NONLINEARITIES IN EXCHANGE RATE: EVIDENCE FROM SMOOTH TRANSITION REGRESSION MODEL

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Abstract

The purchasing power parity puzzle, exchange rate disconnection to macroeconomic fundamentals and pricing to market are central issues of international macroeconomics. Recent research has suggested that these issues can be presented by nonlinear behaviour. In this dissertation, we examine and explain the nonlinearities in the form of regime switching behaviour in real exchange rate series, exchange rate and macroeconomic fundamentals relation and exchange rate pass-through into consumer and import prices. Overall, we find evidence that nonlinearities are important in analysing empirical exchange rate models. The dissertation consists of four self-contained empirical studies.

In chapter 2 we examine whether the Markov switching models and exponential smooth transition autoregressive models can give any additional insights into real exchange rate behaviour for several OECD countries. The results show that there are long swings in the real exchange rate series, which can be characterize as a depreciation and an appreciation regime. These regimes are very persistent, although the processes are eventually mean reverting.

We estimate a multivariate smooth transition autoregressive model for the euro/dollar exchange rate in chapter 3. The significant point of our analysis is the possibility that a nonlinear specification for the exchange rate series might reveal aspects of the exchange rate dynamics that cannot be picked up by linear models. We find that the euro/dollar exchange rate may display random walk or near random walk behaviour within a certain range but the ability of the exchange rate to wander without any bound is limited by long-term government bond interest rate differentials.

In chapter 4 we examine nonlinear relationships between macroeconomic fundamentals and exchange rate for G-7 countries. We estimate a smooth transition error correction model that allows for parameter variation in the error correction form and interest rate differentials. The nonlinearity is determined by the inflation rate differentials between countries. We find significant error correction terms in monetary models. Our findings suggest the importance of nonlinear dynamics for examining deviations from the long-run equilibrium.

We examine whether the degree of exchange rate pass-through is dependent on importing country inflation rate in chapter 5. Our model shows that import prices respond differently to exchange rate changes when we are in a high inflation regime compared to a low inflation regime. We also present empirical evidence by estimating pass-through elasticises for several OECD countries. We find that consumer prices are not very sensitive to exchange rate changes. For aggregate import prices, we find partial or full exchange rate pass-throughs.

The tested nonlinear regime specific models proved appropriate for testing exchange rate dynamics for several currency pairs. Furthermore, we were able to present that macroeconomic fundamentals are important predictors of exchange rates.

Keywords: exchange rate economics, monetary model, nonlinear estimation, pricing to market, purchasing power parity

To my daughter, Noora

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Oulu, November 2005 Marko Korhonen

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1 Introduction

1.1 Background and the purpose of the study

The question whether exchange rates are connected to macroeconomic variables is an old one. As such, a great deal of attention has been focused on explaining the behaviour of exchange rates and the view of the exchange rate determination has changed in history. During the Bretton Woods system[1] from the late 1940s to the early 1970s, the exchange rate adjustment was widely understood in terms of restoring equilibrium in the balance of payments (see Obstfeld 2002). Since the collapse of the Bretton Woods system in the early 1970s, most major currencies have been allowed to float more or less freely. The adoption of freely floating exchange rate regimes by many industrialized countries, in 1973, started a new era of increased exchange rate uncertainty. Not surprisingly, the exchange rate models have not been able to produce good empirical forecasts for exchange rate changes. Furthermore, the euro's launch in 1999 raised the question of its equilibrium level against the other major currencies, most notably against the US dollar. An interesting question is whether the launch of the euro will start a new era of modern exchange rate floating.

In integrated international markets, the exchange rate is the key relative price. However, there is no clear evidence how to examine this key variable or predict its behaviour. Many studies have shown that typical macroeconomic fundamentals, such as relative money supplies, relative output levels, and relative consumption levels, do not explain empirically exchange rates at least for the short-run. In a well-known paper Meese and Rogoff (1983) found that random walk dominates exchange rate behaviour of the 1970s. Their main conclusion was that none of the structural exchange rate models were able to forecast exchange rates better than a simple random walk model. Subsequent research has found similar results. This disconnection between fundamentals and exchange rates has been named "an exchange rate disconnection puzzle" (Obsfeld and Rogoff 2000). This

[1] The establishment of the International Monetary Fund (IMF) at the Bretton Woods Conference in 1944 started the period of fixed exchange rate, which subsequently became known as the Bretton Woods system (Sarno and Taylor 2003).

stylized fact is, however, inconsistent with the theories of exchange rate determination and it is one of the major puzzles in open macroeconomics literature[2].

The previous studies have generally estimated exchange rate models in a linear framework, which implies a constant relationship between macroeconomic variables and exchange rates. Exchange rate models have been traditionally presented as linear models or nonlinear models that have been linearised around equilibrium point. However, many economic processes are inherently nonlinear and linearization might lead to serious misspecifications. This is especially true when the variables of the model and/or parameter estimates are subject to changes during sample periods. Recently, there have been studies, which claim that the relation between exchange rates and macroeconomic fundamentals is asymmetric (see Sarno 2003 for a survey). One explanation of asymmetries in exchange rate behaviour may be the interaction between noise traders and arbitrage traders (McMillan 2005). Interaction between these traders may create a band of inactivity around fundamental equilibrium and arbitrageurs only act after prices have diverged to a significant extent from their equilibrium levels[3]. A variant of the traditional exchange rate models involves the use of time-varying parameter model. This approach can be rationalized, for example, in response to policy regime changes (see Lucas 1976).

The connection between exchange rates and macroeconomic fundamentals is central to the conduct of monetary policy. For example, real exchange rate is a crucial measure of competitiveness and, hence it is important to be able to measure exchange rate misalignment for policy making and policy design. Furthermore, exchange rate equilibria are important to currency unions and selling parities in monetary unions. One of the key questions for policy makers is: How much of the decline in currency pass through to import prices and to overall consumer prices? For example, during the period between February 2002 and May 2004 the real value of the U.S. dollar fell by 19.1% relative to the other major currencies. However, the import prices did not respond by rising 19.1%, and only a part of the dollar depreciation passed through to import prices. The pass-through may be less than full, since foreign exporters may be willing to reduce their profit margins in order to maintain their market shares[4].

In this study we examine structural exchange rate models by using nonlinear estimation. We are interested in testing whether there is a nonlinear relationship between the real and nominal exchange rates and macroeconomic fundamentals. We share the view of many researchers that standard linear models may be misspecified and, thus, unable to fully capture exchange rate dynamics. In all of the chapters we will use the post Bretton

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[2] However, it should be mentioned that the recent studies that use real time data can explain relatively well the direction of exchange rate changes (see Ehrmann and Fratzcher 2005). Furthermore, Groen (2000), Mark and Sul (2001) and Rapach and Wohar (2004) use panel estimation and obtain coefficient estimates that are reasonably consistent with structural exchange rate models.

[3] Taylor and Allen (1992) surveyed market participants in the London foreign exchange rate market and report that 90 percent of those surveyed used some chartism on the short-term horizon and about 60 percent viewed chartism at least as important as fundamentals. When forecasting longer term, most respondents (85 percent) thought fundamentals were more important than chartism.

[4] The empirical studies on pass-through find that an average pass-through for industrialized countries is about 60%, with the greatest effect occurring within four quarters since the change in the exchange rate (see Goldberg and Knetter 1997). However, Campa and Goldberg (2002) note that pass-through rates vary a lot among products and across countries.

Woods data. It turns out that in exchange rate models the empirical linearity can be rejected in several cases.

This study consists of four chapters. Although the chapters are self-contained, the common denominator is that they deal with exchange rates and macroeconomic fundamentals. In chapter 2 we examine the mean reversion of real exchange rate series. Chapter 3 examines whether the exchange rate between the U.S. dollar and the euro can be presented in a nonlinear model that is based on capital and goods market equilibrium. In chapter 4 we examine exchange rate behaviour based on the traditional monetary model for G7 countries by a using time-varying error correction framework. Chapter 4 will use the new open economy macroeconomics framework and examine the U.S dollar exchange rate change's pass-through into consumer and import prices for G7 countries. There we closely follow Taylor's (2000) argument that exchange rate pass-through is dependent on the monetary policy of importing country and we will estimate the exchange rate pass-through parameters that are dependent on the importing country's inflationary environment.

1.2 Exchange rate theories and recent evidence

Purchasing power parity (PPP) is a key building block in determining the equilibrium exchange rate. The key element of the PPP concept is the "law of one price" (LOOP), which states that the purchasing power of a unit of one currency should be the same in all economies[5]. This is termed as the absolute PPP[6]. In reality, we observe that there are several homogeneous commodities sold at different prices in different locations, which is inconsistent with the rationale of the LOOP. Furthermore, the absolute PPP needs two basic assumptions, which are difficult to verify in practise. First, the price indices must cover identical baskets of tradable goods, and, secondly, by the assumption of index variables, the PPP must hold in the base year. Moreover, tariffs, non-tariff barriers, and transportation costs can all prevent the absolute PPP to hold[7]. The relative PPP rests on variations in absolute PPP. The relative PPP holds when the rate of depreciation of one currency relative to another is equal to aggregate price differences between the considered countries. The assumption that the PPP holds in the index base year is, thus, no longer necessary.

A vast amount of empirical research tested the PPP hypothesis by employing the unit root and cointegration methodologies. The early work concentrated to the use of the conventional Dickey-Fuller unit root tests. These studies find that the unit root hypothesis could not been rejected implying the absence of the PPP. Another group of studies uses more powerful panel data unit root tests (see Wu 1996 and Papell 1998 among others).

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[5] Identical goods in different countries must have the same price, when expressed in the same currency. The rationale behind the LOOP is an arbitrage hypothesis, which states that if two homogeneous commodities are traded at different prices in different locations, then there occurs a profitable arbitrage opportunity that arbitrageurs can buy the good in one place and sell it at a higher price in another place.

[6] The absolute PPP exchange rate between two countries is the ratio of the price indices.

[7] Rogoff (1996) notes that most economists believe that the PPP holds in the long run, which implies that real exchange rates time series mean reverting.

Overall, the empirical research has concluded that PPP fails to hold at least in the short run. Rogoff (1996) describes the "purchasing power parity puzzle" as the difficulty to connect high short-term volatility in exchange rate series with very slow adjustment to PPP. However, some recent empirical papers have used more sophisticated time series methods (i.e. cointegration tests) and claim that the PPP holds in the long run. These studies claim that PPP mean reversions are approximately from three to five years. The mystery comes from the fact that the adjustment takes much longer than can be explained by sticky prices (see Froot and Rogoff 1995). Recently *e.g.* Obstfeld and Taylor (1997), Taylor (2001), and Kilian and Taylor (2001) have examined the nonlinear adjustment of the real exchange rate series. These studies provide empirical evidence that there is an effect of transaction and distribution costs, which prevents the occurrence of the LOOP in all markets (see Burstein et al. 2005). According to this view the real exchange rate dynamics should be seen as mean reverting only when the price differentials are larger than the no arbitrage transaction band (Dumas 1992).

The monetary model of exchange rate determines exchange rate changes as a function of traditional fundamentals such as relative output and relative money stocks, and other fundamentals, such as interest rate differentials and relative trade balances[8]. The monetary model was the dominant exchange rate model at the start of the floating exchange rate era in the early 1970s. Although the monetary model is very attractive, it concentrates solely on equilibrium conditions in one market, namely the money market. The other markets, such as goods, labour, foreign exchange, domestic and foreign bond, are assumed to be in equilibrium. Hence, in the monetary model, the exchange rate adjusts freely to achieve equilibrium in the foreign exchange market.

A huge literature has examined the relationship between the nominal exchange rate and the macroeconomic fundamentals suggested by the monetary model of exchange rate determination. The economic fundamentals appear to be more important at longer horizons, while the short-run deviations from the fundamental level of exchange rate are attributed to excess speculation (Taylor and Allen 1992). Cheung et al. (2004) examine exchange rate prediction by using a wide set of models that have been proposed in the last decades. They find that no model consistently outperforms a random walk by the means of squared error measure. However, they find that some model specifications that work well in one period do not necessarily work well in another period. The lack of empirical evidence on the exchange rate models has led some researchers to propose nonlinearity in the relation of exchange rate and macroeconomic fundamentals. That is, the larger the deviation from fundamental equilibrium is, the faster the exchange rate will be driven back to its equilibrium level.

The monetary model or the PPP are theoretically and intuitively appealing models for exchange rates, but empirically explain very poorly the exchange rate dynamics in the linear form. Several theoretical models of exchange rate determination imply nonlinear functional forms of smooth adjustment towards long-run equilibrium. On the theoretical level, nonlinearity in exchange rate dynamics is predicted at least by the target-zone (Krugman 1991) approach, within the concept of bubbles with self-fulfilling expectations

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[8] We should also note the argument in a recent paper by Engel and West (2005), that for countries and data where exchange rates and fundamentals appear to be linked by a long-run relationship, it is possible that exchange rates help predict fundamentals, rather than the other way around.

(Flood and Garber 1980), and intrinsic bubbles[9] models (Froot and Obstfeld 1991). Also, fads or noise trading may create a nonlinear relationship between fundamentals and exchange rate series as was shown for example in De Long et al. (1990). Furthermore, government policy (i.e. either allowing the exchange rate to float or devalue the domestic currency from one fixed rate to another) might lead to nonlinearities in the relationship between exchange rate and fundamentals (see Flood and Marion 1998). In general, these types of exchange rate models imply that the speed of adjustment in exchange rate dynamics responds to fundamental variables in a nonlinear way.

A number of authors have also developed theoretical models that examine the nature of adjustment process when the agents face transaction costs, the sunk costs of arbitrage or there are "institutional" or other limits to arbitrage (see e.g. Coleman 1995, Dixit 1989, Dumas 1992, Uppal 1993, Sercu et al. 1995, O'Connell 1998). These authors demonstrate that transaction costs, sunk costs etc. induce nonlinear adjustment towards equilibrium. The whole data generating process can be globally mean reverting, but this kind of nonlinearity has a property that processes exhibit near unit root behaviour for small deviations from equilibrium. The reason is that small deviations are left uncorrected if they are not large enough to cover transaction costs or the sunk costs of international arbitrage. However, according to Kilian and Taylor (2001) the transaction costs could not provide a compelling explanation of long swings in nominal exchange rate like the large and persistent overvaluation of the US dollar during the mid-1980s, nor do they explain the observed volatility in both real and nominal exchange rates. Kilian and Taylor suggest a model in which uncertainty about the fundamental values of the exchange rate deters agents from speculating against small deviations from fundamentals[10]. One possibility for the nonlinear dynamics in exchange rate behaviour may rise because small deviations may be considered unimportant by the market and policy makers but when the deviations become large enough the pressure will be strong from both market makers and policy makers to bring the exchange rate at least close to the fundamental equilibrium (Taylor and Peel 2000).

There are also several empirical papers that have examined the evidence of nonlinear adjustment in the exchange rate deviation from fundamental equilibrium level (e.g. Balke and Fomby 1997, Taylor and Peel 2000, Taylor, Peel and Sarno 2001). Meese and Rose (1989) note that one important source of nonlinearity might be that the data generation process may be intrinsically nonlinear. For example, the models of rational expectations are intrinsically nonlinear, since in these models forward looking agents forecast the future time path of fundamentals. However, if agents expect that government reaction functions are subject to stochastic change or that authorities will regulate the fundamentals driving the exchange rate when the latter approaches or reaches the band of a target zone, then the appropriate functional form may be nonlinear. Kräger and Kugler (1993) find evidence that exchange rates might show regime-switching behaviour, with the outer regimes exhibiting much higher variances than the inner regime. They argue that this finding is probably due to central banks' interventions aimed at avoiding excessive appreciation or depreciation of a currency. Intuitively, monetary authorities may intervene in the

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[9] Intrinsic bubbles are driven by fundamentals alone in a nonlinear way (Ma and Kanas 2000)

[10] Their model assumes that the fundamental value of exchange rate series is only a function of relative prices between domestic and foreign countries.

foreign exchange market as a reaction to large depreciations or appreciations of a currency, which lead to different behaviour for moderate and large changes of the exchange rate. Similar behaviour may be observed for an exchange rate, which is constrained to lie within a prescribed band or target zone, as was the case in the ERM in Europe. In this case, the level of the exchange rates, rather than its change determines the regimes (Franses and van Dijk 2001).

Recent research on open economy macroeconomics uses dynamic intertemporal approaches in analyzing exchange rates. The models synthesize older sticky-price models of macroeconomic fluctuations with dynamic intertemporal approaches. The foundation of these models is microeconomic optimisation, which allows rigorous welfare analysis of economic policies. This new era of open economy macroeconomics literature has become known as the new open economy macroeconomics[11]. This literature has allowed several forms of market imperfections, such as price and wage rigidity, within fully specified dynamic general-equilibrium models.

The econometric results of the LOOP reveal evidence that market imperfections might be associated with high persistence of real exchange rate mean reversion. It is well known that with constant elasticity of demand, a monopolist supplier may charge a different price in different places, but exchange rate changes do not cause the price differentials (Obstfeld and Rogoff 2000). According to this view the monopoly producers have very broad scope to price to market by charging different prices in different markets. Broadly speaking, there are two alternative ways for firms to set their goods prices in international markets. First, prices can be set in the currency of producer (producer currency pricing, PCP). This implies that there is full pass-through of exchange rate changes to consumer prices. Obstfeld and Rogoff (2002) have noted that under the PCP it is not desirable to target exchange rates and, thus, an appropriate monetary policy can be achieved without any coordination. Second, firms can set their prices in the importing country currency (local currency pricing, LCP), which implies that exchange rate changes have only a partial or no effect at all to consumer prices. For example, Devereux and Engel (2003) note that it is reasonable to consider final consumer prices of commodities as being sticky in domestic currency terms. According to them, an optimal monetary policy is to fix the nominal exchange rates under the LCP.

1.3 Nonlinear estimation and time-varying error-correction

In recent years there has been an ample of empirical evidence that economic time series and their relations contain nonlinearities. The use of nonlinear models in explaining economic phenomena implies that the behaviour of economic variables depends on different states of the world or regimes. This means that properties of economic time series are dependent on the regime which prevails during a certain period of time and, thus, the statistical properties of economic time series, such as its mean, variance and autocorrela-

[11] This literature is stimulated primarily by the seminal work of Obstfeld and Rogoff (1995). For a surveys see, for example, Lane (2001) and Sarno (2001)

tion might vary in different regimes. For example, the dynamics of several macroeconomic variables depend on the phase of the business cycle.

Several theories in exchange rate economics suggest a nonlinear relationship between exchange rates and different macroeconomic variables. However, different theories produce also a different shape of nonlinear functional forms. The problem in the nonlinear econometrics is the large number of possible nonlinear specifications. In this thesis we consider two types of nonlinearities. First, we will use the Markov Switching (MS) model, which was popularised in economics by Hamilton (1989)[12]. The motivation of the MS is to model long swings in the data and discrete switches in the dynamics of the series. The models of the MS assume that the regime cannot actually be observed but is determined by an underlying unobservable stochastic process. Second, we will use smooth transition autoregressive (STAR) models. The seminal works of the STAR models are Teräsvirta and Anderson (1992), Granger and Teräsvirta (1993) and Teräsvirta (1994)[13]. The STAR framework makes it possible to find and estimate a nonlinear model without making any strong assumptions about the functional form. In the literature the STAR models have been successfully applied in a wide range of econometric models.

Several economic variables are linked by long-run equilibrium relationships, or in other words, these variables are cointegrated. The concept of cointegration, introduced by Granger (1981), Engle and Granger (1987) and Johansen (1988), is widely used in econometrics or applied economics literature. However, a standard cointegration model presumes that the error correction mechanism towards the long-run equilibrium is always present and that the strength of error correction does not change during the sample period. A number of studies have recognised that adjustment to long-run equilibrium in financial markets may depend on the size of the disequilibrium. This means that a linear error correction model may be inappropriate to model, for example, the exchange rate series dynamics. The recent literature has shifted towards the use of so called smooth transition error correction models (STECM). These models use linear long-run relationships between nonstationary variables, but the adjustment towards equilibrium is dependent on some disequilibrium or some exogenous variable (see *e.g.* Granger and Swanson 1996, Anderson 1997, Weise 1999 and van Dijk et al. 2000).

1.4 Main findings of the study

The main result of this study is that by allowing for nonlinearities in the form of regime switching behaviour in the underlying data-generating process for the exchange rate macroeconomic fundamentals models we are able to capture the impact of fundamentals on the dynamics of exchange rates. Our econometric analysis shows that the regime switching approach produces a statistically significant error correction model for both real and nominal exchange rates. This is an extension of the relevant literature. Furthermore, by allowing nonlinearities in exchange rate dynamics we are able to argue that macroeconomic fundamentals, indeed, help to predict the exchange rates, and not vice versa as

[12] The original idea comes from Goldfeld and Quandt (1973).

[13] The model was originally suggested by Chan and Tong (1986)

suggested recently by Engel and West (2005). Below, we describe the contents of each chapter in more detail together with a brief summary of the main findings.

