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The Swedish Real Exchange Rate under Different Currency Regimes*

by

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Abstract

This paper presents evidence on the behavior of the Swedish real exchange rate relative to Germany under different currency regimes 1973:1-2001:4. The results suggest that the real exchange rate is cointegrated with Swedish and German productivity, which is consistent with Balassa (1964) and Samuelson (1964). In the short run, the regime has mattered for the dynamics of the real exchange rate. Deviations from equilibrium have been adjusted more quickly when the nominal exchange rate has been allowed to float freely, which indicates that it acts as a shock absorber.

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1 Introduction

In light of the recent launch of the third step of the Economic and Monetary Union (EMU) there has been a vivid debate on exchange rate regimes. Many countries have abandoned fixed exchange rate regimes during the 1990s and in many cases combined floating exchange rates with inflation targets. Opponents to floating exchange rates argue that rather than absorbing shocks and stabilizing the economy, the nominal exchange rate reflects noise from financial markets that may in fact be destabilizing. Obviously, if the nominal exchange rate causes shocks rather than offsets them, the cost of relinquishing it by joining the EMU, is low. [1]

At the same time, the *real* exchange rate is perhaps the most common measure of overall firm competitiveness, and to understand its behavior over different horizons is to understand the conditions faced by firms engaging in international trade.[2] For small open economies in particular, the real exchange rate is therefore clearly a key variable. In this paper we try to shed some light on the behavior of the real exchange rate under different exchange rate regimes by addressing the following questions: What mechanisms ensure that the real exchange rate returns to equilibrium after a distortion? How are these mechanisms affected by the exchange rate regime?

The literature on long-run behavior of real exchange rates is quite extensive. Since the introduction of cointegration in the empirical literature, interest in the real exchange rate has experienced a renaissance through its

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[1] Thomas (1997) finds evidence that, provided that demand shocks are controllable by economic policy, Sweden would face a low cost of relinquishing the nominal exchange rate. Artis and Ehremann (2002) suggest that the Swedish exchange rate is a source of shocks rather than a shock absorber.

[2] The real exchange rate is defined as $Q = SP^*/P$,where S is the nominal exchange rate in domestic currency per units of foreign currency (spot rate), P is the price level, and * henceforth denotes a foreign country.

close connection to Purchasing Power Parity (PPP). A stationary real exchange rate implies that relative PPP must hold, while a non-stationary real exchange rate is a suitable object for cointegration analysis. Numerous tests for cointegration between the nominal exchange rate and various price combinations have therefore been presented in the literature, see Froot and Rogoff (1995) for a survey. In addition to tests for PPP, the literature on real exchange rates comprises various structural models of economic fundamentals. Seminal work by Balassa (1964) and Samuelson (1964) shows how the real exchange rate should be determined by the productivity growth differential between the traded and non-traded sector. Extensions to the Balassa-Samuelson model have also been made and tested empirically. Evidence on how long run real exchange rates depend on fundamentals is given in among others Alexius and Nilsson (2000) and, more recently, Bergvall (2002).

An interesting property of the real exchange rate is that it consists of variables exhibiting contrasting dynamic behavior. As is well known, price levels tend to be sticky, while typically volatile nominal exchange rates can jump instantaneously under a floating rate regime, as in for instance Dornbusch (1976). Modeling the real exchange rate consequently implies modeling the slow adjustment of long run relative prices, and the fast (instantaneous) adjustment of nominal exchange rates.

The research effort dedicated to investigating the short-run dynamics of the real exchange rate is far less extensive than the effort made to examine its long run behavior. Perhaps as a consequence, the impact of the exchange rate regime on the real exchange rate is often neglected.[3] The choice between

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[3] One exception, although a little dated, is Mussa (1986), who finds that the regime indeed matters for the real exchange rate. Since post-Bretton Woods time series only recently have become sufficiently long for results to be reliable, it is of great interest to examine the issue of regimes using recent data. Taylor (2002) distinguishes between

a fixed exchange rate or a floating exchange rate (perhaps combined with a price level or inflation target) should be highly relevant in explaining the dynamic adjustment of the real exchange rate to its long run path.

This paper analyses the long run behavior and the short-run dynamics of the Swedish real exchange rate relative to Germany during the period 1973:1-2001:4, stretching from the collapse of the Bretton Woods system in March 1973 through various exchange rate regimes including the launch of the third step of the EMU on January 1 1999.

The contribution of the paper is as follows: Instead of focusing on the long run, we emphasize the analysis of dynamic models of the real exchange rate where the exchange rate regime is taken into account. Moreover, we investigate the question as to whether it is the Swedish price level, the German price level or the nominal exchange rate that has adjusted to deviations from long run equilibrium, the hypothesis being that Sweden as a small country has been forced to adapt to German conditions rather than vice versa.

The main findings are that (i) there is support for the Balassa-Samuelson hypothesis saying that the real exchange rate is driven by productivity growth in the long run, (ii) the regime has mattered for the dynamics of the real exchange rate and (iii) the Swedish price level and the nominal exchange rate account for most of the adjustment following a distortion to the real exchange rate.[4]

The rest of the paper is organized as follows: Section 2 presents a simple theoretical framework for analyzing the real exchange rate over different horizons and Section 3 describes the data. Section 4 discusses modeling strategies

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regimes in a study of PPP reversion.

[4] The results in this paper differ somewhat from the results in the previous version (FIEF WP 180). This is probably due to a significant improvement in data quality. In this version productivity data is used instead of GDP which was used as a proxy in the previous version. Moreover the series have been prolonged by three years.

and empirical issues. Results are presented in Section 5 and Section 6 concludes.

