L)\Theta(L)$ has all its roots outside the unit circle.

If the ARMAX model is invertible it can be written as an ARX(∞) model, as can be derived easily from (2.1.1).

$$\begin{aligned} \varepsilon_t &= \frac{\phi(L)}{\theta(L)\Theta(L)}y_t - \frac{\alpha(L)}{\theta(L)\Theta(L)}x_t - \frac{1}{\theta(L)\Theta(L)}\mu = \\ &= y_t - \frac{\phi(L)}{\theta(L)\Theta(L)}y_t - \frac{\alpha(L)}{\theta(L)\Theta(L)}x_t - \frac{1}{\theta(L)\Theta(L)}\mu, \end{aligned}$$

where $\Phi(L) = \phi(L) - \theta(L)\Theta(L)$. Hence, since $\Phi(L)/L$ and $\alpha(L)/L$ are finite lag polynomials, we can write

$$y_t = \frac{\mu}{\theta(1)\Theta(1)} + \frac{\Phi(L)/L}{\theta(L)\Theta(L)}y_{t-1} + \frac{\alpha(L)/L}{\theta(L)\Theta(L)}x_{t-1} + \varepsilon_t$$

Consequently, denoting

$$\begin{aligned} z_t &= (y_t, x_t)' \\ \beta &= (\phi_1, \dots, \phi_p, \alpha'_1, \dots, \alpha'_r, \theta_1, \dots, \theta_p, \Theta_1, \dots, \Theta_Q)' \\ \mu(\beta) &= \frac{1}{\theta(1)\Theta(1)}\mu \\ \sum_{j=1}^{\infty} \eta_j(\beta)L^j &= \frac{1}{\theta(L)\Theta(L)} \begin{bmatrix} \Phi(L)/L \\ \alpha(L)/L \end{bmatrix} \end{aligned}$$

we can write ARMAX model (2.1.1) as an ARX(∞) model

$$y_t = \mu(\beta) + \sum_{j=1}^{\infty} \eta_j(\beta)' z_{t-j} + \varepsilon_t \quad (2.1.2)$$

Thus we see that an ARMAX model allows for an infinite lag structure with a parsimonious parameterization.

The ARMAX specification enables us to test for Granger causality in a straightforward way. The concept of Granger causality implies that we are better able to predict the dependent variable using both its past and the past of the variables that are Granger causing the dependent variable, than merely the past of the dependent variable itself. Hence if one of the components of the parameter vector α' differs significantly from zero, then the corresponding X-variable is Granger causing y_t (cf. Granger (1969)).

Our ARMAX modeling strategy is in the same spirit as the dynamic modeling strategy of Hendry and Richard (1982, 1983). The main difference is that we start with a general ARMAX model, whereas they start with an ARX model with a number of lags for the variables.

2.2. Specification of the ARMAX model

It is common practice in present day econometric methodology to start with a fairly general model, capturing as much information as possible, and submit it to a number of model specification tests to see whether the model is adequate.

If the model specification is accepted, we may then simplify this model. On the other hand, if this initial model is rejected its specification should be changed, for example by including more lagged dependent or lagged explanatory variables, and then the new specification has to be tested again. This approach of moving from a general to a simple model, instead of moving from a simple to a more general model, is advocated by e.g. Hendry and Mizon (1978) and Mizon and Hendry (1980).

Our initial model specification aims to capture the dynamic properties of the economic time series under review as much as possible. On the other hand, too rich a parameterization may cause estimation and identification difficulties. These demands can be satisfied by imposing a MA and seasonal (S) MA structure on the residuals of the model and by taking only the first lag of both the dependent and the explanatory variables into account. In this way the dynamics of the time series is captured by the MA and SMA structure of the residuals. We have already argued that the (S)MA part of the model allows an ARMAX model to be written as an ARX(∞) model. Hence, beginning with a lag of one period for y_t and x_t and a MA and SMA structure to capture the dynamics is equivalent to an infinite distributed lag model. Another reason for taking a lag of only one period of the X-variables is to avoid multicollinearity. Thus we set $p=r=1$ in model (2.1.1). Moreover, the initial specification of the MA part in the model is $q=6$, $Q=3$, which hopefully captures the dynamic structure in the economic time series as much as possible. Thus, our initial model becomes

$$y_t = \mu + \phi_1 y_{t-1} + \alpha_1' x_{t-1} + (1 + \theta_1 L + \theta_2 L^2 + \dots + \theta_6 L^6)(1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36}) \varepsilon_t \quad (2.2.1)$$

If this initial model is not accepted by the model specification tests (see section 2.4.3 and 2.4.4), we adjust it according to the information provided by the autocorrelation and partial autocorrelation function of the residuals of the model. The latter indicates whether more lagged dependent variables should be included in the model (cf. Box and Jenkins (1976)). A change in the lag of the explanatory variables is not considered, because there is no straightforward mechanism which can determine what this lag should be. Moreover, as is already argued, inclusion of a MA and SMA structure allows for an infinite number of lags on y_t and x_t . Another approach in case this initial specification is rejected is to change the functional form of the

model, e.g., by taking the logarithm of the dependent variable.

As usual the choice of the X-variables in the model is based on economic theory as well as on availability. Also the transformations of the variables in order to attain stationarity should in first instance be based on economic theory. If economic theory does not prescribe a specific transformation, intuitive ideas about the logical consistency of the variables in combination with exploration of their time series properties are used to determine the correct transformation.

Next we discuss how these time series properties are investigated. We start with the unit root test developed by Phillips and Perron (1988) and we also apply a stationarity test developed by Bierens (1990a). These tests enable us to determine the order of integration without using the sample autocorrelation function. As shown by Hasza (1980) and Bierens (1990b) the sample autocorrelation function can be a misleading indicator of a unit root. The Phillips-Perron test is particularly useful in case of more complicated time series models, such as ARMAX models, because it does not require to estimate or even identify the model in order to test for a unit root. An advantage of this test over the unit root test developed by Phillips (1987) is that the Phillips-Perron test considers a more general alternative, namely stationarity with a possibly nonzero mean rather than zero-mean stationarity. Specifically when a time series has a high nonzero mean (relative to the error variance) the Phillips test performs very poorly. Cf. Bierens (1990a).

Stationarity test number III, developed by Bierens (1990a) also has a number of advantages, compared to the commonly applied unit root tests. First, it has a known distribution under the null hypothesis of a stationarity, namely a standard Cauchy distribution. Recent studies like the ones of Schwert (1989) and Bierens (1990a) have indicated that unit root tests should not be applied automatically. Especially in case of a near-unit root, the size of the commonly applied unit root tests is far out of tune with the theoretical size. Thus these tests do not do a very good job in distinguishing a near-unit root from a genuine one. The size and power properties of the test developed by Bierens (1990a) are, however, better as compared to the much applied Phillips (1987) and Phillips and Perron (1988) tests, especially in case of a near-unit root. Cf. Bierens (1990a).

From studies of Perron (1989) and Hendry and Neale (1989) it is known that unit root tests are sensitive to structural breaks. By looking at the plot for unemployment, the dependent variable in our models, we might infer

that indeed there are structural breaks in unemployment in some countries and that if we correct the series for these breaks they will show stationary behavior. However, these structural breaks may just correspond to structural breaks in other variables, such as inflation and the interest rate, so that by correcting for these structural breaks by means of dummy variables we would just pour out the baby with the bath water. Moreover, it is also not clear how many 'known' structural breaks should be modeled. It appears that the specification of a model, and hence the outcome of the procedure of Perron (1989), depends on the way structural breaks are represented by the dummies, as was shown by Broersma and Franses (1990).

In case the data have a short time span, e.g., monthly or even weekly or daily series, it appears from a study of Schiller and Perron (1985) that the power of unit root tests is low. All these issues imply that we should not only rely on the results of unit root tests to determine stationarity. Intuitive ideas about the logical consistency of the variables may serve as an important alternative. Examination of the sample autocorrelation function of the variables can also be an additional check on stationarity, although the results of Hasza (1980) and Bierens (1990b) indicate that the sample autocorrelation function is not conclusive in indicating a unit root. Also the plots of the time series involved may give additional information about possible stationarity of the time series, as they might reveal possible structural breaks or large outliers.

Apart from unit root tests, intuitive ideas, correlograms and plots, we can also look at the value of the AR coefficient. If this value is very close to unity, we will set the AR parameter equal to one, even if for reasons of logical consistency, the variable cannot contain a unit root. In that case we approximate the near-unit root by a unit root. Thus, in terms of initial model (2.2.1), if $\phi_1 \approx 1$, then we change the specification to

$$\Delta_1 y_t = \mu + \alpha_1' x_{t-1} + (1 + \theta_1 L + \theta_2 L^2 + \dots + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36}) \varepsilon_t, \quad (2.2.2)$$

where $\Delta_k z_t = z_t - z_{t-k}$, $k \in \mathbb{N}$.

In case of a higher order AR structure, we look at the value of the sum of the AR coefficients. If this sum gets close to unity, this implies the presence of a near-unit root. In this case we repeat the whole specification stage, but now with the dependent variable set equal to $\Delta_1 y_t$. The new model

specification then becomes

$$\Delta_1 y_t = \alpha_0 + \alpha_1 \Delta_1 y_t + \alpha_1' x_{t-1} + (1 + \theta_1 L + \theta_2 L^2 + \dots + \theta_6 L^6)(1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36}) \varepsilon_t. \quad (2.2.3)$$

If this specification is not accepted, we try to repair it in the same way as in the case of model (2.2.1).