1.4.1 Chapter 2: An empirical investigation of two nonlinear models in real exchange rate series

One reason why prices are not equivalent in different geographical locations may be that they imply the existence of transaction and transportation costs and other trade impediments. The objective of this chapter is to examine whether the deviations from PPP are mean reverting in the presence of nonlinearity. The chapter examines the possibility that deviations of the real exchange rate series from equilibrium level affect to the short-run dynamics in a nonlinear way. Two different nonlinear models are presented to solve the slow convergence to long-run equilibrium level. Particularly, we examine whether the Markov switching (MS) models and exponential smooth transition autoregressive (ESTAR) nonlinear models can give any additional insights into real exchange rate behaviour during the period 1974-1998. The theoretical bases of our examination are Dumas (1992), Uppal (1992), Sercu et al. (1995) and Obstfeld and Rogoff (2000). They show that the adjustment of real exchange rate towards the purchasing power parity in the presence of market frictions is necessarily a nonlinear process. There are market frictions that imply a band of inaction, within which the deviations from long-run equilibrium are left uncorrected. The key theoretical idea is that the deviations from the LOOP will not be mean reverting as long as they are smaller than the band of arbitrage costs. However, when the deviations from the LOOP cross the band of inaction, the real exchange rate series are mean reverting.

The results from the MS models show that there are clearly long swings in the real exchange rate series, which can be characterized as a depreciation regime and an appreciation regime. Furthermore, the probability of staying in the regime is relatively high, although the process is eventually mean reverting. The ESTAR model implies random walk behaviour for small deviations from PPP and fast adjustment for large deviations, which are in line with the MS model. Our empirical results suggest that, in general, the real exchange rate series can be characterized rather well by using nonlinear estimation. Overall, this chapter provides strong evidence for the real exchange rate mean reversion.

1.4.2 Chapter 3: Nonlinear equilibrium correction model and the euro/dollar exchange rate dynamics

This chapter examines the euro/dollar exchange rate. We use synthetic euro data from January 1990 to July 2002. We estimate an ESTAR model specification for the euro/dollar exchange rate. Specifically, we investigate whether the euro/dollar exchange rate disequilibrium can be presented in a nonlinear error correction form. We argue that the long-term government bond differential between the U.S. and the eurozone governs the speed of adjustment to exchange rate equilibrium level. The significant point of our analysis is the possibility that a nonlinear specification for the exchange rate series might reveal aspects of the exchange rate dynamics that cannot be picked up by linear models. An advantage of this specification is that it reveals a possible instability in the exchange rate and macroeconomic fundamentals relation.

Our results confirm that the use of nonlinear error-correction specification is a better empirical model for the exchange rate dynamics in relation to linear models. We find that the euro/dollar exchange rate may display random walk or near random walk behaviour within a certain range, but the ability of the exchange rate to wander without any bound is limited by long-term government bond interest rate differentials between the eurozone and the USA. The euro/dollar exchange rate is unrelated to fundamentals long periods, but during some other periods there is a very strong mean reverting property towards the equilibrium level. The view of the euro/dollar exchange rate presented in this chapter has interesting implications for the future of the euro/dollar rate and, especially, for the way how macroeconomic factors affect to this widely followed exchange rate.

1.4.3 Chapter 4: Analysing the monetary model of exchange rates in a time-varying error-correction framework

The purpose of this paper is to re-examine the monetary model of the exchange rate. We examine nonlinear relationships between macroeconomic fundamentals and exchange rates for G-7 countries. Several papers have investigated the predictability of exchange rates by cointegration methods. There is evidence of the long-run relationship between nominal exchange rates and monetary fundamentals in the post Bretton Woods era. However, several papers argue that these cointegration methods assume falsely a linear long-run relationship between the variables, and that this relationship is always present. Goldberg and Frydman (2003), for example, note that during some sub-periods of floating exchange rates the macroeconomic models explain exchange rate movements quite well and during other sub-periods quite badly. This temporal inconsistency of monetary models provides support for the view that exchange rate models with fixed parameter coefficients might not be valid.

We use a time-varying error correction framework to improve the empirical results from the flexible price monetary model for the post Bretton Woods period. We estimate a smooth transition error correction model that allows for parameter variation in the error correction term and interest rate differentials. The nonlinearity is determined by the inflation rate differentials between countries. In contrast to previous studies we find significant error correction terms in the monetary model. Overall, our findings suggest that it is important to allow nonlinear dynamics for examining deviations from the long-run equilibrium.

1.4.4 Chapter 5: Exchange rate pass-through: Dependence on inflation regime

This chapter examines the exchange rate pass-through into consumer prices and aggregate import prices for several industrialized countries. We examine whether the degree of exchange rate pass-through is dependent on importing country inflation rate as suggested in Taylor (2000). Our model gives further explanation for this phenomenon. We extend standard monopolistic competition mark-up pricing model to importing country inflation rate regime dependent setting. The model shows that import prices respond differently to exchange rate changes in a high inflation rate regime than in a low inflation regime. We also estimated the pass-through elasticises for several OECD countries. We use nonlinear smooth transition estimation methodology, which takes into account the possibility of pass-through that is dependent on inflation rate regimes. Estimates show that the pass-through is incomplete and positively correlated with the importing country's inflation rate for several OECD countries. We find that consumer prices are not very sensitive to exchange rate changes and local currency pricing might be a good pricing assumption at the consumer level. For aggregate import prices, we find partial exchange rate pass-through.

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2 An empirical investigation of two nonlinear models in real exchange rate series

2.1 Introduction

In this chapter we examine the nonlinear dynamics of adjustment to the long-run PPP. Meese and Rogoff (1983) documented the failure of linear nominal exchange rate models. Some recent studies have provided evidence of the empirical failure of the linear real exchange rate models (see for example Michael et al. 1997, Sarantis 1999, Sarno 2000, and Baum et al. 2001). The number of both theoretical and empirical extensions of the nonlinear real exchange rate models have recently been growing. One possible source for nonlinear dynamics in the real exchange rate series is that market frictions, such as transportation costs and tariffs, permit deviations from the law of one price to be quite large without precipitating goods arbitrage (O'Connell 1998). A number of researchers have argued that transaction costs are a key explanation for relatively low adjustment speed in PPP (Benninga and Protopapadakis 1988 and Sercu, Uppal and Van Hulle 1995). The theoretical support for transaction cost approach have been provided by Dumas (1992), Williams and Wright (1991), Uppal (1993) and Coleman (1995). They show that in the presence of transaction costs, the adjustment of real exchange rates toward PPP is not necessarily a linear process. Heckscher (1916) made the same observation a long time ago, but there have been only a few empirical analyses of this observation. Another possible source of nonlinearities is the chartist behaviour of market participants, which may imply large movements in the real exchange rate series that are unrelated to fundamentals (see e.g. Taylor and Allen 1992 and McMillan 2005).

Dumas (1992) provides a nonlinear model structure for the PPP deviations. In his one good and two countries model, the deviations from PPP (i.e. the dynamics of the real exchange rate) should follow a nonlinear mean reverting process if we assume spatially separated markets and proportional transaction costs. He also shows that the speed of adjustment towards equilibrium varies nonlinearly and depends on the extent of the deviation from PPP. According to Dumas' model the process is divergent if we are inside the transaction band and the deviations will be rapidly extinguished if we are outside it. The existence of other factors, such as the uncertainty of the shock persistence or sunk costs of the activity of arbitrage, may widen transaction bands (Krugman 1989 and Dixit

1989). Dumas' model implies that deviations from PPP can last a very long time, although they do not necessarily follow a random walk (Michael et al. 1997). Although the whole real exchange rate process displays mean reversion, the probability of a move away from the parity value is greater than the probability to move towards parity. This asymmetric behaviour suggests the possibility that the real exchange rate can be characterized by long swings. Dumas' theoretical model is one way to explain why the empirical studies for mean reversion in real exchange rates have been largely unsuccessful.

Many authors have shown that the European exchange rate dynamics is typically characterized by periods of relatively calm phases, when the exchange rate arrangements are credible, and by periods of sudden and short-lived phases of speculative attacks (Engel and Hakkio 1996, Peria 2002, and Bessec 2000). For example, Bergman and Hanson (2005) suggest that the real exchange rate between the major currencies in the post Bretton Wood period can be described by a stationary, two state Markov switching regime AR(1) model. They find that Markov models outperform the nonstationary real exchange rate models. Bessec (2000) also uses Markov models and finds that the European exchange rates of the ERM members display mean reversion in the credible exchange rate regime, and that they adjust to the PPP during the more volatile period.

Michael et al. (1997) applies a smooth transition autoregressive model to analyze long-run PPP for long period data. Their study shows that the nonlinear model provides strong evidence of mean reverting behaviour for PPP deviations. An interesting feature in their work is that the estimated model is consistent with Dumas' hypothesis (i.e. real exchange rates behave like random walks for small deviations, but are strongly mean reverting for large deviations). Obstfeld and Taylor (1997) uses a band threshold autoregressive (B-TAR) model to capture the nonlinear behaviour of the real exchange rates. They assume that nonlinearity of the real exchange rate comes from the transaction costs. In their model, the equilibrium value of the real exchange rate can be anywhere in the band and not necessarily in any fixed equilibrium point of PPP. Their model provides evidence of mean reversion for real exchange rate series, which are as low as two months mean reversion in some cases. O'Connell (1998) tests a threshold autoregressive (TAR) model for the post Bretton Woods data and finds that there are no differences between large and small deviations from PPP. He argues that both deviations are equally persistent. Kuo and Mikkola (2000) compare the forecasting performances of a linear AR model and TAR model for US/DEM real exchange rate. They find that the TAR model does not give consistently better forecasts than the linear AR model.

We extend the recent studies of linear real exchange rate models to the nonlinear models. The traditional empirical studies show that deviations from PPP last very long time and it is difficult to reject the real exchange rate nonstability in linear models. The potential reason is that the traditional linear unit root tests have very low power for near unit root processes. We show that there is significant evidence of regime dependency in the real exchange rate series. This is an important finding since it uncovers stability of the real exchange rate series. Our theoretical background is the Dumas' model for the real exchange rate series. We use two alternative nonlinear models, which can detect the mean reversion in real exchange rate series. First, we apply the methodology of the Markov switching regime models to describe the real exchange rate process in the post Bretton Woods period. The Markov models allow for two states to exist with a series of shifts between the states occurring in a probabilistic fashion. Thus, the shifts occur endogenously rather than being imposed by the researcher. We show that the level of real exchange rate is generated by a stationary regime shifting Markov two state processes, where the states can be named as a depreciation regime and an appreciation regime. Second, we use smooth transition autoregressive (STAR) type of nonlinear econometric modelling technique. The adjustment of the STAR model is smooth instead of discrete as in the Markov model. We apply especially exponential STAR (ESTAR) models to exploit nonlinear dependencies in real exchange rates. These models were originally developed by Teräsvirta and Anderson (1992), and their statistical properties and estimation were examined in Granger and Teräsvirta (1993). We show that linear nonstationarity tests may not be able to detect real exchange rate mean reversion if the true real exchange rate series is a stationary nonlinear process. Also, the speed of adjustment towards long-run real exchange rate level varies more than proportionately with the size of shock.

2.2 Nonlinear adjustment to PPP

The last ten years have witnessed an explosion of interest among econometricians in testing, estimation, and specification of nonlinear models (Potter 1999). The successful modelling of nonlinearities in macroeconomic time series will help to understand the data generation mechanism behind the series. According to Potter, it is important that at least two conditions are fulfilled in order to properly model the nonlinearities in the macroeconomic time series. Firstly, economic time series must actually contain nonlinearities, and secondly we need a statistical method, which reliably notifies the possible nonlinearities in the data. The first condition is usually easy to fulfil but the second one is much more problematic.

The long run PPP can be written as

$$s_t = \beta_0 + \beta_1 p_t + \beta_2 p_t^* + q_t \tag{2.1}$$

where s_t is the log of the nominal exchange rate (expressed as domestic price of foreign currency), p_t and p_t* are the logarithms of the domestic and foreign price indices, β_0 is a constant, and q_t is under the null hypothesis a stationary error term representing the real exchange rate series. The strong form PPP imposes the joint restrictions of symmetry ($\beta_1 = \beta_2$) and proportionality ($\beta_1 = -\beta_2 = 1$). These restrictions are not, however, consistent with the empirical data due to for example measurement errors (Taylor 1988 and Cheung and Lai 1993), differential composition of price indices (Patel 1990) and differential productivity shocks (Fisher and Park 1991) as pointed out (Baum et al. 2001). However, a necessary condition for the weak form PPP to hold in the long run is that the real exchange rate (q_t) in (2.1) without any parameter restrictions is stationary. Thus, we estimate the PPP relationship parameters in (2.1) by ordinary least squares regression. The estimated parameters are presented in the Appendix (Table 6). Cheung and Lai (1993) and Baum et al. (2001) have used similar method in their studies. Our parameter estimates are in line with these studies in that the estimated coefficients are not consistent with the strong form PPP. We conclude that it would be rather unreasonable in our nonlinear model examination to assume the strong form PPP restrictions in model (2.1) and by estimating the parameters we avoid the imposition of incorrect coefficient restrictions, which could affect the true mean reversion property of real exchange rate series.

The linear cointegration methodology is a useful way of testing the long-run PPP theory. While cointegration implies mean reverting behaviour in the relationships between analyzed variables, the conventional cointegration methodology assumes a linear process (i.e. the adjustment process to the equilibrium is both continous and has a constant speed). However, the failure of linear cointegration could be a consequence of the nonlinear behaviour of the system (2.1). In addition, as pointed out by Pippenger and Goering (1993), the possibility of nonlinear behaviour or transaction boundaries in the real exchange rate series brings into question the usefulness of the linear cointegration methodology and the unit root test methodology. They find that the power of conventional unit root and cointegration tests falls dramatically under a threshold process. In addition, an interesting line of research considers the importance of nonlinearity in the error correction mechanism of the real exchange rate towards its long run equilibrium value. This type of nonlinear error correction mechanism (NECM) can not be detected by any linear nonstationarity tests (Sarno 2000).

Economic theory offers several possible explanations for the presence of nonlinearities in real exchange rates. Already, according to Lucas critique policy shifts may lead to breaks in the trend of the macroeconomic series and thus might generate long swings for the series. Regular random walk models do not take into account of any effects of observed policy changes. For example, Kaminsky (1993) shows theoretically that a change from a contractive monetary police to an expansive monetary policy increases the exchange rate depreciation. If we assume that prices are not as flexible as exchange rates the same reasoning applies also to the real exchange rate series. Moreover, heterogeneity of participants in the foreign exchange market is often cited as the major source of nonlinearities in the real exchange rate process (Sarantis 1999). According to Peters (1994) and Guillaume et al. (1995) investors' heterogeneity comes from different investment horizons, geographical location, various types of risk profiles and institutional constraints.

Dumas (1992) argued that the real exchange rate adjusts nonlinearly towards its PPP equilibrium value. He shows that the existence of transaction costs is an important explanation for the failure of the law of one price (LOP) to hold. If there are transaction costs, c_i, the price of good i in location j, p_j^i, may not be equal with its price in location k, p_k^i. The relative price will fluctuate in a range $1/c_i \le p_j^i / p_k^i \le c_i$. Nonlinearity arises because transaction costs, c_i, make an arbitrage nonprofitable in response to small deviations from the LOP. This implies that small deviations do not revert to mean, while sufficiently large price differential lead to arbitrage and thus to mean reversion.

2.2.1 The Markov switching model for the real exchange rate series

The Markov switching regime model differs from the models with exogenously imposed breaks in that the timing of the breaks is entirely stochastic. In the Markov models, the inferences are drawn on the basis of probabilistic estimates of the most likely state prevailing at each point of series time during the observation period. The major advantage of the Markov switching regime model is the flexibility in modelling time series with respect to regime shifts.

The basic idea in the two-state Markov regime switching model for the real exchange rates is that the real exchange rates can be divided to an appreciation or a depreciation regime and a random process governs the switches between the regimes (Klaassen 1999). Engel and Hamilton (1990) modelled such long swings in the exchange rate data and suggested a Markov switching random walk model with drift for exchange rate series. They found that mean squared errors from the Markov switching regime model were lower than the mean squared errors for the single regime random walk model. Engel and Kim (1999) examine the behaviour of UK/USA real exchange rate under the assumption that the real exchange rate series is integrated of order one. They suggest that the deviation from the permanent component can be modelled as a three state Markov switching regime model. The states can be characterized in their model by low, medium and high variance.

In the Markov regime switching process the probability of being in a particular state is only dependent in which state the process was in the previous period. The regime-switching model is originally based on Hamilton's (1989,1990) model for exchange rates. We use the model of Bergman and Hanson (2005), where the size of the autoregressive parameter is also estimated and do not presume it to be unity as in Hamilton exchange rate models. We use only one autoregression term, because it is generally believed that the short-run autocorrelations in exchange rates are small (see for example West and Cho 1995). The restrictions for the observed variables are ergodicity of the regime process and weak stationarity of the real exchange rate variables. We assume also that the probabilities of switching from one regime to the other are constant over time. The model has two elements. The first element is the regime process, which is based on two unobservable regime paths s_1 and s_2 and the second is the mean equation. Within these two regimes, the mean real exchange rate change is μ_1 or μ_2, which are assumed to be constant over time. If the mean regimes are different then persistence of these two regimes leads to long swings.

Bergman and Hanson estimated a model for the real exchange rate (q_t) that is generated by the following Markov regime-switching model:

$$q_t = \mu_{s_t} + \alpha_{s_t} q_{t-1} + \varepsilon_t \,, \tag{2.2}$$

where ε_t is $N(0,\ \sigma^2_{s_t})$ distributed and the initial value q_0 is fixed. In this formulation, the changes in the log of the real exchange rates are normally distributed with mean μ_i and variance σ_i^2 in each of the two possible regime states of the world (i = 1,2). Thus when s_t = 1 the change in the real exchange rate is μ_1, and when s_t = 2 the change is μ_2. The unobserved random regime state variable s_t is independent of the past q_t conditional on s_{t-1}. Real depreciation and real appreciation are modelled as switching regimes of the stochastic process generating the growth rate of the real exchange rate series. The regimes are associated with different conditional distributions of the growth rate, and mean is assumed to be positive in first regime and negative in the second regime. We assume that regime path (s_{t-1}, s_{t-2},…) follows a first order Markov process with time homogeneous transition probabilities,

$$P\{s_t = j \mid s_{t-1} = i\} = p_{ij} \ , \tag{2.3}$$

for i and j =1,2. This process is completely described by the following constant transition probabilities:

$$\begin{aligned} &P(s_t = 1 \mid s_{t-1} = 1) = p_{11} \\ &P(s_t = 2 \mid s_{t-1} = 1) = p_{21} = 1 - p_{11} \\ &P(s_t = 2 \mid s_{t-1} = 2) = p_{22} \\ &P(s_t = 1 \mid s_{t-1} = 2) = p_{12} = 1\text{-}p_{22}. \end{aligned} \tag{2.4}$$

Thus p_{ij} is equal to the probability that the Markov chain model moves from state i at time t-1 to state j at time t. The regime path process s_t depends on past realizations of the real exchange rate series q_t and regime path s_t only through previous value of regime path s_{t-1}. It should be noted that we do not assume that the real exchange rate series follow long swing process. The regime persistence can also be asymmetric. The regime means μ_1 and μ_2 should be opposite in signs and the values for both regime staying probabilities p_{11} and p_{22} should be large under the hypothesis of long swing. The unconditional probabilities of the stationary distributions that the process is in each of the regimes are given by

$$P(s_t = 1) = \frac{1 - p_{22}}{2 - p_{11} - p_{22}} \text{, and} \tag{2.5}$$

$$P(s_t = 2) = \frac{1 - p_{11}}{2 - p_{11} - p_{22}} \tag{2.6}$$

see Hamilton (1994) for derivation of this result.

Equations (2.5) and (2.6) are important for our Markov switching regime model, because they are related directly to the long swing models. To obtain estimates of the parameters and the transition probabilities governing the Markov chain of the unobserved state, we need an iterative estimation technique. Estimates of parameters for the two most likely regimes are calculated using maximum likelihood estimation techniques. The likelihood is calculated for each possible state, and the probability that a particular state is prevailing is obtained by dividing the likelihood of particular state by the total likelihood for both states. Hamilton (1990) used an expectations maximisation (EM) algorithm, which is useful for the case where all parameters change. The EM algorithm introduced by Dempster et al. (1977) is designed for a general class of models where the observed time series depend on some unobservable stochastic variable. For analysing and inferencing the Markov chain states we need to smooth the estimates so that it contains information from the whole sample. After smoothing, the estimate of probabilities is commonly explained so that a state is prevailing when the probability estimate for the particular state is greater than 0.5.