2 Theoretical Framework

In two pioneering papers Balassa (1964) and Samuelson (1964) find that the productivity growth differential between the traded and non-traded sectors within countries, relative to other countries, should affect long run real exchange rates.[5] Alexius (2001) reports that 86 percent of the long run variance of the Swedish real exchange rate is determined by supply-side factors. In an attempt to verify this result, we follow Balassa and Samuelson and think of an economy which is determined from the supply side in the long run.

Consider a small, open economy, consisting of two sectors: One sector producing tradeable goods, T, and a sheltered sector producing non-tradeable goods, indexed N. Since the economy is small, the rental cost of capital R is exogenously given and in the traded sector the law of one price implies $P_T = SP_T^*$, where P_T and P_T^* are domestic and foreign T-sector prices and S is the nominal (spot) exchange rate. There is perfect capital mobility between sectors as well as between countries, and perfect labor mobility domestically implying that the nominal wage is equal in the two sectors: $W_N = W_T \equiv W$. Finally, there is perfect competition in the goods market and the labor market. The production technology is Cobb-Douglas and there are constant returns to scale in both sectors, so that production is given by

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[5] De Gregorio & Wolf (1994) expand the theory of Balassa and Samuelson by considering the terms of trade, i.e. the ratio between export- and import prices. This may however yield a simultaneity problem. Moreover, Chinn (1997) reports that the terms of trade are insignificant in the long run and they are therefore excluded in this study.

$$Y_i = A_i \left(L_i\right)^{\alpha_i} \left(K_i\right)^{1-\alpha_i}$$

where Y_i is output, A_i total factor productivity (TFP), L_i and K_i are labor and capital input respectively, $\alpha_i \in (0,1)$ and $i = N, T$. The first order conditions for profit maximization are

$$\begin{aligned} W_r &= \alpha_T A_T \left(\frac{K_T}{L_T}\right)^{1-\alpha_T} \\ W_r &= \alpha_N A_N P_r \left(\frac{K_N}{L_N}\right)^{1-\alpha_N} \\ R_r &= \left(1-\alpha_T\right) A_T \left(\frac{K_T}{L_T}\right)^{-\alpha_T} \\ R_r &= \left(1-\alpha_N\right) A_N P_r \left(\frac{K_N}{L_N}\right)^{-\alpha_N} \end{aligned}$$

where $W_r \equiv \frac{W_T}{P_T} = \frac{W_N}{P_T}, P_r \equiv \frac{P_N}{P_T}, R_r \equiv \frac{R}{P_T}$. The system has four endogenous variables $W_r, P_r, \left(\frac{K_T}{L_T}\right), \left(\frac{K_N}{L_N}\right)$ and a recursive solution. Taking logs and differentiating the first order conditions yields the central equation of the Balassa-Samuelson hypothesis:

$$\frac{\dot{P}_{rel}}{P_{rel}} = \left(\frac{\alpha_N}{\alpha_T}\right)\frac{\dot{A}_T}{A_T} - \frac{\dot{A}_N}{A_N} \tag{1}$$

where $\dot{X} = \frac{dX}{dt}$and $\frac{\dot{X}}{X}$ is the growth rate of X. The intuition behind hypothesis (1) is as follows: A productivity rise in the T-sector raises the aggregate wage level and prices on N-goods for a given level of productivity in the N-sector. Equally labor intensive production in the two sectors, $(\alpha_T = \alpha_N)$ generates a growth rate in the relative price of non-tradeables that is equal to the inter-sector TFP growth differential.

Assume that the aggregate price level in the economy, i.e. the consumer price level, can be written as a weighted mean of the price levels in the two sectors $P = \left(P_N\right)^{\beta}\left(P_T\right)^{1-\beta}$ where $\beta \in (0,1)$. Finally assume that the conditions stated above apply also to the foreign country so that there is complete symmetry in production and prices. By using the definition of the real exchange rate $Q = SP^*/P$, the aggregate price level and by imposing the law of one price for T-goods, we obtain the following cost-push hypothesis of the real exchange rate:

$$\frac{\dot{Q}}{Q} = \beta^* \left(\frac{\alpha_N^*}{\alpha_T^*} \frac{\dot{A}_T^*}{A_T^*} - \frac{\dot{A}_N^*}{A_N^*} \right) - \beta \left(\frac{\alpha_N}{\alpha_T} \frac{\dot{A}_T}{A_T} - \frac{\dot{A}_N}{A_N} \right) \tag{2}$$

Equation (2) states that higher relative TFP-growth in the traded sector at home than abroad induces a real appreciation, due to the effect of increased wage costs on domestic prices.

Short-run Dynamics

In the short run, we expect changes in the nominal exchange rate to transmit fully to the real exchange rate. The time span is not long enough for prices to adjust. Since the nominal exchange rate is notoriously difficult to model, we use the simple uncovered interest rate parity (UIP) condition to capture some basic ideas. Letting R_t and R_t^* be the nominal domestic and foreign short interest rates respectively, S_t the nominal exchange rate in domestic currency per unit of foreign currency in period t and $E_t(S_{t+1})$ the expectation at time t of the nominal exchange rate in the next period, UIP states that $R_t = R_t^* + \frac{E_t(S_{t+1}) - S_t}{S_t}$. A positive interest rate spread therefore reflects an expected depreciation of the nominal exchange rate. Rearranging the UIP-condition yields

$$S_t = \frac{E_t\left(S_{t+1}\right)}{1 + \left(R_t - R_t^*\right)}$$

If UIP holds, the nominal exchange rate is determined by the expected nominal exchange rate discounted by the interest rate spread $(R_t - R_t^*)$. We treat Sweden as a small country taking German variables as given and assume that depreciation expectations are determined by Swedish unemployment, U and the Swedish Budget Balance $BDef$. The intuition is that the Swedish central bank is more likely to let the exchange rate depreciate in a situation with high unemployment or a large budget deficit, since depreciating the exchange rate may stimulate the economy in such a way that these problems are offset. Historically, the Swedish government has also chosen to depreciate the exchange rate during periods of high unemployment in order to improve firm competitiveness and stimulate employment.