2.3. Estimating the ARMAX model

Estimation of the parameters of our model is conducted by nonlinear least squares (NLS), with the pre-sample data set equal to the mean value of the corresponding observable values. As we already mentioned earlier, ARMAX model (2.1.1) can be written as an ARX(∞) model if the MA polynomial is invertible. We can then write (2.1.2) as

$$y_t = g_t(\beta) + \varepsilon_t, \quad (2.3.1)$$

The nonlinear least squares estimator $\hat{\beta}$ of β is a solution of

$$Q(\hat{\beta}) = \inf_{\beta \in B} \frac{1}{n} \sum_{t=1}^n (y_t - g_t(\beta))^2,$$

where B is the parameter space. The consistency and asymptotic normality of this estimator are proved by Bierens (1988).

2.4. Testing the ARMAX model

Once the ARMAX models for the various countries have been estimated, they are being subjected to a number of diagnostic tests. We apply four kinds of tests. First we test if the errors of our models are normally distributed, and then we test for autoregressive conditional heteroskedasticity (ARCH). However, it should be stressed that neither nonnormality nor ARCH exclude model correctness. These tests are merely conducted to provide additional information about the data generating process. Next we test whether the residuals are serially correlated and we test the functional specification of the ARMAX model. The

most crucial tests are the autocorrelation and the specification tests. The latter is in fact a direct and consistent test of the hypothesis of the residuals being a mean innovation process (MIP) in the sense of Hendry and Richard (1982). There the importance of white noise residuals and the residuals being a MIP is also pointed out. This model specification test is consistent, which implies that as the sample size grows to infinity, any misspecification will be detected. That is one of the reasons why we have used monthly data, because then the asymptotic properties are (hopefully) better approximated.

If the specification of a model can no longer be rejected, we use a Wald test on parameter restrictions to simplify the model by deleting variables with insignificant parameters.

2.4.1. *The normality test*

To test the null hypothesis of normally distributed errors, we apply the test developed by Kiefer and Salmon (1983). This is a joint test of the hypotheses that the skewness equals zero and the kurtosis equals three. This test is based on the fact that under the null hypotheses the random vector

$$\hat{b} = \left[\frac{1}{\sqrt{n}} \sum_{t=1}^n \hat{\varepsilon}_t^3, \frac{1}{\sqrt{n}} \sum_{t=1}^n (\hat{\varepsilon}_t^4 - 3(E(\hat{\varepsilon}_t^2))^2) \right];$$

with $\hat{\varepsilon}_t = y_t - g_t(\hat{\beta})$, is asymptotically $N(0, \Psi)$ distributed with covariance matrix

$$\Psi = \begin{pmatrix} \sigma^6(15 - 9d'\Omega d) & 0 \\ 0 & 24 \end{pmatrix},$$

where $d' = E(\partial \varepsilon_t / \partial \beta')$ and $\Omega = E(\partial \varepsilon_t / \partial \beta')(\partial \varepsilon_t / \partial \beta)$ and β is the vector of model parameters. A Wald test statistic T_{norm} can now be constructed, which has a $\chi^2(2)$ distribution under the null:

$$T_{norm} = n\hat{b}'\hat{\Psi}^{-1}\hat{b}, \quad (2.4.1)$$

where $\hat{\Psi}$ is based on the corresponding estimates $\hat{\sigma}$, $\hat{d}$ and $\hat{\Omega}$.

In standard econometrics the model errors are often assumed to be

normally distributed, in order to accommodate application of maximum likelihood estimation. The only requirement for model correctness, however, is that the errors are martingale differences with respect to the σ -algebra generated by past y and X variables. Moreover, we do not need the normality assumption, because we estimate the model by least squares and asymptotic normality results for the parameter estimators involved can be derived on the basis of the central limit theorem for martingale difference sequences due to McLeish (1974). On the other hand, the asymptotic approximations will likely be better if the errors are normally distributed. Anyhow, for relatively large samples the outcome of the normality test is not crucial. Since ARCH is compatible with the martingale difference hypothesis, the same applies to ARCH.

2.4.2. *The ARCH test*

Traditionally the errors in econometric time series models are assumed to be homoskedastic, that is, their conditional variance, relative to all past y and X variables, is constant. The assumption of heteroskedastic errors has mainly been considered in the context of cross-section data. An important exception is Engle's (1982) autoregressive conditional heteroskedasticity (ARCH) hypothesis, with corresponding test. This test is conducted by regressing the squared residuals $\hat{\varepsilon}_t^2$ on a constant and on p lagged values of $\hat{\varepsilon}_t^2$. Under the null hypothesis of homoskedastic errors, the test statistic nR^2 , where n is the number of observations and R^2 is the coefficient of determination of the regression, follows a $\chi^2(p)$ distribution.

In case the disturbances show an ARCH structure, we should take this into account if one wish to estimate the model efficiently by maximum likelihood. However, in practice this requires distributional assumptions, like normality, about the errors that we are not willing to make a priori. Heteroskedasticity can however not be completely ignored as it may lead to inconsistent covariance matrix estimators, and tests based on that matrix are then no longer valid. To avoid all these complications we apply the heteroskedasticity-consistent covariance estimator proposed by White (1980), so that our test statistics remain asymptotically valid. This implies that also the ARCH test is not crucial for the model specification. We do however report the ARCH test results, because they give some information about the data generating process.

2.4.3. Some autocorrelation tests

We next apply a number of autocorrelation tests, which test the null hypothesis of no serially correlated errors. It is not our intention to test against a specific alternative. That is why the familiar LM test of Godfrey (1979) is not applied. This test also relies heavily on the assumption of independent and normally distributed errors (cf. Godfrey *et.al.* (1988)). Since we do not have specific alternatives, our null hypothesis

$$H_0: \text{cov}(\varepsilon_t \varepsilon_{t-k}) = 0 \quad \text{for } k \in \{1, 2, \dots, s\} \quad (2.4.2)$$

is tested against the general alternative hypothesis

$$H_1: \text{cov}(\varepsilon_t \varepsilon_{t-k}) \neq 0 \quad \text{for a } k \in \{1, 2, \dots, s\}. \quad (2.4.3)$$

First we use the well-known portmanteau test of Box and Pierce (1970) and the modified test of Ljung and Box (1978). These tests are based on the autocorrelation function of the disturbances of the model

$$\hat{\rho}_k = \frac{\sum_{t=k+1}^n \hat{\varepsilon}_t \hat{\varepsilon}_{t-k}}{\sum_{t=1}^n \hat{\varepsilon}_t^2}, \quad k=1, 2, \dots$$

where $\hat{\varepsilon}_t = y_t - g_t(\hat{\beta})$. The Box-Pierce test statistic is defined as

$$T_{B-P}(s) = n \sum_{k=1}^s \hat{\rho}_k^2 \quad (2.4.4)$$

and the Ljung-Box test statistic as

$$T_{L-B}(s) = n(n+2) \sum_{k=1}^s (n-k) \hat{\rho}_k^2 \quad (2.4.5)$$

Both T_{B-P} and T_{L-B} have a $\chi^2(s-p-q-Q)$ distribution under the null, where s is the order of autocorrelation and p , q and Q are the number of AR, MA and seasonal MA parameters.

A disadvantage of these tests is that, although the size of the Ljung-Box test appears to be robust to departures from normality in case of AR(1) disturbances (cf. Ljung and Box (1978)), the robustness in case of higher order AR or MA disturbances is unknown. Also the power of these two tests is

questionable as can be concluded from a great number of studies, like Davies and Newbold (1979), Clarke and Godolphin (1983) and Hall and McAleer (1989), especially if the model contains lagged dependent variables (cf. Kiviet (1986)).

Because of these drawbacks we also apply a test which does not depend on the assumption of normally distributed and independent errors. It only assumes the errors to be a martingale difference sequence. This test is a straightforward further elaboration of the autocorrelation test of Godolphin (1980), cf. Bierens (1988), so that it can now be applied to ARMAX models. The test is based on the statistic

$$\hat{c} = \frac{1}{n-s-p} \sum_{t=s+p+1}^n \hat{\varepsilon}_t \hat{\vartheta}_t, \quad (2.4.6)$$

where s is the order of autocorrelation, p is the number of AR parameters in the model and $\hat{\vartheta}_t = (\hat{\varepsilon}_t, \hat{\varepsilon}_{t-1}, \dots, \hat{\varepsilon}_{t-s})'$ and $\hat{\varepsilon}_t = y_t - g_t(\hat{\beta})$. The statistic $\hat{c}$ is asymptotically $N(0, \Delta)$ distributed. The actual test statistic is

$$T_{ac}(s) = (n-s-p) \hat{c}' \hat{\Delta}^{-1} \hat{c}, \quad (2.4.7)$$

where $\hat{\Delta}$ is a consistent estimator of Δ (cf. Bierens (1988)). $T_{ac}(s)$ has a $\chi^2(s)$ distribution under the null. The power properties of this test can be compared with the ones of the test of Godolphin (1980), which are studied by Clarke and Godolphin (1983). It is especially this last test that is of importance for the specification of our models.