2.2.2 STAR-models

In the Engle-Granger (1987) and Johansen (1988) linear cointegration methodologies, the speed of adjustment to equilibrium is independent of the magnitude of disequilibrium. One way of modelling nonlinear adjustment involves application of the threshold autoregressive (TAR) model (Tong 1993). This would be appropriate for example if there is a threshold level of the absolute deviation from parity beyond which real exchange rate becomes mean reverting. The TAR model might be appropriate in modelling deviations from the law of one price for individual goods for which a given discrete threshold is relevant. We can think that the threshold value is the size of transaction costs in arbitraging the good internationally. According to Granger and Teräsvirta (1993) the nonlinear adjustment process can also be characterized in terms of a smooth transition autoregressive (STAR) model. In the STAR framework, the fixed thresholds of a standard threshold autoregressive (TAR) model are replaced with a smooth function, which need to be continuous and non-decreasing (Tong 1993). We think that a smooth transition (STR) model might be more appropriate than TAR model, when we are examining movements of the consumer price indices based on real exchange rates, which involves a range of goods with different costs of arbitrage.

Although the STAR models look very much like the Markov switching regime model, there is a crucial difference between these two models. In the threshold model, the regimes are defined by the past values of the time series itself. In the Markov switching regime models, the model regimes are defined by the exogenous state of the Markov chain (Potter 1999). Smooth transition regression models are also more preferable for post Bretton Woods data than for example unit root tests, since the modelling strategy does not require very long time series (Teräsvirta 1994). The STAR models have interesting properties. First of all the STAR models do not assume a sharp switch from one regime to the other like the TAR or the Hamilton’s Markov switching regime models. In foreign exchange markets with a large number of investors, each switching at different times due to for example different heterogeneous beliefs, a smooth transition seems to be more appropriate. Even if economic agents make only dichotomous decisions, it is unlikely that they change their behaviour simultaneously. Hence, as Teräsvirta (1994) notes, for aggregated processes the change in regime may be smooth rather than discrete. Another feature of the STAR models is that they nest linear regression model, and we can thus use linear Lagrange multiplier (LM) tests for testing the null of linearity before fitting any nonlinear model (Teräsvirta 1994). We can also use LM tests for choosing between the alternative STAR specifications. The need for symmetry in the response to positive and negative deviations from PPP leads to the exponential STAR (ESTAR) model described by Teräsvirta (1994). The ESTAR model can be viewed as a generalization of the double threshold TAR model. It is particularly attractive in the PPP context, as the strength of equilibrating force increases when the degree of absolute disequilibrium grows.

Following Michaet et al. (1997), the STAR model for the real exchange rate series (q_t) can be written as

$$q_t = \alpha + \sum_{i=1}^{p} \pi_i q_{t-i} + (\alpha^* + \sum_{i=1}^{p} \pi_i^* q_{t-i}) F(q_{t-d}) + \varepsilon_t \tag{2.7}$$

where α and α^* are regime constants, q_t is assumed to be stationary and ergodic process, ε_t is independently, identically and normally distributed with mean zero and constant variance, d is the delay parameter, and F is a transition function which is bounded by zero and one. In these models, nonlinearities arise through conditioning on lagged real exchange rates. The adjustment takes place in every period, but the speed varies with the extent of the deviation from parity. In contrast with TAR model, the regime of the STAR model changes gradually.

A most commonly used specification for transition function is

$$F(q_{t-d}) = 1/\{1 + \exp[-\gamma(q_{t-d} - c^*)]\}, \gamma > 0. \tag{2.8}$$

This is the logistic function and, c^* is a threshold value indicating the half way point between two regimes. The gamma (γ) parameter measures the speed of transition from one regime to the other. It determines the smoothness of the change from one regime to the other. The logistic function (2.8) is not the most plausible for the modelling of the real exchange rate series, since it is not symmetric for positive and negative deviations from PPP. The need of symmetry leads to exponential STAR (ESTAR) model described by Teräsvirta (1994). The transition function in the ESTAR model is

$$F(q_{t-d}) = \left\{1 - \exp\left[-\gamma(q_{t-d} - c^*)^2\right]\right\}, \gamma > 0. \tag{2.9}$$

The ESTAR model suggests that the two regimes have rather similar dynamics, while the transition period can have different dynamics. The transition function is U-shaped, and symmetric around c^* in the sense that the local dynamics are the same for high and for low values of the real exchange rate series (q_t). It should be noted that for the globally stable model the middle regime does not necessarily have to be locally stable.

In the ESTAR model, the inner regime corresponds to $q_{t-d} = c^*$ when $F = 0$ and thus the model (2.7) becomes a linear $AR(p)$ model

$$q_t = \alpha + \sum_{i=1}^{p} \pi_i q_{t-i} + u_t . \tag{2.10}$$

The outer regime corresponds to $q_{t-d} = \pm \infty$ when $F = 1$ and (2.7) becomes a different type of $AR(p)$ model

$$q_t = \alpha + \alpha^* + \sum_{i=1}^{p} (\pi_i + \pi_i^*) q_{t-i} + u_t . \tag{2.11}$$

We prefer the choice of the ESTAR model, since adjustment to PPP deviations can be expected to be same for both positive and negative deviations from equilibrium. The ESTAR model can also be viewed as a generalization of a particular form of the two thresholds TAR model. A symmetric TAR model can be specified in which the outer regimes are the same. Such a model is the limiting case of (2.7) when $\gamma \to \infty$.

The ESTAR model (2.7) can be reparameterized as follows

$$\Delta q_t = k + \lambda q_{t-1} + \sum_{i=1}^{p-1} \phi_i \Delta q_{t-i} + (k^* + \lambda^* q_{t-1} + \sum_{i=1}^{p-1} \phi_i^* \Delta q_{t-i}) F(q_{t-d}) + u_t \, , \tag{2.12}$$

where $\phi_i = 1 - \pi_i$ and $\phi_i^* = 1 - \pi_i^*$. The most important parameters of (2.12) are λ and λ^*. Our previous discussion of nonlinearity in the real exchange rates suggests that the larger the deviation from PPP, the stronger the tendency to move back to equilibrium. This implies that while $\lambda \geq 0$ is possible, we must have $\lambda^* < 0$ and $\lambda + \lambda^* < 0$ for the model to be globally stable. That is, for small deviations real exchange rate series may follow a unit root or even explosive behaviour, but for large deviations the process is mean reverting. Thus, for small deviations from equilibrium value, the real exchange rate process may adjust very slowly or not at all, but for large deviations, the process adjusts very quickly towards equilibrium level. The model may be viewed as a nonlinear error correction model in the form of a smooth transition autoregressive process.

In the linear error correction model (2.12) the error correction coefficient l must be significantly negative for the stable model. The conventional linear cointegration tests of PPP are based on a linear AR(p) model, such as

$$\Delta q_t = k' + \lambda' q_{t-1} + \sum_{i=1}^{p-1} \phi_i' \Delta q_{t-i} + u_t \, . \tag{2.13}$$

If the true real exchange rate series is not linear, then the parameter λ' in regression (2.13) will be between λ and $\lambda + \lambda^*$. Hence the null hypothesis of no cointegration may not be rejected against the stationary alternative, even though the true process is globally stable. Thus, the failure to find linear cointegration relationship does not necessarily invalidate the long-run PPP hypothesis.

2.2.3 Linearity tests for STAR-models

It is reasonable to execute linearity tests if we assume that the real exchange rate series might be nonlinear. In testing linearity, we follow the suggestion by Granger and Teräsvirta (1993) and Teräsvirta (1994), and estimate an auxiliary regression. The first step is to specify the linear part of the model. The linear part of the real exchange rate series (q_t) can be presented as

$$q_t = \beta_0 + \sum_{i=1}^{p} \beta_i q_{t-i} + \varepsilon_t = \beta_0 + \beta' w_t + \varepsilon_t , \tag{2.14}$$

where $\beta = (\beta_1, \ldots, \beta_p)'$, $w_t = (q_{t-1}, \ldots, q_{t-p})'$ and ε_t is a white noise residual term. The order of the autoregressive part (p) is usually determined by the Akaike information criteria (AIC). In the second step, the null hypothesis of linearity is tested against a smooth transition autoregressive alternative. We use the auxiliary regression

$$q_t = \beta_0 + \beta' w_t + \delta_1 ' z_t^1 + \delta_2 ' z_t^2 + \delta_3 ' z_t^3 + u_t , \tag{2.15}$$

where $z_t^k = w_t q_{t-d}^k = (q_{t-1} q_{t-d}^k, \ldots, q_{t-d}^{k+1}, \ldots, q_{t-p} q_{t-d}^k)'$, and k = 1,2, and 3. The transition variable is chosen from the linearity test H_0: $\delta_1 = \delta_2 = \delta_3 = 0$. The null hypothesis may be tested by the LM test (for details see Teräsvirta 1994). It should be noted that when we are using the LM test, we do not need to estimate the model under alternative specification. The order of autoregression is chosen on the basis of serial correlation tests on the residual vectors from alternative autoregressive representations. As Teräsvirta (1994) points out, neglected autocorrelation structure may lead to false rejections of the linearity hypothesis in favour of the presence of nonlinearities. He also warns against the use of automatic selection criteria for choosing the autoregressive lag order without testing for residual autocorrelation. To specify the value of the delay parameter d, the estimation of (2.15) is carried out for a wide range of values, $1 \leq d \leq D$. In cases where linearity is rejected for more than one value of d, d is set equal to value, which minimizes the p-value of the linearity test[14].

When choosing between LSTAR and ESTAR models for those real exchange rates where linearity is rejected we can use the following sequence of nested tests (Sarantis 1999)

$$H_{01}: \delta_3 = 0 \tag{2.16}$$

$$H_{02}: \delta_2 = 0 \mid \delta_3 = 0 \tag{2.17}$$

$$H_{03}: \delta_1 = 0 \mid \delta_2 = \delta_3 = 0, \tag{2.18}$$

If we reject the hypothesis (2.16), we select the LSTAR model. If we accept (2.16) but reject (2.17), we have evidence to choose the ESTAR model specification. If we accept (2.16) and (2.17) but reject (2.18), we choose the LSTAR model again. However, Granger and Teräsvirta (1993) and Teräsvirta (1994) argue that this sequence of tests application may lead to wrong conclusions, if the higher order terms of the Taylor expansion used in deriving these tests are disregarded. They recommend that we should compute the p-values for all F-tests of (2.16) to (2.18) and make the choice of the STAR model on the basis of the lowest p-value.

[14] It should be noted that no test can detect nonlinearity and the tests can only suggest the form of the nonlinearity. Furthermore, a problem the LM test is that parameters must be identified under the null hypothesis of linearity.

2.3 Empirical analysis

The empirical analysis is based on quarterly observations of average spot exchange rates and consumer price indices for the United States, the United Kingdom, Germany, Finland, and Sweden. The data period ranges from the first quarter of 1974 to the last quarter of 1998.

2.3.1 The Markov switching regime model

We specify a univariate two regime AR(1) Markov switching model for all the nine real exchange rate changes. The Markov switching regime model that we consider is based on Hamilton (1989). We assume that the real exchange rate series is allowed to follow a similar autoregressive process in both regimes. Thus, we think that the data generation process for the real exchange rate series is similar in both regimes, but differs in variances and/or regime staying probabilities. We do not think that this is an unreasonable assumption for the real exchange rate series. We use the Markov switching regime model to examine whether the real exchange rate series include long swings. The parameter vector $\theta = (\mu_1, \mu_2, \sigma_1, \sigma_2, p_{11}, p_{22})$ is estimated independently for each series. The difference between our model and the model used in Engel and Hamilton (1990) and in Engel (1994) is that we allow the autoregressive parameter to differ from unity. The parameter vector is estimated by maximum likelihood method and the sample likelihood is a function of the observed values of the changes in the logs of real exchange rates. The maximum likelihood estimation is performed by using the EM algorithm. The EM algorithm is particularly useful for real exchange rate series since all parameters are allowed to switch. The states s_1 and s_2 are unobserved, and we must make inferences about the probability of the state based on the observed data.

Our maximum likelihood estimation results are reported in Table 1. The results show that almost all real exchange rate series are well characterized by an appreciation and a depreciation state. The estimates show that the states differ not only in means but also in variances. The estimates suggest that the first regime (positive) can be considered as a rising real exchange rate (real depreciation) and the second regime (negative) as a declining real exchange rates (real appreciation). For all country pairs the estimated means are positive in the first regime and negative in the second regime. We can see that real exchange rates have depreciated more sharply than appreciated during the sample period. Also, the DEM has relatively lower mean values in the appreciation regime than other currencies. The UK has also very high mean depreciation value during the sample period. This is probably due to the UK peak devaluation in 1992. It is interesting to see that for the FIN/SWE case we cannot detect significant mean regimes. This is probably due to these countries similar exchange rate policy during the sample period. The test results show that there is more variability in the real exchange rate in the depreciation regime. The probability of staying in one state is large for all country pairs. The point estimates of p_{11} range from 0.553 to 0.913, while the estimates of p_{22} range from 0.809 to 0.937. These estimated probabilities show that if the system is in either state one or two, it is likely to stay in that state.

Table 1. Markov switching regime model estimates for real exchange rates

	UK/US	GE/US	FI/US	SW/US	UK/GE	FI/GE	SW/GE	UK/FI	SW/FI
μ_1	4.63	3.06	2.59	3.46	4.73	3.65	1.79	3.07	0.17
	(1.60)	(1.31)	(0.85)	(1.25)	(2.38)	(0.92)	(0.83)	(1.04)	(0.21)
μ_2	-2.05	-3.28	-3.13	-2.84	-1.21	-1.15	-1.44	-1.34	-0.39
	(0.70)	(0.76)	(0.53)	(0.42)	(0.54)	(0.25)	(0.29)	(0.59)	(0.53)
σ_1^2	21.70	17.67	20.65	24.85	18.56	10.17	18.56	10.11	1.11
	(7.99)	(6.46)	(4.25)	(7.31)	(8.36)	(3.44)	(4.53)	(3.31)	(0.41)
σ_2^2	13.17	10.24	6.39	6.13	10.52	3.52	2.62	9.09	10.22
	(3.405)	(2.86)	(1.71)	(1.79)	(2.10)	(0.63)	(0.67)	(1.93)	(3.46)
p_{11}	0.74	0.85	0.87	0.83	0.55	0.79	0.84	0.77	0.91
	(0.14)	(0.11)	(0.07)	(0.08)	(0.20)	(0.10)	(0.08)	(0.11)	(0.06)
p_{22}	0.88	0.81	0.84	0.85	0.89	0.94	0.89	0.91	0.89
	(0.08)	(0.09)	(0.07)	(0.06)	(0.08)	(0.03)	(0.06)	(0.07)	(0.10)
AR	0.921	0.95	0.94	0.94	0.93	0.93	0.86	0.88	0.92
	(0.039)	(0.03)	(0.04)	(0.03)	(0.04)	(0.04)	(0.05)	(0.05)	(0.04)

Note: Standard errors are in parentheses.

In Table 2 we test first the hypothesis H_0: $\mu_1 = \mu_2$ using Wald test. Under this null hypothesis, the two states differ only by their variances. If the null hypothesis is true the real exchange rates follow a random walk model with heteroscedastic errors. The null hypothesis can be rejected for all the real exchange rate series except the SWE/FIN series at 5 per cent significance level. The failure to model SWE/FIN real exchange rate dynamics by Markov switching approach can be explained by their macroeconomics similarities with respect to each other. Both countries display strong interdependence regarding monetary policy during the past 25 years. When Finland or Sweden changed their exchange rates through competitive devaluations, the other country was also forced to change its' exchange rate policy. Thus, the exchange rate changes had quite similar effects for both countries in the long run.

The test problem under the above null hypothesis is that if one regime governs the real exchange rate series data, the parameters for the second regime are not identified. This makes the asymptotic distribution of the Wald test not χ^2 as Hansen (1992) shows. Garcia (1998) tries to solve this problem by deriving the correct asymptotic distribution of the likelihood ratio statistic. We circumvent this problem as Hamilton (1989) by expecting that all parameters are identified. Indeed, our estimation algorithm (EM) results show in Table 1 that all parameters are identified, and we can circumvent the singularity problem of the information matrix. However, we should note that for the case of FIN/SWE both regimes are insignificant and thus we have no identification problem.

The second hypothesis is that H_0: $p_{11} + p_{22} = 1$. Under this null hypothesis the distribution for the regime path process s_t is independent of s_{t-1}. According to Engel and Hamilton (1989) large values of p_{11} and p_{22} indicate long swings in the series. Table 2 shows that the null hypothesis is strongly rejected for all country pairs at the 5 per cent level. We conclude that movements in the real exchange rate series are well described by long swings and when the real exchange rate series enters the state one or two it stays there for years. The expected duration of the state i can be calculated by $1/(1-p_{ii})$ (Hamilton 1989).

On average, state one (i.e. real depreciation) lasts from 1/2 to 3 years, and state two (i.e. real appreciation) lasts from 1 year to 4 years. Thus, the estimated probabilities show that the stay in the real depreciation regime has been shorter than the stay in the real appreciation regime.

Table 2. The Wald tests of the null hypothesis for stochastic process and independency of the regime path

Series	H_0: $\mu_1 = \mu_2$	H_0: $p_1 + p_2 = 1$
UK/US	19.96	24.98
GER/USA	32.62	39.02
FIN/USA	32.02	10.09
SWE/USA	23.19	55.51
UK/GER	7.36	4.16
FIN/GER	28.82	54.20
SWE/GER	12.96	79.33
UK/FIN	20.71	23.92
SWE/FIN	0.87	69.17

Notes: A Wald statistic for testing H_0: $\mu_1 = \mu_2$ is given by

$\frac{(\hat{\mu}_1 - \hat{\mu}_2)^2}{Var(\hat{\mu}_1) + Var(\hat{\mu}_2) - 2Cov(\hat{v}_1, \hat{\mu}_2)} \approx \chi^2(1)$, and a Wald test for H_0: $p_1 + p_2 = 1$ is given by

$\frac{[\ (p_{11} - (1 - p_{22})\ ]^2}{Var(\hat{p}_{11}) + Var(\hat{p}_{22}) - 2Cov(\hat{p}_{11}, \hat{p}_{22})} \approx \chi^2(1)$. The 5 per cent critical value is 3.84.

Another way to examine persistence of regimes is by inspecting estimated regime probabilities. We use smoothed regime probabilities, which utilize the complete data set and thereby smoothing the *ex ante* probabilities (i.e. the conditional probability that the process is in a particular regime at time *t* using only information available at time *t*-1). Thus, smoothed probabilities give the most informative answer to the question in which regime the variable is most likely to be at time *t*. The smoothed probabilities with the real exchange rate series are shown in Figure 1. The shaded areas show the smoothed probabilities when the probability of staying in regime one is above 0.5 representing the real depreciation regime. The real appreciation regime is not shaded. The probabilities show that regime classification is as one might expect. As Figure 1 shows, the smoothed probabilities identify correctly the movements of the real exchange rates. For example, the appreciation of the US dollar against the European currencies in the first half of 1980s and depreciation after that is well captured by the smoothed regime switch probabilities.

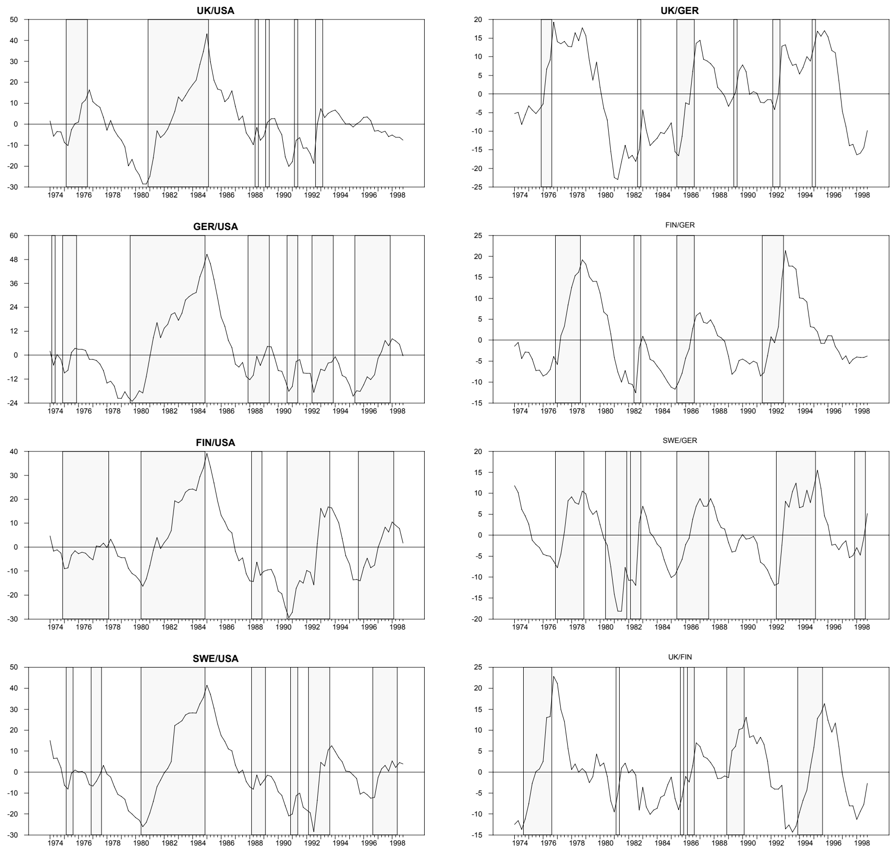

Fig. 1. Smoothed probabilities of being in regime 1.