In addition to variables operating through the nominal exchange rate, the demand-side may have a short-run influence on the dynamics of the real exchange rate through the aggregate price levels. Chinn (1997) suggests that log relative government consumption, $(g - g^*)$, is significant over this horizon, since increased demand may cause demand-pull inflation and thereby a real appreciation. Another hypothesis tested by Chinn (1997), is the possibility that the oil price (p_{oil}) affects the real exchange rate, representing a cost-push hypothesis. The sign of the effect is uncertain since it depends on how production technologies differ between countries. These variables will be included as short-run explanatory variables in the study at hand. The short-run dynamics suggested by theory may therefore be summarized by the following function

$$\Delta q = f\left(R - R^*, U, BDef, g - g^*, p_{oil}, ECM\right)$$

where we expect $f_1 < 0, f_2 > 0, f_3 < 0, f_4 < 0, f_5 \gtrless 0, f_6 < 0$.

3 Empirical Issues

We start by briefly considering the Engle-Granger one-step, see Engle and Granger (1987) and Banerjee *et al* (1993), and proceed by modifying the specifications in an attempt to capture some aspects of the exchange rate regime. The section is ended with a description of the method used to test how the three components, s, p and p^*, react to deviations from long run equilibrium under different exchange rate regimes. Throughout the paper small letters denote logs unless otherwise stated.

(i) Estimating the Cointegrating Vector and obtaining the Error Correction Mechanism (ECM)

The Engle-Granger one-step method suggests estimating the short-run and the long-run relationships in one step. Suppose we suspect that there is cointegration between the real exchange rate, q_t, and Swedish and German productivity, a_t and a_t^* respectively. Let x_j, $j = 1, ..., J$ be the J short-run explanatory variables to be used in first differences if they are found to be non-stationary.[6] The one-step procedure is then to estimate.

$$\Delta q_t = \gamma_0 + \sum_{j=1}^{J} \gamma_j \Delta x_j + \gamma_{J+1} q_{t-1} + \gamma_{J+2} a_{t-1} + \gamma_{J+3} a_{t-1}^* + \varepsilon_t \qquad (3)$$

The error correction term is obtained from (3) as $ECM_t = q_t + \frac{\overset{\wedge}{\gamma}_{J+2}}{\overset{\wedge}{\gamma}_{J+1}} a_t + \frac{\overset{\wedge}{\gamma}_{J+3}}{\overset{\wedge}{\gamma}_{J+1}} a_t^*$, where we expect $\overset{\wedge}{\gamma}_{J+1} < 0$. The reason $\overset{\wedge}{\gamma}_{J+1}$should be non-positive is that if the real exchange rate at time $t-1$ was above its equilibrium value it has to appreciate, i.e. decrease, in period t in order to eventually return to equilibrium. Since first differences of all $I(1)$ variables are stationary and

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[6] Throughout the paper weak (covariance) stationarity is intended. The stochastic process X_t is weakly stationary if $E(X_t) = \mu$, $Var(X_t) = \sigma^2$, where μ, σ^2 are constants, and $Cov(X_t, X_{t+j}) = \sigma_j \ \forall j$.

linear combinations of cointegrated variables are stationary by definition, the residuals in (3) should be stationary as well. If the deviation from long run equilibrium adjusts proportionally to the current *level* of the deviation, 90 percent of the deviation has been adjusted after $t = \ln(0.10)/\overset{\wedge}{\gamma}_{J+1}$periods.[7]

(ii) Modeling the Short run Dynamics taking into Account the Exchange Rate Regime

Under a floating regime, the nominal exchange rate can adjust to shocks instantaneously, while the relative price has to absorb all shocks under an irrevocably fixed rate regime such as a monetary union. In the long run, the regime should not affect the real exchange rate, but the mechanisms through which the real exchange rate reaches its equilibrium value, i.e. the short-run dynamics, are likely to be affected by the exchange rate regime. In this paper, an attempt to capture these features is made by using dummy variables.