An important difference between this test, $T_{ac}(s)$, and the Box-Pierce and Ljung-Box tests, is that this test uses the exact calculation of the covariance matrix Δ , whereas the other two use approximations. From the power study of Clarke and Godolphin (1983) it appeared that $T_{ac}(s)$ is superior as compared to the Box-Pierce and Ljung-Box tests. A disadvantage of using the estimated Δ in the test statistic is the fact that $\hat{\Delta}$ has to be nonsingular. Near-singularity of $\hat{\Delta}$ can be recognized if $T_{ac}(s)$ takes on an excessively high value, while the plots of the autocorrelation function do not correspond to this value in the sense that autocorrelations move closely about zero. Hence, examination of the correlograms of the residuals is of great importance.

2.4.4. A consistent model specification test

The model specification test we apply, is the consistent ARMAX model specification test of Bierens (1987). This test is consistent in the sense that any misspecification will be detected as the sample size goes to infinity. This test is especially developed for testing the ARMAX specification. The null hypothesis to be tested is

$$H_0: E[\varepsilon_t | (y_{t-1}, x_{t-1}), (y_{t-2}, x_{t-2}), \dots] = 0 \quad a.s. \quad \forall t. \quad (2.4.8)$$

This test amounts to testing whether the response function of the ARX(∞) representation of our ARMAX model equals the conditional expectation of y_t relative to the entire past of the vector time series process under review, or in other words, to testing first-order model correctness in the sense of Domowitz and White (1982). Hence (cf. (2.3.1))

$$H_0: g_t(\beta) = E[y_t | (y_{t-1}, x_{t-1}), (y_{t-2}, x_{t-2}), \dots] \quad a.s. \quad \forall t.$$

The alternative hypothesis is that H_0 does not hold. Note that accepting H_0 implies that the errors of the model are a mean innovation process (MIP) in the sense of Hendry and Richard (1982).

This test does not assume that the errors of model (2.1.1) are Gaussian white noise nor that they are homoskedastic (cf. Bierens (1987)). This consistent model specification test is a more general test than T_{ac} . The specification test $T_{BN(01)}$ tests whether

$$E(\varepsilon_t | z_{t-1}, z_{t-2}, \dots) = 0 \quad a.s.,$$

which implies

$$E(\varepsilon_t | \varepsilon_{t-1}, \varepsilon_{t-2}, \dots) = 0 \quad a.s. \quad (2.4.9)$$

If (2.4.8) can not be rejected, then also (2.4.9) and hence (2.4.2) can not be rejected. So we may use T_{ac} as a pretest of model misspecification as severe misspecification in our model will likely be covered by (2.4.2). If a model exhibits autocorrelated disturbances according to T_{ac} , it is no longer necessary to conduct the model specification test $T_{BN(01)}$.

For the exact form of the test statistic $T_{BN(01)}$ we refer to Bierens (1987). This test statistic depends on nuisance parameters. In conducting this test we replace these nuisance parameters by random drawings from a continuous distribution. Under the null hypothesis this statistic converges in distribution to the $N(0,1)$ distribution. As argued in Bierens (1987), it is better to conduct this test several times for different sets of random drawings of the nuisance parameters involved. Therefore we shall run this test 20 times. The autocorrelation test T_{ac} and this consistent model specification test are the two most crucial tests concerning the specification of our ARMAX models. According to Hendry and Richard (1982) models using past information should have at least errors that are white noise and a MIP, and that just is what we want to verify with these two tests.

If an ARMAX model does not pass these two tests, we have to adjust our specification, say by including more lagged dependent variables, or by changing the functional form of the model (see section 2.2).

2.4.5. A test on parameter restrictions

If our specification is accepted by T_{ac} and $T_{BN(01)}$, then we may simplify our model, if possible, by deleting some variables from the model. It is not enough to simply look at the t -values of the parameters to simplify the model. There might be multicollinearity between variables resulting in low t -values, while these variables should not be deleted from the model jointly. To overcome this problem we apply the Wald test of joint parameter restrictions. In particular, we use the Wald test to test whether some of the parameters of the model are jointly equal to zero

$$H_0: \beta_* = 0, \quad (2.4.1)$$

where β_* is a subvector of model parameters. The test is based on the fact that the estimator of β is asymptotically normally distributed. The corresponding Wald test statistic is

$$T_{par} = n\hat{\beta}_*\hat{\Sigma}^{-1}\hat{\beta}_*, \quad (2.4.11)$$

where $\hat{\beta}_*$ is the estimate of β_* and $\hat{\Sigma}$ is the estimated covariance matrix of

$\sqrt{n}(\hat{\beta}_* - \beta_*)$. T_{par} has a χ^2 distribution under the null, with degrees of freedom equal to the dimension of β_* . This test is used after the model specification is accepted by $T_{ac}(s)$ and $T_{BN(01)}$. The advantage of this test over the asymptotically equivalent LM and LR tests is that it does not require to re-estimate the simplified model implied by the null hypothesis.

3. DATA ANALYSIS

We now turn to the monthly data that we use to model the unemployment in the United States of America, Canada, Japan, Germany, Great Britain, France and the Netherlands. These are the main industrial countries. Italy is excluded, because monthly figures of the unemployment were not available. The data are taken from the *Main Economic Indicators (MEI), Historical Data*, provided by the OECD. It concerns monthly data from January 1960 up to and including December 1987. The description of these variables is presented in appendix 1.

The variables that are used are: the unemployment (u), interest rate (r), industrial production (o), wages (w) and price level (p). These variables have been chosen, because, to some extent, they are important in explaining unemployment in the neoclassical and (new) Keynesian tradition. The Keynesian theory, for example, asserts that unemployment is caused by too low an effective demand, which can be represented by the industrial production. Neoclassical theory asserts that unemployment is caused by too high a real wage rate.

A difference with other applied studies concerning unemployment is that we use the number of unemployed rather than the unemployment rate. The reason for this is twofold. First, our theory explains how the *number* of unemployed might increase due to a high interest rate. Second, for some countries the number of monthly observations on the unemployment rate is very limited.

Our point of departure for selecting the model variables is the revenue maximization theory of Baumol (1959), augmented with a flexible labor effort rate as described in Bierens (1987). From this theory we can derive that it is the level of the nominal interest rate that causes the fixed costs, more specifically the interest payments, of firms to increase. According to Bierens (1987) this increase can lead to layoffs and hence unemployment. Similar conclusions are drawn by Farmer (1985) using the real interest rate in a model of the firm exhibiting asymmetric information. This means that we have to consider the level of the interest rate rather than its change.

The level of the interest rate and the level of unemployment are assumed to have zero steady state growth rates, and thus positive steady state levels. Unemployment is bounded from below by zero and from above by the total labor force, hence there cannot be a nonzero steady state growth rate of unemployment. The same argument applies to the interest rate. Moreover, for the same reason these two variables cannot have a unit root (cf.

Bierens (1987)). Furthermore, it can be shown that a time series y_t with a unit root will asymptotically take on negative values with probability $\frac{1}{2}$, i.e., $y_t = y_{t-1} + \varepsilon_t \Rightarrow P(y_t < 0) \rightarrow \frac{1}{2}$. This follows easily from the fact that

$$P(y_t < 0) = P(\sum_{j=1}^n \varepsilon_j + y_0 < 0) = P[(1/\sqrt{n}) \sum_{j=1}^n \varepsilon_j + (1/\sqrt{n}) y_0 < 0] \rightarrow$$

$$\lim_{n \rightarrow \infty} P[(1/\sqrt{n}) \sum_{j=1}^n \varepsilon_j < 0] = \frac{1}{2}, \quad (3.1)$$

provided $(1/\sqrt{n}) \sum_{j=1}^n \varepsilon_j$ is asymptotically $N(0, \sigma^2)$ distributed. For unemployment and the nominal interest rate this would imply that if they contain a unit root, then positive or negative values for both are equally likely. This is impossible, hence both are free from unit roots.

However, in the sample period considered, unemployment seems to exhibit nonstationary behavior for a number of countries, as indicated by the unit root tests, correlograms and plots. The nonstationary behavior of the interest rate seems less pronounced. For the other three variables o , w and p , we assume that firms make their decisions about production, investment, employment and their wage negotiations, etc. on a yearly basis and not from month to month. This is why we transform these variables to annual growth rates, i.e., $\dot{x}_t = 100(x_t - x_{t-12})/x_{t-12}$, which is in accordance with Bierens (1987). A further advantage of these transformations is that we can distinguish between the effects on unemployment of real versus nominal interest rate, and the growth of real versus nominal wages. These transformed variables will also be subjected to both unit root tests and we will also examine their correlograms and plots.

We have to bear in mind that wage increases and inflation were very volatile in the seventies and the beginning of the eighties, due to two oil crises and severe stagnation, so that nonstationarity might be indicated. On the other hand, for those two variables as well as for the percentage change of the industrial production, we may also argue that if a unit root should be present in those series, then wage increases, inflation and production increases would be equally likely as wage decreases, deflation and production decreases. This is clearly not the case in reality. Therefore on logical consistency grounds these variables should be stationary.

Of course, there will be other variables that may partly explain unemployment, for example wage flexibility, wage gaps, capital gaps, mismatch on labor markets, unemployment benefits and other social benefits, labor productivity,

labor supply, participation rates, etc. We cannot use any of these variables, because none of them is available in monthly figures. Lack of availability is also the reason why we have not included any variables representing fiscal policies, such as taxes, government expenditures or budget deficits. A second reason for not using these variables is that some of them have been constructed rather than observed directly, where it is not always clear how they have been constructed. Therefore we use 'raw' data rather than constructed data.

