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Some international evidence

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THE RELATION BETWEEN UNEMPLOYMENT AND INTEREST RATE: SOME INTERNATIONAL EVIDENCE[1)]

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ABSTRACT

In Bierens (1987) a Granger causal relation was found between unemployment and the interest rate for the Netherlands. 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. A number of economic theories can explain these phenomena of which the revenue maximization theory of Baumol (1959) augmented with a flexible labor effort rate seems quite realistic.

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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 do not work anymore. Rising unemployment in the eighties brought about a return to a more classical way of fighting unemployment, based on the monetarist and new classical theory. Cf. Friedman (1968), Phelps (1967), Lucas (1973), Sargent and Wallace (1975). They emphasize that markets and also the labor market do clear, at least in the long run, because of price adjustments. The authorities should conduct a noninterventionist policy aimed only at establishing a constant money growth rate.

Keynesian oriented economists extended the Keynesian theory to account for the phenomena of high and persistent unemployment in combination with sticky wages. Cf. Malinvaud (1977), Stiglitz (1986). In their view markets do not always clear. Quantity rationing might lead to lasting disequilibria.

At the same time a change emerged in the construction of econometric models. The large scale Keynesian macroeconometric models broke down in the second half of the seventies. Monetarists and new classicals used small reduced form equations in support to their theories. Sims (1980) marked a swing back from the deductive modeling strategies of the Keynesian models to a more inductive research strategy. He claimed that the *a priori* restrictions made to identify these large models were incredible and suggested an unconditional VAR model, i.e., none of the parameters in his model as restricted to a certain value determined by economic theory. Thus, no relationship between the variables is ruled out beforehand.

An econometric methodology closely related to that of Sims (1980) is the ARMAX specification approach of Bierens (1987). However, instead of a VAR model, Bierens (1987) starts from a VARMA model. A VARMA model can be considered as a system of ARMAX models, just as a VAR is a system of ARX models. In Bierens (1987) an ARMAX methodology is set out, were also the data determine the specification of the ultimate model. With this ARMAX specification approach Bierens (1987) found a Granger causal relation between unemployment and the interest rate in the Netherlands. This phenomenon was explained by the revenue maximization theory of Baumol (1959) augmented with a flexible labor effort rate. This theory implies that fixed costs, via the interest costs, are important in determining the employment decisions of firms.

In the present study we slightly adjust the ARMAX methodology of Bierens (1987) and apply it to monthly time series of the USA, Canada, Japan, Germany, the UK, France and the Netherlands. These countries are chosen because they are important industrialized countries. Our aim is to investigate whether there also can be found a Granger causal relation between unemployment and the interest rate for those countries.

A cursory look at figures for the USA suggests that such a relation might also be present there. Figures 1 and 2 below give the unemployment and interest figures for the USA. Comparing the unemployment series of figure 1 with the interest rate series of figure 2, suggests a remarkably common pattern between unemployment and the interest rate, lagged about $1\frac{1}{2}$ year, particularly in the seventies and early eighties. These facts are corroborated by the summary statistics of table 1.1.

In section 2 we discuss our ARMAX modeling approach. Section 3 is devoted to a description and preliminary analysis of the data. In section 4 the estimation and test results will be presented and in section 5 the theoretical implications of these empirical results are mentioned. We finally make some concluding remarks in section 6.

Table 1.1. US unemployment (percentage of labor population) and interest rate (official discount rate in percentage per annum)

	unemployment	*interest rate*
1965-69	3.9	4.8
1970-74	5.4	5.9
1975-79	7.0	7.1
1980-84	8.3	10.6
1985-89	6.2	6.0

Source: OECD, *Main Economic Indicators*.

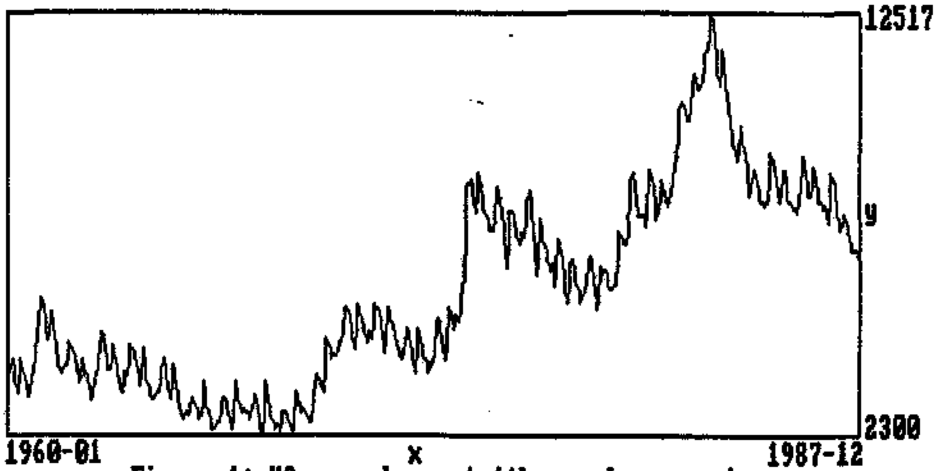

Figure 1: US unemployment (thousand persons)

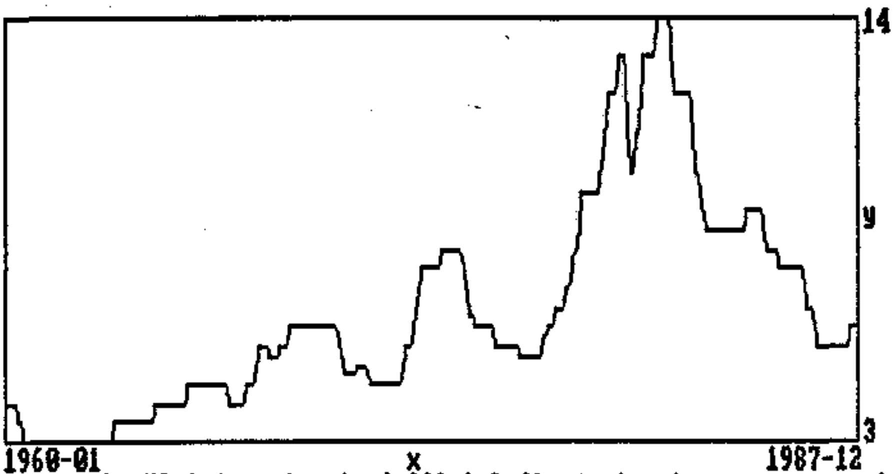

Figure 2: US interest rate (official discount rate, % per annum)

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. This implies that it represents the dynamic properties of the time series and on the other hand includes a role for economic theory. It can also be regarded as an equation of a VARMA model.

Suppose $Z_t \in \mathbb{R}^{k+1}$ is a vector time series process, then the statistical generating mechanism $Z_t = \mu_t + E_t$ specifies a crude approximation to the actual data generating process (DGP), where

$$\mu_t = E(Z_t | \sigma(Z_{t-1})) = E(Z_t | Z_{t-1}, Z_{t-2}, \dots) \tag{2.1}$$

and $\sigma(Z_{t-1})$ is the σ field generated by $Z_{t-1}, Z_{t-2}, \dots$. By definition the error process E_t is

$$E_t = Z_t - E(Z_t | Z_{t-1}, Z_{t-2}, \dots), \tag{2.2}$$

which satisfies the following properties

$$E(E_t | Z_{t-1}, Z_{t-2}, \dots) = 0, \tag{2.3}$$

$$E(E_t E_t' | Z_{t-1}, Z_{t-2}, \dots) = \begin{cases} \Omega(Z_{t-1}, Z_{t-2}, \dots) & k=0 \\ 0 & k \neq 0 \end{cases} \tag{2.4}$$

where $\Omega(Z_{t-1}, Z_{t-2}, \dots)$ is a positive function. The error process (2.3) defines a martingale difference process relative to the increasing sequence of σ-fields $\sigma(Z_1) \subset \sigma(Z_2) \subset \dots \subset \sigma(Z_t) \subset \dots$. This property implies correct specification, in the sense of (2.1). It is easy to derive that $E(E_{t-1}) \subseteq E(Z_{t-1})$, hence from (2.3) we can deduce

$$E(E_t | \sigma(E_{t-1})) = E(E_t | E_{t-1}, E_{t-2}, \dots) = 0, \tag{2.5}$$

i.e., E_t is not predictable from its own past.