First, two regime dummies are specified. The post-Bretton Woods period may in the Swedish case be roughly divided into four currency regimes: (i) The system referred to as *the Monetary Snake*, 1973-1977, (ii) *the Currency Basket*, 1977-1991, (iii) the peg relative the *ECU*, 1991-1992 and (iv) the floating rate regime 1992 onwards. Regimes (i) and (iii) are virtually equivalent to a fixed rate regime relative the D-mark, since the German currency dominated the Snake. Since the Currency Basket prevailed for the longest time period, it will be used as the norm and the following two dummy vari-

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[7] To see this, let $z = z(t)$ be the deviation from long run equilibrium, let $z(0) = z_0$ and assume that the deviation adjusts proportionally to its current level, so that $\frac{dz}{dt} = \overset{\wedge}{\gamma}_{J+1}z$, where $\overset{\wedge}{\gamma}_{J+1} < 0$. This yields an ordinary, separable differential equation with the following solution: $\frac{dz}{dt} = \overset{\wedge}{\gamma}_{J+1}z$ implies $lnz = \overset{\wedge}{\gamma}_{J+1}t + C$, where C is a constant. The initial value implies $C = \ln z_0$, and the solution $z = z_o \exp\left(\overset{\wedge}{\gamma}_{J+1}t\right)$. (100ρ) percent of the deviation has consequently been adjusted after t periods according to the following: $(1-\rho)\, z_0 = z_o \exp\left(\overset{\wedge}{\gamma}_{J+1}t\right)$ implying $t = \ln(1-\rho)/\overset{\wedge}{\gamma}_{J+1}$, where $\rho \in (0,1)$, and $t > 0$ since $lnx < 0$ for all $x < 1$. Note that the standard half-life measure is obtained by letting $\rho = 0.5$.

ables are used to capture regime-specific effects:

$$\begin{aligned} D_{fix} &= \begin{cases} 1, \text{ if 1973:1-1977:2, or 1991:1-1992:4} \\ 0, \text{ otherwise} \end{cases} \\ D_{float} &= \begin{cases} 1, \text{ if 1993:1 onwards} \\ 0, \text{ otherwise} \end{cases} \end{aligned}$$

Hence, D_{fix} indicates a fixed rate regime and D_{float} indicates a floating exchange rate regime. Dichotomous dummy variables will also be added to capture the devaluations during the fixed exchange rate regimes and the German reunion. These variables assume the value zero for all quarters except the quarter defined in each dummy according to the following: $D_1 = \{1, \text{ if } 1977:2\}$, $D_2 = \{1, \text{if } 1977:3\}$, $D_3 = \{1, \text{ if } 1981:3\}$, $D_4 = \{1, \text{ if } 1982:3\}$, $D_5 = \{1, \text{ if } 1992:4\}$ and $D_6 = \{1, 1990:3 \text{ onwards}\}$. The variable D_5 thus captures the interest rate turbulence in November 1992, the interest rate shock and the transition to a floating exchange rate regime. D_6 captures a possible structural break due to the German reunion.

It will prove useful to change the notation slightly by letting superscript r denote the exchange rate regime. Define $D^0 \equiv 1, D^1 = D_{fix}, D^2 = D_{float}$, so that the exchange rate regime is captured by $D^r, r = 0, 1, 2$. The unrestricted general model can then be formulated in the following way:

$$\begin{aligned} \Delta_i q_t &= \mu_0 + \sum_{j=1}^{J} \sum_{r=0}^{2} \Delta_i x_j \mu_j^r D^r + \sum_{r=0}^{2} \mu_{J+1}^r D^r ECM_{t-i} \\ &\quad + \sum_{k=1}^{K} \mu_{J+1+k} D_{k,t} + \sum_{l=1}^{L} \mu_{J+K+1+l} \Delta_i q_{t-l} + \varepsilon_t \end{aligned} \tag{4}$$

where $i = 1, 4$ since we will estimate both quarterly and annual differences. Moreover, $j = 1, ..., J$ for the explanatory variables, $k = 1, ..., K$ for the additive dummies and $l = 1, ..., L$ indicate the L lagged dependent variables added in order to remove any serial correlation in the residuals. The devaluations will therefore be allowed to shift the intercept of the dynamic models, while the regime shifts are assumed to shift the slope coefficients. In this fashion, we allow for the possibility that the impact of the explanatory variables is regime-specific.

(iii) Estimating the Impact of the ECM on the Three Components of the Real Exchange Rate

In order to assess the effect of the *ECM* on the three components of the real exchange rate, we use each component as the dependent variable, deducting the other two, while keeping the right hand side virtually intact. By this method three models are obtained. Maintaining the notation from the previous section we obtain

$$\begin{aligned}\Delta_i p_t &= \lambda_0 + \sum_{j=1}^{J} \sum_{r=0}^{2} \Delta_i x_j \lambda_j^r + \sum_{r=0}^{2} \lambda_{J+1}^r ECM_{t-i} \\ &\quad + \sum_{k=1}^{K} \lambda_{J+k} D_{k,t} + \sum_{l=1}^{L} \lambda_{J+k+l} \Delta_i p_{t-l} + \varepsilon_t \qquad (5)\\ \Delta_i p_t^* &= \varphi_0 + \sum_{j=1}^{J} \sum_{r=0}^{2} \Delta_i x_j \varphi_j^r + \sum_{r=0}^{2} \varphi_{J+1}^r ECM_{t-i} \\ &\quad + \sum_{k=1}^{K} \varphi_{J+k} D_{k,t} + \sum_{l=1}^{L} \varphi_{J+K+l} \Delta_i p_{t-l}^* + \varepsilon_t \qquad (6)\\ \Delta_i s_t &= \psi_0 + \sum_{j=1}^{J} \sum_{r=0}^{2} \Delta_i x_j \psi_j^r + \sum_{r=0}^{2} \psi_{J+1}^r ECM_{t-i} \\ &\quad + \sum_{k=1}^{K} \psi_{J+k} D_{k,t} + \sum_{l=1}^{L} \psi_{J+K+l} \Delta_i s_{t-l} + \varepsilon_t \qquad (7)\end{aligned}$$