We also do not use seasonally adjusted time series, because fitting nonseasonal models to seasonally adjusted data might lead to model misspecification, as was pointed out by Newbold (1980). We explicitly incorporate seasonality in our models by applying the twelfth percentage difference to o , w and p and eleven seasonal dummies to capture the seasonal pattern in the unemployment series. Explicitly incorporating seasonality into the model is recommended by Wallis (1974), Sims (1974) and Newbold (1980).

We now present the results of the data analysis. First we give the results of the Phillips-Perron unit root test applied to the times series under review. Next, we filter the data by the lag operator $(1-0.5L)$ before applying the unit root test. In case there is a unit root in the series prefiltering with $(1-0.5L)$ should not matter for the outcome of the test. However, if there is only a near-unit root, this filter may improve the power of the unit root test (see Bierens (1990a)). The test results, both with and without the prefilter are presented in tables 3.1 and 3.2.

The empirical distribution of the test statistic is provided by Fuller (1976) and will also be presented in appendix 1. The critical value of the 5 % significance level is about -14 and for the 10 % significance level it is about -11, when the number of observations is between 200 and 300. Phillips and Perron (1988) apply a Newey-West type variance estimator with truncation parameter $m = o(n^{1/4})$. We have set $m = [n^{0.2}]$.

For the Netherlands inspection of the plot of $\Delta_1 u$ revealed an outlier in January 1976. From that date on the way of registration unemployment has been changed. Before that date this was done by hand, after that data it was done automatically, hence registration became more efficient. This caused a break as can be observed from publications of the Netherlands Central Planning Bureau.

The values marked with [†] in the tables are larger than the corresponding 5 % critical values and consequently the hypothesis of a unit root can not be rejected. The tests indicate that unemployment, interest rate and inflation

are nonstationary. However for reasons of logical consistency this unit root does not seem plausible, as we have argued above. The results in table 3.2 are more in agreement with common sense.

Table 3.1. The results of the Phillips-Perron unit root test for the variables used in the models of the various countries (monthly data).

	<i>USA</i>	<i>CAN</i>	<i>JPN</i>	<i>DEU</i>	<i>GBR</i>	<i>FRA</i>	<i>NLD</i>
<i>u</i>	-4.808 [†]	-5.081 [†]	-6.929 [†]	-1.589 [†]	-.0029 [†]	.8565 [†]	.3547 [†]
<i>r</i>	-3.500 [†]	-6.949 [†]	-6.124 [†]	-7.655 [†]	-12.12 [†]	-7.440 [†]	-9.896 [†]
$\dot{o}$	-23.00	-20.23	-14.99	-36.77	-40.93	-50.26	-30.69
$\dot{w}$	-17.70	-11.61 [†]	-15.33	-23.07	-10.55 [†]	-10.36 [†]	-8.549 [†]
$\dot{p}$	-5.318 [†]	-4.851 [†]	-8.898 [†]	-9.317 [†]	-5.647 [†]	-3.546 [†]	-10.30 [†]

Table 3.2. The results of the Phillips-Perron unit root test for the variables used in the models of the various countries.
Monthly data prefiltered with $(1-0.5L)$.

	<i>USA</i>	<i>CAN</i>	<i>JPN</i>	<i>DEU</i>	<i>GBR</i>	<i>FRA</i>	<i>NLD</i>
<i>u</i>	-14.18	-11.35 [†]	-19.34	-4.700 [†]	-.9465 [†]	.1065 [†]	-.3651 [†]
<i>r</i>	-8.257 [†]	-13.28 [†]	-27.73	-18.53	-28.62	-13.66 [†]	-20.76
$\dot{o}$	-26.26	-32.69	-24.53	-70.83	-67.62	-66.40	-60.76
$\dot{w}$	-37.36	-27.40	-25.75	-48.02	-22.91	-14.60	-18.87
$\dot{p}$	-9.001 [†]	-7.134 [†]	-15.67	-11.86 [†]	-8.404 [†]	-6.185 [†]	-18.32

Recently Bierens (1990a) proposed four Cauchy tests of the stationarity hypothesis against the unit root hypothesis. It appears that Cauchy test number III (cf. Bierens (1990a, theorem 3)) performs the best in case of a near unit root, especially after filtering the data by the lag operator $(1-0.5L)$. Indeed, when Bierens' (1990a) stationarity test number III is applied, it appears that the results are more in agreement with our intuitive ideas, especially if we first filter the data by $(1-0.5L)$. The results, both with and without the prefilter, are presented in table 3.3 and 3.4.

The critical values for the Cauchy distribution, which is equivalent to the *t* distribution with 1 degree of freedom, are ± 63.66 at 99 % significance, ± 12.71 at 95 % and ± 6.314 at 90 % significance.

Note that even after application of the prefilter $(1-0.5L)$ the tests indicate that unemployment in the European countries is still not stationary, whereas in the USA, Canada and Japan it is. If we take prefilter $(1-0.7L)$ instead of $(1-0.5L)$, the results for Germany indicate that unemployment can be considered stationary. Taking the logarithm of unemployment, instead of the actual level, the tests, with prefilter $(1-0.5L)$, indicates that the logarithm of unemployment is stationary for all countries.

Table 3.3. The results of the Bierens nr.III stationarity test for the variables used in the models of the various countries (monthly data).

	<i>USA</i>	<i>CAN</i>	<i>JPN</i>	<i>DEU</i>	<i>GBR</i>	<i>FRA</i>	<i>NLD</i>
<i>u</i>	3.972	13.96 ^{t)}	5.047	25.76 ^{t)}	201.4 ^{t)}	165.4 ^{t)}	162.9 ^{t)}
<i>r</i>	14.29 ^{t)}	5.527	6.441	1.369	1.573	5.758	.6184
$\dot{o}$	-.2512	-.7283	.0431	-.1723	-.6384	-.2084	-.2764
$\dot{w}$	-.7933	-.9583	-17.07	-1.036	3.208	-.1330	1.904
$\dot{p}$	-2.079	1.620	2.633	.0677	-5.262	-6.932	8.078

Table 3.4. The results of the Bierens nr.III stationarity test for the variables used in the models of the various countries.
Monthly data prefiltered with $(1-0.5L)$.

	<i>USA</i>	<i>CAN</i>	<i>JPN</i>	<i>DEU</i>	<i>GBR</i>	<i>FRA</i>	<i>NLD</i>
<i>u</i>	.9847	3.635	1.286	9.485	56.98 ^{t)}	65.28 ^{t)}	58.39 ^{t)}
<i>r</i>	3.799	1.539	4.924	.8045	.4755	1.644	.2103
$\dot{o}$	-.0690	-.4294	-.0799	-.1732	-.6061	-.2055	-.2050
$\dot{w}$	-.6438	-.7055	-24.68	-.9355	3.032	-.0581	.6157
$\dot{p}$	-.7312	.6000	1.848	.0873	-1.808	-2.006	6.264

4. ESTIMATION AND TEST RESULTS

4.1. Introduction

In this section we will discuss the estimation and test results of the ARMAX models. The actual results are presented in tables 4.1 to 4.7. To capture seasonal effects in unemployment we included eleven seasonal dummies in our models. Although a possible stochastic seasonal pattern may be removed using a Δ_{12} -filter on unemployment, this would actually imply that we assume a unit root in the unemployment series. However, we have argued in the previous sections that unemployment cannot have a unit root. Therefore we have not applied this Δ_{12} -filter on unemployment. For reasons of convenience we do not report the seasonal dummies in tables 4.1 to 4.7.

4.2. The United States

From table 4.1 we observe that none of the crucial specification tests rejects the specification of the initial model. The tests $T_{ac}(s)$, for $s=2, 6, 12, 24$ and 36 do not indicate serial correlation in the residuals, and the test $T_{BN(01)}$ does not indicate misspecification. Note that the normality test statistic has a significant value at 5 %. This is caused by an outlying error term at January 1975, which causes high excess kurtosis. There appears to be no ARCH in the residuals and the model can be simplified to a model containing only the interest rate and the percentage change in production as explanatory variables. This simplified model can also not be rejected. So in this case we find that the interest rate and the percentage change in production are significantly Granger causing unemployment in the USA. The AR coefficient is not disturbingly close to unity. A value of 0.95 for ϕ_1 in (2.2.1) implies that the impact of a unit shock in y_t is less than 0.55 after twelve months.

4.3. Canada

The results for Canada, reported in table 4.2, are very similar to those of the USA. Also in this case the initial model specification cannot be rejected, as indicated by $T_{ac}(s)$ and $T_{BN(01)}$. The normality test indicates normally

distributed errors, but they are not homoskedastic as first and twelfth order ARCH can not be rejected. As for the USA this model can be simplified to include only the interest rate and the percentage change in production as significant Granger causing explanatory variables. Also for Canada the value of AR parameter ϕ_1 is not disturbingly close to unity; after one year the effect of a unit shock is 0.77 and after two years it is 0.6.

4.4. Japan

In table 4.3A the results for Japan are presented. Due to large outliers in the errors of the model, the null hypothesis of normally distributed errors cannot be accepted and there also appears to be first order ARCH in the errors. However the crucial specification tests, $T_{ac}(s)$ and $T_{BN(01)}$, for the correctness of our model do not reject the initial specification. Note that the AR coefficient is 0.9993. Because of this near-unit value we turn from specification (2.1.1) to specification (2.2.2). Hence the near-unit root in unemployment is approximated by a unit root.