If we assume $Z_t = (y_t, X_t')'$ then the systematic part of the statistical generating mechanism with respect to y_t becomes

$$\mu_t = E(y_t | (y_{t-1}, X_{t-1}')', (y_{t-2}, X_{t-2}')', \dots). \tag{2.6}$$

and the corresponding error process is

$$\varepsilon_t = y_t - E(y_t | (y_{t-1}, X_{t-1}')', (y_{t-2}, X_{t-2}')', \dots). \tag{2.7}$$

The properties (2.3)–(2.5) remain valid for ε_t as well.

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

$$\phi(L)y_t \;=\; \mu \;+\; \alpha(L)X_t \;+\; \theta(L)\Theta(L)\varepsilon_t, \tag{2.8}$$

where μ is the constant and

$$\phi(L) = 1 - \Sigma_{j=1}^{p}\phi_j L^j,\;\; \alpha(L) = \Sigma_{j=1}^{r}\alpha_j' L^j,$$

Moreover, we assume a multiplicative MA error structure with

$$\theta(L) = 1 + \Sigma_{j=1}^{q}\theta_j L^j,\;\; \Theta(L) = 1 + \Sigma_{j=1}^{Q}\Theta_j L^{sj}, \tag{2.9}$$

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 is the error process, which is defined to be a martingale difference. If the ARMAX model is to be stationary, $\phi(L)=0$ 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 $\theta(L)\Theta(L)=0$ has all its roots outside the unit circle, i.e., $\theta(L)=0$ and $\Theta(L)=0$ have all their roots outside the unit circle.

If the ARMAX model is invertible it can be written as an ARX(∞) model. If we denote

$$\begin{aligned}
\beta \;&=\; (\phi_1,\ldots,\phi_p,\alpha_1',\ldots,\alpha_r',\theta_1,\ldots,\theta_q,\Theta_1,\ldots,\Theta_Q)' \\
\mu(\beta) \;&=\; \frac{1}{\theta(1)\Theta(1)}\mu \\
\Sigma_{j=1}^{\infty}\eta_j(\beta)L^j \;&=\; \frac{1}{\theta(L)\Theta(L)}\begin{bmatrix}\psi(L)/L \\ \alpha(L)/L\end{bmatrix}
\end{aligned}$$

ARMAX model (2.8) can be written as the ARX(∞) model:

$$y_t \;=\; \mu(\beta) \;+\; \Sigma_{j=1}^{\infty}\eta_j(\beta)'Z_{t-j} \;+\; \varepsilon_t \tag{2.10}$$

The ARMAX specification (2.8) 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).

2.2. Specification of the ARMAX model

In essence our method is based on the assumption that the data generating process (DGP), which gives rise to the actual data, should be represented by a VARMA model, in contrast to Sims (1980) who uses a VAR model. The advantage of a VARMA over a VAR is that a VARMA model allows for an infinite lag structure with a parsimonious parameterization and it therefore is more capable of representing the strong dependence in macroeconomic time series. A VARMA model can be represented as a system of ARMAX models, just is a VAR is a system of ARX models. It is important that our model can represent the DGP of the actual data and that the time series used are stationary for valid inference on them.

Following the strategy of Kiviet and Phillips (1986), we start with a general model capturing as much of the dependence of the time series as possible and next we test whether the property of correct specification (2.3) is satisfied and whether the model can be simplified. It is our intention to represent the dynamics of the time series by means of the MA and seasonal MA parts (2.9) and to avoid multicollinearity among the explanatory variables. Therefore, we set $p=r=1$ and $q=6$ and $Q=3$ in model (2.8), which hopefully captures the dynamic structure of the economic time series as much as possible.

$$y_t = \mu + \phi_1 y_{t-1} + \alpha_1' X_{t-1} + (1+\theta_1 L+ \dots +\theta_6 L^6)(1+\Theta_1 L^{12}+ \dots +\Theta_3 L^{36})\varepsilon_t \tag{2.11}$$

It is tested whether this specification satisfies property (2.1); see section 2.3. If this is not the case this implies that the dependence of the economic time series under review is not adequately represented by (2.11). This misspecification can be repaired by adjusting (2.11) according to the information provided by the (partial) autocorrelation function of the residuals of the model. Cf. Box and Jenkins (1976). This indicates which additional lagged dependent variables should be included in the model. A change in the number of lags of the explanatory X variables is not considered, because there is no straightforward mechanism which can determine what this lag should be.

Next, we turn to the property that our model should be stationary. Stationarity of time series should in first instance be based on economic theory. If this theory does not prescribe a specific transformation, which renders stationary time series, then intuitive ideas about the logical consistency of time series in combination with exploration of the time series properties are used to determine correct transformations.

Data based transformations may be instigated by applying unit root or stationarity tests. Cf. Dickey and Fuller (1981), Phillips (1987), Phillips and Perron (1988), Bierens (1989). Automatic application of these unit root tests has recently been criticized by Schwert (1989), Bierens (1989) and Cochrane (1991). 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 this test does not do a very good job in distinguishing a near unit root from a genuine one. In order to improve the power of those tests, Perron (1989) suggests to correct time series for structural breaks. However, it is not clear how many 'known' structural breaks should be modeled. It appears that the specification of a model depends on the way breaks are represented by dummy variables, as was shown by Broersma and Franses (1990). Moreover, breaks in one series may correspond to breaks in another series and correcting them by dummies would just pour out the baby with the bath water. Another argument against automatic application of unit root tests is provided by Schiller and Perron (1985) who stress that unit root tests have a low power when applied to data with a high frequency, like monthly series.

Apart from these tests, we can also consider the value of the AR coefficient of model (2.11). It it has a value close to unity, this implies the presence of a near unit root, which can be approximated by a unit root. This in terms of model (2.11): if $\phi_1 \approx 1$, then we move to

$$\Delta_1 y_t = \mu + \alpha_1' X_{t-1} + (1+\theta_1 L+ \ldots +\theta_6 L^6)(1+\Theta_1 L^{12}+ \ldots +\Theta_3 L^{36})\varepsilon_t, \tag{2.12}$$

where $\Delta_k y_t = y_t - y_{t-k}$, $k \in \mathbb{N}$. In case (2.11) had to be augmented with additional lags in the dependent variable to accommodate initial misspecification, we consider the sum of the AR coefficients. If this sum is close to unity, this implies the presence of a near unit root. In this case account should be taken of the number and the order of additional AR parameters in (2.11). The new model specification should then either be based on (2.12), where the residual partial autocorrelation determines the number and order of AR coefficients in (2.12), or it should be based on repeating the whole specification stage, but instead of y_t we take $\Delta_1 y_t$ as dependent variable.

Essential for every model is the fact that the systematic part, as represented by (2.10) has to satisfy property (2.1) for the model to be an adequate representation of the actual DGP. In order to test this property we apply a number of misspecification tests. If the initial specification is no longer rejected, we test whether the model can be simplified by restricting parameter values to zero.