where the *ECM* in the above equations is implied by the long run path of the real exchange rate, i.e. derived from equation (3).[8] In order to assess the impact of the *ECM* on the various components we are interested in testing linear restrictions on the parameters. For instance, in the case of the Swedish price level as modeled in equation (5), letting $i = 4$ implies that $\hat{\lambda}_{J+1,1}$ is the annual percentage change in the real exchange rate triggered by a one percent increase in the deviation from long run equilibrium under the Currency Basket. Testing the hypothesis $\hat{\lambda}^{1}_{J+1} + \hat{\lambda}^{2}_{J+1} = 0$ is equivalent to testing the significance of the *ECM* under a fixed regime. Similarly, the null of $\hat{\lambda}^{1}_{J+1} + \hat{\lambda}^{3}_{J+1} = 0$ is equivalent to testing the significance of the *ECM* under a floating regime. Alternatively, the null of $\hat{\lambda}^{2}_{J+1} = 0$ does not imply testing if the *ECM* is insignificant under a fixed regime, merely that the elasticity is the same under the Currency Basket and under a fixed exchange rate (provided, of course, that $\hat{\lambda}^{1}_{J+1} \neq 0$).

4 Data

Data consist of quarterly observations for the period 1973:1-2001:4.[9] Letting small letters denote logs the variables are defined as follows: s is the nominal exchange rate (SEK/D-mark), p and p^* are the CPI:s for Sweden and Germany respectively, p_{oil} is the log price of oil, R and R^* are the nominal short interest rates, $g - g^*$ is the log ratio between Swedish and German government consumption and U is the open Swedish unemployment rate measured in percent. $BDef$ is the Swedish budget balance as a share of GDP. Since data on sectorial TFP is notoriously hard to find, we use as proxies Swedish and German labor productivity as measured by the ratio between industrial

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[8] Enders (1988) uses a similar approach in a study testing for PPP.

[9] All series are from the *International Financial Statistics* (IFS), IMF, except for data on employment obtained from *OECD Economic Outlook 68*, OECD and data on unemployment which was retrieved from *OECD Main Economic Indicators* (MEI), OECD.

production and overall employment. Due to lack of data we are forced to use quarterly data on industrial production while the only data available on employment is annual data. Finally a note on how to handle the German transition to the EMU. We assume that the nominal exchange rate between the D-mark and the EURO has remained unchanged since the German transition to the monetary union, i.e. we create an artificial D-mark/SEK exchange rate where we have fixed the D-mark/EURO exchange rate to the conversion rate of 1.96.

Data Properties

The series were tested for the order of integration using Augmented Dickey-Fuller (ADF) tests, see Engle and Granger (1991).[10] We could not reject the null hypothesis of a unit root, i.e. the series being I(1), for any variables except the nominal interest rate spread ($R - R^*$). The remaining series will therefore, in the following, be treated as containing a unit root. Swedish unemployment behaves like a stationary series up to mid 1990 and is quite possibly exposed to a structural break in the early 1990s. It was, however, not possible to interpret the results of a Perron test for a structural break due to heavy auto correlation. The unemployment series will therefore also be treated as potentially I(1).[11] Graphs of Swedish and German productivity and the log real exchange rate are given in *Figure 1.*

Indeed, the series tend together, and we may suspect that there is in fact a cointegrating relationship between them.

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[10] Test results for the order of integration are available on request.

[11] Note that the we could not reject the hypothesis that the real exchange rate is I(1) and hence that Swedish post-Bretton Woods data provide no evidence in support of the PPP-hypothesis.

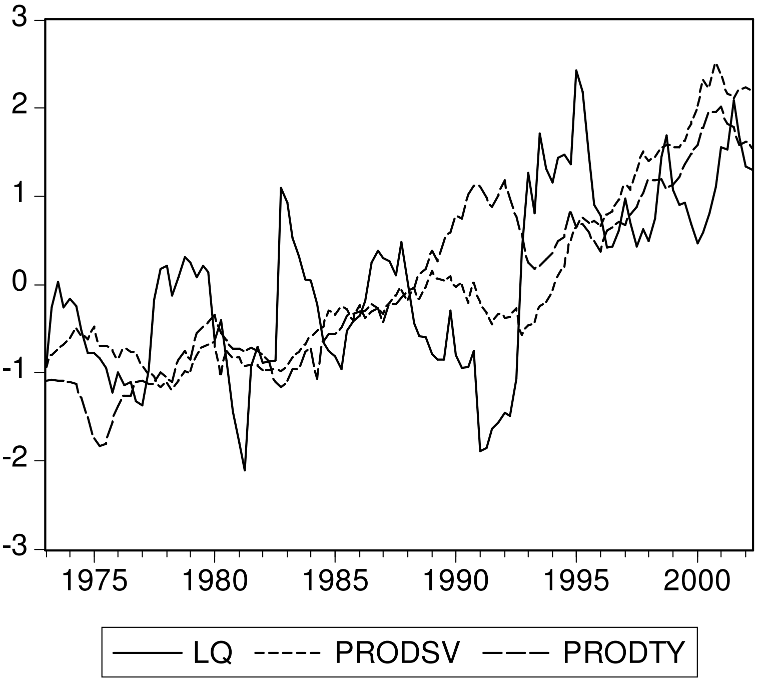

Figure 1: The log real exchange rate (LQ), Swedish Labor productivity ($PRODSV$) and German Labor Productivity ($PRODTY$) over the period 1973:1-2001:4

5 Results

First the long run relationship is estimated by using the Engle-Granger one-step as described by equation (3). The models are then reduced by Likelihood Ratio (LR) tests of linear restrictions and subject to testing until a preferred specification is obtained.

