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Real Variables, Nonlinearity, and European Real Exchange Rates

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ABSTRACT

We carry out an analysis of European real exchange rate behaviour before and after the implementation of Economic and Monetary Union (EMU). In particular, we model real exchange rates for a number of EMU and non-EMU countries against Germany in an explicitly nonlinear framework and allow for variation in the equilibrium level of the long-run equilibrium real exchange rate using either relative productivities or real diffusion indices. The estimated models show that real variables are a significant determinant of long-run real exchange rates when incorporated into a nonlinear framework. We also find that the speed of adjustment is generally faster after the implementation of EMU.

Keywords: real exchange rates; nonlinearity; diffusion indices.
JEL Classification: F31

1. INTRODUCTION

In this paper we carry out an analysis of European real exchange rate behaviour before and after the implementation of Economic and Monetary Union (EMU), i.e. the single European currency, in January 1999. In particular, we model real exchange rates for a number of EMU and non-EMU countries against Germany in an explicitly nonlinear framework and allowing for variation in the equilibrium level of the long-run equilibrium real exchange rate using either relative productivities or real diffusion indices.

The relative productivity specification derives from the well known Harrod-Balassa-Samuelson effect (Harrod, 1933; Balassa, 1964; Samuelson, 1964). According to the Harrod-Balassa-Samuelson effect, countries with rapidly expanding economies should tend to have rapidly appreciating real exchange rates.

While the Harrod-Balassa-Samuelson effect focuses on a few series in order to explain the equilibrium level of the real exchange rate, the long-run equilibrium real exchange rate may, however, be impacted by a wider range of real variables in the macroeconomy. Including the wide range of available real variables in an econometric specification, however, raises a number of practical problems for the a modeler, notably the lack of degrees of freedom as well as potential multicollinearity. One way of circumventing this approach is to construct diffusion indices or factors that capture the core variability in a set of macroeconomic tiems series in a parsimonious fashion (Stock and Watson, 1998, 2002a,b; Bernanke and Boivin, 2003; Bernanke, Boivin, and Eliasz, 2005).

The remainder of the paper is organised as follows. In the next section we examine the underlying rationale for nonlinear real exchange rate adjustment while in Sections 3 and 4 we briefly discuss the Harrod-Balassa-Samuelson and diffusion index methods of capturing variation in the equilibrium real exchange rate, respectively. In Section 5 we discuss how these approaches can be incorporated in a nonlinear empirical model and Section 6 we describe our data set. Our empirical results are reported in Section 7 and in Section 8 we make some concluding remarks.

2. NONLEARITY AND EXCHANGE RATE DYNAMICS

The idea of nonlinear adjustment in real exchange rate adjustment has been put forward in a number of papers over the last decade or so. Proponents of this approach argue that exchange rates are relatively insensitive to fundamentals close to equilibrium values but have a tendency to mean revert strongly as the deviation from equilibrium becomes more pronounced. Transaction costs in international trade, such as shipping costs, are often cited as a source of nonlinear adjustment (Dumas, 1992; Sercu et al; 1995; O'Connell, 1997), although heterogeneity of opinion concerning the equilibrium level of the exchange rate or official intervention may also induce nonlinearity (Kilian and Taylor, 2003; Taylor, 2004; Reitz and Taylor, 2008), as may heterogeneous speeds of adjustment in prices at a disaggregated goods level (Imbs et al, 2005)[1].

Taylor and Peel (2000), Taylor, Peel and Sarno (2001), Kilian and Taylor (2003) and Lothian and Taylor (2008) investigate the plausibility of nonlinear exchange rate adjustment using the smooth transition autoregressive (STAR) family of nonlinear models, which allow the degree of mean reversion of a serially correlated process to be smooth function of the distance from equilibrium. Using the dollar-sterling and dollar-mark nominal exchange rate, for example, Taylor and Peel (2000) find that the exponential smooth transition autoregressive (ESTAR) model parsimoniously describes the deviation from an exchange rate equilibrium determined by monetary fundamentals. Taylor, Peel and Sarno (2001) also fit nonlinearly mean-reverting models to real dollar exchange rates over the post-Bretton Woods period. In their study of the half lives of real exchange rates shocks they find faster adjustment as the size of shocks increases while for smaller shocks the exchange rate becomes more persistent and, in the neighbourhood of equilibrium, unpredictable. The predictability of exchange rates in a nonlinear setting is assessed by Kilian and Taylor (2003), who develop a bootstrap test of the random walk hypothesis in a nonlinear setting and provide evidence of predictability at longer horizons of two or three years. Lothian and Taylor (2008) examine the Harrod-Balassa-Samuelson effect in a nonlinear framework for the US, UK and France over a period spanning nearly two centuries and find significant evidence of nonlinear real exchange rate mean reversion towards an equilibrium that is a function of relative productivity.

In sum, the empirical research reveals that the concept of nonlinearity appears to capture well the salient characteristics of exchange rate dynamics. This nonlinearity may also explain the difficulty researchers have encountered in rejecting the unit root hypothesis for real exchanger rates and the observed slow speeds of adjustment in cases where significant mean reversion has been detected—i.e. the 'Purchasing Power Parity Puzzle' (Rogoff, 1996). In particular, if the true data are generated by a nonlinear process then standard unit root tests have very weak power to reject a false null

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[1] Their study suggests that differing speeds of adjustment at a disaggregated good level can be translated into an 'aggregation bias' that employ linear econometric techniques.

hypothesis and the adjustment speed of real exchange rates may be severely underestimated (Taylor, Peel and Sarno, 2001; Taylor and Taylor, 2004).

3. THE HARROD-BALASSA-SAMUELSON EFFECT

According to the well known Harrod-Balassa-Samuelson effect, the long-run equilibrium real exchange rate should depend on the productivity of the tradable and nontradable sectors in the home and foreign economies. Historically, productivity growth in traded goods sectors has been faster than in nontradable sectors. If the law of one price holds, then the prices of tradables will tend to be equalized across countries, while the prices of nontradables may not. Faster productivity growth raises wages in the tradables sector and, with labour mobile within the economy, wages in the entire economy will rise, leading to a rise in the price of nontradables (since the wage rise is not offset by productivity growth in that sector) an an increase in the overall price index (a function of both tradables and nontradables prices), resulting in an appreciation of the real exchange rate, other things equal.