2.3. Testing the ARMAX model

We argued that (2.6) is crucial for a correct model specification, as the conditional expectation of y_t relative to the entire past of the vector time series process under review is the best predictive scheme in the sense that it has minimal mean squared error. In the terminology of Domowotz and White (1980) this implies first order model correctness. Therefore it is crucial to be able to test (2.6). However, any test relative to $\{y_t | \sigma(Z_{t-1})\}$ can only be tested via the estimated residuals $\hat{\varepsilon}_t$ of the model. Thus, (2.6) can be tested by testing (2.7) applied to $\hat{\varepsilon}_t$. Bierens (1988) has developed a test, $T_{ac}(s)$ of the null hypothesis (2.7) against the alternative (2.5) applied to ε_t, i.e.,

$$E(\varepsilon_t | \varepsilon_{t-1}, \varepsilon_{t-2}, \ldots, \varepsilon_{t-k}) \neq 0, \tag{2.13}$$

for some $k \in \{1,2,\ldots,s\}$. This test is a straightforward further elaboration of the autocorrelation test of Godolphin (1980). The power of this test is studied by Clarke and Godolphin (1982). It appeared that the Godolphin test performed superior compared to the familiar tests of Box and Pierce (1970) and Ljung and Box (1978). The test T_{ac} that we apply has the advantage that it does not depend on the assumption of normally distributed errors, like most applied autocorrelation tests. For reasons of convenience we also apply the Box-Pierce and Ljung-Box tests, even though they depend on the assumption of normality and they are especially unreliable when the model contains lagged dependent variables, as is shown in studies of, e.g., Kiviet (1986) and Hall and McAleer (1989).

Bierens (1987) has developed a consistent test of property (2.7) against the more general alternative that (2.7) does not hold. Consistency implies in his case that any misspecification will be detected as the sample size goes to infinity. This is a more general test than T_{ac}. Rejection of (2.13) does not provide enough evidence to accept (2.7). It may be that (2.13) is rejected while (2.7) is not true. In that case we still have to apply this consistent misspecification test to verify whether (2.7) is true. We use T_{ac} as a pretest of model misspecification as severe misspecification will likely be covered by (2.13). If a model exhibits autocorrelated residuals according to T_{ac} it is no longer necessary to conduct the consistent misspecification test.

For the exact form of the test statistic $T_{BN(01)}$ of this test we refer to Bierens (1987). It is a randomized test and this randomization guarantees the consistency. The test statistic depends on nuisance parameters. In conducting the test these nuisance parameters are replaced by random drawing from a continuous distribution. Under null hypothesis (2.7), this statistic converges in distribution to a $N(0,1)$ distribution. As argued by Bierens (1987), it is better to conduct this test several times for different random drawings of the nuisance parameters involved. We shall run this test 20 times.

These two tests are crucial with respect to the specification of our ARMAX models. If the initial specification (2.11) or (2.12) is rejected by either of the two tests, this misspecification is repaired by including additional lagged dependent variables based on the (partial) autocorrelation function of the residuals of the model. See section 2.2.

Note that the usual assumptions of normality and homoskedasticity are not essential for model correctness (2.7), as it does not imply both properties. The consequence of nonnormality is that maximum likelihood is no longer applicable, but the least squares method can be applied and renders consistent, but not efficient estimators. However, in our approach we are less concerned with the loss in efficiency and more relying on the robustness of the tests we apply. Moreover, asymptotic normality results for the parameter estimators involved can be derived on the basis of the central limit theorem for martingale difference sequences of McLeish (1974). This is one of the reasons for using monthly data so that the asymptotic approximations are likely to be better.

We do however conduct the normality test of Kiefer and Salmon (1983), $T_{norm}(2)$, which is equivalent to the familiar test of Jarque and Bera (1980), because it may provide additional information about the DGP. For the same reason we also apply the ARCH test of Engle (1982), T_{ARCH}. The ARCH test is chosen because that form of heteroskedasticity is connected to time series data. Heteroskedasticity leads to inconsistent covariance matrix estimators. However, we do not assume *a priori* homoskedasticity, as is indicated by (2.4). Instead, the function Ω of (2.4) represents the heteroskedasticity-consistent covariance estimator of White (1980), hence our tests remain valid asymptotically.

As a final misspecification test we conduct a test on predictive failure of the model. Parameter constancy is also an important property of econometric models and is connected with the assumption that the parameters of the distribution of the time series process Z_t are time invariant. We apply the test advocated by Hendry (1979) and Hendry and Ericsson (1991), which simply compares subsample and forecast residual variances. The test statistic is $T_{forec}=(SSR_{n-n_1}/\sigma^2_{n_1})/(n-n_1)$, where SSR_{n-n_1} is the sum of squared forecast residuals and $\sigma^2_{n_1}$ is the residual variance of the model, based on n_1 observations. T_{forec} is an index of parameter constancy for the $n-n_1$ *ex-ante* forecasts and is approximately $F(n-n_1,n_1)$ distributed. The corresponding 5% critical value is about 2.

Thus, we apply a number of misspecification tests of which only T_{ac} and $T_{BN(01)}$ are crucial for correct specification and also T_{forec} is of some importance. T_{ac} is nested within $T_{BN(01)}$ and application of the latter depends on the outcome of the first. This implies that we apply a testing strategy in the sense of Kiviet and Phillips (1986) and the pretesting problem will not be serious. The predictive failure test does depend on normality, but this F-type test is fairly robust against nonnormality. Cf. Hendry and Ericsson (1991).

Finally, we use a Wald test on parameter restrictions, T_{par}, to simplify the initial model specification by testing whether certain coefficients can be restricted to zero.

Note that all tests follow a χ^2 distribution, with the degrees of freedom denoted in the brackets of the statistic, except T_{forec} and $T_{BN(01)}$. The first is an index of parameter constancy with an approximate F distribution. Thus values larger than 2 imply poor *ex-ante* forecasts. The latter is a randomized test with a $N(0,1)$ distribution under the null hypothesis of correct specification. This test is run 20 times. All tests are summarized in table 2.1.

Table 2.1. Some criteria for evaluating and designing ARMAX models.

Alternative	*Statistic*	*Source*
skewness and excess kurtosis	$T_{norm}(2)$	Kiefer and Salmon (1983)
k the order ARCH	$T_{ARCH}(k)$	Engle (1982)
predictive failure over $n-n_1$ observations	T_{forec}	Hendry (1979)
s th order residual autocorrelation	$T_{B\text{-}P}(s-p-q-Q)$	Box and Pierce (1970)
	$T_{L\text{-}B}(s-p-q-Q)$	Ljung and Box (1978)
	$T_{ac}(s)$	Bierens (1988)
error not a martingale difference	$T_{BN(01)}$	Bierens (1987)
m invalid parameter restrictions	$T_{par}(m)$	Spanos (1986)

Our ARMAX modeling strategy is the same in spirit as the dynamic modeling strategy of Hendry and Richard (1982, 1983). Their method boils down to starting with an autoregressive distributed lag model with a large number of lags in order to capture as much of the dynamics of the time series involved. Next, tests are performed whether coefficients of this general model can be restricted to certain values, e.g., zero. They frequently test whether simplification to an error correction model can be established. Another important aspect of their methodology is the application of a number of misspecification tests in order to determine whether the model satisfies the necessary requirements for an adequate representation of the DGP giving rise to the observed phenomena.