The Markov switching regime estimation results are consistent with Dumas' theoretical model, in which the real exchange rates exhibit long swings. Moreover, the results are in line with Mussa (1986) observation, that the real exchange rates are not "nominal exchange rate neutral", since the real exchange rates states appears to shift at points when there are also significant nominal exchange rate events.

2.3.2 Linear models against ESTAR models

To test for real exchange rate linearity against the ESTAR specification we use similar auxiliary regression as Baum et al. (2001) for second order, that is

$$q_t = \beta_{00} + \sum_{j=1}^{k} (\beta_{oj} q_{t-j} + \beta_{1j} q_{t-j} q_{t-d} + \beta_{2j} q_{t-j} q_{t-d}^2) + \varepsilon_t , \tag{2.19}$$

where q_t is the real exchange rate series. The rejection of the null hypothesis (i.e. linearity) H_0: $\beta_{1j} = \beta_{2j} = 0$ for all j suggesting the use of ESTAR models. In carrying out linearity tests we have considered values for the delay parameter (d) over the range $1 \le d \le 16$, and calculated the p-values for the linearity test in each case. The autoregression length parameter (k) for the real exchange rate series is chosen on the basis of the AIC criterion. According to Michael et al. (1997) overspecification of the autoregressive model is not as serious a problem as underspesification since autocorrelated errors may affect the linearity tests. With the proper choice of lag length (k), we vary the delay length (d) in order to provide the strongest probability for nonlinearity. If linearity is rejected for more than one value of d, then the estimate of d is chosen by the lowest p-value of the above linearity test. Thus, we will choose the delay length (d) for which the nonlinearity is strongest according to our test procedure. If nonlinearities can be detected, they imply that standard linear unit-root and cointegration tests might fail to find any significant mean reversion in real exchange rate series.

The linearity test results are shown in Table 3. We use ordinary F-test as an approximation to the LM test since it is generally known to have relatively good size and power properties in finite samples. The test results show that for seven cases out of nine the test classifies the time series as nonlinear ESTAR model at the 5 per cent level of significance. We can therefore proceed to build nonlinear models for these exchange rates. Linearity test is not rejected for country pairs UK/GER and UK/FIN[15].

Table 3. Linearity tests results for real exchange rate series

Countries	AIC based Lag (k)	Delay parameter (d)	p-value
UK/US	8	3	0.033*
GER/US	2	5	0.031*
FIN/US	4	8	0.021**
SWE/US	8	9	0.031*
UK/GER	2	1	0.255
FIN/GER	5	3	0.00019***
SWE/GER	5	3	0.000009***
UK/FIN	4	9	0.284
SWE/FIN	2	10	0.039*

Notes: The F statistic tests the null hypothesis of linearity against the alternative of nonlinearity; d denotes the delay parameter. Marginal significance levels are calculated using the appropriate F-distribution. The test is the minimum p-value over the interval $1 \le d \le 16$. The selection of the maximum lag, k, of the linear AR model is made using the AIC statistic. The *,**,and *** are 5%, 2.5%, and 1% significance levels, respectively.

[15] One reason for this might be the Louvre accord in 1987, which aimed to stabilize exchange rates by introducing target zones in the UK exchange rates. This implies that swings in the UK exchange rates are probably short-lived, and nonlinearity test is not able to detect any nonlinearities after 1987.

2.3.3 The ESTAR specification models

Assuming that linearity is rejected, we can proceed to estimate the ESTAR models for the real exchange rate series. We consider the possibility that the first differences of the logarithmic real exchange rate series are nonlinear. We assume further that if the real exchange rates are nonlinear, we can adequately characterize nonlinearity by the ESTAR models. We estimate an ESTAR model by nonlinear least squares estimation technique, which provides estimators that are consistent and asymptotically normal. Tong (1993) shows that these conditions also hold for STAR models if series is stationary, ergodic and the error terms are independently and identically distributed. In each case we follow the recommendation of Granger and Teräsvirta (1993) and standardize the transition parameter (γ) by the sample variance of dependent variable and using starting value $\gamma = 1$ for the estimation algorithm.

In Table 4, we present the LR test results for the following restrictive hypothesis for the model (2.12)

$$H_0^A : k = k^* = c^* = 0 \tag{2.20}$$

$$H_0^B : 1 + \lambda = -\lambda^* \text{ and } \phi_j = -\phi_j^*\text{, given } H_0^A \tag{2.21}$$

$$H_0^C : \lambda = 0\text{, given } H_0^A \text{ and } H_0^B. \tag{2.22}$$

If the restrictions are true, they are imposed to obtain a more parsimonious model. This is an important concern in the estimation of the STAR models (Granger and Teräsvirta 1993). The ESTAR model specification simplifies considerably if all three restrictions are valid. If we reject one of the restrictions then they are not imposed in the ESTAR models. The first null hypothesis (H_0^A) implies that both constant terms (k and k^*) are in both regimes zero, and that the equilibrium value of the real exchange rate series (q_t) in the model (2.12) is also zero. The second hypothesis (H_0^B) implies that the real exchange rate series (q_t) is in the outer regime (i.e. when $F = 1$) a white noise process. Thus, conditional on the first restriction, the second restriction implies, that when deviation from the PPP is large, the real exchange rate process will be mean reverting. The third hypothesis (H_0^C) implies that in the middle regime (i.e. when $F = 0$) the real exchange rate series (q_t) has a unit root. Thus, conditional on the first and the second restrictions the failure to reject the third hypothesis implies unit root behaviour for small deviations from the PPP.

Results from Table 4 indicate that all three restrictions are relevant for three bivariate real exchange rate series, namely FIN/USA, SWE/GER, and SWE/FIN. For five cases out of seven the first hypothesis (H_0^A) is valid. Only for cases GER/USA and FIN/GER we cannot reject the first restriction (2.20) implying that PPP equilibrium value is not zero. However, the constant terms (k and k^*) may be expected to be zero, because the real exchange rate series is estimated as the residuals of the cointegrating regression. The ESTAR model for these cases is estimated without any restrictions (2.20), (2.21), and (2.22). For two cases, namely UK/USA and SWE/USA real exchange rate series, the H_0^B is not rejected implying that the real exchange rate series is not white noise in the outer regime. However, the stability condition ($\lambda + \lambda^* < 0$) is satisfied at a 1 per cent significance level.

Table 4. The LR test results for the restriction hypothesis

Statistics	UK/US	GE/US	FI/US	SW/US	FI/GE	SW/GE	SW/FI
LR1	1.1507	12.25^c	0.002	1.651	6.672^c	0.237	0.007
LR2	225.1^c		0.000	303.1^c		0.003	0.005
LR3			0.120			0.359	0.362
λ	0.103 (0.201)			0.033 (0.072)			
λ^*	-0.282 (0.197)			-0.215 (0.080)			
F-test for $\lambda+\lambda^* = 0$	10.520^c			13.146^c			
Concl.	ESTAR	ESTAR with constant	ESTAR	ESTAR	ESTAR with constant	ESTAR	ESTAR

Notes: LR_1, LR_2 and LR_3 are likelihood ratio test statistics corresponding tests of (2.20), (2.21), and (2.22). The stability hypothesis H_0: $\lambda_1+\lambda_2 = 0$ is tested using F-test, with degrees of freedom 1 and 99-k, where k is the lag length. (c) denotes significance at 1 % level.

Table 5 presents estimated parameter values and residual diagnostics of ESTAR models for seven country pairs. The ESTAR models are estimated for those real exchange rate series, which nonlinear behaviour was found by the test results in Table 3 at the 5 per cent significance level. The results show that the transition parameter (γ) varies across country pairs. The interpretation of the *t*-statistics for transition parameter (γ) in Table 5 should be done carefully, since under the null hypothesis of model (2.12) the real exchange rate series follow a random walk process. Hence, we cannot assume that the "*t*-test" statistics are normal (Taylor and Peel 1998). However, our *t*-ratio test results in Table 5 are sufficiently large according to Dickey-Fuller critical values to find significant transition parameters for five cases out of seven. For GER/USA we have not significant *t*-value for the g parameter according to Dickey-Fuller table, but we have reported the test results because the *t*-test value differs markedly from those that are insignificant. Thus, only for the cases UK/US and SWE/US we find that the transition parameter cannot be distinguished from zero. The residual diagnostic tests are satisfactory for most of the cases. There does not seem to be any autocorrelations or any *ARCH*-effects in any of the residuals. Only the normality test (*Jarque-Bera*) shows slight non-normality in some cases. This is probably due to some countries separate devaluations under the fixed exchange rate regimes. The variance ratio (V) between nonlinear (NL) and linear (L) model variances shows a reduction in the residual variance in all cases for the nonlinear model specification.

Table 5. Estimates of the ESTAR model for deviations from PPP

	UK/US	GE/US	FI/US	SW/US	FI/GE	SW/GE	SW/FI
d	3	5	8	9	3	3	10
γ	1.30	0.01	0.01	0.34	0.01	0.05	0.02
t-value	(0.54)	(2.29)	(3.37)	(0.45)	(3.25)	(3.62)	(3.04)
ϕ_1	0.13	0.30	0.27	<0.01	0.26	0.25	0.42
s.e.	(0.06)	(<0.01)	(0.09)	(0.06)	(0.09)	(0.07)	(0.09)
DW	1.91	1.92	1.94	1.97	1.95	1.77	1.93
Q(1)	0.03	0.01	0.02	<0.01	0.03	1.03	0.21
Q(4)	0.65	1.80	0.59	0.39	0.43	1.43	1.18
ARCH(1)	0.86	0.02	3.75*	0.40	1.52	2.85*	2.66
ARCH(4)	1.50	0.92	4.58	1.37	3.50	9.06*	7.00
JB	4.10	1.99	8.77*	29.13***	21.62***	5.46*	1.97
SSR	0.05	0.04	0.05	0.04	0.03	0.03	0.03
$V=s^2_{NL}/s^2_L$	0.97	0.96	0.96	0.94	0.99	0.97	0.94
c		0.13			0.06		
s.e.		(0.04)			(0.02)		

Notes: Standard errors and t-test values are given in parentheses. Q_1 and Q_4 are the Box-Ljung *Q*- statistics for the absence of serial correlation up to first and fourth order, respectively. *ARCH(1)* and *ARCH(4)* are tests for the *ARCH* effects up to first and fourth order, respectively. Both Q_n and $ARCH_n$ tests have a χ^2 distribution with n degree of freedom. SSR is the standard error of regression and *DW* is Durbin-Watson autocorrelation statistics. *JB* is the Jarque-Bera normality test. (***),(**), and (*) denote significance at 1%, 5%, and 10% levels, respectively.

In Figure 2 we have plotted the significant ESTAR transition functions $F(y_{t-d})$ against lagged CPI based PPP deviations series. The horizontal axis present real exchange rate deviations from the threshold levels c^*, and the vertical axis presents the correspondent value of transition function. In all cases the figures show a reasonable number of observations for both sides of c^*, indicating ESTAR type of nonlinearity. The figures show also much more observations for smaller real exchange rate values, which indicates very slow, albeit significant reversion towards the long-run PPP. The amplitudes of transition functions are much smaller than for example in the study of Michael et al. (1997). This might reflect the fact that deviations from PPP are much more persistent for post Bretton Woods period than for two century period. Similar conclusions have also drawn by Baum et al. (2001). Furthermore, the slow speeds of adjustment are consistent with the difficulties to reject the unit root hypothesis in real exchange rate series.

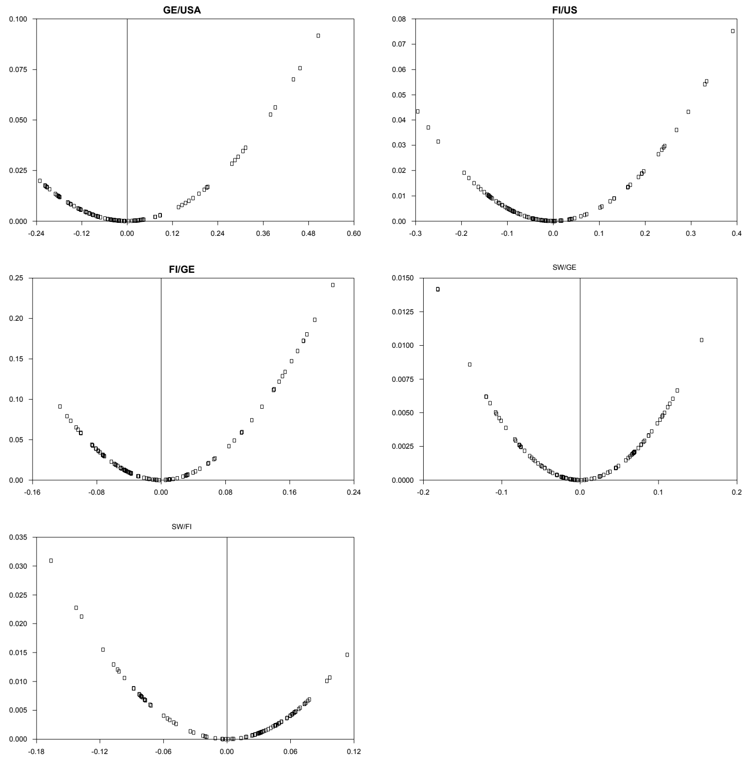

Fig. 2. Estimated transition functions

The ESTAR estimation results are in line with the theoretical adjustment process described in Dumas (1992). In the middle regime, the deviations from the PPP equilibrium value might follow in some cases unit root process, but in the outer regime the deviations from parity are diminished. In all models, where the ESTAR transition function is significant, the real exchange rate series satisfy also the stability condition. The transition parameter estimates show that the speeds of adjustment back to PPP is rather slow for post Bretton Woods period.

2.4 Conclusions

A random walk model often characterizes the real exchange rate series. Dumas (1992) showed that the equilibrium model of real exchange rate determination in the presence of transaction costs implies a nonlinear adjustment towards the law of one price. We provided empirical evidence that, for a number of OECD countries, the real exchange rate is well characterized by a nonlinear but stationary process. We use two alternative nonlinear models to model the real exchange rate series. The difference between the autoregressive Markov switching regime model and the ESTAR model is that they differ in how they model the movement between regimes. For the Markov model the movements between regimes are unrelated to the past realizations of the process, and the estimate results are modelled unobserved Markov chain process. The ESTAR specification moves between regimes depending on the past realizations of the process.

The central question of the Markov switching regime models is whether long swings exist in the real exchange rate series. The results suggest the real exchange rates are reasonably well represented by two state Markov switching AR(1) regime models. The real exchange rate movements can be characterized by long swings of an appreciation and a depreciation regime for the post Bretton Woods period. The results also show that the 1980s can be characterized by a real depreciation regime for the US dollar based real exchange rate series with relatively high persistence, although the process is eventually mean reverting. The 1990s can be characterized as a real appreciation process with high persistence for the US dollar based real exchange rate series.

We evaluate the linear autoregressive model against the ESTAR alternative using nine different bivariate real exchange rate series. The ESTAR models are applied to the real exchange rate series for which the linearity hypothesis is rejected, and display local instability and global stability. Thus, the real exchange rate can be characterized by a unit root or even explosive process in the neighbourhood of its long-run equilibrium, but it adjusts faster when the size of the deviation from real exchange rate equilibrium value increases. The empirical evidence exhibits ESTAR type of nonlinearity for the seven series. The transition parameter estimates for the ESTAR models show that the speed of transition from one regime to the other is quite slow for most of the country pairs. The parameter estimates implies random walk behaviour for small deviations from PPP in some cases, but fast adjustment for large deviations from PPP.

Overall, we find strong evidence of nonlinearities in the dynamics of real exchange rates. These results may help to explain the previous empirical regularity that the PPP deviations follow a random walk in linear models.

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Appendix 1 Table

Table 6. Ordinary least squares estimates for the model (2.1)

Series	$\beta' = (\beta_0, \beta_1, \beta_2)$
UK/US	(1, -1.76, 2.78)
GER/USA	(1, -3.40, 2.33)
FIN/USA	(1, -2.49, 4.05)
SWE/USA	(1, -1.67, 2.68)
UK/GER	(1, 0.10, 0.88)
FIN/GER	(1, 0.02, 0.97)
SWE/GER	(1, -0.50, 0.48)
UK/FIN	(1, -4.94, 5.91)
SWE/FIN	(1, 4.89, 5.88)

3 Nonlinear equilibrium correction model and the euro/dollar exchange rate dynamics

3.1 Introduction

The purpose of this chapter is to examine empirical validity of a model based on purchasing power parity and uncovered interest rate parity fundamentals of exchange rate determination as an appropriate framework for analysing movements of the euro-dollar exchange rate during the period 1990-2002. We presuppose that new currency euro and its somewhat extraordinary behaviour against the US dollar in recent years need a deeper understanding of the factors driving its dynamics. Hence, the first question is whether the euro is driven by macroeconomic fundamentals and whether the dynamics of the euro/dollar exchange rate can be modelled by a structural model. The closest reference is La Cour and MacDonald (2000), which examines the variant of the monetary model of ECU-U.S. dollar exchange rate over the period 1982 to 1994. Their analysis reveals that the value of ECU was sensitive to the US monetary shocks. Furthermore, they find that the out-of-sample forecasting performance of their model can beat a simple random walk model.

There have been many explanations for the euro depreciation against the US dollar in 1999-2002. Some authors suggest, that the productivity differential between Europe and the United States has been a factor determining the euro-dollar exchange rate (see Alquist and Chinn 2002 and Bailey et al. 2001 among others). They suggest that an improvement of the US productivity increases the capital flows out of Europe to the United States. However, Schnatz et al. (2003) used four different measures of productivity and did not find any evidence that productivity differential could explain the weakness of the euro against the dollar in 2000-2001. Another explanation is based on the surge in the US equity markets since mid-1990s, which raised market capitalization relative to GDP to unprecedented level and boosted both consumption and investment (Meredith 2001). The anticipated future impact of positive demand shock caused the long-term interest rate move up sooner than the short-term interest rates, which according to uncovered interest rate parity leads to the appreciation of the dollar against the euro. However, this explanation applies only for the period until 1999, since thereafter the interest rate differentials have narrowed and the euro has remained weak.

We examine a variant of monetary model, where the exchange rate dynamics is assumed to depend on long-term government bond rate differentials between euro zone and the USA. Despite the failure of empirical exchange rate models, we argue that macroeconomic fundamentals can explain euro/dollar exchange rate movements reasonably well if we allow nonlinear dynamics in the equilibrium correction models. We think that dynamics based on a nonlinear model in analysing euro/dollar exchange rate movements makes sense since for example monetary policy interventions may take place only when a significant degree of undervaluation of euro might cause a risk to the inflation stabilisation policy, while the same consideration does not necessarily hold for the euro's overvaluation case[16].

We also examine the presence of cointegrating relationships among the variables of the exchange rate models based on fundamentals. At the beginning, we use the cointegrating VAR framework of Johansen (1988) and test whether fundamentals can explain exchange rate movements in a linear way. Although cointegration might suggests support for linear model we are not able to accept any structural model based on economic theory for the long-run relations. As a possible remedy for this problem, we will assume that there might be some nonlinear dynamics in the exchange rate movements. Furthermore, we argue that nonlinearity may be due to different adjustment speeds for smaller and larger long-term interest rate differentials. Next, we apply smooth transition (STR) econometric methods popularised by Granger and Teräsvirta (1993) and Teräsvirta (1994,1998) to capture nonlinearity, since they allow a varying adjustment speed depending on the extent of interest rate differential. Our study differs from most of the earlier literature on nonlinear exchange rate modelling at least in two important ways. First, we base our exchange rate equilibrium conditions on economic theory. Second, we are especially interested in testing for a nonlinear relationship between the exchange rate changes and PPP fundamentals and/or PPP fundamentals augmented by short-term interest rate differential between euro zone and the USA in the multivariate context, rather than nonlinear specification of serial dependence in the univariate exchange rate series.

We will show that the deviations from the nominal euro/dollar exchange rate equilibrium display nonlinear behaviour. The first model can be characterised by a simple parametric PPP model. The second model is PPP fundamental model augmented by short-run interest rate differences. One of the main ideas in the models is that the nonlinear equilibrium correction of the exchange rate towards long-run equilibrium is captured by long-term government bond differences. Our examinations reveal a very striking result that a nonlinear type of equilibrium correction depending on the long-term government bond spreads is relevant for euro/dollar exchange rate changes during the sample period 1990-2002.

We proceed as follows. The essential elements of the fundamentals based exchange rate models and a brief literature review on exchange rate nonlinearities are presented in section 3.2. In section 3.3 we present the multivariate nonlinear equilibrium correction methodology and our model based on it. In section 3.4 we present and discuss details of

[16] Also, the euro seems to be more vulnerable when GDP differentials are unfavourable (Corsetti and Pesenti 1999) while in the opposite case the positive factors may amplify their upward influence by reinforcing their cross effects. Furthermore, an extraordinarily positive stock market performance may temporarily cause asymmetries in exchange rate dynamics.

the data. In section 3.5 we present our empirical findings. Finally, section 3.6 concludes with a summary and conclusions.