The standard Harrod-Balassa-Samuelson effect may be illustrated in a simple two-country, two-goods context (Froot and Rogoff, 1991) as follows. A standard Cobb-Douglas technology is assumed:

$$Y_I = A_I K_I^{\phi_I} \qquad I=T, N \tag{1}$$

where T_I, K_I and A_I denote, respectively, the output-labor ratio, the capital-labor ratio and productivity in sector T (tradables) or N (nontradables), respectively. Suppose factors are perfectly mobile across the tradable and nontradable sectors and the two sectors are characterized by perfect competition, so that there is long-run real rate of return equalization in capital across economies and long-run real wages are equalized in the long run across sectors within economies:

$$R = \phi_T A_T K_T^{\phi-1}, \qquad R = \frac{P_N}{P_T} \phi_N A_N K_N^{\phi-1} \tag{2}$$

$$W = (1-\phi_T) A_T K_T^{\phi_T}, \qquad W = \frac{P_N}{P_T}(1-\phi_N) A_N K_N^{\phi_N} \tag{3}$$

where R denotes the real long-run return to capital and W is the long-run real wage rate (measured in tradables) and P_{NT}/P_T is the relative price of nontradables to tradables. Taking logarithms and totally differentiating (2) and (3), we derive:

$$d\ln K_N = dInK_T = dInW = d\ln A_T (1-\phi_T)^{-1} \tag{4}$$

$$dIn(\frac{P_N}{P_T}) = (1-\phi_N)(1-\phi_T)^{-1} dInA_T - dInA_N \tag{5}$$

Equation (5) incorporates the Harrod-Balassa-Samuelson condition that relatively higher productivity growth in the tradables sector will tend to generate a rise in the relative price of nontradables. Integrating (5), we can obtain the following logarithmic form of the relative price of nontradables (indicating logarithms by the use of lower case letters):[2]

$$p_N - p_T = (1-\phi_N)(1-\phi_T)^{-1} a_T - a_N \tag{6}$$

The overall price level, p, is a geometric average of its decomposed into its tradable and nontradable components:

$$p = p_T + (1-\lambda)(p_N - p_T)\,. \tag{7}$$

Assuming that the law of one price holds among tradable goods:

$$p_T^* = s + p_T \tag{8}$$

where s is the logarithm of the nominal exchange rate (foreign price of home currency). Provided that equations similar (6) and (7) hold in the foreign economy, the following expression for the long-run real exchange rate is obtained:

$$s - p^* + p = (1-\lambda)[(\frac{1-\phi_N}{1-\phi_T})a_T - a_N] - (1-\lambda)[(\frac{1-\phi_N^*}{1-\phi_T^*})a_T^* - a_N^*] \tag{9}$$

For simplicity, if we assume that productivity in the nontradables sector in each country is close to zero, equation (9) simplifies to

$$s - p^* + p = (1-\lambda)[(\frac{1-\phi_N}{1-\phi_T})a_T] - (1-\lambda)[(\frac{1-\phi^*_N}{1-\phi^*_T})a_T^*] \tag{10}$$

Equation (10) illustrates the Harrod-Balassa-Samuelson effect: relative high levels of productivity in tradables will generate a real exchange rate appreciation. Equivalently, rich countries will tend to have a higher exchange rate-adjusted price level on average.

The empirical evidence provides mixed results on the Harrod-Balassa-Samuelson effect, as surveyed by Froot and Rogoff (1995) and Taylor and Taylor (2004). There are several explanations for such mixed results. Bergin, Glick and Taylor (2006), for example, propose models with endogenous tradability of products, and suggest that this effect has been variable over time. Devereux (1999) also argues that the positive

[2] For simplicity, we ignore the constant of integration.

link between relative productivity and real exchange rates can be reversed in some cases with the presence of endogenous productivity gain in distribution services.[3] In this framework, the formal presentation of the reversed Harrod-Balassa-Samuelson effect may be introduced with the price of tradables as a sub-price index given by wholesale price of tradables, $P_{W.}$ and the price of distribution services, P_D;

$$P_T = P_W^{\beta} P_D^{1-\beta} \tag{11}$$

In addition, assume the distribution sector is comprised of a continuum of monopolistically competitive firms of of total measure θ and the technology for the production of the distribution sector is

$$X(i) + \delta = A_x K_x^{\phi_T} \tag{12}$$

where *X(i)* is output of distribution services of firm *i* and δ is a fixed cost. The consumption distributional sector services is given by

$$D_T = (\int_0^{\theta} X(j)^{\rho} dj)^{\frac{1}{\rho}} \tag{13}$$

where $0 \leq \rho \leq 1$. Equation (13) implies there are increasing returns to specialization in the distribution sector. θ represents an endogenous productivity effect in distributional services and ρ is the elasticity of substitution between distributional services (The lower ρ, the greater is the strength of this specialization effect).

The individual firms that produce specialized distribution services are monopolists, and set price at a mark-up over marginal cost. In a symmetric equilibrium, this gives

$$P_x = \frac{1}{\rho}\frac{W}{A_x}, \quad P_w = \frac{W}{A_T}. \tag{14}$$

The efficient pricing of the distribution composite implies

$$P_D = \theta^{1-\frac{1}{\rho}} P_x \tag{15}$$

Now suppose that productivity in the distribution sector grows at the same rate as for tradables; a_T=a_W and endogenous productivity, θ, grows as

$$d \ln \theta = (\frac{1}{1-\phi_T}) da_T \tag{16}$$

[3] In particular, Devereux (1999) explains the lack of strong real appreciation in the Hong Kong and Singapore dollars as attributable to fast productivity growth in the service sector. Supporting empirical evidence is also presented by Muscatelli, Spinelli and Trecroci (2007).

Then, taking logarithms of equation (11), (14), and (15), the domestic price level is determined by[4]

$$p = [(\frac{1}{1-\phi_T})(1-\frac{1}{\rho})(1-\beta)\lambda + (\frac{1-\phi_N}{1-\phi_T})(1-\lambda)]a_T - (1-\lambda)a_N \qquad (17)$$

From equation (17), keeping productivity levels in the foreign country constant, the real exchange rate may tend to appreciate or depreciate. If β=1, the Harrod-Balassa-Samuelson effect applies since the growth of productivity of a_T increases the price level. But when β<1, the first term in equation (17) is negative, decreasing the domestic price level, and the real exchange rate possibly depreciates as a result of the endogenous productivity "deepening" in the distribution sector.[5]

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[4] For simplicity, we take the wholesale price of tradables as numeraire ($P_W=1$).

[5] This is more likely with smaller ρ (i.e. a greater specialisation effect).

4. DIFFUSION INDICES

The main advantage of diffusion index analysis is to summarise the information of large data sets in a small number of factors. Suppose that X_t denotes an n-dimensional vector of $I(1)$ time series at time t, where n may be very large. If X_t is described by a factor model, it can be written as the sum of two orthogonal components:

$$X_t = \Lambda_t F_t + \xi_t \tag{18}$$

where $X_t=(x_{1t}\ldots x_{nt})'$ and ξ_t is the (n×1) idiosyncratic disturbance, which may be serially correlated. F_t is the (r×1) vector of integrated process latent common factors such that

$$F_t = F_{t-1} + u_t \tag{19}$$

where u_t is a vector zero-mean $I(0)$ process. F_t is, however, unobservable. There are several indirect ways to estimate unobservable factors. In this study, we use a two-step principal components method, following Stock and Watson (2002a) and Bai (2004). This involves solving the eigenvalue problem for the sample covariance matrix:

$$\Xi_0 Y_r = Y_r D_r \tag{20}$$

where Ξ_0 , estimated as $XX'/(T^2N)$, is the covariance matrix of the standardised data matrix for a sample size T, $X=[X_1\ X_2 \ldots X_T]'$ and $Y_r=[y_1 \ldots y_r]$ is the $(n\times r)$ matrix whose columns are the r eigenvectors that correspond with the first r largest eigenvalues of the covariance matrix, Ξ_0. The matrix D_r then denotes a diagonal matrix representing the first r largest eigenvalues. From the solution to this problem, the first r principal components are defined as:

$$\hat{F}_t = \mathrm{Y}_r' \hat{X}_t \tag{21}$$

Since the principal components are by construction orthogonal to one another, there is no redundant information in individual factors. The estimated common factors are $I(1)$ processes since the $I(1)$ process X_t is represented in levels.