The main difference between our method and the one of Hendry and Richard lies in the fact that we allow a MA structure in our models, which implies a better representation of the dependence in time series, as argued earlier. This is especially true for models based on monthly time series. The Hendry and Richard method is more appropriate for quarterly or yearly data, as in that case the number of lags included in the initial model can be limited to avoid computational difficulties. In case of monthly data this number of lags could easily become too large to able to represent the dependence of the economic time series. Our ARMAX modeling approach is then more suitable.

Another difference is that we emphasize the properties of the statistical generating mechanism, i.e., property (2.3) or (2.7). They also stress the importance of normality and homoskedasticity.

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), an index of industrial production (o), a wage index (w) and a price index (p). These variables have been chosen, because, to some extend, 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 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 main reason is that for some countries the number of monthly observations on the unemployment rate is very limited or even absent. Nevertheless, we have conducted the same specification analysis for the unemployment *rate* for the four countries (USA, Canada, Japan and Germany) for which these figures were available. The results were very similar to those below. Cf. Broersma (1991a).

The interest rate that is being used is the official discount rate, when available. In case of the UK and France we used the call money rate. This discount rate is taken because it can be considered the interest rate on which all other interest rates are founded. We also used other available interest rates, like treasury bill rates, mortgage rates, government bond yields, business loan rates, *etc.*, in the same specification analysis for the various countries. The results we found were very similar to those below. Cf. Broersma (1991b).

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). We can argue that the other three variables o, w and p contain a persistent exponential trend. Therefore these variables are transformed to annual growth rates, $\dot{x}_t = 100(x_t - x_{t-12})/x_{t-12}$, which is in agreement 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 and stationarity tests.

However, in the sample period considered, unemployment has risen considerably, which might indicate nonstationary behavior for a number of countries, as indicated by, e.g., unit root tests. Also the interest rate has shown some volatility, but its possible nonstationary behavior seems less pronounced. We have to bear in mind that wage and price 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. However, for reasons of logical consistency these variables cannot contain a unit root. Possible nonstationary behavior merely indicates the presence of a near unit root.

We now present the results of the data analysis. First we give the results of the Phillips–Perron unit root test applied to the time 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. Cf. Bierens (1989). The 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, when the number of observations is between 200 and 300. Phillips and Perron (1988) apply a Newey and West (1987) type variance estimator with truncation parameter $m=o(n^{1/4})$. We have set $m=[n^{0.2}]$.

The values marked with † in the tables are larger than the corresponding 5% critical values so the hypothesis of a unit root cannot be rejected. The test indicates 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, where the data were filtered with $(1-0.5L)$, are more in agreement with our intuitive ideas and common sense.

Table 3.1. **The results of the Phillips-Perron unit root test for the time series of the various countries.**

	USA	*CAN*	*JPN*	*DEU*	*GBR*	*FRA*	*NLD*
u	-4.808†	-5.081†	-6.929†	-1.589†	-.0029†	.8565†	.3547†
r	-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†

†*unit root hypothesis not rejected at 5% significance.*

Table 3.2. **The results of the Phillips-Perron unit root test for the prefiltered time series of the various countries.**

	USA	*CAN*	*JPN*	*DEU*	*GBR*	*FRA*	*NLD*
u	-14.18	-11.35†	-19.34	-4.700†	-.9465†	.1065†	-.3651†
r	-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

†*unit root hypothesis not rejected at 5% significance.*

Bierens (1989) proposed four Cauchy tests of the stationarity hypothesis against the unit root hypothesis. A Monte Carlo study favored test number III in case of a near unit root. Cf. table 3.3.

Table 3.3. **The results of Bierens' nr. III stationarity test for the time series of the various countries.**

	USA	*CAN*	*JPN*	*DEU*	*GBR*	*FRA*	*NLD*
u	3.972	13.96†	5.047	25.76†	201.4†	165.4†	162.9†
r	14.29†	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

†*stationarity rejected at 5% significane.*

Note that the results of this test are more in agreement with our intuitive ideas about logical consistency, than the results of the Phillips-Perron test. The results of the latter test with the prefiltered data are not reported.

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. For reasons of convenience we do not report the seasonal dummies in tables 4.1 to 4.7.

When applying the predictive failure test, we estimated the model based on the sample until December 1985 and used it to predict the remaining two years, from January 1986 to December 1987. These forecasts are compared with the realizations over that period. So in the sense of the predictive failure test, we set the forecast period $n-n_1=24$. This test is only applied to the simplified models that were chosen. The critical value we use is 2, which is an approximate 5% significance level. In addition we give the coefficient of determination R^2, the residual standard error $S.E.$ and the number of observations n. In the brackets are the t values based on the heteroskedasticity-consistent covariance matrix estimator of White (1980).

4.2. The United States of America

The estimation and test results of the initial model (2.11) for the USA are presented in table 4.1. Note that none of the important misspecification tests, T_{ac} and $T_{BN(01)}$, reject the specification of this model. However, the normality test rejects the null hypothesis of normally distributed residuals at 5%. This is due to an outlier in January 1975, which causes high excess kurtosis. There appears to be no first and twelfth order ARCH in the residuals of the model and the Wald test on parameter restriction does not indicate that the initial model is invalidly simplified, as $T_{par}(5)=6.91$, which is well within the 95% confidence interval of the $\chi^2(5)$ distribution.

Table 4.1. Estimation and test results for the USA.

$$u = \mu+\phi_1 u_{-1}+\alpha_1 r_{-1}+\alpha_2 \dot{o}_{-1}+\alpha_3 \dot{w}_{-1}+\alpha_4 \dot{p}_{-1}+(1+\theta_1 L+\theta_2 L^2+\theta_3 L^3+\theta_4 L^4+\theta_5 L^5+\theta_6 L^6)(1+\Theta_1 L^{12}+\Theta_2 L^{24}+\Theta_3 L^{36})\varepsilon$$

	general model					*simplified model*				
μ	42.69	(.5512)				15.91	(.2352)			
ϕ_1	.9539	(86.88)				.9568	(108.9)			
α_1	40.56	(3.309)				35.90	(4.659)			
α_2	-10.89	(-4.294)				-10.87	(-4.263)			
α_3	-7.612	(-1.300)								
α_4	.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)								
$S.E.$	193.6					195.5				
R^2	.9935					.9933				
n	323					323				
$T_{norm}(2)$	39.68†					35.68†				
$T_{ARCH}(1)$	1.386					1.664				
$T_{ARCH}(12)$	5.004					5.307				
s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	1.01	4.26	9.75	24.77†	33.42†	1.74	5.55	13.97	33.50	42.28†
$T_{B\text{-}P}(s)$			3.84	24.58†	43.73†			6.91	25.94	51.21†
$T_{L\text{-}B}(s)$			3.98	26.07†	47.28†			7.14	27.41	55.49†
$T_{BN(01)}$: (*general model*)										
	.7333	1.125	.8217	-.4186	1.184	-.0018	1.320	.4314	1.358	1.001
	1.363	2.093	.5802	1.067	1.248	.5656	.9104	.7038	1.029	1.741
$T_{par}(5)$	6.910	$(\alpha_3=\alpha_4=\theta_5=\theta_6=\Theta_3=0)$								
$T_{BN(01)}$: (*simplified model*)										
	.8815	.2181	.4581	-.2219	.8342	.7402	.3724	.5057	.4199	.4291
	.9668	.3071	1.468	.3940	-.1477	.3640	.2913	.6687	.0904	.2617
T_{forec}						1.460				

† *significant at a 5% level*

This simplified model contains only the interest rate and the growth rate of industrial production as significant Granger causing variables. This specification can also not be rejected by any of the important misspecification tests. Note also that this model does not suffer from predictive failure, as indicated by T_{forec}. The AR coefficient is rather close to unity, but a value of $\phi_1 = 0.95$ in model (2.11), implies that the impact of a unit shock in y_t is less then 0.55 after twelve months. If a model with the first difference of unemployment, specified as (2.12), was estimated, a very similar model can be found, which can also not be rejected and which also has the interest rate and growth of industrial production as explanatory variables. However, we believe that specification (2.11) is better capable of representing the persistence in the unemployment series and it is in agreement with the premise made in the previous section that unemployment cannot contain a unit root.