The estimation and test results for this model are presented in table 4.3B. We observe that this new specification can also not be rejected by any of the crucial model specification tests (cf. table 4.3B). Note that in this case, after simplification, only the interest rate and the percentage change in output remain in the model as explanatory variables.

4.5. Germany

For Germany the initial specification was not accepted because of serially correlated disturbances. After inspection of the (partial) autocorrelation function and some experimentation, this misspecification could be repaired by including a number of lagged dependent variables. The estimation and test results for the augmented model are presented in table 4.4A. In this case the sum of the AR coefficients was close to unity. Because of outliers, especially in 1967, the hypothesis of normally distributed errors is not accepted. However, both crucial specification tests, $T_{ac}(s)$ and $T_{BN(01)}$ do not indicate any severe misspecification. Because the sum of AR coefficients is close to unity we turn to specification (2.2.3).

This initial model was also not accepted by the most important model specification tests. After inspection of the partial autocorrelation function, we included some lags of the first difference of unemployment in the model and this specification could not be rejected. The estimation and test results of this model are presented in tables 4.4B. Note also how the consistent specification test improved. Whereas for the model in table 4.4A, three of the 20 values of $T_{BN(01)}$ exceeded the 10 % significance level, in the model of table 4.4B none does. After simplification it appears that only the interest rate and the percentage change in output are significantly Granger causing unemployment.

4.6. United Kingdom

The initial specification for the UK was not accepted because of serially correlated disturbances. Inspection of the (partial) autocorrelation function and some experimentation resulted in the specification presented in table 4.5A, which could not be rejected by $T_{ac}(s)$. In fact none of the diagnostic tests firmly rejects the specification of this model. As could be expected from earlier arguments in chapter 3, in this case the sum of the AR coefficients is close to unity. That is why we turn to specification (2.2.3).

This specification could however not be accepted due to autocorrelated disturbances. Therefore it was repaired by including some lagged dependent variables in the model. The estimation and test results of this augmented model are reported in table 4.5B. Note the improvement of the model specification test $T_{BN(01)}$: for the model in table 4.4A it exceeded the 10 % significance level twice out of 20 times, whereas for the model in table 4.5B this level was not exceeded at all. After simplification it appeared that the interest rate is the only significant Granger causing variable.

4.7. France

For France something peculiar occurred. Observation of the plot of unemployment and especially the plot of the first difference of unemployment reveal a dramatic change in the seasonal pattern during the early seventies. This is probably the reason why we could not find, even after trying several different

model specifications, a specification that was not rejected either by $T_{ac}(s)$ or by $T_{BN(01)}$. In order to get rid of this change in seasonal pattern, we exclude data before the change. Thus instead of the full sample we estimated and tested the model for France using the observations from January 1971 to December 1987, which amounts to 204 observations. A plot of this change of seasonal pattern in unemployment is reported in appendix 1.

Using the reduced sample period, the initial model for France could not be rejected by any of the diagnostic tests, although there appears to be twelfth order ARCH in the residuals (cf. table 4.6A). The value of the AR coefficient is very close to unity, which is why we turn to specification (2.2.2). The estimation and test results for this model are presented in table 4.6B. Note again the dramatic improvement of $T_{BN(01)}$. For the specification in table 4.6A the 10% significance level was exceeded six out of 20 times, whereas for the specification of table 4.6B this level was not exceeded once. There appears to be multicollinearity between the interest rate and inflation. We can find simplified models in which each appear as explanatory variables, in combination with the percentage change in output. However the model with inflation and percentage change in production appears to have disturbingly high 24-th order serial correlation in its residuals; $T_{ac}(24) = 36.32$, whereas the 5 % critical value is 36.4. We therefore choose the simplified model in which the interest rate and the percentage change in output are significantly Granger causing unemployment.

4.8. The Netherlands

Finally we present the results for the Netherlands in tables 4.7A and 4.7B. As was the case for Germany and the UK, the initial model specification (2.2.1) could not be accepted. This result is in accordance with that of Bierens (1987). Rather than transforming unemployment by a log transformation, as done by Bierens (1987), we have added additional lagged dependent variables in order to repair the misspecification. This resulted in an acceptable specification (cf. table 4.7A). However, the sum of the AR coefficients was greater than unity. That is why we turned to specification (2.2.3).

The initial model specification (2.2.3) was also rejected because of serially correlated disturbances. After inspection of the (partial) autocorrelation function of the errors of the model and inclusion of some lagged

dependent variables, the specification of the augmented general model in table 4.7B could not be rejected by $T_{ac}(s)$. Due to severe outliers in the early eighties the hypothesis of normally distributed errors was not accepted. There also appears to be ARCH in the residuals. However the crucial specification tests we apply indicate correct specification for both the general and the simplified model (cf. table 4.7B). Compared to the model of table 4.7A in this case the consistent specification test improved. For the model of table 4.7A $T_{BN(01)}$ exceeded the 90 % confidence bound three out of 20 times, whereas for the model of table 4.7B this bound was exceeded only once. After simplification it appeared that the interest rate, the percentage change in production and the percentage change in the wage rate are significantly Granger causing unemployment in the Netherlands. Note that this result differs from Bierens (1987). There the final model was a simple ARX(1) model in the log of unemployment, with the interest rate (lagged one month) as the only X-variable.

5. CONCLUDING REMARKS

The purpose of this paper was to develop an ARMAX modeling strategy, with which a possible relation between unemployment and interest rate, as found by Bierens (1987), was investigated for the major industrial countries. The specification of the ARMAX models is based on preliminary data analysis, where not only time series techniques play a role, but also information given by the relevant economic theory and information about logical consistency of certain variables is important. Another important feature of our ARMAX modeling approach concerns testing the specification of the various models with a number of diagnostic tests. Starting from a general model capturing as much dynamics as possible and, when the specification is not rejected, testing if it can be simplified by deleting variables with insignificant coefficients, is nowadays a frequently propagated modeling strategy.

It appeared that with our ARMAX modeling approach, using monthly data, we can confirm this Granger causal relation between unemployment and interest rate. For the major industrial countries the interest rate in combination with the percentage change in output had a significant impact on unemployment. For the UK only the interest rate was important and for the Netherlands, apart from interest rate and percentage output change, also the percentage change in the wage rate played a role.

This study challenges the assertions made by the neoclassical theory of the firm. Not the marginal costs of labor, i.e., the wage rate, determine the employment decisions of firms, but instead the interest rate, via the fixed costs. According to the neoclassical theory of the firm, fixed costs do not have any influence on employment; only marginal costs have. From our ARMAX models it appeared that the wage rate had no significant impact on unemployment, except for the Netherlands, whereas the interest rate did have a significant influence for all countries. Thus these empirical phenomena refute the neoclassical theory of the firm. Moreover, our empirical results confirm the managerial theory of the firm developed by Baumol (1959) and augmented by Bierens (1987) with a flexible labor effort rate.

Table 4.1. Estimation and test results for the USA

$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 r_{-1} + \alpha_3 \dot{o}_{-1} + \alpha_4 \dot{w}_{-1} + \alpha_5 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6) (1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36}) \varepsilon$									
general model					simplified model				
α_0	42.69	(.5512)			15.91	(.2352)			
α_1	.9539	(86.88)			.9568	(108.9)			
α_2	40.56	(3.309)			35.90	(4.659)			
α_3	-10.89	(-4.294)			-10.87	(-4.263)			
α_4	-7.612	(-1.300)							
α_5	.2072	(.0286)							
θ_1	-.0252	(-.4075)			-.0452	(-.7087)			
θ_2	.1432	(2.465)			.1366	(2.414)			
θ_3	.0502	(.8515)			.0240	(.4216)			
θ_4	.1517	(2.393)			.1201	(1.946)			
θ_5	-.0437	(-.8046)							
θ_6	.0751	(1.230)							
Θ_1	.1905	(3.146)			.1788	(3.114)			
Θ_2	.1985	(3.490)			.1852	(3.421)			
Θ_3	.0913	(1.550)							
$S_2 E.$	193.6				195.5				
R^2	.9935				.9933				
n	323				323				
T_{norm}	39.68				35.68				
$ARCH(1)$	1.386				1.664				
$ARCH(12)$	5.004				5.307				
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$		$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$		
2	1.006				1.741				
6	4.263				5.553				
12	9.753	3.842	3.979		13.97	6.912	7.141		
24	24.77	24.58	26.07		33.50	25.94	27.41		
36	33.42	43.73	47.28		42.28	51.21	55.49		
$T_{BN(01)}$: (general model)									
	.7333	1.125	.8217	-.4186	1.184	-.0018	1.320	.4314	1.358
	1.363	2.093	.5802	1.067	1.248	.5656	.9104	.7038	1.029
T_{par}	6.910	$(\alpha_4 = \alpha_5 = \theta_5 = \theta_6 = \Theta_3 = 0)$							
$T_{BN(01)}$: (simplified model)									
	.8815	.2181	.4581	-.2219	.8342	.7402	.3724	.5057	.4199
	.9668	.3071	1.468	.3940	-.1477	.3640	.2913	.6687	.0904