In practice, n, the number of time series making up X_t, may be very large — perhaps of the order of fifty or a hundred. If most of the variation in X_t is contained in the first r common factors, however, then the dimensionality of the data problem may be reduced from n to r, and r will typically be a small number. Insofar as the resulting factors or diffusion indices can be used to augment an econometric model, the applied modeller may then be able to reflect the complexities of the economy in a parsimonious fashion with a workable number of degrees of freedom.

5. EMPIRICAL SPECIFICATION OF NONLINEAR REAL EXCHANGE RATE ADJUSTMENT

Smooth transition autoregressive (STAR) models were first generalised into the econometrics literature by Granger and Teräsvirta (1993). In a STAR model the speed of adjustment towards equilibrium takes place in every period but the speed varies smoothly with the extent of the deviation from equilibrium. As discussed briefly above, there are sound theoretical reasons for expecting nonlinear real exchange rate adjustment and STAR models have indeed been successfully applied to real and nominal exchange rates (e.g. Michael, Nobay and Peel, 1997; Taylor, Peel and Sarno, 2001; Taylor and Peel, 2000; Kilian and Taylor, 2003).[6] In line with this previous exchange rate research, our paper focuses on the exponential smooth transition autroregressive or ESTAR specification, which implies symmetric adjustment above and below equilibrium and which has been found to be particularly applicable to exchange rates. The ESTAR specification for a time series process q_t may be written:

$$q_t = \sum_j^p \alpha_j q_{t-1} - \sum_j^p \beta_j [q_{t-j} - \mu]\Phi[\gamma; q_{t-d} - \mu] + e_t . \qquad (22)$$

where $\sum_j^p \alpha_j = 1$ and $\sum_j^p [\alpha_j - \beta_j] < 1$. In the present application, q_t is the real exchange rate defined as

$$q_t = s_t + p_t - p_t^* \qquad (23)$$

and $\Phi[\gamma; q_{t-1}\text{-}\mu]$ is the exponential transition function specified as

$$\Phi[\gamma; q_{t-1} - \mu] = 1 - \exp[-\gamma(q_{t-1} - \mu)^2] \qquad (24)$$

The exponential transition function determines the degree of mean reversion and is itself governed by the nonlinear adjustment parameter, γ, which effectively determines the speed of mean reversion and the parameter μ which is the equilibrium level of the real exchange rate. $\Phi[\gamma; q_{t-1}\text{-} \mu]$ is bounded by zero and unity and is

[6] An alternative nonlinear model, the threshold autoregressive model, allows for a transactions costs band within which no adjustment take place. But many of theoretical studies suggest that, for aggregate real exchange rates, smooth rather than discrete adjustment may be more appropriate in the presence of proportional transaction costs, time aggregation and nonsynchronous adjustment by heterogeneous agents (Taylor, 2004).

symmetrically inverse bell-shaped around zero. The transition parameter $\gamma>0$, determines the speed of transition between the two extreme regimes of random walk and a mean-reverting AR(p) specification, with lower absolute values of gamma implying slower transition.

When the real exchange rate is very close to the equilibrium level μ, for instance, the nonlinear part of equation (22) disappears as $\Phi(0)\approx 0$, and the real exchange rate is close to a unit root process:

$$q_t = \sum_j^p \alpha_j q_{t-1} + e_t, \quad \sum_j^p \alpha_j = 1 \tag{25}$$

As departures from the equilibrium increase, however, the argument of the transition function gets larger and larger and, in the limit, the transition function approaches unity: $lim_{|x|\to\infty}\ \Phi[\gamma;\ q_{t-1} - \mu]=1$. In the limit, therefore, (22) becomes a mean-reverting AR(p) model:

$$q_t = \sum_j^p [\alpha_j - \beta_j] q_{t-1} + e_t, \quad \sum_j^p [\alpha_j - \beta_j] < 1 \tag{26}$$

For intermediate deviations of the real exchange rate from equilibrium, the process will display an intermediate speed of adjustment.

Equation (24) implicitly assumes that the real exchange rate has a constant equilibrium, μ, but we can relax this assumption by allowing for a time-varying long-run equilibrium. In this paper, two types of long-run equilibrium are examined; one determined by the Harrod-Balassa-Samuelson effect by including productivity differentials in the specification, the other augmented by diffusion indices.

First, we allow the Harrod-Balassa-Samuelson (HBS) effect as a determinant of long-run equilibrium in the real exchange rate by modelling time-variation in the equilibrium level as a function of relative productivity:

$$\mu_t = \mu + \mu_1 (a_t - a_t^*) \tag{28}$$

where a_t and a_t* represent the productivity of the home and foreign economy, respectively.

If diffusion indices are incorporated into the long-run equilibrium level of q_t, μ may be represented as;

$$\mu_t = c + bF_t \tag{27}$$

By combining the factors in a nonlinear function, the equilibrium real exchange rate becomes time varying and dependent on macroeconomic fundamentals estimated by the common factors or diffusion indices. We term ESTAR models of this kind, in which diffusion indices or macroeconomic factors have been added into the specification, 'factor-augmented ESTAR' or FESTAR models.

A key innovation of the research reported in this paper is the incorporation of macroeconomic variables as the determinants of the long-run equilibrium real exchange rate. Previous attempts to allow for time-variation of the equilibrium real exchange rate in nonlinear framework have typically focused on a very small number of variables such as a relative productivity (e.g. Paya and Peel, 2007; Lothian and Taylor, 2008).

6. DATA

Our data sets consist of two parts: real exchange rates and macroeconomic real variables. Given our European focus, we concentrated on a set of intra-European real exchange rates against Germany. In particular, we considered German exchange rates against five major European economies—the UK, Switzerland, Denmark, France and Austria. Among these, the UK, Switzerland and Denmark have not participated in EMU while, in common with Germany, France and Austria have adopted the single European currency, the Euro as, their official currency since January 1999. For the EMU countries, we therefore use a floating nominal exchange rate against the German mark before 1999 and a fixed nominal rate afterwards. Domestic and foreign price levels were approximated by monthly observations on consumer price indices (CPI). Then, the real exchange rate was constructed with these data in logarithmic form as in equation (23), with s_t, taken as the logarithm of the nominal exchange rate against the German mark, p_t, as the logarithm of the German consumer price level, and p_{t^*} as the logarithm of the consumer price level of the second European country. Nominal exchange rates after January 1999 were constructed based upon a floating Euro rate adjusted by the fixed conversion rate at the time of conversion to EMU, although this in fact affects only the fixed mean of the real exchange rate. The real exchange rates were then normalized on the beginning of each observation period.