For the sake of reasoning, we also estimated the same model as in table 4.1 with a dummy included to represent the outlier in January 1975. The estimated model we got in that case was very similar to the one without the dummy and its specification could not be rejected by any of the misspecification tests. Hence, the estimation and test results are robust against nonnormality.

The results of table 4.1 imply that the simplified model is an adequate representation of the DGP giving rise to the monthly data on US unemployment. Moreover, the interest rate and growth rate of industrial production are significantly Granger causing unemployment.

4.3. Canada

The ARMAX model of unemployment for Canada, as reported in table 4.2, is very similar to the one for the USA. Also in this case the initial model (2.11) cannot be rejected by the crucial misspecification tests, T_{ac} and $T_{BN(01)}$. The null hypothesis of normally distributed errors can also not be rejected, but the residuals are not homoskedastic as absence of first and twelfth order ARCH can not be accepted by the ARCH test. However, we have argued that homoskedasticity is not crucial for model correctness.

Table 4.2. Estimation and test results for Canada.

$$u = \mu+\phi_1 u_{-1}+\alpha_1 r_{-1}+\alpha_2 \dot{o}_{-1}+\alpha_3 \dot{w}_{-1}+\alpha_4 \dot{p}_{-1}+(1+\theta_1 L+\theta_2 L^2+\theta_3 L^3+\theta_4 L^4+\theta_5 L^5+\theta_6 L^6)(1+\Theta_1 L^{12}+\Theta_2 L^{24}+\Theta_3 L^{36})\varepsilon$$

	general model		*simplified model*	
μ	16.25	(1.319)	20.85	(1.865)
ϕ_1	.9785	(109.0)	.9782	(123.4)
α_1	2.947	(2.134)	2.724	(2.773)
α_2	-.6785	(-1.596)	-.8564	(-2.373)
α_3	.8131	(1.171)		
α_4	-.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)	.2276	(3.616)
Θ_2	.1657	(2.474)	.1762	(2.655)
Θ_3	.1162	(1.788)	.1258	(1.838)
$S.E.$	28.07		28.29	
R^2	.9948		.9947	
n	311		311	
$T_{norm}(2)$	3.779		6.581†	
$T_{ARCH}(1)$	7.905†		8.460†	
$T_{ARCH}(12)$	32.80†		32.89†	

s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	.288	1.48	2.66	20.88	35.89	1.47	2.36	2.72	21.60	38.81
$T_{B\text{-}P}(s)$			.950	16.88	34.75		2.84	3.43	20.01	39.04
$T_{L\text{-}B}(s)$			.983	17.96	37.79†		2.89	3.50	21.18	42.31

$T_{BN(01)}$: (*general model*)

.6520	.1716	-.1410	.3695	.6948	.3252	-.5729	.1377	.2925	-.3120
1.143	-.1070	-.8269	.1144	.8543	.2349	.1357	.2172	.8755	.3340

$T_{par}(8)$ 4.460 $(\alpha_3=\alpha_4=\theta_1=\theta_2=\theta_3=\theta_4=\theta_5=\theta_6=0)$

$T_{BN(01)}$: (*simplified model*)

-.1970	-1.146	-2.081	-.5365	-1.758	-.4505	-2.141	-.6300	-.6859	-1.645
-1.080	-1.353	-.5922	-1.405	-.7898	-.8099	-.8964	-1.600	.0741	-.4954

T_{forec} .6661 (simplified model)

†*significant at a 5% level*

Essential for model correctness is the property that the errors are a martingale difference process as indicated in (2.3) or (2.7). Moreover, we apply the heteroskedasticity consistent covariance matrix of White (1980), so asymptotically, the t values and test statistics remain valid. Estimation of a model of the logarithm of the unemployment did not remove the residual ARCH.

Testing the null hypothesis that $\dot{w}_{-1}$, $\dot{p}_{-1}$ and the MA(1) through MA(6) parameters can be deleted from the model yields a $\chi^2(8)$ test statistic with value $T_{par}(8)=4.46$ and can therefore not be rejected at a 5% significance level. None of the important misspecification tests rejects the simplified model. Thus, also in this case we find that only the interest rate and the growth of industrial production are Granger causing unemployment. The AR parameter of this model is close to unity, but the impact of a unit shock is about 0.75 after one year. Therefore, this model is capable to represent the strong dependence in unemployment and still lead to valid inference, as for the USA.

4.4. Japan

The model for Japan based on (2.11) does not seem to be misspecified, as it is not rejected by any of the crucial misspecification tests. However, the AR coefficient is very close to unity, as $\phi_1=0.9993$. This near unit value might cause invalid inference because of near nonstationarity. That is why we move from specification (2.11) to (2.12), which implies that we have approximated the near unit root in unemployment by a unit root.

The estimation and test results of specification (2.12) are reported in table 4.3. There is hardly any difference between the general model based on (2.11) and the one of table 4.3, both in parameter values and test results. Also in this case the specification is not rejected by any of the crucial misspecification tests. Nonnormality is caused by a very large outlier in March 1967. The only source of concern is the fact that the simplified model seems to suffer from parameter nonconstancy for the 24 predictions, as T_{forec} exceeds 2. However, if we set $n-n_1=48$, we find $T_{forec}=1.552$. Hence, it does not seem to be a severe problem. But taking

Table 4.3. Estimation and test results for Japan.

$$\Delta_1 u = \mu+\alpha_1 r_{-1}+\alpha_2\dot{o}_{-1}+\alpha_3\dot{w}_{-1}+\alpha_4\dot{p}_{-1}+(1+\theta_1L+\theta_2L^2+\theta_3L^3+\theta_4L^4+\theta_5L^5+\theta_6L^6)$$
$$(1+\Theta_1L^{12}+\Theta_2L^{24}+\Theta_3L^{36})\epsilon$$

	general model		*simplified model*	
μ	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	
$T_{norm}(2)$	102.8†		92.64†	
$T_{ARCH}(1)$	4.508†		4.959†	
$T_{ARCH}(12)$	16.48		20.77	

s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	.030	3.13	8.96	21.99	26.27	.048	6.12	10.25	29.84	37.81
$T_{B\text{-}P}(s)$			3.22	16.41	19.63		3.24	7.10	19.44	26.34
$T_{L\text{-}B}(s)$			3.34	17.41	20.99		3.31	7.32	20.49	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

$T_{par}(8)$ 12.24 $(\alpha_3=\alpha_4=\theta_3=\theta_4=\theta_5=\theta_6=\Theta_2=\Theta_3=0)$

$T_{BN(01)}$: (*simplified model*)

.5274	-.1610	-.1436	-.4163	-.1339	-.1465	-.3135	-.0309	-.1210	-.3940
.4490	.7300	.2242	.4523	.3097	.6978	1.048	.6972	-.0297	.7115

T_{forec} — simplified model: 2.137

†*significant at a 5% level*

account of the nonnormality, this problem may aggravate.