3.2 Exchange rate fundamentals and nonlinearities

The purchasing power parity (PPP) hypothesis can be seen as a cornerstone of determination of equilibrium exchange rate level, since it is, for example, a central hypothesis of flexible price monetary approach to exchange rate determination. The PPP hypothesis implies that movements in the nominal exchange rate should be proportional to the ratio of national price levels or, in other words, that the real exchange rate should be constant[17] (e.g. Sarno and Taylor 2002). While very few economists agree that the PPP holds continuously, most believe that some variant of PPP can be seen as an anchor for the long-run behaviour of real exchange rates (Rogoff 1996). This is true for both the traditional international macroeconomic analysis and of the "new open economy macroeconomics" (Obstfeld and Rogoff 1995).

We define the real exchange rate series as the deviation from the strong form PPP condition:

$$q_t = p_t - p_t^* - e_t, \tag{3.1}$$

where p_t and p_t^* are the logarithms of the domestic and the foreign price levels, respectively. The nominal exchange rate e_t is the spot exchange rate in logarithms measured in domestic currency units. Theoretically, the PPP is a stationary steady-state relation and suggests a long-run equilibrium relationship between exchange rates and relative national price levels. Thus, the real exchange rate series (q_t) can be seen as a deviation of the nominal exchange rate from the PPP fundamental value. The real exchange rate series may have considerable short-run variations, but a necessary condition for PPP to hold is that the real exchange rate is stationary over time and not driven by any permanent shocks. However, it is a very common empirical finding that there exist large, persistent deviations from PPP during the post Bretton Woods period and it is still an unsettled question whether economic fundamentals are the driving forces for deviations in real exchange rate series.

Over the last thirty years, a large part of empirical research on modelling the purchasing power parity (PPP) hypothesis has concentrated mainly on using linear cointegration and equilibrium correction[18] techniques that allow joint modelling of the equilibrium level of PPP and the dynamic adjustment towards this equilibrium level. One of the main results in the literature is that even if we use a longer data spans or panel data estimation methods we have not been able to solve the PPP puzzle[19]. According to Rogoff (1996)

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[17] If the PPP were actually the only driving force behind the fluctuations in nominal exchange rate, then real exchange rate would have to be constant.

[18] In the recent literature several authors have started to use the term 'equilibrium correction' instead of using the more traditional 'error correction'.

[19] See for example the statement of "How one can reconcile the enormous short-term volatility of real exchange rates with the extremely slow rate at which shocks appear to damp out?" by Rogoff (1996).

there seems to be a consensus that the size of half-life of deviations from PPP is about 3 to 5 years. In recent years a widely accepted result in the literature has been that transaction costs may create a band of inaction in real exchange rate series where it is unprofitable to make arbitrage even if trading opportunities were potentially profitable (MacDonald 1999). Furthermore, a number of papers have shown that transportation and distribution costs are also a very significant explanation for the PPP deviations (see Burstein et al. 2005). Obstfeld and Taylor (1997) show that the adjustment to the long-run equilibrium for some major real exchange rate series is as low as 12 months.

The real exchange rate series equation (3.1) reveals the significant restrictiveness of PPP hypothesis as a measure of equilibrium exchange rate, since it ignores the influence of capital flows[20] (MacDonald 1999). A further assumption in the traditional exchange rate models is that the uncovered interest rate parity (UIP) holds continuously (i.e. the interest rate differential between domestic and foreign is just equal to the expected rate of depreciation of the domestic currency). That is, if home and foreign bonds are perfect substitutes, and international capital is fully mobile, the two bonds can only pay different interest rates if agents expect that there will be a compensating movement in exchange rates. International capital market equilibrium is given by uncovered interest rate parity (UIP), which states that

$$E_t(e_{t+k}|I_t) - e_t = i_t^k - i_t^k* \tag{3.2}$$

where E_t is the expectation operator given the available information at time t, e_{t+k} is the k-period exchange rate value, i_t^k and i_t^{k*} are domestic and foreign k-period interest rates, respectively. The UIP suggest that investors should expect to earn the same return on similar assets from any country.

The standard textbook view in the international finance is, that higher interest rates at the home country relative to the foreign country, indicate home currency depreciation. However, the UIP condition (3.2) predicts only that an increase in interest rates at home country relative to those in the foreign country is associated with an expected future depreciation of home currency. Thus, the UIP condition does not tell whether this future depreciation take effect by an immediate depreciation of home currency or by reduction in the expected future value of home currency or is the currency depreciation some combination of both effects (Mark and Moh 2001). We can see from the UIP equation (3.2) that interest rate differences and exchange rate changes are functionally related and hence, it is possible to consider also nonlinear relations between exchange rate and interest rate differential series. This possibility arises especially if the interest rates are used as a monetary policy tool and hence there might exist policy interventions only when the interest rate differential is large but not when it is small (Mark and Moh 2001).

An obvious empirical fact in recent literature is that the PPP and the UIP do not work very well by themselves (e.g. Engel 1999, 2000). It is claimed that volatile expectations or departures from rationality are likely to account for the failure of these two important building blocks of exchange rate models (Neely and Sarno 2002). However, the PPP condition from the goods market and the UIP condition from the capital market are con-

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[20] Furthermore, the traditional view of PPP relies on the assumption of the law of one price and thus it also assumes absence of international arbitrage.

nected through the expected exchange rate. Juselius and MacDonald (2001) find evidence supporting the interdependence of the PPP and UIP conditions by using Johansen cointegration analysis in the linear context. Their conclusion was that significant but very slow adjustment towards real exchange rate parity was compensated by short-term interest rate spreads. They further showed that the UIP is empirically satisfied if the real exchange rate is at the PPP level. An important empirical question is also whether fundamental variables reflect asymmetric adjustment speeds in goods and asset markets.

3.3 Modelling

3.3.1 Nonlinear vector equilibrium correction methodology

The nonlinear models are very useful in explaining economic time series, when the behaviour of economic variables depends on different states of the world. The most popular regime dependent nonlinear model over the last few years has been the smooth transition autoregressive (STAR) methodology (Granger and Teräsvirta 1993), where the transition between regimes occurs in a smooth way. However, in recent empirical literature only univariate STAR models have been mostly used in empirical analysis. The analysis of multivariate nonlinear models has not been very extensive. The general ideas on how to extent univariate STAR models to a vector autoregression (VAR) framework has been discussed in Granger and Swanson (1996), Weise (1999), van Dijk et al. (2002) and Rothman et al. (2001) among others.

Many economic variables are not only nonstationary individually but also linked by long-run equilibrium relationships, such that they tend to move together in the long run. The concept of linear cointegration together with equilibrium correction models were first introduced by Granger (1981) and Engle and Granger (1987). In the standard equilibrium correction models the adjustment towards the long-run equilibrium is linear in the sense that equilibrium correction is always present and of the same strength under all circumstances. In many cases it is important to model several time series jointly in such a way that it is possible to exploit linkages that exist between these variables. The interest in multivariate nonlinear modelling has started to develop only recently and it has been mainly application oriented. The relevant statistical theory has not been fully developed yet and it is a topic of much current research.

Linear vector equilibrium correction model (VECM) is a common and widely used methodology to examine individually nonstationary economic variables, which are linked together by long-run relationship. In the standard form VECM assumes that variables follow a linear adjustment process towards their long-run equilibrium. However, there seem to be relevant examples where also the nonlinear adjustment may take place. For example, policy interventions may take place only when the economic variables deviate significantly from equilibrium. Also arbitrageurs enter the market only if the deviations from the no-arbitrage fundamental equilibrium are sufficiently large to compensate their transaction costs. Thus, it would be worthwhile to consider nonlinear models in the VECM context.

Smooth transition vector equilibrium correction models (STVECM) are extensions of the STAR models. Following Rothman et al. (2001) we define a *k*-dimensional smooth transition vector error correction model (i.e. a k-dimensional analogue of the univariate 2-regime STAR model) as

$$\begin{aligned}\Delta y_t &= \left(\mu_1 + \alpha_1 z_{t-1} + \sum_{j=1}^{p-1}\Phi_{1,j}\Delta y_{t-j} + \sum_{j=0}^{p-1}\Phi_{2,j}\Delta x_{t-j}\right)\left(1 - G(s_t;\gamma,c)\right) \\ &+ (\mu_2 + \alpha_2 z_{t-1} + \sum_{j=1}^{p-1}\Phi_{3,j}\Delta y_{t-j} + \sum_{j=0}^{p-1}\Phi_{4,j}\Delta x_{t-j})G(s_t;\gamma,c) + \varepsilon_t\end{aligned} \tag{3.3}$$

where y_t is a (k x 1) vector of I(1) endogenous vector time series, x_t is an (m x 1) vector of I(1) exogenous variables, α_i (i=1,2) are (k x r) matrices, z_t is the (qxr)-matrix with $q = k + m$ of cointegrating vector(s) or equilibrium correction terms, and $\varepsilon_t = (\varepsilon_{1t},...\varepsilon_{kt})'$ is a k-dimensional vector white noise process with mean zero and positive definite covariance matrix Σ. The transition function $G(s_t;\gamma,c)$ is assumed to be continuous and bounded between zero and one. The transition variable (s_t) can be a function of delayed components of endogenous variables, a lagged exogenous variable, a linear combination of the k series, a function of a deterministic trend t (i.e. a model with parameters that change smoothly over time).

The STVECM can be seen as a regime switching model that allows for two regimes associated with the extreme values of the transition function, $G(s_t) = 0$ and $G(s_t) = 1$, where the transition between regimes is smooth. It is assumed that the regimes are common to the k endogenous variables in the sense that one and the same transition function determines the prevailing regime and the switches between regimes in all k equations of the model. A model of our particular interest is the one in which the components of y_t are linked by a linear long-run equilibrium relationship, whereas adjustment towards this equilibrium is nonlinear and can be characterised as regime switching, with the regimes determined by the size and/or sign of the deviation from equilibrium (i.e. so called nonlinear equilibrium correction models).

In the literature there are at least three potential choices for the transition function $G(s_t)$. The logistic transition function can be stated as

$$G(s_t;\gamma,c) = \frac{1}{1+\exp\left\{-\gamma(s_t - c)/\sigma^2(s_t)\right\}}, \gamma > 0, \tag{3.4}$$

where $\sigma^2(s_t)$ is the sample variance of s_t. The parameter c can be interpreted as the threshold or border between the two regimes, in the sense that the logistic function changes monotonically from 0 to 1 as s_t increases, and $G(s_t;\gamma,c) = .5$. The parameter γ determines the smoothness of the change in the values of the logistic function and thus the smoothness of the transition from one regime to the other. As γ becomes very large, the change in $G(.)$ from 0 to 1 becomes almost instantaneous at $s_t = c$. When $\gamma \to 0$, the logistic function becomes equal to a constant (0.5) and when $\gamma = 0$, the STVECM model reduces to a linear VECM.

The second potential transition function is the exponential function

$$G(s_t;\gamma,c) = 1 - \exp\left\{-\gamma(s_t - c)^2 / \sigma^2(s_t)\right\},\ \gamma > 0. \tag{3.5}$$

It assumes symmetric adjustment for both positive and negative deviations from the equilibrium. The strength of the equilibrium correction is also changing gradually as the deviations become larger. In the STVECM presentations of equation (3.3) with the exponential transition function we can see that adjustment parameter changes from $\alpha_1 + \alpha_2$ to α_1 and back again symmetrically around c with increasing values of transition parameter (s_t).

The third popular choice is the so called 'quadratic logistic' function proposed by Jansen and Teräsvirta (1996), where

$$G(s_t;\gamma,c) = \left\{1 + \exp\left[-\gamma(s_t - c_1)(s_t - c_2)\right] / \sigma^2(s_t)\right\}^{-1},\ \gamma > 0. \tag{3.6}$$

This choice comes from the fact that the exponential model becomes linear if either $\gamma \to 0$ or $\gamma \to +\infty$, but the 'quadratic logistic' function becomes linear if $\gamma \to 0$, whereas the transition function is equal to 1 for $s_t < c_1$ and $s_t > c_2$, and equal to 0 when $c_2 < s_t < c_1$. This third function form is particularly attractive since it implies the existence of an interval band (c_1, c_2) and outside the band there is a strong tendency for the exchange rate to revert to its equilibrium level. For example, Dumas (1992) identifies this band for real exchange rate in terms of the costs of trading goods.

3.3.2 Modelling nonlinear equilibrium correction in the euro/dollar exchange rate

We follow partly MacDonald and Marsh (1997) and Juselius and MacDonald (2001). They propose augmenting the traditional PPP hypothesis with an interest rate differential. This augmentation can be considered as a balance of payments equilibrium condition, where the current account (*ca*) and the capital account (*ka*) sum to zero. We can write the long-run equilibrium in the current account as a function of competitiveness (PPP in equation 3.1) and capital account depends on nominal interest rate differentials adjusted for expected exchange rate changes (UIP in equation 3.2). Hence, for the overall balance of payments we have

$$ca_t + ka_t = \alpha(p_t - p_t^* - e_t) + \mu(i_t - i_t^* - E_t\Delta e_{t+k}) = 0^{21} \tag{3.7}$$

[21] It should be noted that we assume that net interest payments on net foreign assets are zero in the current account of balance of payments and that the measure of competitiveness (i.e. real exchange rate) is assumed to be positively related with net exports satisfying the Marshall-Lerner condition

where α is the elasticity of net exports with respect to competitiveness, and μ captures the mobility of international capital. By assuming that capital is less than perfectly mobile ($\mu < \infty$) we can solve the above equation for the long-run exchange rate equilibrium relationship in a statistical form[22] as

$$e_t = p_t - p_t^* + \mu/\alpha(i_t - i_t^* - E_t\Delta e_{t+k}) + \varepsilon_t \tag{3.8}$$

where ε_t is an error term which should be I(0). MacDonald and Marsh labelled this relationship as the Casselian PPP, since Cassel (1916) was the first to recognise the role of interest rate in keeping an exchange rate away from its PPP equilibrium value[23]. If the long-run equilibrium relation between goods and asset markets can be constructed from the variable vector (e_t, p_t, p_t^*, i_t^s, i_t^{s*}) and the long-run relation for the nominal exchange rate is given by equation (3.8), then following Juselius and MacDonald (2001) we can define the hypothetical adjustment relation for the nominal exchange rate in statistical form as

$$\Delta e_t = w_1(e - p + p^*)_{t-1} + w_2(i^s - i^{s*})_{t-1} + w_3\Delta(p - p^*)_t + \eta_t , \tag{3.9}$$

where η_t denotes the white noise error process. From the adjustment relation (3.9) we can see that the actual nominal exchange rate depreciation is related to adjustment to PPP, to the spread in short-term interest rates (monetary policy effect), and to a change in inflation differential with adjustment parameters, w_1, w_2 and w_3, respectively.

We argue as MacDonald and Marsh (1997)[24] and Juselius and MacDonald (2001) do that there are possible linear long-run relations between variables, but we also argue that the long-term government bond rate differential between the euro zone and the USA might affect the short-run exchange rate dynamics. This is consistent with Juselius and MacDonald (2001), where the US and German long-term bond rates were found to be the main driving forces in the system and short-term interest rates were important for the determination of the exchange rate changes. One of the ideas behind our argument is also that long-term government bond rates can be seen as a proxy for inflation expectations and that the difference in inflation expectations between the USA and the euro zone drives the nominal exchange rate back to its equilibrium level determined by fundamentals. However, we argue further that the adjustment process is asymmetric, or in other words, nonlinear, since agents in the foreign exchange markets are heterogeneous in their

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[22] One that captures the reasons why an exchange rate might not be in equilibrium (MacDonald 1999).

[23] MacDonald and Marsh estimate the above model for the mark, pound, and yen against the US dollar for the period 1974-1992. For each country they were able to find two cointegrating vectors or long-run relationships, which they identify as relation between the exchange rate and relative prices and relationship between interest rates, prices and exchange rate. Their result can be summarised as follows. For the one-month horizon the random walk dominates across all the currencies but for the longer horizons the relative accuracy of the structural exchange rate models improves. The results of MacDonald and Marsh are in contrast with much of the literature on forecasting exchange rates, since they suggest that the forecastibility of exchange rates with respect to fundamentals starts at around two months rather than after years like in the studies of Chinn and Meese (1995) and Mark (1995).

[24] We should note that we differ MacDonald and Marsh (1997) in that they use long-term interest rates in analysing long-run cointegration relationships.

expectations about how the future inflation rate[25] might affect the nominal exchange rate series. Hence, we argue that the degree of equilibrium correction for both real exchange rate series and short-term interest rate differential depends on the size of the long-run government bond rate difference. Here we follow Frankel's (1979) real interest rate differential (RID) model and differentiate the impact on the exchange rate of short and long term interest rate differentials. We modify Frankel's model by assuming that short-term interest rate differentials might affect to exchange rate changes nonlinearly depending on the future expectations of long-term bond differentials (i.e. expected inflation differentials). This argument makes sense, since according to the UIP an increase in relative interest rate predicts only an expected future appreciation of currency but does not tell whether it happens immediately or some time in the future. By assuming short-term interest rate differential to be dependent on the future inflation expectations in a nonlinear way allows a combination of both immediate and long-term effects to exchange rate. This means that we should examine the following smooth transition autoregressive equilibrium correction model for the euro/dollar exchange rate adjustment:

$$\begin{aligned}\Delta e_t &= (\mu_{11} + \alpha_{11} z_{t-1} + \sum_{j=1}^{p-1} \phi_{1j} \Delta e_{t-j} + \sum_{j=0}^{p-1} \kappa_{1j} \Delta X_{t-j}) \\ &+ (\mu_{12} + \alpha_{12} z_{t-1}) * G(s_t;\gamma,c) + \varepsilon_{1t}\end{aligned}, \tag{3.10}$$

where ε_{1t} are white-noise innovations. The coefficients α_{11} and α_{12} are the error correction parameters for linear and nonlinear part of the model, respectively. The transition function has the form

$$G(s_t;\gamma,c) = 1 - \exp\left\{-\gamma\left[(i^l - i^{l*})_{t-1} - c\right]^2\right\},^{26} \tag{3.11}$$

where i^l and i^{l*} denote the long-term government bond rates in domestic and foreign countries, respectively. The parameter c can be interpreted as the threshold value between two regimes in the sense that the exponential function reverts to a linear autoregressive model if the long-term government bond differential is in the equilibrium point c. The parameter γ determines the smoothness of the transition between the regimes. In our model we use both the PPP and the PPP augmented by a short-term interest rate differential as our equilibrium correction term z_{t-1}. In the first case the error-correction term has a form, $z_{t-1} = e_t - p_t + p_t^*$, and in the second case the form is $z_{t-1} = w_1(e_t - p_t + p_t^*) + w_2(i_t - i_t^*)$, where w_1 and w_2 are parameters.

We will use exponential transition function, since it is symmetric for both positive and negative deviations from equilibrium and as a transition variable we will use the previous

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[25] Also market frictions, transaction costs etc. might affect to the asymmetric behaviour.

[26] A common feature of nonlinear error correction model is that large deviations from equilibrium level are corrected, while small are not. Balke and Fomby (1997) examined so called threshold cointegration, where adjustment towards equilibrium takes effect only if the equilibrium error gets larger than a certain threshold value. An alternative view is to allow the strength of adjustment increase gradually as the equilibrium error gets larger. In this paper, we use exogenous variable as our adjustment effect variable.

period long-term government bond differences. From the above model we can see that the nonlinear equilibrium correction model allows a different adjustment process for large and small deviations from bond difference. We see also that in the above model the exogenous structure, ΔX_t, is similar in both regimes and thus only the equilibrium correction parameter changes between regimes. The exogenous structure includes lagged dependent (Δs_{t-j-1}) and independent (Δp_{t-j}, Δp_{t-j}^{*}, i_{t-j}^{s} and $i_{t-j}^{s^*}$) variables.

Our empirical analysis in section 3.5 concentrates mainly on two questions. First, we examine whether the PPP and the interest rate differences between the US and euro zone can serve as a long-run equilibrium relation for the euro/dollar exchange rate series. We examine this by testing whether there are cointegrating relationships between variables. Second, we examine whether there are smooth transition types of nonlinear relationships in the nominal euro/dollar exchange rate equilibrium correction dynamics.

3.4 Data

Our data set consists of monthly observations from January 1990 to May 2002 (148 observations) for the USA and the euro zone. The composition of the sample was determined by the availability of the data for the euro zone. The use of longer samples may also be inappropriate because of regime shifts, structural changes, which requires extreme care in empirical testing. The actual exchange rate data for dollar/euro is very short[27] and thus it is impractical to estimate any structural model with such a few observations. This means that we are forced to use synthetic euro exchange rate data from the pre-EMU period. Although, the basket currency ECU can be seen as a forerunner of the euro, the use of ECU/dollar exchange rate might not be appropriate in the euro/dollar exchange rate modelling, because the ECU basket contained currencies which are not merged into the euro and it did not include some currencies which are participating in the euro system (Clostermann and Schnatz 2001)[28].