The other part of data sets is comprised of macroeconomic time series for all economies, taken from DATASTREAM. The composition of the monthly data sets was determined by the data availability which represents the real variables such as real production, employment, consumption, orders, earnings, retail sales. The total number of time series collected was 19, 22, 14, 16, 21 and 17 for Germany, the UK, Switzerland, Denmark, France and Austria, respectively.[7] In addition, the productivity of one country

[7] If series are only available in quarterly basis, they are interpolated to the monthly frequency. In the case

is measured by real income data divided by total employment.[8]

Taking post German reunification as a starting point for the data set, our sample period runs from January 1991 to June 2007. The data sample period is further split into a period of pre-and post-implementation of EMU in January 1999 in order to see the impact of introduction of the euro on real exchange rate adjustment. For the pre-implementation period, the data therefore ranges from January 1991 to December 1998 (96 observations) and the EMU period covers from January 1999 to June 2007 (102 observations).

7. EMPIRICAL RESULTS

7.1 Movement of real exchange rates since January 1991

As a preliminary examination of the data, the real exchange rates of the five European economies for whole sample period are plotted in Figure 1. Since real exchange rates of these economies are calculated against German mark, a downward shift means the depreciation of the home currency while an upward movement means appreciation. During the first half of 1990s there was a strong tendency for real depreciation for all currencies against the mark. From the second half of the 1990s to 2007, however, the real exchange rates against the mark show a hump-shape for the UK and Switzerland, indicating a notable appreciation until the early 2000s followed by steady depreciation. In the case of Denmark, France and Austria, they exhibit a common pattern of steady appreciation since the mid 1990s. When we compare the movement of real exchange rates depending on the status of EMU membership, non-EMU countries like the UK and Switzerland show the more volatile real exchange rates during 2000s than those of EMU countries (France and Austria). Real exchange rate movements for Denmark look similar to those of EMU countries, however.[9]

7.2 Harrod-Balassa-Samuelson Effect

In this section, we report the results of real exchange rate specification with an allowance for time-varying long-run equilibrium based on the Harrod-Balassa-Samuelson effect. Examination of the partial autocorrelation functions indicated that a first-order autoregressive model would be adequate in every case and a first delay

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of Switzerland, the majority of data sets are converted to monthly series. More details on the data set are given in the appendix.

[8] Ideally, one would like to obtain data on tradable sector output and employment, but data on capital inputs is notoriously unreliable as Froot and Rogoff (1995) indicate, hence we use aggregate productivity.

[9] Currently, Denmark is a member of the second stage ERM (European Exchange Rate Mechanism).

parameter was also chosen, based on a set of nested likelihood ratio tests.[10] Further, we tested the restrictions on the autoregressive parameters α=1, β=-1 and in no case could we reject at the five percent significance level. These restrictions imply an equilibrium of the real exchange rate, in the neighborhood of which q_t is close to a random walk, becoming increasingly mean reverting with the absolute size of the deviation from equilibrium. Thus, the estimated ESTAR model was:

$$q_t = q_{t-1} - (q_{t-1} - \mu)\Phi[\gamma; q_{t-1} - \mu] + e_t \tag{29}$$

When the relative productivity, measured by the aggregate term (real income per employee) is included in the model, the model is specified as

$$\begin{aligned} q_t &= q_{t-1} - [q_{t-1} - \mu_0 - \mu_1(a_{t-1} - a^*_{t-1})] \times \\ &\{1 - \exp[-\gamma(q_{t-1} - \mu_0 - \mu_1(a_{t-1} - a^*_{t-1})]^2 / \sigma_q^2\} + e_t \end{aligned} \tag{30}$$

where a_t* indicates German productivity and a_t is the productivity of the corresponding country. In order to pick up the Harrod-Balassa-Samuelson effect, the sign of μ_1 should be positive. We also found that the μ_0 terms were in every case estimated as insiginificantly different from zero and so were also excluded. The final parsimonious, estimated form of equation (30) is reported in Table 3.

From Table 3, nonlinear estimation with allowance for the Harrod-Balassa-Samuelson effect can be summarized by the following key features. In both periods, '*t-statistics*' for the estimated value of γ suggest that it is significantly different from zero. However, the ratio of this estimated coefficient to its standard error cannot be referred to the Student-*t* or normal distribution for purposes of inference, because under the null hypothesis H_0: γ=0 the process becomes a linear unit-root process.[11] Since the distribution of γ is unknown under the null hypotheses, we calculated the empirical significance levels using Monte-Carlo simulation under the null hypothesis of a unit root AR(1) process, i.e, γ=0, inorder to generate the empirical significance level.[12] This is reported in square brackets in Table 3. The estimated transition parameter is significantly different from zero in most cases except in the first period in the UK case and in both periods for the Swiss franc.[13] In the EMU period especially, strong evidence of nonlinear mean reversion is accompanied by faster adjustment speeds in term of higher estimated values of γ for the UK, Denmark, France and Austria.

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[10] See Granger and Teräsvirta (1993) or Taylor et al. (2001).

[11] Analogous to the way in which the distribution of the Dickey-Fully statistic cannot be assumed to be Student's *t*.

[12] The method of estimation of the empirical significance levels follows the procedure described in Taylor et al (2001), with 5,000 simulations.

[13] For Austria, the empirical significance level lies on the ten percent border line of the rejection region.

When we look at the Harrod-Balassa-Samuelson effect, the estimated coefficient for the relative productivity term, μ_1 is strongly significantly different from zero for the case of the UK (both periods), Switzerland (second period), Denmark (second period) and Austria (both periods). Among them, the sign of coefficient in relative productivity is correctly signed for Denmark and Austria: relative higher productivity generates a real appreciation of the equilibrium value. But the sign is reversed to the Harrod-Balassa-Samuelson effect for the UK, which may possibly be attributed to the growth of service sector in Germany post reunification. For Denmark (first period) and France (both periods), however, there is no significant evidence of such an effect.

6.3. Diffusion Indices as Determinants of Long-Run Equilibrium

As specified in Section 2, real factors for each country were estimated using two-step principal components for each period (pre- and post- implementation of EMU). From the estimated factors based on real variables, we calculated the percentage of total variability of the data sets explained by each factor and found that the first principal component explains the major portion of total variability. The actual contribution of the first principal component to total variability is reported for each country in Table 1. For all economies, the first principal components explains around a half to three quarters of variability of the data set, ranging from 45 to 77 percent. Given this result, we fixed the number of factor included in both periods at one.

The relationship between the estimated factor and real variables was further examined by simple OLS regression in order to shed light on which real variables are most important for the estimated factor. For example, each of the series is first differenced and regressed against the first differenced empirical factor and the R^2 of the regressions are graphed in Figure 2. As Figure 2 shows, real output is crucial for most countries except the UK where retail trade looks most important in the first estimated factor. For Germany, the volume of retail trade in particular plays an important role.

If a long-run relationship exists between the estimated factor and the real exchange rates, these variables should move in a common direction in the long run. In order to detect this trend, we conducted Johansen's cointegration test in for a VAR involving the real exchange rate and the diffusion indices, $y=[q_t, f_t, f_t^*]$, and the results are reported in Table 2. According to both the trace and the maximum eigenvalue statistics, we can reject at standard significance levels the null hypothesis of no cointegration between the real exchange rate and the diffusion indices. This evidence of la ong-run relationship motivated us to construct the nonlinear real exchange rate specification which includes diffusion indices as determinants of long-run equilibrium.