(i) The Long Run Relationship

The results obtained when estimating model (3) are displayed in column (1) in *Table 1.* All parameters have expected signs, except for log relative government consumption which is significant with a positive coefficient. As it turns out, this is true for a majority of the models, and we therefore briefly consider another possibility than the hypothesis that demand-pull mechanisms generate a change in the relative price. It is possible that an increase in $(g - g^*)$ gives rise to inflation expectations in Sweden, causing a nominal depreciation. Since we are estimating quarterly growth rates, the short horizon makes the nominal exchange rate a more probable channel than relative prices.

The long-run relationship is obtained by reducing the model by LR-tests of linear restrictions. The result is displayed in column (2) in *Table 1.* The residuals are clearly stationary with a t-value from the ADF-test of -4.34 (the critical value being equal to -2.89 at the five percent level),[12] suggesting that the variables are cointegrated with a long run relationship equal to

$$q_t = -.015a_t + .034a_t^* \tag{8}$$

Since all variables, except for $(R - R^*)$, U and $BDef$, are expressed in logs, the coefficients of the model in *Table 1* are elasticities.

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[12] Critical values in testing for stationary residuals from MacKinnon (1991).

Table 1: Engle-Granger one-step estimation of model (3) by OLS, quarterly differences 1973:1-2001:4.

Column	(1)	(2)	(3)
Dependent variable	Δq	Δq	Δq
$Intercept$	.111	.	.
	(.068)	.	.
$\Delta(R - R^*)$	-.003**	-.003**	-.003**
	(.001)	(.001)	(.001)
ΔU	-.008	.	.
	(.016)	.	.
$\Delta BDef$	-.002	.	.
	(.003)	.	.
$\Delta(g - g^*)$	.288	.338*	.338*
	(.193)	(.190)	(.184)
Δp_{oil}	.016	.	.
	.021	.	.
q_{-1}	-.118**	-.042**	.
	(.049)	(.022)	.
a_{-1}	-.000	-.001	.
	(.001)	(.001)	.
a^*_{-1}	.001	.001**	.
	(.001)	(.001)	.
ECM_{-1}	.	.	-.042**
	.	.	(.018)
AR-components	0	0	0
Additive Dummies	No	No	No
Seasonal Dummies	No	No	No
R^2_{adj}	.071	.094	.090
N	115	116	115
Durbin-Watson	1.729	1.809	1.843

Standard errors in parenthesis
Significance codes: ***=1%, **=5%, *=10%

In order to formally test for cointegration, we re-run the model in column (2), using the obtained *ECM* as an explanatory variable. To test the significance of the *ECM* is then a test for cointegration as long as the conditioning variables are weakly exogenous to the cointegration parameters, see Banerjee et al (1993) and Kremers et al (1992). The results are shown in column (3) in *Table 1*. The *ECM* is highly significant with a p-value of .018, and we reject the null of no cointegration between the real exchange rate and Swedish and German productivity with long-run elasticities implied by (8).

The long run path given by (8) is plotted against the actual log real exchange rate in *Figure 2*. We see how the German reunification provided some space for a real depreciation of the Swedish real exchange rate or conversely a real appreciation of the German real exchange rate.

(ii) Short-run Dynamics

Throughout the analysis both quarterly differences using adjusted series, and annual differences using unadjusted series and seasonal dummies were used.[13] Estimating quarterly differences implies that there may be a lot of short-run noise in the data that suppresses genuine relationships. This is confirmed by the empirical results. The goodness of fit is generally much higher when estimating annual rather than quarterly differences and the fitted values correspond better to the actual values. In what follows we therefore focus on annual difference estimation using unadjusted series and seasonal dummies in the main text. [14]

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[13] When estimating quarterly differences, series showing seasonal patterns (unemployment and government consumption) were seasonally adjusted using the X11-filter.

[14] Tables of quarterly differences are available on request.

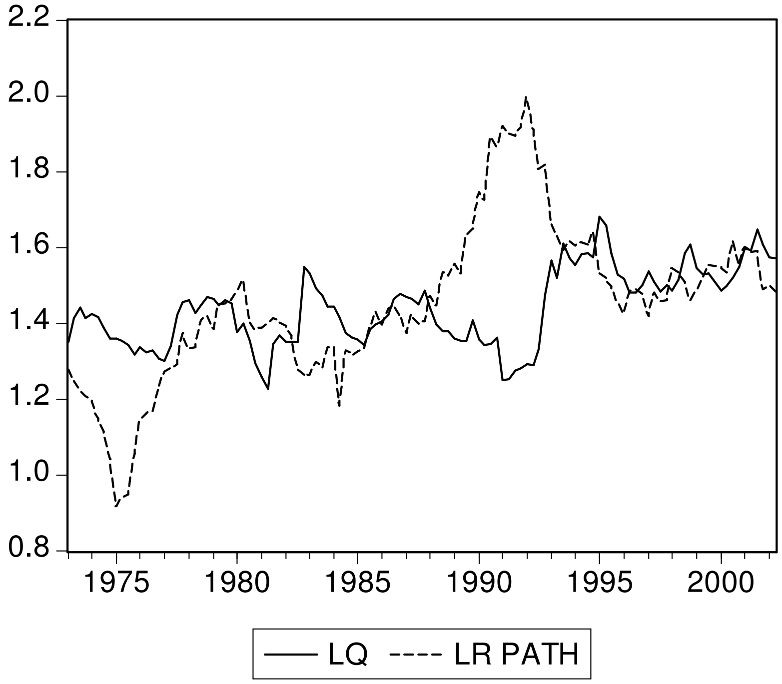

Figure 2: The log real exchange rate (*LQ*) and the estimated long run path (*LR PATH*) over the period 1973:1-2001:4.