Table 4.2. Estimation and test results for Canada

$$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 r_{-1} + \alpha_3 \dot{o}_{-1} + \alpha_4 \dot{w}_{-1} + \alpha_5 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6) (1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36}) \varepsilon$$

general model			simplified model			
α_0	16.25	(1.319)	22.15	(1.963)		
α_1	.9785	(109.0)	.9776	(120.7)		
α_2	2.947	(2.134)	2.897	(3.037)		
α_3	-.6785	(-1.596)	-.7895	(-2.259)		
α_4	.8131	(1.171)				
α_5	-.7295	(-1.412)				
θ_1	.0125	(.1877)				
θ_2	.0531	(.8085)				
θ_3	-.0278	(-.3962)				
θ_4	.0206	(.2913)				
θ_5	-.0416	(-.6421)				
θ_6	.0183	(.2522)				
Θ_1	.2242	(3.607)	.2551	(4.017)		
Θ_2	.1657	(2.474)	.1996	(3.127)		
Θ_3	.1162	(1.788)	.1396	(2.148)		
$S.E.$	28.07		28.49			
R^2	.9948		.9945			
n	311		323			
T_{norm}	3.779		4.415			
$ARCH(1)$	7.905		15.52			
$ARCH(12)$	32.80		35.76			
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$
2	.2881			2.426		
6	1.482			3.456	4.257	4.317
12	2.660	.9495	.9826	4.872	5.357	5.450
24	20.88	16.88	17.96	26.77	25.85	27.28
36	35.89	34.75	37.79	46.53	45.32	48.81
$T_{BN(01)}: (general\ model)$						
	.6520	.1716	-.1410	.3695	.6948	.3252
	1.143	-.1070	-.8269	.1144	.8543	.2349
T_{par}	4.460	$(\alpha_4 = \alpha_5 = \theta_1 = \theta_2 = \theta_3 = \theta_4 = \theta_5 = \theta_6 = 0)$				
$T_{BN(01)}: (simplified\ model)$						
	-1.105	-.2515	-1.690	-.2007	-1.733	.6099
	-.8173	.2019	-.5604	-.4538	.1048	-.8711
						-.5360
						.7401
						.0722
						-.9171
						-.6398
						-.0329
						1.280
						-.9302

Table 4.3A. Estimation and test results for Japan

$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 r_{-1} + \alpha_3 \dot{o}_{-1} + \alpha_4 \dot{w}_{-1} + \alpha_5 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6) \\ (1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36}) \varepsilon$				
<i>general model</i>				
α_0	4.063	(.3043)		
α_1	.9993	(261.6)		
α_2	1.430	(1.450)		
α_3	-.6620	(-2.409)		
α_4	.2097	(.7254)		
α_5	-.3305	(-1.237)		
θ_1	-.4944	(-6.652)		
θ_2	-.2448	(-3.363)		
θ_3	.0175	(.2911)		
θ_4	.0291	(.5272)		
θ_5	-.1068	(-2.015)		
θ_6	-.0535	(-1.121)		
Θ_1	.2545	(2.889)		
Θ_2	-.0326	(-.5475)		
Θ_3	.0599	(.8280)		
$S_2 E.$	53.14			
R^2	.9832			
n	323			
T_{norm}	102.3			
$ARCH(1)$	4.622			
$ARCH(12)$	16.50			
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$	
2	.0238			
6	3.293			
12	9.434	3.173	3.294	
24	22.64	16.42	17.42	
36	27.04	19.67	21.02	
$T_{BN(01)}$: (<i>general model</i>)				
	-1.419	-.7097	-1.930	-1.505
	-1.052	-1.384	-1.136	-1.567
	-.7547	-1.310		
	-.7920	-1.434	-1.246	-1.097
	-.8810	-1.286	-.4626	-.8837
	-1.160	-.8038		

Table 4.3B. Estimation and test results for Japan

$$\Delta_1 u = \alpha_0 + \alpha_1 r_{-1} + \alpha_2 \dot{o}_{-1} + \alpha_3 \dot{w}_{-1} + \alpha_4 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6) \times (1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36}) \varepsilon$$

	general model			simplified model		
α_0	3.082	(.2403)		-.1491	(-.0111)	
α_1	1.430	(1.527)		1.938	(2.448)	
α_2	-.6217	(-5.369)		-.7578	(-5.769)	
α_3	.2372	(1.278)				
α_4	-.3247	(-1.235)				
θ_1	-.4951	(-6.744)		-.4840	(-6.689)	
θ_2	-.2455	(-3.346)		-.2391	(-3.673)	
θ_3	.0172	(.2875)				
θ_4	.0286	(.5183)				
θ_5	-.1077	(-2.018)				
θ_6	-.0547	(-1.147)				
Θ_1	.2558	(2.881)		.2693	(3.193)	
Θ_2	-.0319	(-.5374)				
Θ_3	.0594	(.8180)				
$S.E.$	53.14			53.89		
R^2	.7049			.6965		
n	323			323		
T_{norm}	102.8			92.64		
$ARCH(1)$	4.508			4.959		
$ARCH(12)$	16.48			20.77		
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$
2	.0299			.0479		
6	3.131			6.122	3.240	3.306
12	8.955	3.217	3.339	10.25	7.098	7.320
24	21.99	16.41	17.41	29.84	19.44	20.49
36	26.27	19.63	20.99	37.81	26.34	28.23

$T_{BN(01)}$: (general model)										
	-1.107	-1.095	-.2751	-.9470	-1.250	-.8936	-1.047	-.8091	-1.020	-1.277
	-.5791	-.8224	-1.031	-.8218	-.8960	-1.048	-.7229	-.7413	-.8388	-.8343

$$T_{par} \quad 12.24 \quad (\alpha_3 = \alpha_4 = \theta_3 = \theta_4 = \theta_5 = \theta_6 = \Theta_2 = \Theta_3 = 0)$$

$T_{BN(01)}$: (simplified model)										
	.5274	-.1610	-.1436	-.4163	-.1339	-.1465	-.3135	-.0309	-.1210	-.3940
	.4490	.7300	.2242	.4523	.3097	.6978	1.048	.6972	-.0297	.7115

Table 4.4A. Estimation and test results for Germany

$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 u_{-11} + \alpha_3 u_{-15} + \alpha_4 u_{-16} + \alpha_5 u_{-18} + \alpha_6 u_{-19} + \alpha_7 r_{-1} + \alpha_8 \dot{o}_{-1} + \alpha_9 \dot{w}_{-1} + \alpha_{10} \dot{p}_{-1} +$ $(1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\varepsilon$										
<i>general model</i>										
α_0	57.93	(2.833)								
α_1	1.001	(51.22)								
α_2	.0755	(1.595)								
α_3	-.2776	(-4.579)								
α_4	.1533	(2.424)								
α_5	.2135	(3.401)								
α_6	-.1661	(-3.595)								
α_7	7.491	(3.366)								
α_8	-.7323	(-1.698)								
α_9	-.2923	(-.6037)								
α_{10}	.7612	(.7297)								
θ_1	.3322	(5.089)								
θ_2	.1162	(1.615)								
θ_3	.0515	(.8036)								
θ_4	-.1032	(-1.672)								
θ_5	-.0972	(-1.421)								
θ_6	.0682	(1.082)								
Θ_1	.3004	(4.251)								
Θ_2	.3005	(4.323)								
Θ_3	.1566	(2.221)								
$S.E.$	26.26									
R^2	.9989									
n	299									
T_{norm}	11.89									
$ARCH(1)$	2.019									
$ARCH(12)$	17.17									
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$							
2	.7768									
6	3.869									
12	15.50									
24	29.39	20.72	22.20							
36	41.61	27.86	30.20							
$T_{BN(01)}$	<i>(general model)</i>									
	1.218	1.002	.3630	1.064	1.708	.5742	1.778	-.5605	1.962	.0943
	.5596	-.0189	-.0486	.0284	.5454	1.060	.3174	.4489	.7757	.6576

Table 4.4B. Estimation and test results for Germany

$$\Delta_1 u = \alpha_0 + \alpha_1 \Delta_1 u_{-15} + \alpha_2 \Delta_1 u_{-18} + \alpha_3 r_{-1} + \alpha_4 \dot{o}_{-1} + \alpha_5 \dot{w}_{-1} + \alpha_6 \dot{p}_{-1} +$$

$$(1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\varepsilon$$

general model			simplified model			
α_0	84.27	(6.624)	83.07	(7.027)		
α_1	-.1853	(-3.388)	-.1803	(-3.370)		
α_2	.1753	(3.798)	.1862	(3.867)		
α_3	7.991	(3.548)	8.080	(6.539)		
α_4	-.9987	(-2.697)	-.8064	(-2.680)		
α_5	-.5547	(-1.156)				
α_6	.6717	(.6425)				
θ_1	.3462	(5.214)	.3542	(5.339)		
θ_2	.1279	(1.811)	.1267	(1.809)		
θ_3	.0478	(.7759)	.0372	(.6197)		
θ_4	-.1002	(-1.677)	-.1212	(-2.049)		
θ_5	-.1011	(-1.644)	-.1286	(-2.266)		
θ_6	.0777	(1.413)				
Θ_1	.3837	(5.765)	.3676	(5.356)		
Θ_2	.3345	(4.907)	.3298	(4.865)		
Θ_3	.1725	(2.363)	.1721	(2.394)		
$S.E.$	26.65		26.79			
R^2	.9018		.9008			
n	299		299			
T_{norm}	14.06		12.86			
$ARCH(1)$	2.414		2.831			
$ARCH(12)$	22.30		21.98			
s	$T_{ac}(a)$	$T_{BP}(s)$	$T_{LB}(s)$	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$
2	1.229			1.891		
6	4.857			5.624		
12	12.99	5.494	5.710	13.13	6.893	7.149
24	24.65	23.73	25.38	29.48	25.05	26.72
36	32.35	31.32	33.86	36.56	32.54	35.10
$T_{BN(01)}$: (general model)						
	-.1034	.5714	.9563	.5799	.4906	.5405
	.3341	1.268	.5337	.6945	1.060	1.576
				.7810	1.464	.5344
				.9319		
T_{par}	4.075	$(\alpha_5 = \alpha_6 = \theta_6 = 0)$				
$T_{BN(01)}$: (simplified model)						
	1.603	-.4268	.4058	.6133	1.524	.4766
	.2923	-.5943	1.171	.0528	.2209	-.6623
					1.058	.0071
					1.119	.0246