To be specific, one factor or diffusion index from each economy is included into the nonlinear system as

$$\begin{aligned} q_t &= q_{t-1} - (q_{t-1} - b_1 f_{t-1} - b_2 f^*_{t-1} - c)\times \\ &\{1 - \exp[-\gamma(q_{t-1} - b_1 f_{t-1} - b_2 f^*_{t-1} - c]^2 / \sigma_q^2\} + e_t \end{aligned} \tag{31}$$

where f_{t-1} and f^*_{t-1} represent the home and foreign factors respectively. With the inclusion of factors, we can allow for the effect a number of macroeconomic time series on the real exchange rate in a parsimonious manner.[14] Following a general-to-specific procedure, insignificant factors were excluded and the final results of the estimated FESTAR models, estimated by nonlinear least squares, are reported in Table 4.

The estimation results can be summarized by the following key features. First, the '*t-statistics*' of the nonlinear transition parameter appear large enough to ascertain the significance of γ. When we calculated the empirical significance level as described in the previous section, most real exchange rates show evidence of nonlinearity except during the pre-EMU for the UK and Switzerland. Second, among the significant estimated transition parameters, the size increases during the EMU period for the UK, Denmark, France and Austria. The higher estimated value of γ in the EMU period suggests that the adjustment of real exchange rates in the presence of shocks shows faster mean reversion. In the case of Switzerland, the transition parameter is insignificant and has a smaller value in the second period.[15] In the choice between ESTAR and FESTAR, the estimated factors or diffusion indices are shown to be a significant determinant of real exchange rates for the UK, Switzerland and Austria while the Danish and French data prefer the simpler ESTAR specification.

Overall, the results of our analysis strengthen the evidence for nonlinearities in real exchange rate adjustment and suggest that the speed of adjustment is faster for the EMU period. Faster mean reversion of real exchange rates after the implementation of EMU can be understood in the context of the rapid change of economic and monetary integration taking place in Europe.

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[14] Kim and Taylor (2008) calculate the speed of adjustment of real exchange rates for dollar-sterling, dollar-yen and dollar-Australian dollar in the Bretton-Woods period. They find that ignoring those influences in a standard univariate evaluation of nonlinear specification leads to biased results in measuring real exchange rate adjustment.

[15] In Figure 1, the Swiss real exchange rate shows an unusual degree of volatility given the observation period, which look notably different from the other real exchange rates.

8. CONCLUSION

A key contribution of the research reported in this paper is the incorporation of real variables into a nonlinear framework as determinants of a number of intra-European real exchange rates. We found strong evidence of nonlinearity of most of the European real exchange rates examined and evidence of time variation in the equilibrium level of the real exchange rate. When the equilibrium real rate was modeled as a function of relative productivities, the Harrod-Balassa-Samuelson effect found some empirical support, although some of results showed the reversed sign for the effect, which may suggest that the relative productivity gains came from a "deepening" of the distributional service sector. Using real diffusion also allowed us to capture time variation in the real exchange rates. In both cases, a faster speed of adjustment of real exchange rates in the post EMU period was indicated.

Figure 1. Real Exchange Rates (January 1991 – June 2007)

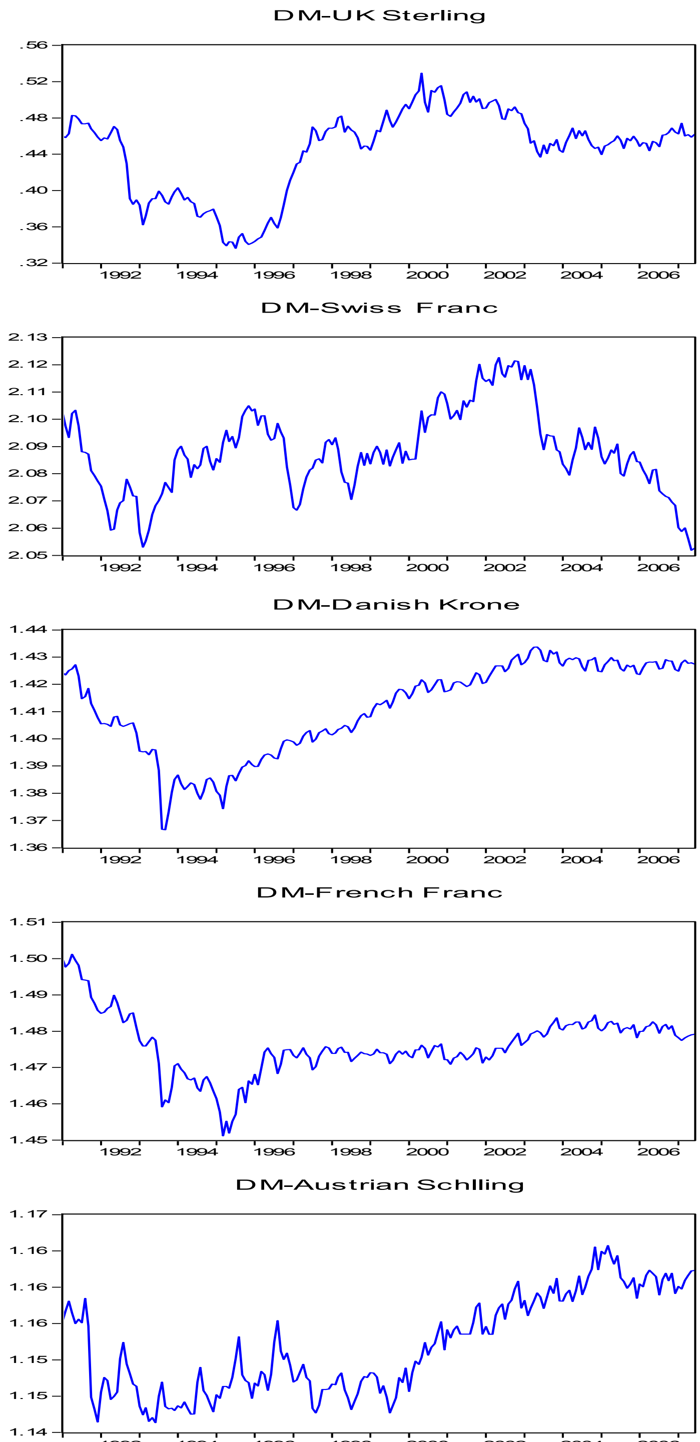
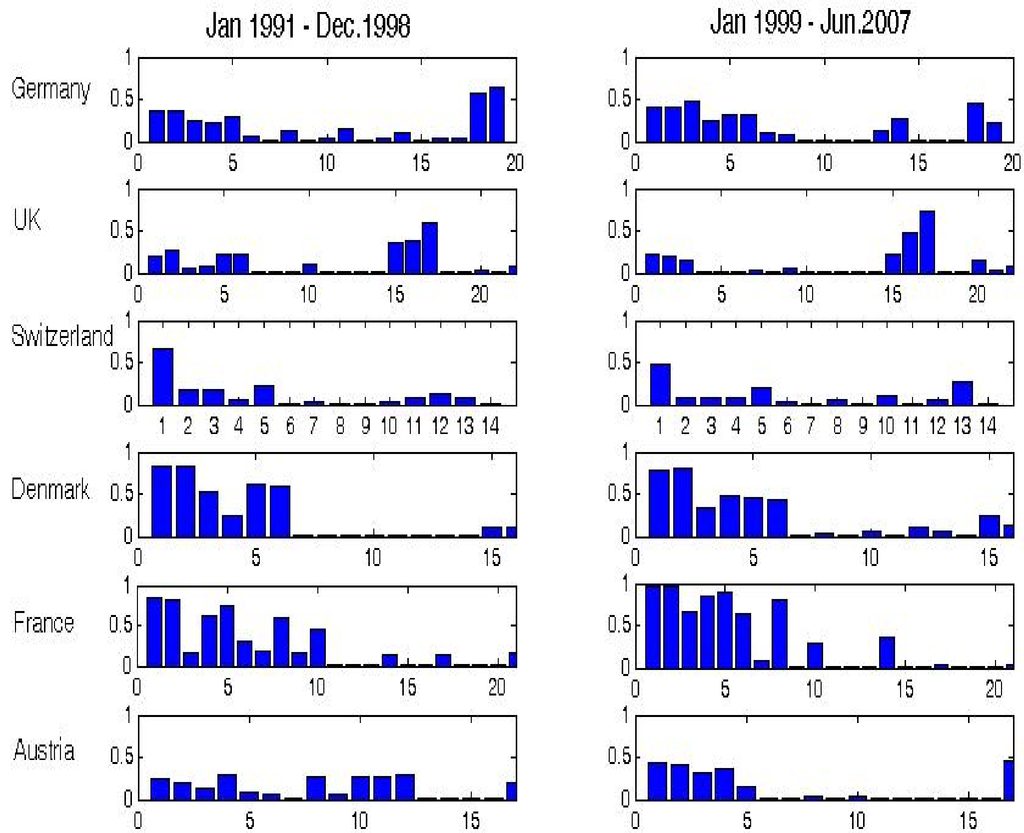