Estimation of the general unrestricted model (4) renders many variables insignificant. The model is therefore reduced using Lagrange Multiplier tests for linear restrictions on single variables and groups of variables, until an interpretable model of significant variables is obtained. *Table* 2 presents reduced versions of model (4) using annual differences, $i = 4$.

The model in column (1) shows that the dynamics of the real exchange rate is determined by the feedback from previous periods as captured by the autoregressive (AR)components, the change in the budget balance, and the interest rate spread. A positive interest rate spread causes a real appreciation of the Swedish real exchange rate under the Currency Basket and under a floating exchange rate. The model indicates that the interest rate spread does not affect the real exchange rate when the exchange rate is fixed, albeit on the rather low 12 percent level of significance. This is consistent with our theory that the interest rate affects the real exchange rate through the nominal exchange rate and thus only matters under a floating regime.

We note that while log relative government consumption played a certain role when estimating quarterly differences in *Table 1,* it seems that the budget balance matters for annual differences in the real exchange rate. Recall our theoretical prediction that changes in log relative consumption might cause inflation expectations and therefore a nominal depreciation , while an improvement in the budget balance causes deflation and hence a real depreciation. Once again it seems that while the nominal exchange rate is the mechanism through which the real exchange rate is affected when estimating quartarly differences while the Swedish price level accounts for real exchange rate movements on an annual basis.

Table 2: Estimation of dynamic real exchange rate models by OLS, annual differences, 1973:1-2001:4

Column	(1)	(2)	(3)
Dependent variable	$\Delta_4 q$	$\Delta_4 q$	$\Delta_4 q$
$\Delta_4 (R - R^*)$	-.004*	.	-.004**
	(.002)	.	(.002)
$\Delta_4 (R - R^*) \cdot D_{fix}$	.004	.	.005*
	(.003)	.	(.003)
$\Delta_4 BDef$	.005**	.004**	.004**
	(.002)	(.002)	(.002)
ECM_{-4}	-.051*	-.049*	-.050*
	(.030)	(.030)	(.030)
$ECM_{-4} \cdot D_{float}$	-.142**	-.158***	-.142**
	(.062)	(.061)	(.062)
AR-components	3	3	3
Additive Dummies	Yes	Yes	Yes [a]
Seasonal Dummies	Yes	No	Yes
R^2_{adj}	.728	.731	.728
N	107	107	107
Breusch-Pagan LM-statistic (nR^2)	4.939	2.622	3.046

Notation as in *Table 1*.

[a] Excluding D_4

The error correction term has a significant negative effect on real exchange rate growth. A positive deviation from equilibrium by 1 percent causes a real appreciation of .051 percent under the Currency Basket and a fixed exchange rate, and an additional .142 percent under a floating exchange rate. In terms of long term adjustment it means that under the Currency Basket it takes about 4.9 years for 90 percent of the deviation to adjust (half-life 1.5 years), while it takes only 1.3 years under a floating regime (half-life .4 years). Hence it seems that the nominal exchange rate indeed acts as a shock absorber.

Removing the interaction term in column (2) renders both interest rate terms insignificant, which is rather re-assuring considering that the interest rate spread should be unable to affect the nominal exchange rate under a fixed exchange rate regime. Despite the low significance of the interaction term, we conclude that the model in column (1) is a good candidate for a preferred specification and we examine its properties in more detail. The fitted values, displayed in *Figure 3,* correspond rather well to the actual values.The graph shows that the largest deviation between actual and fitted values occurs around 1983. Incidently, removing the dummy for the large devaluation 1982:3, i.e. D_4 in column (3) renders the interest rate term significant on the 10 percent level. Moreover, plotting recursive parameters indicates that the estimated coefficients are rather stable, converge as the sample size is increased and that none of the parameters change their signs.

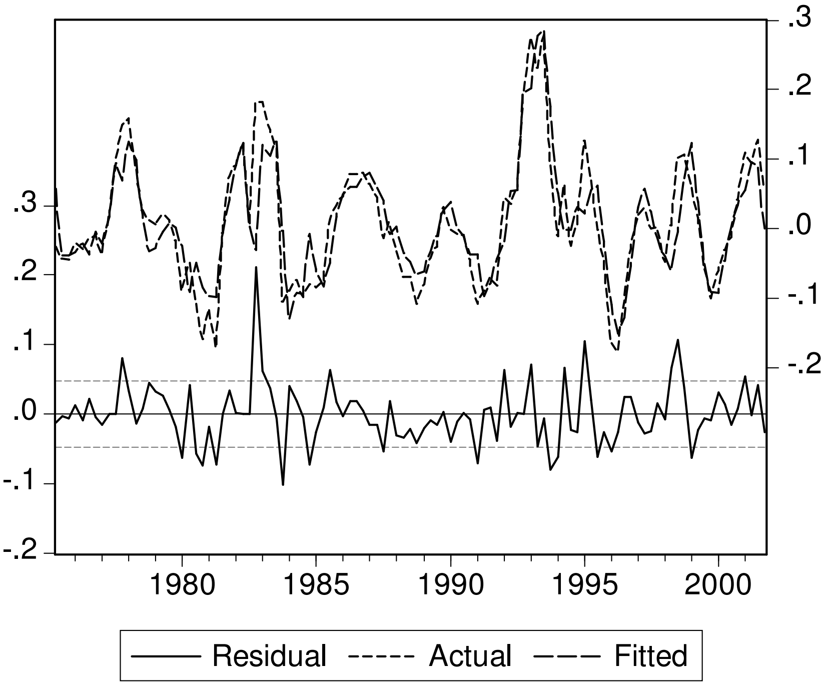

Figure 3: Actual and Fitted values for the model in *Table* 2, column (2)