Table 4.5A. Estimation and test results for the United Kingdom

$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 u_{-5} + \alpha_3 u_{-18} + \alpha_4 u_{-19} + \alpha_5 r_{-1} + \alpha_6 \dot{o}_{-1} + \alpha_7 \dot{w}_{-1} + \alpha_8 \dot{p}_{-1} +$ $(1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\varepsilon$									
<i>general model</i>									
α_0	.0791	(1.137)							
α_1	1.196	(40.18)							
α_2	-.2186	(-6.188)							
α_3	.1776	(3.245)							
α_4	-.1599	(-3.164)							
α_5	1.136	(2.485)							
α_6	-.8249	(-3.344)							
α_7	.2170	(.8684)							
α_8	.3202	(1.522)							
θ_1	.0791	(1.137)							
θ_2	.0165	(.2464)							
θ_3	-.1168	(-1.834)							
θ_4	-.1555	(-1.863)							
θ_5	.0806	(1.249)							
θ_6	.0594	(.8108)							
Θ_1	.4830	(7.185)							
Θ_2	.4533	(6.775)							
Θ_3	.1379	(1.711)							
$S.E.$	18.48								
R^2	.9997								
n	287								
T_{norm}	3.761								
$ARCH(1)$	0.129								
$ARCH(12)$	11.43								
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$						
2	1.494								
6	4.025								
12	15.01								
24	24.41	17.02	18.17						
36	37.88	29.30	32.12						
$T_{BN(01)}: (general model)$									
	-.8047	1.659	.0372	-.2836	-.2052	-1.054	-.0351	-.7172	.5903
	-.8455	.4098	.3460	-.3273	-1.227	-1.906	-.2502	-.0654	-.2732

Table 4.5B. Estimation and test results for the United Kingdom

$$\Delta_1 u = \alpha_0 + \alpha_1 \Delta_1 u_{-1} + \alpha_2 \Delta_1 u_{-13} + \alpha_3 \Delta_1 u_{-18} + \alpha_4 \Delta_1 u_{-19} + \alpha_5 r_{-1} + \alpha_6 \dot{o}_{-1} + \alpha_7 \dot{w}_{-1} + \alpha_8 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\epsilon$$

	general model			simplified model						
α_0	-1.855	(-2.740)		-1.354	(-.2163)					
α_1	.7964	(10.83)		.9469	(38.04)					
α_2	-.0718	(-2.066)		-.1067	(-4.222)					
α_3	.2083	(3.318)		.2605	(4.078)					
α_4	-.1803	(-2.930)		-.2555	(-4.042)					
α_5	.5266	(1.644)		.5449	(2.985)					
α_6	-.3900	(-1.956)								
α_7	.2439	(1.348)								
α_8	.2045	(1.398)								
θ_1	-.5094	(-5.205)		-.6179	(-11.69)					
θ_2	.0539	(.8212)								
θ_3	.0127	(.1906)								
θ_4	.0209	(.2885)								
θ_5	.1071	(1.529)								
θ_6	.0548	(.8296)								
Θ_1	.4567	(6.597)		.3728	(6.334)					
Θ_2	.4805	(7.287)		.4614	(7.598)					
Θ_3	.1059	(1.301)								
$S.E.$	18.43			18.87						
R^2	.8267			.8184						
n	287			287						
T_{norm}	4.300			4.011						
$ARCH(1)$	.2059			.0110						
$ARCH(12)$	10.42			9.185						
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$				
2	1.504			.1207						
6	6.434			2.265						
12	13.83			14.16	13.10	13.62				
24	22.86	14.80	15.75	26.86	22.82	24.12				
36	38.62	27.44	30.10	36.96	39.86	43.56				
$T_{BN(01)}$: (general model)										
	-.5866	.1334	.7515	.9142	.8015	.3557	.4760	.8291	-.7951	.2739
	.0890	-.9033	.5822	.4919	.5738	.0549	.3580	.0582	-.4192	.6768
T_{par}	10.69	$(\alpha_6=\alpha_7=\alpha_8=\theta_2=\theta_3=\theta_4=\theta_5=\theta_6=\Theta_3=0)$								
$T_{BN(01)}$: (simplified model)										
	.1618	-.1970	-.2337	.7176	.7381	.8080	1.187	-.5367	.5301	.3827
	.2854	.7854	.6250	1.605	.2878	1.534	-.0687	-.9854	-.4641	.9777

Table 4.6A. Estimation and test results for France

$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 r_{-1} + \alpha_3 \dot{o}_{-1} + \alpha_4 \dot{w}_{-1} + \alpha_5 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6) \\ (1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36}) \varepsilon$				
<i>general model</i>				
α_0	.8523	(.0503)		
α_1	.9921	(159.8)		
α_2	1.159	(1.356)		
α_3	-.6988	(-2.636)		
α_4	-.8519	(-1.081)		
α_5	1.874	(1.418)		
θ_1	.5364	(6.073)		
θ_2	.3554	(3.824)		
θ_3	.1698	(1.639)		
θ_4	.0993	(.9236)		
θ_5	-.0148	(-.1702)		
θ_6	-.0018	(-.0271)		
Θ_1	.4311	(5.242)		
Θ_2	.4605	(5.600)		
Θ_3	.2071	(2.497)		
<i>S.E.</i>	15.90			
R^2	.9996			
<i>n</i>	204			
T_{norm}	3.161			
<i>ARCH</i> (1)	6.331			
<i>ARCH</i> (12)	24.96			
<i>s</i>	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$	
2	.2994			
6	7.424			
12	10.98	6.775	7.148	
24	34.56	20.97	22.75	
36	42.37	35.09	39.57	
$T_{BN(01)}$: (<i>general model</i>)				
	-.2156	-.4734	-2.361	-.6570
	-.5670	-.6277	-1.823	-.8045
			-1.108	-2.697
			-1.099	-1.982
				-1.785
				1.237

Table 4.6B. Estimation and test results for France

$\Delta_1 u = \alpha_0 + \alpha_1 r_{-1} + \alpha_2 \dot{o}_{-1} + \alpha_3 \dot{w}_{-1} + \alpha_4 \dot{p}_{-1} + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6) (1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36}) \varepsilon$									
general model					simplified model				
α_0	-12.13	(-.8819)			-11.44	(-1.246)			
α_1	.6272	(.7252)			1.591	(2.629)			
α_2	-.6039	(-2.176)			-.7548	(-2.839)			
α_3	-.5465	(-.7006)							
α_4	2.143	(1.632)							
θ_1	.5391	(6.070)			.5210	(6.493)			
θ_2	.3520	(3.812)			.2747	(4.098)			
θ_3	.1617	(1.573)							
θ_4	.0942	(.8786)							
θ_5	-.0175	(-.2003)							
θ_6	-.0653	(-.0952)							
Θ_1	.4314	(5.286)			.3640	(4.595)			
Θ_2	.4590	(5.719)			.4100	(5.469)			
Θ_3	.2143	(2.651)			.1691	(2.015)			
$S.E.$	15.97				16.29				
R^2	.8989				.8948				
n	204				204				
T_{norm}	2.497				3.981				
$ARCH(1)$	6.902				2.598				
$ARCH(12)$	25.34				18.66				
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$		$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$		
2	.0869				3.392				
6	6.896				4.418	5.012	5.141		
12	9.928	6.644	7.008		10.30	9.366	9.731		
24	35.95	22.05	23.92		33.80	31.53	34.12		
36	43.73	36.88	41.62		37.57	47.60	53.24		
$T_{BN(01)}$: (general model)									
	-.9625	.0974	-.0247	-.4448	-.0984	.4867	.2249	-.1651	.1814
	-.9223	.4519	.2515	-.1522	.1103	.4243	.6039	.0924	-.1616
T_{par}	5.836	$(\alpha_3 = \alpha_4 = \theta_3 = \theta_4 = \theta_5 = \theta_6 = 0)$							
$T_{BN(01)}$: (simplified model)									
	.5626	-.3989	.3401	-.3024	.4902	1.360	.0215	.8399	1.181
	.0784	-.4729	.5225	.4043	-1.082	-.1691	-.4003	.6464	.8872