Figure 2. R^2 of Each Real Variable against the First Factor



Note: Since series and factors are nonstationary, they are first-differenced before applying the regression. Figures on the horizontal axis correspond to the numerical series ID given in the appendix.

Table 1. Percentage of Total Variability Explained by First Principal Component

	Jan. 1991-Dec.1999	Jan. 2000-Jun.2007
German	51.6	51.0
UK	65.5	58.8
Switzerland	45.2	53.8
Denmark	77.1	53.4
France	47.0	47.3
Austria	70.4	66.1

Table 2. Cointegration tests between factors and real exchange rates

	Jan.1991 – Dec. 1998			Jan.1999 – Jun. 2007		
		Statistics			Statistics	
	Lags	Trace	Max	Lags	Trace	Max
UK	3	44.01*	29.06*	5	34.32*	29.43*
		(0.000)	(0.000)		(0.002)	(0.000)
Switzerland	2	45.93*	37.40*	8	31.57*	20.71*
		(0.000)	(0.000)		(0.005)	(0.017)
Denmark	4	43.10*	29.45*	5	34.50*	23.90*
		(0.000)	(0.002)		(0.013)	(0.019)
France	5	51.02*	27.12*	5	32.74*	20.36*
		(0.000)	(0.019)		(0.003)	(0.016)
Austria	2	55.72*	42.56*	3	22.43*	18.9*
		(0.000)	(0.000)		(0.08)	(0.034)

Note: * indicates significance at the 5% level for the null of no cointegration. () indicates MacKinnon-Haug-Michelis p-values.

Table 3. Estimated Nonlinear Models: HBS effect

$$q_t = q_{t-1} - (q_{t-1} - \mu_1(a_{t-1} - a^*_{t-1}) - c)\{1 - \exp[-\gamma(q_{t-1} - \mu_1(a_{t-1} - a^*_{t-1}) - c)^2 / \sigma^2]\} + e_t$$

		91-98	99-07
UK	c	-0.050 (3.171)	0.043(11.300)
	γ	0.016 (1.305) [0.45]	0.1288(3.796)[0.000]
	μ_1	-	-1.242(2.875)
		z1=[0.183]; z2=[0.006]	z1=[0.161]; z2=[0.243]
Switzerland	c	-0.027(3.126)	-
	γ	0.033(2.020)[0.186]	-0.007(1.331) [0.47]
	μ_1	0.1177(0.580)	-
		z1=[0.583]; z2=[0.000]	z1=[0.237]; z2=[0.825]
Denmark	c	-0.025(7.011)	0.013(8.250)
	γ	0.022(2.424)[0.072]	0.060(2.735)[0.036]
	μ_1	0.019(0.052)	0.377(2.010)
		z1=[0.322]; z2=[0.000]	z1=[0.581]; z2=[0.203]
France	c	-0.022(6.872)	0.004(8.035)
	γ	0.018(2.119) [0.095]	.083(2.715)[0.028]
	μ_1	-	0.020(0.131)
		z1=[0.213]; z2=[0.064]	z2=[0.805]; z3=[0.614]
Austria	c	-0.006(1.678)	0.004(4.041)
	γ	0.057(2.258)[0.100]	0.063(2.488)[0.110]
	μ_1	0.330(2.012)	0.486(3.477)
		z2=[0.100];z3=[0.004]	z2=[0.245]; z3=[0.013]

Notes: Figures in round parentheses are t-statistics;. z1= White (1980) heteroscedasticity test; z2=Eitrheim-Teräsvirta serial correlation test; figures in square brackets are marginal significance levels. The marginal significance levels for the estimated transition parameters were calculated by Monte-Carlo simulation under the null hypothesis of a unit root AR(1) process.

Table 4. Estimated Nonlinear Models: FESTAR.

$$q_t = q_{t-1} - (q_{t-1} - \beta_1 F_{t-1} - \beta_2 F^*_{t-1} - c)\{1 - \exp[-\gamma(q_{t-1} - \beta_1 F_{t-1} - \beta_2 F^*_{t-1} - c)^2 / \sigma^2]\} + e_t$$

		91-98	99-07
UK	c	-0.052(4.873	0.029(9.169) [0.001]
	γ	0.039(1.314)[0.668]	0.128(3.632)-
	β_1	0.688(4.498)	-
	β_2	0.799(3.203)	0.079(2.692)
		z1=[0.481]; z2=[0.009]	z1=[0.837]; z2=[0.19]
Switzerland	c	-0.016(4.218)	-
	γ	0.031(1.790) [0.400]	-0.023(2.240) [0.09]
	β_1	-	-
	β_2	0.061(1.851)	0.232(5.435)
		z1=[0.846]; z2=[0.09]	z1t=[0.882]; z2=[0.833]
	c	-0.024(6.780)	0.014(11.010)
Denmark	γ	0.019(2.380) [0.05]	0.032(2.670) [0.024]
	β_1	-	-
	β_2	-	-
		z1=[0.49]; z2=[0.000]	z1=[0.525]; z2=[0.056]
France	c	-0.022(6.872)	0.004(8.120)
	γ	0.018(2.120) [0.095]	0.080(3.280) [0.00]
	β_1	-	-
	β_2	-	-
		z1=[0.213]; z2=[0.064]	z1=[0.608]; z2=[0.606]
Austria	c	-0.007(9.091)	0.007(12.060)
	γ	0.093(2.733)[0.05]	0.095(3.071)[0.014]
	β_1	0.012(2.397)	0.041(7.631)
	β_2	-	-
		z1=[0.130]; z2=[0.004]	z1=[0.433]; z2=[0.019]

Notes: Figures in round parentheses are t-statistics;. z1= White (1980) heteroscedasticity test; z2=Eitrheim-Teräsvirta serial correlation test; figures in square brackets are marginal significance levels. The marginal significance levels for the estimated transition parameters were calculated by Monte-Carlo simulation under the null hypothesis of a unit root AR(1) process.