Note that we initially allowed for regime-specific intercepts in all estimates, but they did not matter in any of the estimations. In addition some sensitivity analysis was made by estimating over 1973:1-1998:4 only, i.e. by eliminating the data for which Germany is a member of the EMU. The main results were unaffected by this exercise. To sum up, it seems that the regime has indeed mattered for the dynamics of the real exchange rate. More specifically, deviations from long run equilibrium are corrected more quickly when the nominal exchange rate has been allowed to float freely. This is a key result.

(iii) The Component-specific Impact of the ECM

Table 3 presents unrestricted and restricted versions of models (5) – (7) using annual differences. According to *Model 1* in *Table 3*, a positive deviation from long run equilibrium causes inflation in Sweden under a fixed exchange rate which is consistent with theory. Under a floating exchange rate however, the *ECM* does not affect prices (the hypothesis that the coefficients cancel is not rejected at the 5 percent level). Other determinants of Swedish inflation include the oil price and log relative government consumption.

Using the German price level as the regressand as suggested in equation (6) yields *Model 2*. As for the interpretation, our main concern is the impact of the *ECM* under different regimes. According to theory, a positive deviation from long run equilibrium should cause deflation in Germany, and thereby a real appreciation. We see from the results in *Table 3* that this holds true regardless of regime.

Table 3: Estimation of dynamic models of the three components by OLS, annual differences 1973:1-2001:4

Model 1	Dependent variable	$\Delta_4 p$		
	$\Delta_4 p_{oil}$	.010***		
		(.003)		
	$\Delta_4 (g - g^*)$	.195***		
		(.047)		
	$\Delta_4 (g - g^*) \cdot D_{fix}$	-.322***		
		(.080)		
	$\Delta_4 (R - R^*)$	.001**		
		(.000)		
	$ECM_{-4} \cdot D_{fix}$	.031***	AR components	1
		(.009)	N	111
	$ECM_{-4} \cdot D_{float}$	-.036***	R^2_{adj}	.940
		(.010)	Breusch-Pagan LM-statistic, nR^2	.065
Model 2	Dependent variable	$\Delta_4 p^*$		
	$\Delta_4 p_{oil}$	.006***	AR components	1
		(.002)	N	113
	ECM_{-4}	-.008***	R^2_{adj}	.909
		(.003)	Breusch-Pagan LM-statistic, nR^2	4.606
Model 3	Dependent variable	$\Delta_4 s$		
	$\Delta_4 (g - g^*) \cdot D_{float}$	1.056**		
		(.508)		
	$\Delta_4 (R - R^*)$	-.003*		
		(.002)		
	$\Delta_4 (R - R^*) \cdot D_{fix}$	.004*		
		(.003)		
	ECM_{-4}	-.136***	AR components	3
		(.043)	N	107
	$ECM_{-4} \cdot D_{fix}$	.152***	R^2_{adj}	.692
		(.052)	Breusch-Pagan LM-statistic, nR^2	.368

Notation as in *Table 1*.
Significant seasonal and additive dummies included.

However, when comparing the reaction of the Swedish and the German price level respectively we see that the effect on the Swedish price level is almost four times as large (.031 compared to .008). It therefore seems that the Swedish price level responds much more forcefully to short-run deviations from long run equilibrium than the German price level, which indicates that Sweden, being a small country, has to adapt to German conditions rather than the other way around.

Finally, the nominal exchange rate in *Model 3*, is affected by the change in the interest rate spread under the Currency Basket and a floating regime, but not under a fixed exchange rate regime. This is completely in line with what one would expect. Moreover, under a floating rate, higher growth in log relative government consumption causes a nominal depreciation, probably due to inflation expectations. A positive deviation from long run equilibrium triggers a nominal appreciation under the Currency Basket and a floating regime but not under a fixed exchange rate regime (the hypothesis that the coefficients cancel is not rejected at the 5 percent level). Hence, it seems that a floating nominal exchange rate acts as a shock absorber.

We conclude that all three components have contributed to the adjustment of the real exchange rate to its long run path when the regime has allowed for such adjustments.

6 Concluding Remarks

This paper provides empirical evidence on the long run behavior and short run dynamics of the Swedish real exchange rate relative to Germany during the post-Bretton Woods period. The results show that there is cointegration between the real exchange rate and Swedish and German labor productivity, supporting the Balassa-Samuelson hypothesis.

We conclude that the most important short-run explanatory variables are the interest rate spread, the feedback from previous periods and the deviation from long run equilibrium. The regime has mattered for the dynamics of the real exchange rate. It seems that disturbances are adjusted more quickly when the exchange rate has been allowed to float freely. Thus is seems that the nominal exchange rate acts as a shock absorber.

Finally, the results provide some evidence that all three components have contributed to correcting distortions, but that the Swedish price level and the nominal exchange rate have responded more forcefully than the German price level to deviations from long run equilibrium. Therefore, it appears that Sweden, being a smaller country than Germany, is forced to adapt to German conditions and respond to distortions rather than the other way around.

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