The outlier for March 1967 in the model of the first difference of unemployment could be removed by inclusion of the first difference of a dummy for March 1967, even though we are not sure if this outlier represents a structural break. The first difference of the dummy was needed, because we also had to take the first difference of unemployment in (2.12). The estimation and test results of this model are very similar to the ones of table 4.3. Inclusion of this dummy yields normally distributed errors, as $T_{norm}(2)=4.298$ for the simplified model. Hence, as was found for the USA, also in this case the results seem to be robust against nonnormality. First order ARCH remained present and the problem of predictive failure remained for 24 periods. However, if $n-n_1=48$, then $T_{forec}=1.753$, which implies absence of serious parameter nonconstancy.

Summarizing we may conclude that also in case of Japan, unemployment is significantly Granger caused by the interest rate and the growth rate of production. Moreover, the results are fairly robust against the presence of nonnormality, which implies that fat tail distributions of the residuals do not matter much.

4.5. Germany

For Germany the initial specification (2.11) was not accepted due to serially correlated disturbances. After inspection of the (partial) autocorrelation function and some experimentation, this misspecification was repaired by inclusion of some additional lagged dependent variables. However, the sum of the AR coefficients is very close to unity, as $\Sigma_{j=1}^{6}\phi_j=0.9996$, which points towards near nonstationarity and possibly invalid inference.

Therefore, we move to specification (2.12) and base the additional lags of the dependent variable on the (partial) autocorrelation function of the residuals of (2.12) and the additional lags of y_t in the initial model specification (2.11). The estimation and test results of this model are given in table 4.4. None of the important misspecification tests rejects this specification. In fact, comparison of the results of the

Table 4.4. Estimation and test results for Germany.

$$\Delta_1 u = \mu + \phi_1 \Delta_1 u_{-15} + \phi_2 \Delta_1 u_{-18} + \alpha_1 r_{-1} + \alpha_2 \dot{o}_{-1} + \alpha_3 \dot{w}_{-1} + \alpha_4 \dot{p}_{-1} + (1+\theta_1 L+\theta_2 L^2+\theta_3 L^3+\theta_4 L^4+\theta_5 L^5+\theta_6 L^6)(1+\Theta_1 L^{12}+\Theta_2 L^{24}+\Theta_3 L^{36})\varepsilon$$

	general model		*simplified model*	
μ	84.27	(6.624)	83.07	(7.027)
ϕ_1	-.1853	(-3.388)	-.1803	(-3.370)
ϕ_2	.1753	(3.798)	.1862	(3.867)
α_1	7.991	(3.548)	8.080	(6.539)
α_2	-.9987	(-2.697)	-.8064	(-2.680)
α_3	-.5547	(-1.156)		
α_4	.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	
$T_{norm}(2)$	14.06†		12.86†	
$T_{ARCH}(1)$	2.414		2.831	
$T_{ARCH}(12)$	22.30†		21.98†	

s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	1.23	4.86	12.99	24.65	32.35	1.98	5.62	13.13	29.48	36.56
$T_{B\text{-}P}(s)$			5.49†	23.73†	31.32			6.98†	25.05†	32.54
$T_{L\text{-}B}(s)$			5.71†	25.38†	33.86			7.15†	26.72†	35.10

$T_{BN(01)}$: (*general model*)

-.1034 .5714 .9563 .5799 .4906 .5405 1.404 .2859 .2981 .3624
.3341 1.268 .5337 .6945 1.060 1.576 .7810 1.464 .5344 .9319

$T_{par}(3)$ 4.075 $(\alpha_3=\alpha_4=\theta_6=0)$

$T_{BN(01)}$: (*simplified model*)

1.603 -.4268 .4058 .6133 1.524 .4766 -.5900 .1554 .5617 -.0217
.2923 -.5943 1.171 .0528 .2209 -.6623 1.058 .0071 1.119 .0246

T_{forec} — simplified model: 1.590

†*significant at a 5% level*

consistent model misspecification test $T_{BN(01)}$ between the model based on (2.11) and the one of table 4.4, learns that it has improved compared to the first model. For the model based on (2.11), $T_{BN(01)}$ exceeded the 10% significance level three out of 20 times, whereas it does not exceed this level once for the model of table 4.4.

The general model can be simplified into a model where only the interest rate and the growth rate of production are significantly Granger causing unemployment. This simplified specification cannot be rejected by any of the important misspecification tests and there is no evidence of predictive failure. Thus, the simplified model of table 4.4 is an adequate representation of the DGP giving rise to the monthly unemployment figures in Germany.

4.6. The United Kingdom

The initial specification (2.11) could not be accepted for UK data, because of serially correlated disturbances. Additional lagged dependent variables, based on the (partial) autocorrelation function of the residuals of this model, were included in the model to repair this misspecification. However, the sum of the AR coefficients of this model is very close to unity, $\Sigma_{j=1}^{4}\phi_j = 0.9951$, indicating near nonstationarity and hence possibly invalid inference.

We therefore turn to specification (2.12). This model could initially not be accepted because of residual autocorrelation. Inspection of the (partial) autocorrelation function of the residuals and taking account of the additional lags of the model based on (2.11) resulted in the model specification of table 4.5. This model cannot be rejected by neither T_{ac} nor $T_{BN(01)}$. In particular the model misspecification test $T_{BN(01)}$ improved when moving to specification (2.12): for the model based on (2.11) it exceeded the 10% significance level twice out of 20 times, whereas for the model in table 4.5 this level is not exceeded.

After simplification, the growth rate of industrial production yielded an insignificant parameter value. The same is true when wage inflation is included in the simplified model. The Wald test on parameter

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

$$\Delta_1 u = \mu + \phi_1 \Delta_1 u_{-1} + \phi_2 \Delta_1 u_{-13} + \phi_3 \Delta_1 u_{-18} + \phi_4 \Delta_1 u_{-19} + \alpha_1 r_{-1} + \alpha_2 \dot{o}_{-1} + \alpha_3 \dot{w}_{-1} + \alpha_4 \dot{p}_{-1} + (1+\theta_1 L+\theta_2 L^2+\theta_3 L^3+\theta_4 L^4+\theta_5 L^5+\theta_6 L^6)(1+\Theta_1 L^{12}+\Theta_2 L^{24}+\Theta_3 L^{36})\varepsilon$$

	general model		*simplified model*	
μ	-1.855	(-.2740)	-1.354	(-.2163)
ϕ_1	.7964	(10.83)	.9469	(38.04)
ϕ_2	-.0718	(-2.066)	-.1067	(-4.222)
ϕ_3	.2083	(3.318)	.2605	(4.078)
ϕ_4	-.1803	(-2.930)	-.2555	(-4.042)
α_1	.5266	(1.644)	.5449	(2.985)
α_2	-.3900	(-1.956)		
α_3	.2439	(1.348)		
α_4	.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	
$T_{norm}(2)$	4.300		4.011	
$T_{ARCH}(1)$	.2059		.0110	
$T_{ARCH}(12)$	10.42		9.185	

s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	1.50	6.43	13.83	22.86	38.62	.121	2.27	14.16	26.86	36.96
$T_{B\text{-}P}(s)$				14.80	27.44			13.10†	22.82	39.86
$T_{L\text{-}B}(s)$				15.75	30.10			13.62†	24.12	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

$T_{par}(9)$ 10.69 $(\alpha_2=\alpha_3=\alpha_4=\theta_2=\theta_3=\theta_4=\theta_5=\theta_6=\Theta_3=0)$

$T_{BN(01)}$: (*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

T_{forec} (simplified model) 2.914

† *significant at a 5% level*

restrictions cannot reject the hypothesis that $\dot{o}_{-1}$, $\dot{w}_{-1}$ and $\dot{p}_{-1}$ and some MA parameters are restricted to zero, since $T_{par}(9)=10.69$. Only the interest rate is significantly Granger causing unemployment and this simplified model is not rejected by any of the crucial misspecification tests. However, this model does suffer from predictive failure, as indicated by T_{forec}.