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Appendix: Data used in constructing the diffusion indices

All series are obtained from DATASTREAM with the codes given below. The bold characters with asterisks denote quarterly data interpolated to monthly frequency. The data are seasonally adjusted and put into logarithmic form as appropriate. All series are further standarised as zero mean and unit variance over the sample period in order to remove any scaling effect in the construction of the diffusion indices.

GERMANY

ID	Code	Name
Real Output		
1	BDOPRI35H	BD PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION) VOL (2000=100)
2	BDOPRI38H	BD PRODUCTION IN TOTAL MANUFACTURING VOL (2000=100)
3	BDOPRI61H	BD PRODUCTION OF TOTAL MANUFACTURED INTERMEDIATE GOODS VOL (2000=100)
4	BDOPRI50H	BD PRODUCTION OF MANUFACTURED DURABLE CONSUMER GOODS VOL (2000=100)
5	BDOPRI51H	BD PRODUCTION OF MANUFACTURED NON-DURABLE CONSUMER GOODS VOL (2000=100)
6	BDIPTOT%G	BD INDUSTRIAL PRODUCTION INCLUDINGCONSTRUCTION (%YOY) VOL
7	BDOPRI08P	BD PRODUCTION OF MANUFACTURED CRUDE STEEL VOL (,etric tonnes, thou)
8	BDPRODVTQ	BD PRODUCTIVITY: OUTPUT PER MAN-HOUR WORKED IN INDUSTRY SADJ (2000=100)
Employment		
9	BDOEM047P	BD EMPLOYMENT - PART-TIME (ECONOMIC REASON) VOL
10	BDOUN008P	BD REGISTERED UNEMPLOYED VOL (thou)
11	BDOUN013R	BD REGISTERED UNEMPLMT. (PERCENT OF CIVILIAN LABOUR FORCE) (%)
12	BDOUN014Q	BD STANDARDIZED UNEMPLOYMENT RATE SADJ (%)
Orders		
13	BDOL0583G	BD COMPOSITE LEADING INDICATOR: VOLUME NET NEW ORDERS (MFG. VOL, 2000=100)
14	BDOL0268Q	BD COMPOSITE LEADING INDICATOR: ORDERS INFLOW SADJ (%)
Exchange Rates		
15	BDOCC011	BD REAL EFFECTIVE EXCHANGE RATE - CPI BASED VON (1995=100)
Productivity		
16	BDJAB024D	BD PRODUCTIVITY & LABOURCOSTS: LABOUR COSTS PER UNIT OF OUTPUT (2000=100)
17	BDJAC000D	BD PRODUCTIVITY & LABOUR COSTS GDP DEFLATOR CONA (2000=100)
Retail Trade		
18	BDOSLI15H	BD TOTAL RETAIL TRADE VOL (2000=100)
19	BDRETTOTG	BD RETAIL SALES EXCL.CARS-X-12-ARIMA(EXPANDED SAMPLE FROM 0106) (2000=100)

UNITED KINDOM

	Code	Name
Real Output		
1	UKOPRI35G	UK PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION) VOLA (2000=100)
2	UKOPRI38G	UK PRODUCTION IN TOTAL MANUFACTURING VOLA(2000=100)
3	UKOPRI61G	UK PRODUCTION OF TOTAL MANUFACTURED INTERMEDIATE GOODS VOLA (2000=100)
4	UKOPRI08P	UK PRODUCTION OF MANUFACTURED CRUDE STEEL VOLN (metric tones, thou)
5	UKOPRI13P	UK PRODUCTION OF PASSENGERS CARS VOLN (thou)
6	UKOPRI16P	UK PRODUCTION OF COMMERCIAL VEHICLES VOLN (thou)
Employment		
7	UKMGSW..	UK LFS: EMPLOYMENT RATE, FEMALE, AGED 16-59 % SADJ
8	UKMGSS..	UK LFS: EMPLOYMENT RATE, MALE, AGED 16 & OVER SADJ
9	UKYBSE..	UK LFS: IN EMPLOYMENT, ALL, AGED 16-59/64 VOLA (thou)
10	UKOUN008O	UK REGISTERED UNEMPLOYEDVOLA (thou)
11	UKOUN015Q	UK UNEMPLOYMENT RATE (% OF TOTAL LABOUR FORCE) SADJ
Consumption		
12	**UKCNHLD.D***	UK FINAL CONSUMPTION EXPENDITURE - HOUSEHOLDS (CVM) CONA (mil, 2003 CHND PRC)
13	**UKCNPER.D***	UK CONSUMER SPENDING (CVM) CONA (mil, 2000 CHND PRC)
14	**UKI96F.CB***	UK PRIVATE CONSUMPTION CURA (pound, mil)
Orders		
15	UKOODI54G	UK ORDERS FOR EXPORTED MANUFACTURED GOODS (VOLUME) VOLA (2000=100)
16	UKOODI53G	UK ORDERS FOR MANUFACTURED GOODS FROM DOM. MARKET VOLA (2000=100)
17	UKOODI45G	UK ORDERS FOR TOTAL MANUFACTURED GOODS (VOLUME) VOLA (2000=100)
Exchange Rates		
18	UKI..RECE	UK REAL EFFECTIVE EXCHANGE RATE INDEX - CPI BASED SADJ (1995=100)
Earnings		
19	UKOCFRLCG	UK LABOUR COST INDEX (REAL) (AR)(DISC.) VOLA (2000=100)
Retail trade		
20	UKRETTOTG	UK RETAIL SALES: ALL RETAILERS - ALL BUSINESS VOLA (2000=100)
21	UKRTHOUSG	UK RETAIL SALES: HOUSEHOLD GOODS STORES - ALL BUSINESS VOLA (2000=100)
22	UKRTONFDG	UK RETAIL SALES: OTHER NON-FOOD STORES - ALL BUSINESS VOLA (2000=100)

SWITZERLAND

	Code	Name
Real Output		
1	SWI66..IG	SW INDUSTRIAL PRODUCTION VOLA (2000=100)
2	**SWOPRI35G***	SW PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION) VOLA (2000=100)
3	**SWOPRI38G***	SW PRODUCTION IN TOTAL MANUFACTURING VOLA (2000=100)
Employment		
4	**SWOEM040H***	SW EMPLOYMENT - CIVILIAN VOLN (2000=100)
5	**SWOEM026H***	SW CIVIL EMPL IN INDUSTRY NADJ (2000=100)
6	**SWOEM063G***	SW CIVIL EMPL IN SERVICES VOLA (2000=100)
7	**SWOUN015Q***	SW UNEMPLOYED % TOTAL LABOUR FORCE SADJ
8	**SWOPL035O***	SW TOTAL LABOUR FORCE VOLA (thou)
9	**SWOUN013Q***	SW REGISTERED UNEMPLOYMENT (PERCENT OF TOTAL LABOUR FORCE) SADJ
10	**SWUN%TOTR***	SW UNEMPLOYMENT RATE NADJ
Consumption		
11	**SWCNPER.D***	SW PRIVATE FINAL CONSUMPTION EXPENDITURE CONA (swiss franc, mil, 2000 chnc proc))
12	**SWCNGOV.D***	SW GOVERNMENTFINAL CONSUMPTION EXPENDITURE CONA (swiss franc, mil, 2000 chnc proc))
Orders		
13	**SWCNORDCH***	SW NEW ORDERS - CONSTRUCTION (VOLUME, LAST 3 MONTHS) VOLN
Exchange Rates		
14	SWI..RECE	SW REAL EFFECTIVE EXCHANGE RATE INDEX - CPI BASED SADJ (1995=100)