The full set of variables is defined as

e_t = The end-of-period exchange rate (US$/euro)

[27] The common currency euro was introduced on January 1, 1999 as a new currency on the international financial markets representing eleven member states, which irrevocably fixed their domestic currencies against the newly created euro. Greece joined the European Monetary Union at the beginning of 2001 as its 12th member. When the euro was born, the European Central Bank (ECB) became solely responsible for monetary policy in the euro area, and European financial markets started to quote prices and operate in euros. In January 2002 twelve European countries took the final step in establishing a monetary union, EMU. Of the 15 countries in the European Union (EU) three stayed out, namely the UK, Sweden and Denmark. Officially the monetary union began already in January 1999. At that point the governing council of the European Central Bank (ECB) in Frankfurt became the sole decision maker of monetary policy for the participating countries. While the common euro currency only existed in electronic form and not as paper currency at that point, each of the national currencies was irrevocably fixed to the euro and defined as a sub-denomination of the euro. Europe had previously a system of fixed exchange rate, with promise to keep currencies near a certain parity within a range, but they were permitted to periodically reset these parities. A monetary union can be viewed as a more extreme version of fixed exchange regime, i.e. one that can not be adjusted or abandoned.

[28] The basic assumption in many previous studies has been that before the introduction of the single currency the substitute of the euro area was Germany because the fact that German played a major role in the EMU area.

p_t = The US consumer price index

p_t^* = The euro zone harmonized consumer price index

i_t^s = The US 3 month interest rate (Federal Treasury bill rate)

i_t^{s*} = The euro zone 3 month interest rate (Euribor)

i_t^l = The US 10 year government bond rate

i_t^{l*} = The euro zone 10 year government bond rate.

All of the data have been extracted from the OECD main economic indicators database. All variables, excluding interest rates are in natural logarithms. The graphs of the variables in levels and in differences are given in the Appendix (Figures 5 to 10).

We have also depicted various parity series in the Appendix. The series are shown in Figures 11 and 12 and they show clear asymmetric behaviour and large persistence during the sample period. For example, the behaviour of the real exchange rate series can be characterised as slower increases before 1999 and much faster increases after 1999. The most probable reason for this phenomenon is that euro/dollar nominal exchange rate has extraordinary behaviour after the launch of common currency euro[29]. Overall, the series are very persistent and show very slow adjustment or no adjustment at all back to the parity level. We can see a positive relationship between the short-term interest rate difference and the real exchange rate series and this suggest that there might be a cointegration relationship between these fundamental parity conditions. We can also see that the PPP determined exchange rate series does not closely mirror the actual exchange rate series. However, there seems to be a tendency to follow same long-run movements before the launch of common currency. The long-term government bond rate difference seems to be

[29] Over the first years the euro almost steadily depreciated against the major currencies, especially against the US dollar, and this has lead to an extensive discussion about the sources of that widely unexpected behaviour of the euro on foreign exchange rate markets. Many explanations have been given for weakness of euro but most of these explanations have invoked fundamental variables, in particular the stronger economic growth performance of the US economy in comparison with the Euro area growth performance. For example, in their study Corsetti and Pesenti (1999) presented evidence, which indicate that the euro-dollar exchange rate during 1999-2000 was very well explained by revisions of the forecasts of the growth rates of output in the USA and in Euroland. They show that when the growth rate in the USA was expected to increase relative to Euroland the value of the dollar relative to the euro increased. According to Corsetti and Pesenti the expected growth rate differentials were the most important variable explaining the euro-dollar rate during the period 1999-2000. However, they also found that this relative growth performance link between the US and Euroland does not hold for other currencies or other sample periods. De Grauwe (2000) explains this Corsetti and Pesenti finding so that at the end of 1998 the markets were increasingly influenced by positive beliefs about the growth potential of the US economy. Because of the se beliefs, agents have focused on the one variable (i.e. the growth rate of output) that provides evidence for their beliefs and agents are almost totally disregarded other fundamental variables, such as inflation differentials, current accounts, and other variables that might be equally important in determining the euro-dollar exchange rate. Thus, De Grauwe concludes that the tight fit between growth forecasts and the exchange rate is not the result of a law that links the exchange rate to economic growth differentials, but rather to a passing belief that this is the only variable that matters. In other exchange markets or in other period the beliefs are different and this strong link between output growths and exchange rate disappears. For example, during the 1970s and in early 1990s the markets were mostly interested in inflation differentials.

negative when the nominal exchange rate is appreciating and it seems to be positive when the nominal exchange rate is depreciating. This gives preliminary empirical support for the use of the bond differential as a transition variable. From the figure 12 in the appendix we can see a close comovement in the long run between the real exchange rate and bond differential.

3.5 Empirical results

3.5.1 Cointegration, identification and nonlinearity

The first step in our empirical analysis[30] was to test whether there are any linear cointegration relations between our system of five variables (e_t, p_t, p_t^*, i_t^s, i_t^{s*}). Escribano and Mira (1996) and Balke and Fomby (1997) show that the cointegrating vectors can still be estimated super consistently regardless of possibly nonlinearities in the adjustment processes. We used the Johansen cointegration technique, which is based on a maximum likelihood estimation of vector autoregressive models, since this methodology is particularly suited to examine the long-run equilibrium relationships on which our theoretical considerations in section 3.2 were based. In applying the Johansen test procedure, we set the lag length of 4 for the VAR model. We use both Johansen test statistics, namely λ_{trace} and λ_{max}.

Table 7 presents the results from the linear Johansen cointegration trace and maximum eigenvalue tests for our five variable system. The results of both tests reject the hypothesis of no cointegration in favour of cointegration. In other words, we can reject the hypothesis that there are no linear long-run relationships between variables. Both test statistics also show the existence of two significant cointegrating vectors. According to Juselius and MacDonald (2001), the cointegration rank can be seen as an indication how well the market adjusts in the long run and hence, the finding of two cointegrating vectors might reflect the fact that goods markets and capital markets markets between the US and euro zone are quite well integrated during the sample period.

[30] As preliminary tests, we tested for unit root behaviour of each variable and also for the interest rate differentials. In each case the number of lags in the augmented Dickey-Fuller procedure were chosen such that no residual autocorrelation was evident in the auxiliary regressions. We were in each case unable to reject the unit root null hypotheses at conventional levels of significance and concluded that all the variables appear to be nonstationary in levels and stationary in first differences. Thus, in the rest of the analysis we assume that all the series under consideration are in level realizations of I(1) processes.

Table 7. Tests for the number of cointegration relationships

Hypotheses H_0	Eigenvalue	λmax	trace
r = 0	0.279	*47.00*	*126.79*
r ≤ 1	0.273	*45.95*	*79.79*
r ≤ 2	0.113	17.11	33.84
r ≤ 3	0.082	12.31	16.63
r ≤ 4	0.030	4.32	4.32
ci rank		2	2

Note: The variables included in the cointegration test are euro/$ exchange rate, the CPI price levels in the euro zone and in the US, and the 3 month short-term interest rates in the euro zone and in the USA. The parameter *r* denotes the number of cointegrating vectors. If the computed test statistic is below the critical value, then we cannot reject the H_0 hypothesis. A significant (5%) test statistic is given in italics face. The critical values are taken from Osterwald-Lenum (1992) tables. The number of lags in the corresponding VAR is set equal to 4. This number was established on the basis of the Schwartz information criteria.

Although the results in Table 7 reveal the presence of two cointegrating relationships, they do not indicate whether any of these cointegration relations enter the exchange rate equation in the system of equations. We can start to analyse this question by testing whether the exchange rate is a weakly exogenous variable in our five variables system. The test of long-run weak exogeneity tests whether there is absence of long-run level feedback in the system. If the null hypothesis is accepted then the variable in the question can be considered as a driving variable in the system. This means that to interpret the equilibrium correction mechanism in an exchange rate equation all the variables except the exchange rate must meet the condition of weak exogeneity. In other words, the deviations from the long-run equilibrium level are corrected solely through exchange rate responses.

Table 8. Test for long-run weak exogeneity

e	p*	p	i_s*	i_s	i_s* and i_s
3.22	14.54	27.88	9.51	5.12	14.50 (0.01)
(0.20)	(0.00)	(0.00)	(0.01)	(0.08)	

The results in Table 8 show that the euro/dollar exchange rate and euro zone interest rate can be considered weakly exogenous variables at 10 per cent significance level. We also tested for weak exogeneity of both short-term interest rates and in addition, the p-values in Table 8 show that the interest rate spread is not weakly exogenous in this system of equations a 5 per cent significance level. The results in Table 8 strongly suggest that there might be linear cointegrating relationships in the system that are not relevant for the short-run movements of the euro/dollar exchange rate series at 5 per cent significance level. This result can be seen as a strong evidence against the linear error correction model for the exchange rate dynamics.

Although there is some evidence in Tables 7 and 8 that the PPP and monetary fundamentals might affect the short-run exchange rate movements, this does not necessarily imply that they matter in a way suggested by the linear equilibrium correction model. It is important to note that the finding of cointegration is only a necessary condition for any possible long-run stationary relationship and that the Johansen procedure only identifies the cointegration space, but it does not identify the cointegrating vectors. The reason for this is the fact that any linear combinations of two or more cointegrating vectors are also cointegrating vectors. Therefore, it is crucial for further analysis to test whether any specified structural relationship is contained in the cointegration space. Thus, we need to impose some conditions from economic theory to give a meaningful explanation for our cointegrating vectors. There are potentially many different types of fundamentals that could drive the nominal euro/dollar exchange rate. However, according to the economic theory given above in our five variable model there are only three meaningful cointegration or fundamental equilibrium relations: the PPP relationship, the interest rate spread, and the PPP relationship augmented by two interest rates. Next, we try to identify the cointegration space by imposing these restrictions on the cointegrating vectors.

In Table 9 we give the results from testing whether the above discussed three long-run relationships are supported by the data. The first restriction hypothesis (H_1) tests whether the variables s_t, p_t, and p_t^* have a cointegration relation with normalised coefficient form as (-1, 1, -1, *, *) for the vector of cointegrating variables, where the symbol (*) indicates unrestricted coefficients. Thus, the test hypothesis is formulated so that the same linear restriction is valid in all cointegration relations. The second restriction hypothesis (H_2) test, whether variables i_t and i_t^* have a cointegration relations in all cointegrating vectors with the normalized coefficient form (*, *, *, 1, -1), where symbol (*) indicates unrestricted coefficients. This restriction tests whether the interest rate differential relation enters in all cointegration vectors. The results in Table 9 shows that both hypotheses (H_1) and (H_2) can be rejected at 5 per cent significance level. The hypothesis (H_3) tests whether there exist any significant linear combination between the variables s_t, p_t, and e_t that is stationary. Thus, we test whether a vector (-1,1,-1,0,0) (i.e. strong form PPP) is a valid hypothesis on itself. The test results in Table 9 show that the hypothesis is not valid and we conclude that the variables s_t, p_t, and p_t^* do not form a long-run cointegration relationship. The hypothesis (H_4) is the so-called interest rate parity (IRP) condition for short-term interest rates (i.e. to test whether $i_t^s - i_t^{s*} \sim I(0)$), which is a simplification of the UIP relation. The IRP condition does not take into account any depreciation or appreciation effects on the exchange rate. It reflects only the arbitrage effects between domestic and foreign interest rate markets, and the short run deviations from IRP can be interpreted as the effects of transaction costs or risk premiums (see for example Mitchell 2000). The IRP condition can be rejected in our data, which implies the fact that there are no linear cointegration relations between the US and euro zone interest rates. The hypothesis (H_5) examines whether the strong form PPP is valid in the linear context if we take account also the interest rate differentials. Thus, we test the restriction that the PPP augmented with the short-term interest rate differential may be a cointegrating relationship. The results in Table 9 show that the hypothesis (H_5) can be rejected.

Table 9. Restrictions on cointegration vectors: LR tests

Hypotheses	LR	p-value
$H_1(3)$	19.29	0.00
$H_2(1)$	12.71	0.00
$H_3(3)$	36.18	0.00
$H_4(3)$	9.77	0.02
$H_5(2)$	9.53	0.01

Notes: The hypotheses are set in the form $\beta_i = H_i\phi_i$, where H_i is *(p x s_i)* matrix defined by economic hypothesis, ϕ_i is *(s_i x r)* parameter matrix to be estimated, and $s_i = p - k_i$, where k_i is the number of restrictions imposed on the cointegration relation. *p* is the number of variables in the system and *r* is the number of cointegrating relations. The significance of the restricted model is tested against the unrestricted model by a likelihood ratio test.The degrees of freedom for the χ^2 distribution are given in the parentheses. The null hypotheses for the vector (*e,p*,p,i*,i*) are following:

H_1: (-a,a,-a,,*) PPP in the cointegration space*

H_2: (,*,*,b,-b) short-term interest rate differential in the cointegration space*

H_3: Is (-1,1,-1,0,0) in Sp(β)

H_4: Is (0,0,0,1,-1) in Sp(β)

H_5: PPP together with short-term interest rate differential (-1,1,-1,m,-m)

Overall, the results for the linear cointegrating relation hypotheses in Tables 7 and 8 show a strong evidence for linear long-run relationships between our five variables in the model, but we could not find any meaningful linear structural presentations for our long-run relationships. It should be noted that tests for number of cointegrating vectors and them restrictions in previous section are based on the assumption that adjustment process driving the variables towards the equilibrium is linear. This implies the rejection of the linear long-run equilibrium correction model in five variable system model.

We proceed by hypothesising that there might be nonlinear short-run dynamics for exchange rate deviation series. We argue as in Killian and Taylor (2001) that economic models of exchange rate determination imply linear long-run equilibrium conditions but that the adjustment towards equilibrium might be nonlinear. So, for the further analysis we set the cointegration rank $r = 2$ and pre-specify the two cointegrating vectors as (-1, 1, -1, 0, 0)' and (-1, 1, -1, *m, -m*)'. Thus we set the first cointegrating vector as a strong form PPP and second cointegrating vector as the strong form PPP augmented by interest rate differential. However, to compare our estimation results between the linear and the nonlinear equilibrium correction specifications we estimate also the linear equilibrium correction model now with the restrictions on two cointegrating vectors according to the above-mentioned pre-specifications.

Estimation of linear equilibrium correction model for the PPP model now yields

$$\Delta e_t = \underset{(0.06)}{0.04} - \underset{(0.02)}{0.02}\, q_{t-1} + \underset{(0,09)}{0.38}\, \Delta e_{t-1} - \underset{(0,09)}{0.16}\, \Delta e_{t-2} + \hat{\beta}_1' \Delta X_t + \varepsilon_{1t}\,, \tag{3.12}$$

R^2 = 0.15, DW = 1.94, σ_ε = 0.022, *SK* = -0.11, *Kur* = 0.41, *JB* = 1.33 (0.51), *ARCH*(1) = 0.01 (0.94), *ARCH*(4) = 2.53 (0.64), *Q*(1) = 0.02 (0.88), *Q*(4)= 4.77 (0.31), *T* =146, *AIC* = -362.3

and for the PPP augmented with interest rate difference gives

$$\Delta e_t = \underset{(0.06)}{0.09} - \underset{(0.02)}{0.03}\left[q_{t-1} - \underset{(0.03)}{0.04}(i_s^{usa} - i_s^{euro})_{t-1} \right] + \underset{(0.09)}{0.36}\,\Delta e_{t-1}$$
$$- \underset{(0.09)}{0.162}\,\Delta e_{t-2} + \hat{\beta}_2'\,\Delta X_t + \varepsilon_{2t} \qquad (3.13)$$

R^2 = *0.16, DW* = 1.94, σ_ε = 0.022, *SK* = -0.12, *Kur* = 0.55, *JB* = 1.35 (0.51), *ARCH*(1) = 0.015 (0.90), *ARCH*(4) = 2.971 (0.56), *Q*(1) = 0.073 (0.78), *Q*(4)= 4.285 (0.37), *T* =146, *AIC* = -362.4

where ΔX_t denotes the exogenous structural variables of the model, σ_ε is residual standard deviation, *SK* is skewness, *Kur* is kurtosis, *JB* is the Jarque-Bera test of normality of the residuals, *ARCH(q)* is the LM test of no ARCH effect up to order *q*, *Q(j)* is Ljung-Box test for no residual autocorrelation up to lag *j*, and *T* is the number of observations. The figures in parentheses following the test statistics are *p*-values.

From linear equilibrium correction estimations we can see that the equilibrium correction parameters are not significant and this confirms our hypothesis that there is no linear adjustment in exchange rate movements during the sample period. However, the model structure seems to be quite satisfactory for all the other coefficients and there is no autocorrelation, normality or heteroskedasticity problems in the estimation.

3.5.2 Nonlinear equilibrium correction models

In this section we investigate the empirical usefulness of smooth transition equilibrium correction models (STECMs) for the euro/dollar exchange rate adjustment towards its long-run equilibrium. The STECMs are estimated for both the PPP fundamental and PPP fundamental model augmented by interest rate differential. We estimate parameters of our STECM by using nonlinear least squares (NLS) estimation technique and follow the suggestion by Teräsvirta (1994) and standardise the exponent of transition function by dividing it by the variance of the transition variable[31]. This makes the smoothing parameter a scale-free parameter. We start the nonlinear estimation with an unrestricted AR model with 5 lagged first difference in the first regime as we started also in the linear specification and removed all the insignificant lagged first differences until all autoregressive parameters were statistically significant. The estimation results for the PPP fundamental model gives

[31] We tried also other transition functions, but when we estimated the parameters of our STECM with the logistic and quadratic logistic function, we find that they do not fit the data well and do not render sensible results. Therefore, in the rest of our analysis the exponential function is our form for transition function.

$$\Delta e_t = \underset{(0.02)}{0.03} - \underset{(0.06)}{0.16}\, q_{t-1} + \underset{(0,08)}{0.33}\, \Delta e_{t-1} - \underset{(0,09)}{0.19}\, \Delta e_{t-2} + \hat{\beta}_3'\, \Delta X_t$$
$$+ (-\underset{(0.01)}{0.04} + \underset{(0.06)}{0.18}\, q_{t-1}) \times \left\{ 1 - \exp\left[-\underset{(0.87)}{2.64}^2 (i_l^{usa} - i_l^{euro})_{t-1}^2 / \hat{\sigma}_{s_t}^2 \right] \right\} + \varepsilon_{3t} \quad , \qquad (3.14)$$

R^2 = 0.21, DW = 1.94, σ_ε = 0.022, SK = 0.15, Kur = -0.20, JB = 0.68 (0.71), $ARCH$(1) = 0.19 (0.66), $ARCH$(4) = 4.53 (0.34), Q(1) = 0.09 (0.76), Q(4) = 2.11 (0.72), T = 146, $V = \sigma_{NL}/\sigma_L$ = 0.94, AIC = -366.0

and for the PPP augmented by interest rate difference gives

$$\Delta e_t = \underset{(0.03)}{0.07} - \underset{(0.09)}{0.27}\left[q_{t-1} - \underset{(0.01)}{0.03}(i_s^{usa} - i_s^{euro})_{t-1} \right] + \underset{(0,09)}{0.28}\, \Delta e_{t-1} - \underset{(0.09)}{0.23}\, \Delta e_{t-2} + \hat{\beta}_4'\, \Delta X_t$$
$$+ (-\underset{(0.49)}{0.07} + \underset{(0.09)}{0.25}\left[q_{t-1} - \underset{(0.02)}{0.04}(i_s^{usa} - i_s^{euro})_{t-1} \right]) \qquad (3.15)$$
$$\times \left\{ 1 - \exp\left[-\underset{(3.00)}{8.83}^2 (i_l^{usa} - i_l^{euro} + \underset{(0.02)}{0.13})_{t-1}^2 / \hat{\sigma}_{s_t}^2 \right] \right\} + \varepsilon_{4t}$$

R^2 = 0.30, DW = 1.89, σ_ε = 0.020, *Skewness* = 0.24, *Kurtosis* = -0.09, *Jarque-Bera* = 1.43 (0.49), $ARCH$(1) = 0.04 (0.83), $ARCH$(4) = 2.44 (0.65), Q(1) = 0.29 (0.59), Q(4)= 2.13 (0.71), T =146, $V = \sigma_{NL}/\sigma_L$ = 0.89, AIC = -377.0

where V is the relative residual standard deviation between the nonlinear STECM and linear VECM models. To compare our models relative to linear model, we check the reduction in the estimated residual variances delivered by the nonlinear specifications. The residuals standard deviation of nonlinear PPP equilibrium correction model 3.14 is about 94% of the linear models and corresponding ratio for the PPP augmented by interest rate differential is about 89% (in other words the gain in terms of residual variances for models 3.14 and 3.15 are 6% and 11%, respectively). The errors appear to be normal and according to LM-tests there seems to be no heteroscedasticity and the errors are also free from autocorrelations.

The same results for parameter estimates were obtained in both nonlinear equilibrium correction models by varying the starting values indicating the global maximums for both nonlinear least square functions. The parameter γ provides an indication about the speed at which the exchange rate reverts to its equilibrium level. The estimated parameters are of the expected sign and size and especially the crucial smooth transition parameter coefficients are significant also by the Dickey-Fuller critical values[32]. When viewing the equation for nonlinear short-run dynamics, we can see that the equilibrium correction terms are now significant in both nonlinear equilibrium correction specifications. We can

[32] It is important to note that the "t-ratios" of smooth transition parameters should be interpreted with caution, since under the null hypothesis H_0: $\gamma = 0$, each of the exchange rate series follow a unit root process (Taylor et al. 2001).

see that the estimated parameter estimates for equilibrium correction terms affect positively in the first regime but negatively in the second regime. This result implies that the exchange rate is an unpredictable process when the long-run government bond difference (i.e. proxy of inflation expectations difference) is small but equilibrium correction terms affect negatively and thus correct the disequilibrium when the bond difference becomes larger according to exponential function form. Especially in model 3.15 the magnitude of smooth transition parameter indicates very rapid adjustment when the long-term interest rate spread becomes large enough.