This might be caused by invalid conditioning, e.g., because the growth of industrial output was not included. However, if this variable, which had an insignificant parameter, was included in the model the predictive performance did not improve. Inclusion of wage inflation could also not improve the predictive performance of the model. We also tried several other interest rates, but the results were very similar to the ones of table 4.5. Cf. Broersma (1991b).

A solution to this problem might be to take account of the introduction in the UK of retail sight deposits in 1984. These are high interest rate bearing accounts where money can be stored for a short period. Correcting the interest rate on debts with a learning adjusted version of this retail sight deposit interest rate might repair this misspecification. Cf. Hendry and Ericsson (1991), where the introduction of this interest rate yielded a stable money demand relation. Unfortunately, monthly figures on this interest rate are not available.

It should however be noted that the simplified model does pass all other important misspecification tests. The interest rate is the only explanatory variable, which is significantly Granger causing unemployment in the UK.

5.2.6 France

In the case of France something peculiar occurred. Observation of the plot the first difference of unemployment, reveals a dramatic change in the seasonal pattern of the series during the early seventies. This is probably the reason why we could not find a specification that was not rejected by either T_{ac} or $T_{BN(01)}$. In order to remove this change, we exclude data before January 1971 from the sample.

Table 4.6. Estimation and test results for France.

$$\Delta_1 u = \mu + \alpha_1 r_{-1} + \alpha_2 \dot{o}_{-1} + \alpha_3 \dot{w}_{-1} + \alpha_4 \dot{p}_{-1} + (1+\theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1+\Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\varepsilon$$

	general model		*simplified model*	
μ	-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	
$T_{norm}(2)$	2.497		3.981	
$T_{ARCH}(1)$	6.902†		2.598	
$T_{ARCH}(12)$	25.34†		18.66	

s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	.087	6.90	9.93	35.95	43.73	3.39	4.42	10.30	33.80	37.57
$T_{B\text{-}P}(s)$			6.64	22.05	36.88		5.01†	9.37	31.53†	47.60†
$T_{L\text{-}B}(s)$			7.01	23.92	41.62†		5.14†	9.73	34.12†	53.24†

$T_{BN(01)}$: (*general model*)

-.9625	.0974	-.0247	-.4448	-.0984	.4867	.2249	-.1651	.1814	-.4746
-.9223	.4519	.2515	-.1522	.1103	.4243	.6039	.0924	-.1616	.3886

$T_{par}(6)$ 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	.5467
.0784	-.4729	.5225	.4043	-1.082	-.1691	-.4003	.6464	.8872	.4327

T_{forec} (simplified model): 1.646

†*significant at a 5% level*

Estimation and testing specification (2.11), yields a model which none of the crucial misspecification tests firmly rejects. Only the AR coefficient is very close to unity, $\phi_1 = 0.9921$, which might yield invalid inference. Thus, we turn to specification (2.12).

The estimation and test results for this model are presented in table 4.6. Again $T_{BN(01)}$ improves dramatically when this model is compared with the results for specification (2.11). For the latter specification the 10% significance level was exceeded six out of 20 times, which already points towards possible misspecification, whereas for the specification of table 4.6 this level was not exceeded once. Note also that, apart from first and twelfth order ARCH, none of the other misspecification tests rejects specification (2.12) for France.

There appears to be near collinearity between the interest rate and inflation. We can simplify the general model of table 4.6 into models in which each appears as explanatory variable, in combination with the percentage growth of industrial production. However, the model with inflation and growth of industrial output, has disturbingly high 24th order residual autocorrelation; $T_{ac}(24) = 36.32$, whereas the 5% critical value is 36.4. We therefore simplify the initial model to the one where the interest rate and the growth rate of industrial production are the only significantly Granger causing variables for unemployment. This simplified model is not rejected by any of the important misspecification tests and the model does not seem to suffer from predictive failure. Thus, it is a good approximation of the DGP giving rise to monthly unemployment in France.

5.2.7 The Netherlands

Finally, we give the estimation and test results for the Netherlands in table 4.7. In January 1976 the way of registration of unemployment changed, which caused a structural break in the unemployment series. Thus, a dummy was included in order to represent this break.

Specification (2.11) could not be accepted because of residual autocorrelation. This result is in agreement with Bierens (1987). Rather than taking the logarithm of unemployment, as was done by Bierens (1987), we included additional lagged dependent variables in the model, based on inspection of the (partial) autocorrelation function of the residuals of this initial model. Its specification passes the misspecification tests, T_{ac} and $T_{BN(01)}$, but large outliers cause excess kurtosis and first and twelfth order ARCH. Moreover, the sum of the AR coefficients points towards a near unit root as $\Sigma_{j=1}^{7}\phi_j = 1.001$. To avoid invalid inference due to near nonstationarity, we move to specification (2.12).

This initial specification was rejected due to serial correlation in the error process. After inspection of the (partial) autocorrelation function of the errors of this model, we included some additional lagged dependent variables. This model is given in table 4.7. Outliers in the early eighties, cause excess kurtosis and the hypothesis of first and twelfth order ARCH cannot be rejected. However, the two crucial misspecification tests, T_{ac} and $T_{BN(01)}$, do not reject the specification. When $T_{BN(01)}$ of the model of table 4.7 is compared with the one based on (2.11), we find a remarkable improvement. In the latter case it is rejected three out of 20 times at a 5% significance level, which already points towards possible misspecification. However, for the specification of table 4.7, this bound is not exceeded once. Hence, this misspecification might be due to near unit root behavior of unemployment and could be repaired by approximating this near unit root by a unit root.

Testing whether the coefficient of $\dot{p}_{-1}$ and a number of MA parameters may be restricted to zero, yields $T_{par}(8) = 8.433$, which cannot be rejected at the usual 5% significance level. Hence, the interest rate, the growth rate of production and the wage inflation are significantly Granger causing unemployment. The specification of this simplified model is not rejected by the crucial misspecification tests, T_{ac} and $T_{BN(01)}$, and there is no evidence of predictive failure over the last 24 periods, as $T_{forec} = 1.357$.

Thus, for the Netherlands not only the interest rate and the growth rate of industrial production Granger cause unemployment, but also the wage inflation. The model we ultimately find seems an adequate representation of the DGP of monthly unemployment in the Netherlands.

Table 4.7. Estimation and test results for the Netherlands.

$$\Delta_1 u = \mu + \phi_1 \Delta_1 u_{-1} + \phi_2 \Delta_1 u_{-5} + \phi_3 \Delta_1 u_{-6} + \phi_4 \Delta_1 u_{-11} + \phi_5 \Delta_1 u_{-23} + \alpha_1 r_{-1} + \alpha_2 \dot{o}_{-1} + \alpha_3 \dot{w}_{-1} + \alpha_4 \dot{p}_{-1}$$
$$+ \alpha_5 D76 + (1 + \theta_1 L + \theta_2 L^2 + \theta_3 L^3 + \theta_4 L^4 + \theta_5 L^5 + \theta_6 L^6)(1 + \Theta_1 L^{12} + \Theta_2 L^{24} + \Theta_3 L^{36})\varepsilon$$

	general model		*simplified model*	
μ	5.787	(2.137)	5.679	(2.463)
ϕ_1	.1334	(2.450)	.2379	(3.652)
ϕ_2	.0511	(1.533)	.0880	(2.411)
ϕ_3	.1157	(1.767)	.1209	(2.586)
ϕ_4	.1368	(1.989)	.1053	(1.677)
ϕ_5	.1333	(2.410)	.1031	(1.993)
α_1	.7845	(2.448)	.7858	(3.526)
α_2	-.1356	(-2.162)	-.1952	(-3.663)
α_3	.3035	(2.353)	.1947	(2.532)
α_4	-.1555	(-.7904)		
α_5	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)		