DENMARK

Real Output

1	DKOPRI35G	DK PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION) VOLA (2000=100)
2	DKOPRI38G	DK PRODUCTION IN TOTAL MANUFACTURING VOLA (2000=100)
3	DKOPRI50G	DK PRODUCTION OF MANUFACTURED DURABLE CONSUMER GOODS VOLA (2000=100)
4	DKOPRI51G	DK PRODUCTION OF MANUFACTURED NON-DURABLE CONSUMER GOODS VOLA (2000=100)
5	DKOPRI61G	DK PRODUCTION OF TOTAL MANUFACTURED INTERMEDIATE GOODS VOLA (2000=100)
6	DKOPRI70G	DK PRODUCTION OF TOTAL MANUFACTURED INVESTMENT GOODS VOLA (2000=100)
Employment		
7	DKOUN013Q	DK REGISTERED UNEMPLOYMENT (PERCENT OF TOTAL LABOUR FORCE) SADJ
8	**DKOEM019P***	DK EMPLOYEES VOLN (thou)
9	DKUN%TOTQ	DK UNEMPLOYMENT RATE SADJ (%)
10	DKOUN014Q	DK STANDARDIZED UNEMPLOYMENT RATE SADJ (%)
11	DKOUN013Q	DK REGISTERED UNEMPLOYMENT (PERCENT OF TOTAL LABOUR FORCE) SADJ
Orders		
12	DKESINDMH	DK INDL. ORDERS - MANUFACTURING, WORKING ON ORDERS VOLN (2000=100)
Exchange Rates		
13	DKOCC011	DK REAL EFFECTIVE EXCHANGE RATE - CPI BASED VOLN (1995=100)
Earnings		
14	DKOLC007H	DK HOURLY EARNINGS: MANUFACTURINGNADJ (2000=100)
15	DKRETTOTG	DK RETAIL SALES VOLA (2000=100)
16	DKRTOTHGG	DK RETAIL SALES - OTHER CONSUMPTION GOODS VOLA (2000=100)

FRANCE

	Code	Name
Real Output		
1	FROPRI35H	FR PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION) VOL (2000=100)
2	FROPRI38H	FR PRODUCTION IN TOTAL MANUFACTURING VOL (2000=100)
3	FROPRI30H	FR PRODUCTION OF TOTAL CONSTRUCTION VOL(200=100)
4	FROPRI49H	FR PRODUCTION OF TOTAL MANUFACTURED CONSUMER GOODS VOL (2000=100)
5	FROPRI61H	FR PRODUCTION OF TOTAL MANUFACTURED INTERMEDIATE GOODS VOL (2000=100)
6	FROPRI70H	FR PRODUCTION OF TOTAL MANUFACTURED INVESTMENT GOODS VOL (2000=100)
7	FROPRI44H	FR PRODUCTION OF TOTAL ENERGY VOL (2000=100)
8	FROPRI47H	FR PRODUCTION IN TOTAL AGRICULTURE VOL (2000=100)
9	FROPRI08P	FR PRODUCTION OF MANUFACTURED CRUDE STEEL VOL (meteoric thou)
10	FROPRI58H	FR PRODUCTION OF TOTAL VEHICLES VOL
Employment		
11	FROEM006O	FR EMPLOYMENT INDUSTRY (OLD) VOLA (thou)
12	FROEM012O	FR EMPLOYMENT - MARKET SERVICES VOLA (thou)
13	FROUN008P	FR REGISTERED UNEMPLOYED VOL (thou)
14	FROUN007G	FR NEW UNEMPLOYMENT CLAIMS SADJ (actual thou)
15	FROUN014Q	FR STANDARDIZED UNEMPLOYMENT RATE SADJ (2000=100)
16	FROUN015Q	FR UNEMPLOYMENT RATE (% OF TOTAL LABOUR FORCE) SADJ
Housing Started		
17	FRHOUSE.P	FR HOUSING STARTED VOL (actual)
Consumption		
18	FRCNHLD.D	FR HOUSEHOLD CONSUMPTIONCONA (euro, mil)
19	FRCNPER.D	FR CONSUMER SPENDING CONA (euro, mil)
Exchange Rates		
20	FRI..RECE	FR REAL EFFECTIVE EXCHANGE RATE INDEX - CPI BASED SADJ (1995=100)
Retail trade		
21	FROSLI15H	FR TOTAL RETAIL TRADE (VOLUME) VOLN (2000=100)

AUSTRIA

	Code	Name
Real Output		
1	OEOPRI35H	OE PRODUCTION OF TOTAL INDUSTRY (EXCLUDING CONSTRUCTION) VOL (2000=100)
2	OEOPRI38H	OE PRODUCTION IN TOTAL MANUFACTURING VOL (2000=100)
3	OEOPRI61H	OE PRODUCTION OF TOTAL MANUFACTURED INTERMEDIATE GOODS VOL (2000=100)
4	OEOPRI70H	OE PRODUCTION OF TOTAL MANUFACTURED INVESTMENT GOODS VOL (2000=100)
5	OEOPRI08P	OE PRODUCTION OF MANUFACTURED CRUDE STEEL VOL ((meteoric thou)
Employment		
6	OEOEM019O	OE DEPENDENT EMPLOYMENT VOLA (thou)
7	OEOEM011O	OE EMPLOYMENT - SERVICESVOLA (thou)
8	OEOUN008P	OE REGISTERED UNEMPLOYEDVOLN (thou)
9	OEVACTOTP	OE JOB VACANCIES VOLN (actual)
10	OEUNPTOTP	OE UNEMPLOYED - REGISTERED VOLN (actual)
11	OEUN%TOTR	OE UNEMPLOYMENT RATE % NADJ
12	OEOUN012R	OE UNEMPLOYMENT(PERCENT OF CIVIL LABOUR FORCE) NADJ
Consumption		
13	**OECNPER.D***	OE PRIVATE CONSUMPTION EXPENDITURE CONA (2000 chnc proc)
Exchange Rates		
14	OEI..RECE	OE REAL EFFECTIVE EXCHANGE RATE INDEX - CPI BASED SADJ (1995=100)
Earnings		
15	**OEOLC007H***	OE COMPOSITE LEADING INDICATOUR: MONTHLY EARN-MINING & MFG. NADJ (2000=100)
16	**OEOLC006H***	OE MONTHLY EARNINGS: MINING & MANUFACTURING NADJ (2000=100)
Retail trade		
17	OEOSLI15H	OE TOTAL RETAIL TRADE (VOLUME) VOLN (2000=100)