Figures 13 and 14 in the Appendix show graphs, which illustrate estimated models. The satisfactory goodness of fit of the nonlinear equilibrium correction model is highlighted with plot of actual and fitted values of change in exchange rate series during the sample period in the upper left panels. The plots show that the fitted values are reasonably close to actual values, especially for the augmented model. The residual plots in the upper right panels show that residuals in both models are reasonably well behaved (i.e. white-noise processes). The lower left panels show how the transition function evolves over time. We can see that nonlinearity explains mainly the behaviour of the exchange rate dynamics at the end of the 1990s and at the beginning of 2000s. There are only a few periods after the launch of common currency in the PPP model, which can be interpreted by the linear part of the model.

One way to show the effect of transition speed on the evolution between the two regimes is to plot the estimated transition function against the transition variable. These plots are presented in the lower right panels of Figures 13 and 14 and they show the effect on transition speeds. We can see a rather low speed of adjustment for the PPP model, where the transition function has small values even for large deviations and a rather high speed of adjustment for the interest rate augmented PPP model, where the adjustment takes place even for small deviations. Overall, these plots show that the ESTAR specifications are quite appropriate, since we have both positive and negative deviations and the shapes of the transition functions are symmetric around the equilibrium point.

In figures 3 and 4 we have plotted the time varying error correction parameters for the models (3.14) and (3.15). The figures show that there are periods when the error correction is very rapid, especially in the early 1990s. However, the error correction moves very rapidly from large negative values to small positive or negative values and, thus, suggesting that the error correction dynamics is time-varying. This implies the difficulties for linear models to find any significant error correction parameter values.

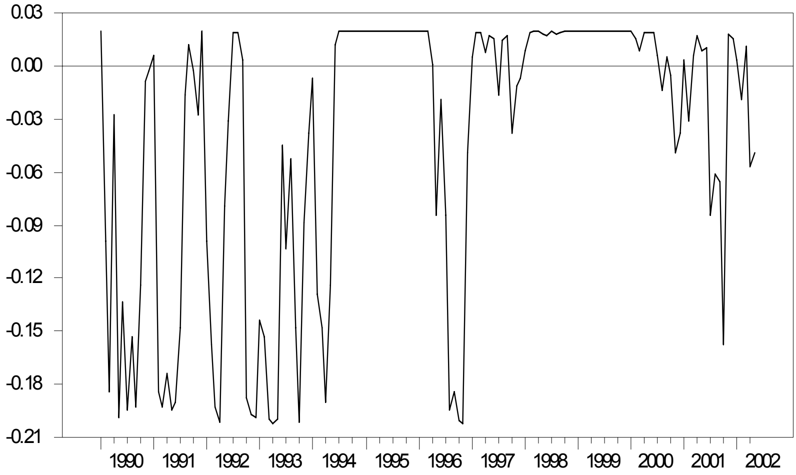

Fig. 3. Time-varying error-correction parameter for the model (3.14)

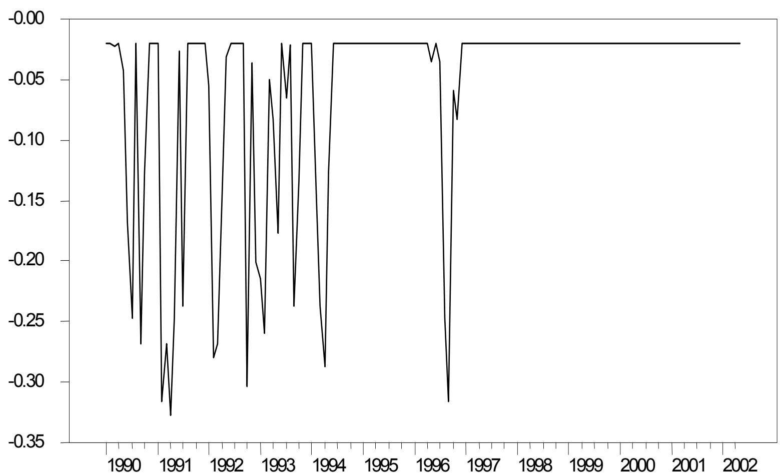

Fig. 4. Time-varying error-correction parameter for the model (3.15)

3.6 Conclusions

We have examined the empirical specification of a nonlinear vector equilibrium correction model for the monthly euro/dollar exchange rate movements during the period 1990 to 2002. We used a model, in which the influence of error correction parameter is allowed to change over time. We find that the euro/dollar exchange rate long-run dynamics is driven by the balance of payments equilibrium condition and its short-run dynamics is dependent on the long-term government bond difference between the euro zone and the USA. There is evidence of an unstable and nonlinear relationship between fundamentals and the euro/dollar exchange rate. The time-varying nature of fundamentals response to the euro/dollar exchange rate improves significantly the fit of the error-correction model. Specifically, by allowing nonlinearities in the short-run dynamics we were able to find an error-correction model for the euro/dollar exchange rate series. This seems to be a new result in the literature.

We find a better in sample fitting for the ESTAR type of nonlinear dynamics relative to the linear model and this seems to confirm the regime dependent hypothesis in the exchange rate movements. We showed empirically that the comovements of euro/dollar exchange rate and fundamentals series cannot be reliable modelled by simply linear models. It seems that fundamentals do not only matter for the exchange rate, but are also related to the switches between significant and no significant error-correction term. A better device is to use nonlinear econometric models, which could cast some light on the deeper relationships between the euro/dollar and fundamental economic variables.

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Appendix 1 Figures

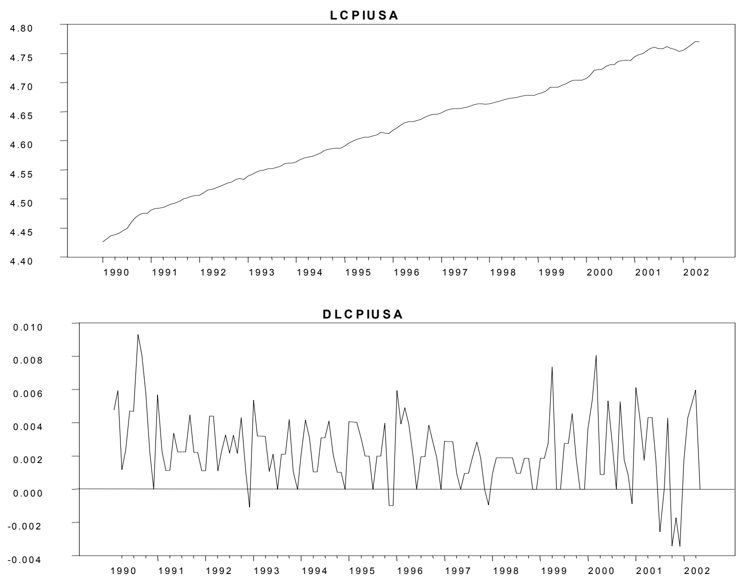

Fig. 5. The log of the USA CPI in levels and differences

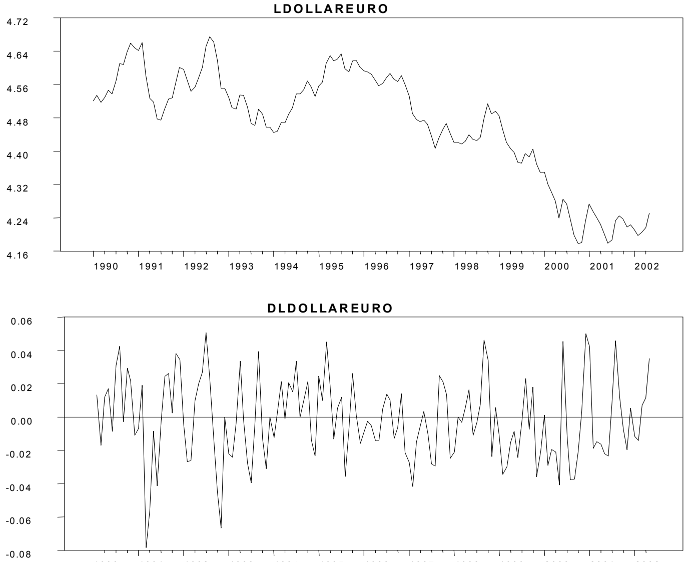

Fig. 6. The log of nominal exchange rate of the euro in dollars in levels and differences

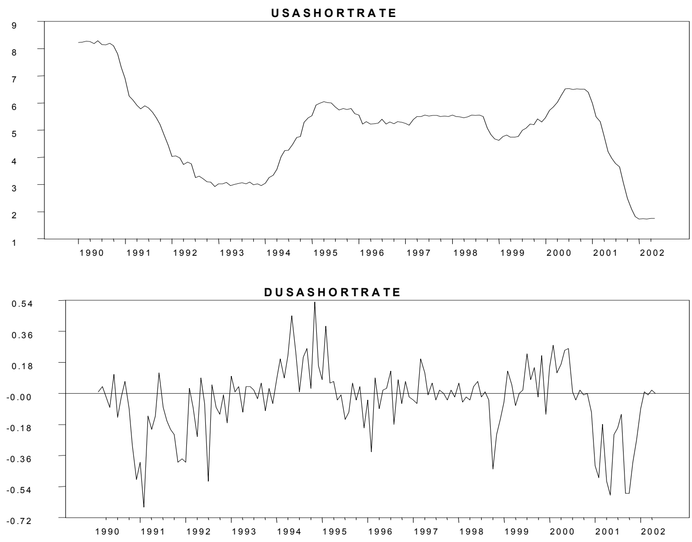

Fig. 7. The USA monthly short-term interest rate in levels and differences

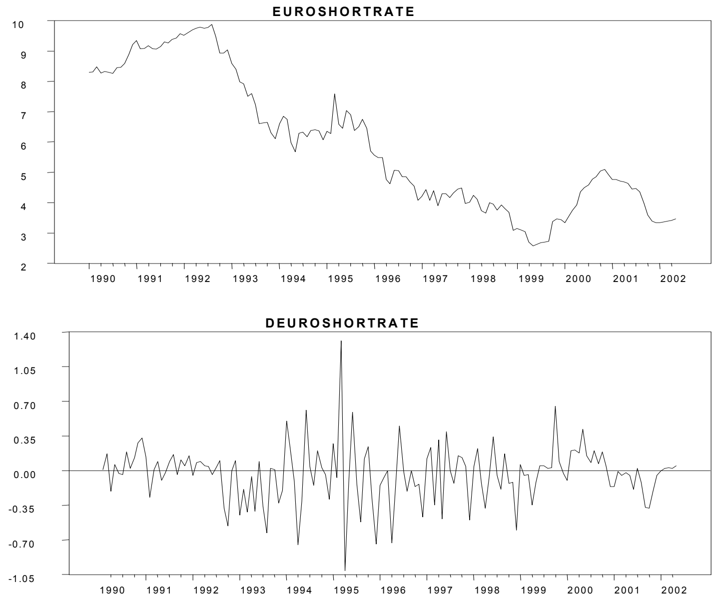

Fig. 8. The eurozone monthly short-term interest rate in levels and differences

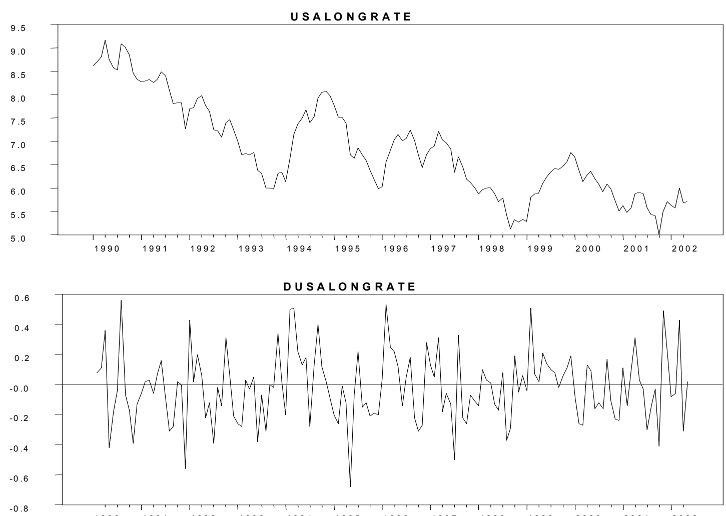

Fig. 9. The USA monthly bond rate in levels and differences

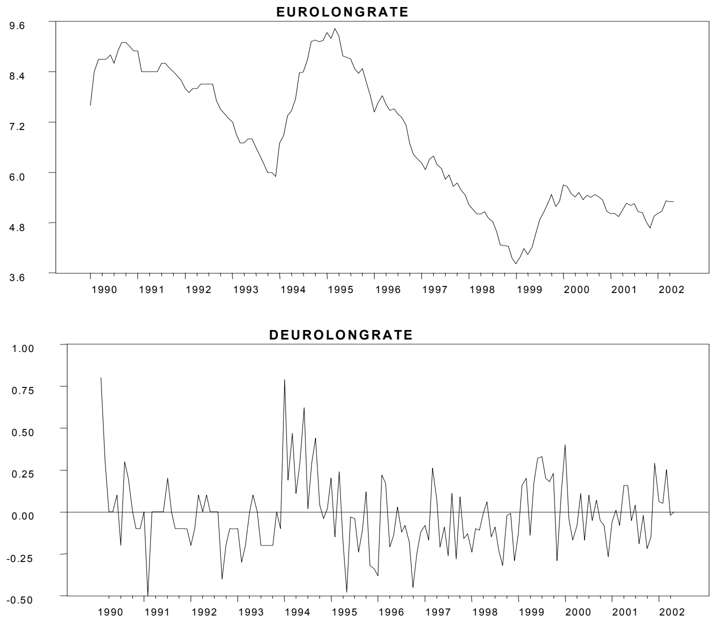

Fig. 10. The eurozone monthly bond rate in levels and differences

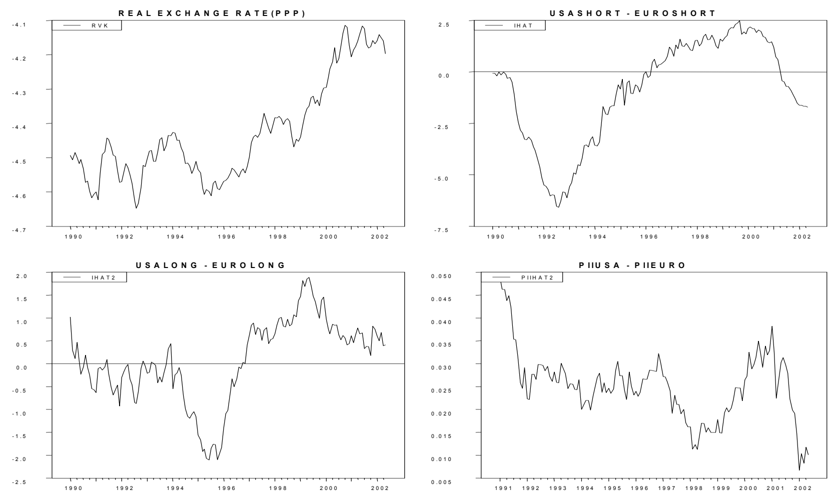

Fig. 11. The series

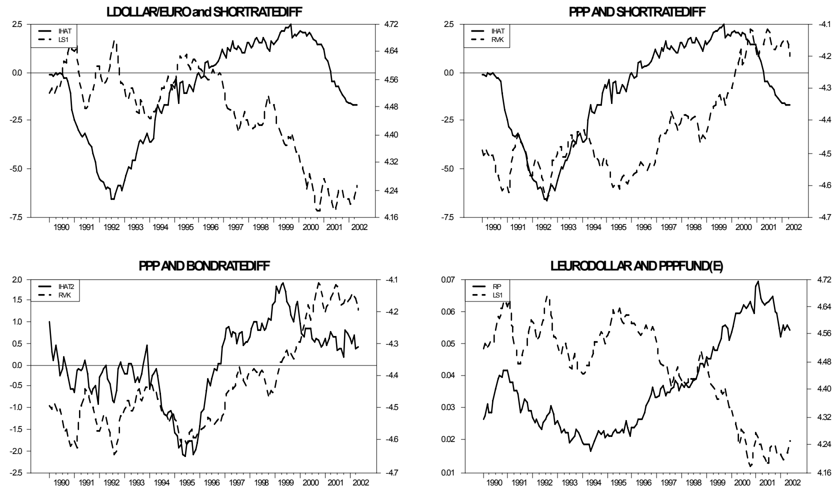

Fig. 12. The parity relations

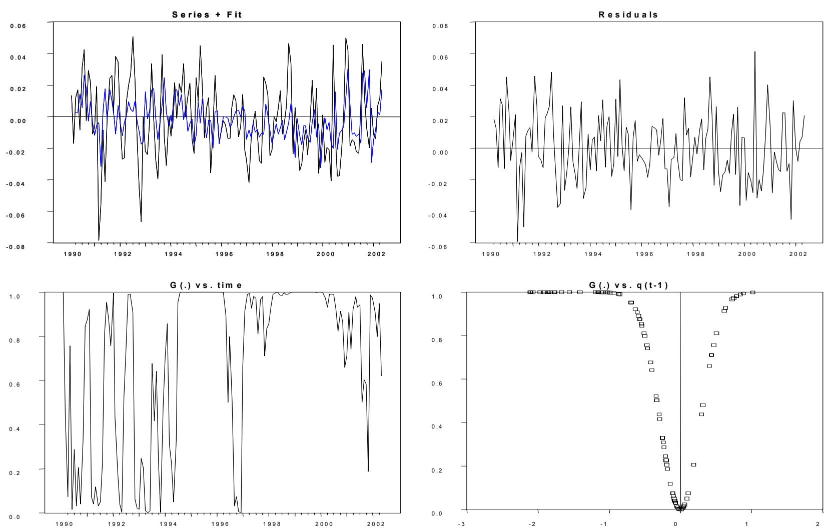

Fig. 13. ESTAR model for PPP fundamental exchange rate series

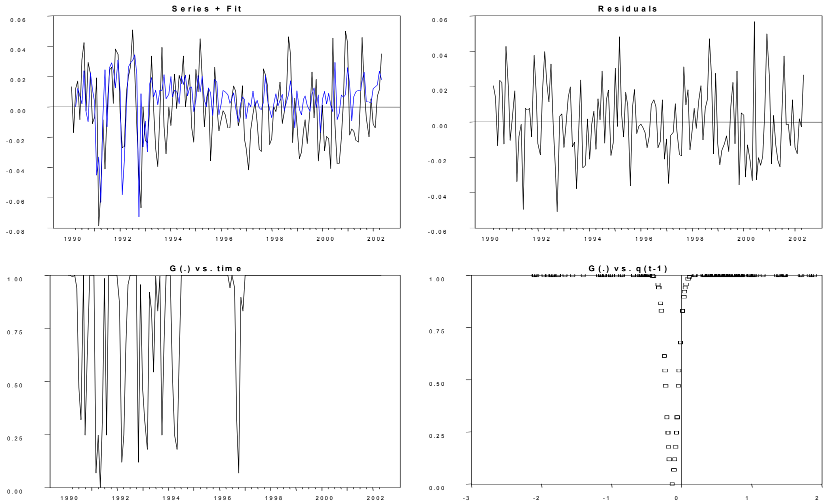

Fig. 14. ESTAR model for PPP fundamental augmented by interest rate differential

4 Analysing the monetary model of exchange rates in a time-varying error-correction framework

4.1 Introduction

The monetary model of exchange rate determination gives a link between exchange rate and a set of monetary fundamentals. Linear cointegration has been one of the most typical estimation frameworks to examine monetary models' empirical performance (e.g. MacDonald and Taylor 1994 and Mark 1995). However, results from these cointegration studies have not been able to present any clear empirical link between monetary fundamentals and exchange rate at least in the short run. Meese and Rogoff (1983a,b) results that the conventional monetary models are unable to outperform the naïve random walk models still remains valid. However, recent studies by Mark (1995), Chinn and Meese (1995), MacDonald (1999) and Mark and Sul (2001) show that deviations from a simple set of monetary fundamentals can be useful in predicting US dollar exchange rates at longer horizons[33]. Berben and van Dijk (1998) and Berkowitz and Giorgianni (2001) have criticised these studies heavily about the assumption of a stable cointegrating relationship. Furthermore, a recent study of Cheung et al. (2002) uses a wide range of linear exchange rate models and comes to the conclusion that no existing exchange rate model specification could be very successful in forecasting the exchange rate behaviour[34].