	general model	*simplified model*
$S.E.$	4.913	4.997
R^2	.8530	.8479
n	312	312
$T_{norm}(2)$	$11.76^{\dagger}$	$10.15^{\dagger}$
$T_{ARCH}(1)$	$13.64^{\dagger}$	$13.85^{\dagger}$
$T_{ARCH}(12)$	$42.27^{\dagger}$	$44.04^{\dagger}$

s	2	6	12	24	36	2	6	12	24	36
$T_{ac}(s)$	1.77	4.15	6.80	27.98	49.10	2.67	4.71	9.96	27.52	47.82
$T_{B\text{-}P}(s)$				$26.98^{\dagger}$	$50.06^{\dagger}$				$34.28^{\dagger}$	$55.77^{\dagger}$
$T_{L\text{-}B}(s)$				$28.69^{\dagger}$	$54.18^{\dagger}$				$36.15^{\dagger}$	$59.93^{\dagger}$

$T_{BN(01)}$: (*general model*)

-.4921	-.8444	-.7949	-.6992	-1.897	-.2934	-.0916	-.0820	-.2561	-1.200
1.187	-.9251	-.8100	-.2474	-.7255	-1.141	-1.430	-1.310	1.190	1.147

$T_{par}(8)$ 8.433 $(\alpha_4 = \theta_1 = \theta_2 = \theta_3 = \theta_4 = \theta_5 = \theta_6 = \Theta_3 = 0)$

$T_{BN(01)}$: (*simplified model*)

-1.771	.1711	.1568	-.6750	.0199	.2401	.3821	.2643	-.4566	-.4448
-.0527	.0822	-.1007	1.773	.1647	.9176	-1.816	.2224	1.010	.5293

T_{forec} (simplified model): 1.357

† *significant at a 5 % level*

5. THEORETICAL IMPLICATIONS

In this section we discuss the theoretical implications of the empirical results of the previous section. All empirical models indicate that the interest rate is Granger causing unemployment, in combination with the annual growth rate of production, except for the UK and the Netherlands. For the first only the interest rate Granger causes unemployment. For the latter also output growth and wage inflation are included.

There are a number of economic theories which provide an explanation of this phenomenon. First, Ashenfelter and Card (1982), in an inductive investigation for the USA, also find a Granger causal relation between unemployment and the interest rate. They explain this phenomenon by extending the labor supply model of Lucas and Rapping (1970). They ultimately derive that unemployment is determined by the difference of the actual and the long run expected real wage rate, where the latter is a combination of lagged expected real wage rates and lagged interest rates. In that way the interest rate appears as Granger causing variable for unemployment.

Second, Farmer (1985) developed a model based on the implicit contract theory augmented with the concept of asymmetric information and limited liability. In this theory the real interest rate plays an important role in causing layoffs. Farmer (1989) provided statistical evidence in favor of this theory, where he confirmed the relation between unemployment and both the real and the nominal interest rate. His theory of asymmetric information and limited collateral boils down to the fact that a firm makes employment contracts with workers and debt contracts with creditors. Payments of the firm to the factors of production can be no greater than the total amount produced and the collateral of the firm.

A neoclassical firm will in first instance hire more workers in response to an interest rate increase. Hence there are more workers associated with one machine, so the marginal unit of employment is less productive in any state of nature. This leads to less favorable employment contracts. In addition, the equilibrium wage rate will fall, but the total factor costs will rise, as the fall in the wage rate is offset by the increase in the number of workers and an increase in the interest costs. The firm is forced to pay a higher expected return to the factors of production, but has limited collateral. If the limited liability constraint is violated, the factors of production will want a bonus in good stats of nature to accommodate the possibility of a loss in bad states. However, this bonus interferes with the firms employment decision by raising the marginal costs of an additional unit of labor above the disutility of employment. The profit maximizing firm makes its employment decision by equating the marginal costs of employment to the marginal revenue. Since the marginal costs schedule is steeper in case of an inefficient contract, the firm will hire less labor than would have been the case in an optimal situation.

Finally, Bierens (1987) explains the Granger causal relation between unemployment and the interest rate by the managerial theory of Baumol (1959) augmented with a flexible labor effort rate. Baumol (1959) asserts that firms, especially large and medium sized firms, are led by managers and are therefore revenue maximizers rather than profit maximizers. Profits only play a role as a constraint: a minimum profit level is requires to safeguard the firms' viability and continuity. The concept of a flexible labor effort rate is in fact very simple: if workers work harder they can produce more. This labor effort rate varies between zero an some upper bound.

A higher interest rate implies more interest payments and hence more fixed costs. If this increase is such that profit falls below its minimum required level, 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 at its maximum level not only employment, but also production decreases up to a point where where the minimum profit constraint is retained. If this is not possible the firm has to shut down and employment is zero.

On the other hand, 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, also demand for nondurables. Thus the drop in profits due to an interest rate increase is aggravated by a drop in demand. An increase in the wage rate has a similar effect on the firm as an interest rate increase, but it has an opposite effect on demand, especially if the wage increase is the result of collective wage bargaining. The increase of demand caused by a wage increase may therefore offset the negative effect on the profits of the firm.

The latter theory seems realistic when the economic history of the last two decades is considered. It is also in agreement with our empirical results. Apart from the Netherlands the wage rate does not have a significant influence on unemployment. However, the interest rate, in combination with the growth rate of industrial production, significantly Granger causes unemployment in all countries under review.

6. CONCLUDING REMARKS

It appeared that with our ARMAX modeling approach, using monthly data, we found a Granger causal relation between unemployment and interest rate in combination with the growth rate of the industrial production. For the UK only the interest rate was important and for the Netherlands also the percentage change in the wage rate played a role.

This study challenges the neoclassical theory of the firm. Instead of the marginal costs, i.e., the wage costs, the fixed costs, via the interest costs determine the employment decisions of firms. According to the neoclassical theory of the firm, fixed costs do not have any influence on employment; only marginal costs have. Our ARMAX models indicate that the wage rate has no significant impact on unemployment, except for the Netherlands, whereas the interest rate has for all countries.

These empirical results also provides evidence against new classicals, as Lucas (1975) and Sargent and Wallace (1975), who claim that monetary policy has no real effect, even in the short run. However, interest rate policies, aimed at controlling money supply, might have dire consequences for employment.

Finally, we mentioned some economic theories, in both neoclassical and Keynesian tradition, which can account for our empirical findings.

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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)

Empirical distribution of Z_{α} under H_0 (cf. Fuller (1976), pp. 371).

	Probability of a smaller value							
n	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
U.K.:	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 PERCNT PA
Canada:	Official discount rate:	CAN.OFF.DISC.RATE PERCNT PA
Japan:	Official discount rate:	JPN.OFF.DISC.RATE PERCNT PA

Germany:	Official discount rate:	DEU.OFF.DISC.RATE PERCNT PA
U.K.:	Call money rate:	GBR.CALL.MON.RAT PERCNT PA
France:	Call money rate:	FRA.CALL.MON.RAT PERCNT PA
Netherlands:	Official discount rate:	NLD.OFF.DISC.RATE PERCNT 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, manufac.:	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
U.K.:	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
U.K.:	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 gds:	USA.PPI.FIN GDS I/80
Canada:	Producers price index, manufac. gds:	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
U.K.:	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

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