Table 4.7A. Estimation and test results for the Netherlands

$u = \alpha_0 + \alpha_1 u_{-1} + \alpha_2 u_{-2} + \alpha_3 u_{-5} + \alpha_4 u_{-11} + \alpha_5 u_{-12} + \alpha_6 u_{-23} + \alpha_7 u_{-24} + \alpha_8 r_{-1} + \alpha_9 o_{-1} + \alpha_{10} w_{-1} + \alpha_{11} p_{-1} + \alpha_{12} D76 + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \theta_1 L^{12} + \theta_2 L^{24} + \theta_3 L^{36})\varepsilon$									
<i>general model</i>									
α_0	8.702	(2.755)							
α_1	1.082	(32.53)							
α_2	-.1406	(-3.307)							
α_3	.0993	(3.152)							
α_4	.1082	(1.557)							
α_5	-.1589	(-2.453)							
α_6	.1410	(2.470)							
α_7	-.1302	(-2.455)							
α_8	.5152	(1.356)							
α_9	-.12.78	(-1.931)							
α_{10}	.3829	(2.709)							
α_{11}	-.0624	(-.2820)							
α_{12}	63.53	(24.62)							
θ_1	.2184	(2.806)							
θ_2	.1857	(2.507)							
θ_3	.1760	(2.402)							
θ_4	.1841	(2.687)							
θ_5	.1534	(2.603)							
θ_6	.1915	(3.416)							
θ_1	.5923	(6.937)							
θ_2	.2518	(2.623)							
θ_3	.0415	(.5223)							
$S.E.$	4.916								
R^2	.9997								
n	312								
T_{norm}	12.66								
ARCH(1)									
ARCH(12)									
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$						
2	3.749								
6	6.133								
12	10.00								
24	26.99	26.35	28.00						
36	47.65	49.60	53.70						
$T_{BN(01)}$: (general model)									
	.3502	1.144	.0317	-.2908	-2.755	.3730	.5195	2.534	.2110 .5722
	-1.250	-1.594	-.7581	-.6446	-.9205	-2.488	-1.255	-1.354	-1.471 -.7960

Table 4.7B. Estimation and test results for the Netherlands

$$\Delta_1 u = \alpha_0 + \alpha_1 \Delta_1 u_{-1} + \alpha_2 \Delta_1 u_{-5} + \alpha_3 \Delta_1 u_{-6} + \alpha_4 \Delta_1 u_{-11} + \alpha_5 \Delta_1 u_{-23} + \alpha_6 r_{-1} + \alpha_7 o_{-1} + \alpha_8 w_{-1} + \alpha_9 p_{-1} + \alpha_{10} D76 + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\varepsilon$$

general model			simplified model			
α_0	5.787	(2.137)	5.679	(2.463)		
α_1	.1334	(2.450)	.2379	(3.652)		
α_2	.0511	(1.533)	.0880	(2.411)		
α_3	.1157	(1.767)	.1209	(2.586)		
α_4	.1368	(1.989)	.1053	(1.677)		
α_5	.1333	(2.410)	.1031	(1.993)		
α_6	.7845	(2.448)	.7858	(3.526)		
α_7	-.1356	(-2.162)	-.1952	(-3.663)		
α_8	.3035	(2.353)	.1947	(2.532)		
α_9	-.1555	(-.7904)				
α_{10}	65.41	(35.07)	64.32	(60.77)		
θ_1	.1787	(2.037)				
θ_2	.1181	(1.577)				
θ_3	.0834	(1.202)				
θ_4	.0741	(1.259)				
θ_5	.1047	(1.624)				
θ_6	.0740	(.9161)				
Θ_1	.5398	(6.413)	.5085	(6.497)		
Θ_2	.2120	(2.219)	.2115	(2.707)		
Θ_3	.0580	(.0757)				
$S.E.$	4.913		4.997			
R^2	.8530		.8479			
n	312		312			
T_{nor}	11.76		10.15			
$ARCH(1)$	13.64		13.85			
$ARCH(12)$	42.27		44.04			
s	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$	$T_{ac}(s)$	$T_{BP}(s)$	$T_{LB}(s)$
2	1.770			2.670		
6	4.150			4.713		
12	6.798			9.962	9.716	9.949
24	27.98	26.98	28.69	27.52	34.28	36.15
36	49.10	50.06	54.18	47.82	55.77	59.93
$T_{BN(01)}: (general model)$						
	-.4921	-.8444	-.7949	-.6992	-1.897	-.2934
	1.187	-.9251	-.8100	-.2474	-.7255	-1.141
T_{par}	8.433	$(\alpha_{11}=\theta_1=\theta_2=\theta_3=\theta_4=\theta_5=\theta_6=\Theta_3=0)$				
$T_{BN(01)}: (simplified model)$						
	-1.771	.1711	.1568	-.6750	.0199	.2401
	-.0527	.0822	-.1007	1.773	.1647	.9176
				</		

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APPENDIX 1. UNIT ROOT TEST AND DATA SOURCES

The empirical distribution of the unit root test statistic of Phillips and Perron (1988)

The unit root test involved is the Phillips-Perron test for

$$H_0: y(t) = y(t-1) + u(t), \quad E[u(t)] = 0, \quad u(t) \text{ is } \alpha\text{-mixing}$$

$$H_1: y(t) = c + u(t),$$

where c is the intercept.

Reference: Phillips and Perron (1988).

Phillips and Perron use a Newey-West type variance estimator with truncation parameter $m = o(n^{1/4})$, where n is the number of observations.

Empirical distribution of Z_α under H_0 (cf. Fuller (1976), pp. 371).

n	Probability of a smaller value							
	0.01	0.025	0.05	0.10	0.90	0.95	0.975	0.99
25	-17.2	-14.6	-12.5	-10.2	-0.76	0.01	0.65	1.40
50	-18.9	-15.7	-13.3	-10.7	-0.81	-0.07	0.53	1.22
100	-19.8	-16.3	-13.7	-11.0	-0.83	-0.10	0.47	1.14
250	-20.3	-16.6	-14.0	-11.2	-0.84	-0.12	0.43	1.09
500	-20.5	-16.8	-14.0	-11.2	-0.84	-0.13	0.42	1.06
∞	-20.7	-16.9	-14.1	-11.3	-0.85	-0.13	0.41	1.04

The monthly data

The variables we use are taken from the OECD, *Main Economic Indicators, Historical Data*. They are presented below in untransformed form, with the corresponding name by which they appear on the OECD diskette.

1. Unemployment (u)

U.S.A.:	Total unemployment:	USA.UNEM.TOT TH PERSONS
Canada:	Total unemployment:	CAN.UNEM.TOT TH PERSONS
Japan:	Unemployment:	JPN.UNEM TH PERSONS

Germany:	Registered unemployment:	DEU.UNEM.REG TH PERSONS
G. Britain:	Registered unemployment:	GBR.UNEM.REG TH PERSONS
France:	Unemployment:	FRA.UNEM TH PERSONS
Netherlands:	Registered unemployment:	NLD.UNEM.REG TH PERSONS

2. Interest rate (r)

U.S.A.:	Official discount rate:	USA.OFF.DISC.RATE PERCENT PA
Canada:	Official discount rate:	CAN.OFF.DISC.RATE PERCENT PA
Japan:	Official discount rate:	JPN.OFF.DISC.RATE PERCENT PA
Germany:	Official discount rate:	DEU.OFF.DISC.RATE PERCENT PA
G. Britain:	Call money rate:	GBR.CALL.MON.RAT PERCENT PA
France:	Call money rate:	FRA.CALL.MON.RAT PERCENT PA
Netherlands:	Official discount rate:	NLD.OFF.DISC.RATE PERCENT PA

Comment: The units are in percentages per annum for all countries. For Great Britain the official discount rate is not reported by the OECD. France has kept its official discount rate at a constant level of 9.5% since August 1977. For both we have taken the call money rate.

3. Output (o)

U.S.A.:	Index of industrial production, total:	USA.IIP.TOT I/80
Canada:	Index of industrial production, manufact.:	CAN.IIP.MFG I/80
Japan:	Index of industrial production, total:	JPN.IIP.TOT I/80
Germany:	Index of industrial production, total:	DEU.IIP.TOT I/80
G. Britain:	Index of industrial production, total:	GBR.IIP.TOT I/80
France:	Index of industrial production, total:	FRA.IIP.TOT I/80
Netherlands:	Index of industrial production, total:	NLD.IIP.TOT I/80

4. Wages (w)

U.S.A.:	Hourly earnings in manufacturing:	USA.HLY.EARN.MFG I/80
Canada:	Hourly earnings in manufacturing:	CAN.HLY.EARN.MFG I/80
Japan:	Unit labour costs:	JPN.UNIT.LAB.CST I/80
Germany:	Unit labour costs:	DEU.UNIT.LAB.CST I/80
G. Britain:	Unit labour costs in manufacturing:	GBR.ULC.MFG I/80
France:	Labour costs in engineering:	FRA.LAB.CST.ENGIN I/80
Netherlands:	Hourly rates in manufacturing:	NLD.HLY.RAT.MFG I/80

5. Prices (p)

U.S.A.:	Producers price index, finished goods:	USA.PPI.FIN GDS I/80
Canada:	Producers price index, manufact. goods:	CAN.PPI.MFD.GDS I/80
Japan:	Consumers price index, total goods:	JPN.CPI.TOT I/80
Germany:	Producers price index, total goods:	DEU.PPI.TOT I/80
G. Britain:	Producers price index, total output:	GBR.PPI.OUT.TOT I/80
France:	Consumers price index, total goods:	FRA.CPI.TOT I/80
Netherlands:	Consumers price index, total goods:	NLD.CPI.TOT I/80

Finally the plot of the first difference of unemployment in France is presented in figure A1 to show the change in the seasonal pattern that occurred during the beginning of the seventies. Because of this, we excluded data before this change from the sample we used for estimation and testing.

Figure A1: First difference of French unemployment (1000 persons)