The rather poor empirical performance of the exchange rate models provides a motivation to find an empirically meaningful way to present a link between exchange rates and monetary fundamentals. A major shortcoming in many previous studies might have been that they have implicitly assumed that exchange rate adjustment towards monetary fundamentals is linear. The recent studies of exchange rate models have examined the possibility of nonlinear relationship between monetary fundamentals and exchange rate[35]. There are a few studies, most notably Taylor and Peel (2000) and Kilian and Taylor (2003), which show that exchange rate and monetary fundamentals link might involve

[33] For a survey of exchange rate modelling see, for example, Sarno and Neely (2002) and Sarno and Taylor (2003).

[34] However, they conclude also that one model might do well for one exchange rate but not for another. Similar arguments can be easily shown to be valid for examining different time periods for one exchange rate series.

[35] For a selective survey on nonlinear exchange rate models see Sarno (2003).

important nonlinearities[36]. These authors have argued that a key explanation of poor empirical fit in a monetary model is that the standard models are theoretically relevant but their empirical analyses during the floating exchange rate regime is inadequate. The argument goes that the data on exchange rate series include nonlinearity's that cannot be sufficiently detected by using only linear estimation methods.

The purpose of this chapter is to present a way to detect the presence of cointegration relationships in a monetary model of exchange rates for major currencies against the US dollar during the post Bretton Woods period. We start by using a conventional linear cointegration model, and embed it by estimating a nonlinear error-correction model in the form of smooth transition regression (STR) model[37]. There are a few papers which have examined nonlinearity in monetary exchange rate models. But to our knowledge our study is the first that uses time-varying error-correction framework. Furthermore, we connect linear cointegration methodology and nonlinear STR method by using a linear long-run relationship or cointegrating vector with nonlinearity in the speed of adjustment to the long-run equilibrium[38]. Thus, we argue that there is a theoretically consistent long-run relationship between exchange rates and monetary fundamentals, but the way this relation affects short-run dynamics might change during the sample period. Especially, we argue that the inflation differential is the key variable in determining the strength of error correction.

The objectives of this chapter can be summarized as follows. First, we try to identify the way that nonlinearities are revealed in monetary models of exchange rate. Specifically, we propose a flexible price monetary model, where the inflation rate differential is a key variable in determining the parameter values of the model. We assume that interest rate differential also depends on the inflation differential. Second, we examine whether there are any linear long-run relationships between exchange rate and monetary fundamentals. The questions we address in linear models are:

1. is the system identified by theoretically consistent parameter values;
2. whether the systems long-run structure changes during the sample period (i.e. whether cointegration is significant for a full sample period or is it significant only for a particular subperiod),;and
3. which variables are weakly exogenous and, thus, which variables are driving the system.

Third, we test whether we can present the short-run dynamics of the model by a nonlinear adjustment process.

Our findings support the nonlinear connection between exchange rates and monetary fundamentals. We find significant cointegration relations between the traditional monetary fundamentals (i.e. relative money and relative output) and exchange rate series for

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[36] However, a number of papers have used nonlinear extensions with little success. For example, Engel and Hamilton (1990) and Engel (1994) find no superior forecasts to the random walk by using the Markov switching regression models.

[37] Similar type of models have most notably been used by Sarno (1999), Teräsvirta and Eliasson (2001) and Sarno et al. (2003) for money demand equations and van Dijk and Franses (2000) for term structure of interest rate models.

[38] For example, Taylor and Peel (2000) examined whether the cointegration vector itself is nonlinear in the monetary model.

several major currencies against the US dollar when the cointegration relations are dependent on inflation rate differentials. In particular, our proposed monetary model variant is able to find significant error correction in majority of cases. This result can be interpreted to support a regime specific validity of monetary model as suggested in recent studies (see Goldberg and Frydman 2001, De Grauwe and Vansteenkiste 2001, and Frömmel et al. 2003). Moreover, we find that interest rate differentials affect to exchange rate changes in a nonlinear way.

The rest of the chapter is structured as follows. In Section 4.2 we present some theoretical background on the exchange rate determination and motivate our model. Our empirical specification is presented in Section 4.2. Section 4.3 presents the data and reports our test results for nonstationarity and cointegration. Section 4.4 analyzes the nonlinear error correction model. Section 4.5 concludes.

4.2 Theoretical framework and empirical specification

4.2.1 The monetary model

The starting point of our analysis is the standard flexible price monetary model of exchange rate introduced by Frenkel (1976). The model has also been a starting point for many recent empirical papers (see MacDonald and Taylor, Mark 1995 and Mark and Sul 2001 among others). The monetary model can be presented by the following four equations:

$$e_t = p_t - p_t^* \tag{4.1}$$

$$m_t - p_t = \beta_1 y_t - \beta_2 i_t \tag{4.2}$$

$$m_t^* - p_t^* = \beta_1 y_t^* - \beta_2 i_t^* \tag{4.3}$$

$$i_t - i_t^* = E(e_{t+1} \mid I_t) - e_t, \tag{4.4}$$

where e_t presents the price of domestic currency in terms of foreign currency, p_t is the price level, m_t is the domestic stock of money, i_t is the interest rate, and β_1, β_2 are positive parameters. An asterisk denotes a foreign country variable and with the exception of interest rate, the lower-case letters denote log-levels. The presentation $E[e_{t+1}|I_t]$ denotes expectations of next period's level of the exchange rate, conditional on information in period t.

The equation (4.1) states that purchasing power parity (PPP) is assumed to hold continuously and it is the good-markets arbitrage that will move the exchange rate to equalize price differentials. In equations (4.2) and (4.3) we assume that money supplies are exogenous and stable. Furthermore, the domestic and foreign demands for real money are simply a function of real income and interest rate, where the income is assumed to be at its full employment level. In equation (4.4) it is assumed that assets are perfectly substitutable and, hence, the uncovered interest rate parity (UIP) condition holds, which implies that risk-neutral arbitrage will equalize the expected returns of domestic and foreign investments.

The PPP and the UIP are the central building blocks of the monetary model for exchange rate. Unfortunately short-run failures of these parities are stylized facts in international macroeconomics literature[39].

Solving from (4.1) to (4.3), we obtain the following relationship:

$$e_t = (m_t - m_t^*) - \beta_1(y_t - y_t^*) + \beta_2(i_t - i_t^*) \tag{4.5}$$

When the UIP condition (4.4) holds we can write the exchange rate equation as

$$e_t = (m_t - m_t^*) - \beta_1(y_t - y_t^*) + \beta_2[E(e_{t+1} \mid I_t) - e_t]^{40}. \tag{4.6}$$

In the steady state the term $E(e_{t+1}|I_t - e_t)$ equals zero[41], so that $i_t = i_t^*$ and we get the following expression

$$e_t = (m_t - m_t^*) - \beta_1(y_t - y_t^*). \tag{4.7}$$

For example, Mark and Sul (2001) view the equation (4.7) as a "generic representation" of the long-run equilibrium exchange rate implied by theories of exchange rate. They restrict also the income elasticities of money demand to be on for both countries (i.e. $\beta_1 = 1$). In further analysis we will use presentation (4.7) as our long-run cointegrating vector of monetary model for two reasons. First, recent papers (see Mark and Sul 2001 among others) have shown that there is a significant long-run predictable component (i.e., a linear long-run cointegrating vector) in monetary fundamentals that can be used in examining exchange rate changes in major currencies at least in longer horizons. However, we argue that the long-run relation expressed in equation (4.7) is not necessarily always valid. Second, we use equation (4.7), since several survey studies (see e.g. MacDonald and Marsh 1996, Cheung et al. 1999 and Cheung and Chinn 2001), have shown that these traditional fundamentals are the two most important variables for market participants in determining the long-run equilibrium level of exchange rate series. However, the impact of these fundamentals affecting in the exchange rates probably changes over time and we take the cointegrating vector as a time-varying variable.

We propose a monetary model that is very similar to the flexible price monetary model presented above. The point we deviate from it is in the formulation of expectations in

[39] However, recent papers (e.g. Michael et al. 1997 and Baum et al. 2001) have presented more encouraging results for PPP in taking into account the possible nonlinearities in the data generating process for several major currencies.

[40] It is typical in rational expectations solutions that the nominal exchange rate depends upon the expected future path of the driving variables, which is in this case the interest rate differential.

[41] However, this is not a general case and it is valid only when the money and output growth rate differentials between countries are zero.

equation (4.6). Instead of assuming perfect foresight we ask the question of what determines the exchange rate expectations at time *t*. To put it differently, we ask what is the best predictor for expectations about future exchange rate behaviour at time *t*. Surely much of the short-term fluctuations in exchange rates are driven by expectations about the future behaviour of macrofundamentals and hence, we should try to find possible future values of these variables. This is probably an impossible task and to get some information we must use both theoretical considerations and survey data. Several survey studies have shown that the most important variables that affect future exchange rate series in addition to traditional fundamentals are the interest rate and inflation rate differentials between countries. We argue first that the short-term interest rate differential between countries might affect the exchange rates in a regime specific way. Specifically, we argue that the short-term interest rate differential does not behave like a traditional fundamental. According to the UIP, countries with high interest yields should be expected to see their currencies depreciate. Empirical results, however, suggest that exchange rates connected to high interest yields seem to appreciate (see e.g. Lewis 1995 and Engel 1996). Our argument is that the short-term interest rate differentials do not move directly in relation to the exchange rate series, but their effect depends on relative inflation differentials. This argument makes sense, since the UIP condition only predicts that an increase in the relative interest rate means an expected future appreciation of the currency and thus do not imply whether this appreciation happens immediately or some time in the future or if it is a combination of both these effects (see Mark and Moh 2001).

The argument that the interest rate differential and exchange rate changes are connected to each other in a nonlinear way is reasonable also for several other reasons. First, since several papers have shown that the exchange rate series contain nonlinearities and because the short-term interest rates and exchange rate changes are functionally related by the interest parity condition it is relevant to assume potential nonlinearities also in relative interest rate series. Second, since the short-term interest rates are also tools of monetary policy, large changes in relative inflation rate series tell us something about future monetary policy (i.e. interest rate) changes[42]. An implication of the monetary policy change is that there are some environments where the standard textbook view of linear connection between exchange rates and interest rates does not hold. For example, interest rate changes in a particular country may occur only when the relative inflation rate is sufficiently large[43]. Third, the relationship between the exchange rate change and the interest rate differential is closely connected to capital mobility. If the capital mobility is not perfect the UIP does not hold, and hence, the interest rate differential might have a very loose or no effect at all for exchange rate changes. However, we do not think that in this environment there is a total lack of connection between the exchange rate and the interest rate difference, but that market participants are agnostic about the true equilibrium value of exchange rate for given level of interest rate differential. In the next subsection we will formulate our version of the empirical specification of the model.

[42] For example, since the bank of England was given responsibility for controlling inflation in 1997, its monetary policy committee has used short-term interest rates as the tool for achieving its inflationary target.

[43] Furthermore, the nature of policy action may also depend on the direction of the deviation from equilibrium.

4.2.2 Error-correction specification

We consider the following nonlinear error correction model

$$\Delta e_t = a_0 + \sum_{i=1}^{p} b_i \Delta e_{t-i} + \sum_{j=0}^{q} c_j \Delta f_t + \lambda_1 z_{t-1} + \sum_{k=0}^{r} \mu_k (i-i^*)_{t-k}$$
$$+ \left[a_0^* + \lambda_1^* z_{t-1} + \sum_{l=0}^{s} \mu_l^* (i-i^*)_{t-l} \right] G(s_t;\gamma,d) + \varepsilon_t \quad , \qquad (4.8)$$

where $f_t = (m_t - m_t^*) - (y_t - y_t^*)$, ε_t is independently and identically distributed error term and z_{t-1} is a linear cointegration vector for the system $[e_t, (m_t - m_t^*), (y_t - y_t^*)]$ and thus gives a long-run equilibrium value of the exchange rate determined by the fundamentals in the form of equation (4.7)[44]. The parameters λ are the error-correction parameters and they govern the adjustment to the long-run equilibrium. The parameter d denotes the threshold value of regimes. The constant parameter a_0, the error correction parameter λ_1, and the interest rate differential parameter μ_k are dependent on state s_t. The parameter values depend on the transition variable, which is defined as an inflation rate differential

$$s_{t-m} = \pi_{t-m} - \pi_{t-m}^* , \qquad (4.9)$$

where m is the lag of transition variable. The choice of relative inflations as the transition variable suggests that in a higher relative inflation regime the monetary model behaves differently than in a lower relative inflation regime. This is consistent with Frankel (1979), who suggested that the inflation environment plays crucial role in monetary models. The error-correction form in (4.8) with the transition variable defined as in (4.9) can be seen as a progressive monetary model, where the error-correction term and interest rate differential are regime dependent on the inflation differential variable. This is in line with Frankel's (1979) assumption that the current spot exchange rate and an equilibrium rate are dependent on inflation differential between domestic and foreign countries. However, we differ from his model in that we assume that magnitude of inflation differential affects nonlinearly to the error-correction and the interest rate differential terms.

The class of transition functions, G, we use are the parametric STAR models and their variations. If we allow a nonlinear adjustment to the long-run equilibrium, an important question is what form of transition functions is best suited to characterize the dynamics of error correction in the exchange rate series. In the literature, the distinction is usually made for example between positive and negative deviations, small and large deviations or symmetric and asymmetric deviations. We use the following transition function forms

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[44] We should note that we have a linear long-run equilibrium relation, which is consistent with monetary model theory in mind when we specify the cointegration vector for our error correction presentation (4.8). This view is motivated by the assumption that exchange rates cannot move independently of macroeconomic fundamentals in the long run.

$$G(s_t;\gamma,d)=\begin{cases}1/(1+\exp\{-\gamma_i(s_{t-d}-d_i\ )\})\\ 1-\exp\{-\gamma_i(s_{t-d}-d_i)^2\}\end{cases}. \tag{4.10}$$

The first functional form is the logistic smooth transition autoregressive (LSTAR) function (see e.g., Teräsvirta 1998). The LSTAR is a monotonically increasing function of deviation from equilibrium and it describes asymmetric nonlinear adjustment process. The second function is the exponential smooth transition autoregressive (ESTAR) function and it is commonly used in nonlinear applications of exchange rate determination (see e.g. Taylor and Peel 2000). This formulation allows symmetric adjustment above and below of the threshold level. However, a drawback in the ESTAR model is that the model will be linear if the adjustment parameter γ is near zero or if it will become very large[45]. The parameter γ is the adjustment parameter, which determines the smoothness of the change in the value of function (4.8). The parameter d in (4.10) denotes the threshold level of relative inflation rate.

We can summarize our empirical presentation as follows. We argue that there is a long-run cointegrating relationship that ties the traditional monetary fundamentals down to nominal exchange rate. This relation acts as an attractor for our dynamic error-correction models. The error-correction dynamics are nonlinear in the sense that the transition variable, inflation differential, determines the strength of attraction towards the long-run equilibrium. In other words, the larger is the transition variable in an absolute sense, the faster it will be driven to its fundamental level. Thus, it might be that for some sample period there is a lack of error-correction and for some sample period there is a very fast movement towards the steady state. However, the short-run coefficients on the traditional fundamentals are kept constant across all regimes. Furthermore, we present the interest rate differential as a nonlinear function of inflation differential in the short-run analysis.

4.3 Data, preliminary tests, and long-run relations

4.3.1 Data

We use quarterly data over the period 1974:1 – 2001:3 for six bilateral exchange rates and monetary fundamentals. We use only the post Bretton Woods period in order to eliminate the possible instability due to transition from fixed to floating exchange rate regime. The reference country is USA and the other countries involved are Canada, France, Germany, Italy, Japan, and the United Kingdom. The data is obtained from Engel and West (2003). The exchange rates are the end-of-period exchange rates, and the prices are consumer

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[45] Jansen and Teräsvirta (1996) show that linearity can be avoided if one uses the quadratic logistic (QLSTAR) function form. In this form the model is linear if γ goes to zero and equals one if γ goes to infinity. Thus, QLSTAR has a three-regime threshold model as the limiting case. We shall also note, as in Kilian and Taylor (2003), that the transition function dimension might be broader than a singleton presented above and thus allows more smoothness in the adjustment process.

price indices. The inflation series are constructed as annualized changes in price indices. We have also used seasonally adjusted money supply (M4 for the UK and M1 for all other countries) and seasonally adjusted real GDP for all countries except Germany and Japan, for which we have used combinations from different data sources (for more information see Engel and West 2003). We have used 3-month nominal interest rates from the euro markets, since we have assumed that interest rates from different countries have similar degree of risk and that there should be no different transaction or information costs associated with different assets. The data is converted by taking logarithms, except for the interest rates.

4.3.2 Testing for stationarity

We first examine the unit root behaviour of spot exchange rate series and relative traditional fundamentals by using the standard augmented Dickey-Fuller unit root test. In each case the number of lags is chosen by the Akaike information criteria (AIC). Table A4.1 in the appendix reports the results from our unit root tests. The test results indicate that we cannot reject a unit root for all the exchange rate series and almost every relative traditional fundamental series for the whole sample. The only exception is the Italy-US relative money series, for which we could reject the unit root null at a conventional five percent significance level. However, there is much more evidence of stationarity in the traditional relative fundamental series for sub-sample periods. We consider this as evidence of nonstability in the data generating processes. However, we should note that both subsample periods are too short to obtain fully reliable test statistics[46]. In order to make conclusions about the nonstationarity for the further analysis we should use only the full sample results. Additional unit root tests were performed to test whether traditional fundamentals are nonstationary in first differences. The test results indicate that the series have one unit root at most. The only problem is the Japan-US relative output series, for which we could not reject the null hypothesis of second unit root for any sample periods.

We also checked the order of integration of the interest rate and inflation rate differentials by using the same test statistics as for the traditional fundamental series. Interestingly, there is some evidence of unit-root behaviour for the interest rate differentials between the analyzed countries. For three of the cases, namely Germany-US, Japan-US, and Canada-US we could not reject the unit root in interest rate differentials for the full sample period at five percent significance level. According to the ADF unit root tests, also the inflation rate differentials seem to be unit root processes in many cases. However, since the test statistics are quite close to five percent critical level and it is a common knowledge that inflation rate series includes structural breaks during the sample period that might affect unit root test results. We conclude for subsequent analysis that the inflation rate differential series are stationary series[47].

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[46] There is probably a small sample problem in analysing unit roots for the sub-sample periods. A major criticism of the ADF unit root test is that it cannot always distinguish between unit-root and near unit-root processes in short samples.

[47] We are aware of the fact that transition variables' stationarity is a crucial aspect for appropriate transition regime switching modelling (see e.g. van Dijk and Franses 2001).

In summary, we conclude that all our long-run model variables are unit root processes in levels and transition variables can be considered as stationary variables. Next, we test the flexible-price monetary model using the Johansen (see Johansen 1995 for a more detailed account of this methodology) cointegration methodology and try to find the best possible linear long-run attractor for each case.

4.3.3 Linear cointegration tests and long-run relations

Based on unit root tests in section 4.3.2 and the model presented in section 4.2.1 we next estimate a standard linear vector-error correction model (VECM) by using the Johansen full-information maximum likelihood cointegration methodology, i.e.,

$$\Delta X_t = \mu + \sum_{i=1}^{p} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + u_t \text{ [48]}, \tag{4.11}$$

where $X_t = [e_t, m_t - m_t^*, y_t - y_t^*]$ is a column vector of variables in the long-run model, Γ_i's are parameters, and $\Pi = \alpha\beta'$ is $m \times m$ matrix of unknown parameters, where α and β are $m \times r$ matrices representing the rate of reversion and cointegrating parameters for the system. The error term u_t is $NID(0,\Sigma)$ and μ is a vector of constants. We set the lag length equal to four as suggested in Engel and West (2005) for the same data set.

Tables 10 - 12 present our cointegration test findings for both the full-sample and sub-sample analyses. Johansen (1995) proposed two test statistics for inferring the number of cointegrating vectors. The trace statistics is used to test the null hypothesis of at most r cointegrating vectors against the alternative of m cointegrating vectors. The maximal eigenvalue statistics is used in testing the null hypothesis of $r - 1$ against r cointegrating vectors. Both Johansen test statistics indicate the non-existence of a unique cointegration vector in most of the examined cases for the full sample periods[49]. The only exceptions are UK-US case, for which both test statistics indicate at least one cointegrating vector, and the Italy-US case, for which the Johansen max-statistic shows cointegration. For each of the sub-periods we can detect cointegration more often, but we can also see that there are significant changes from cointegration to no cointegration and vice versa during the sample periods. For example, in the Germany-US case we cannot detect cointegration for full sample period for the period 1974-1990, but for the period 1990-1998 there is very significant evidence of cointegration in this particular system. The available time series are relatively short for analysing reliably the sub-sample data, but we can see some evi-

[48] Using Monte Carlo-simulations van Dijk and Franses (2000) have shown that the standard cointegration tests based on linearity can be used to test for the presence of cointegration and to estimate the corresponding cointegrating parameters in nonlinear error-correction models. Specifically, they observe that the bias in estimating the cointegration rank and the cointegrating parameters is not larger for asymmetric or nonlinear adjustment than it is for linear symmetric adjustment. Their findings are in line with the theoretical considerations of in Escribano and Mira (1996) and Corradi, Swanson and White (1995). Especially, Escribano and Mira show that cointegrating parameters can be estimated superconsistently even if there are nonlinearities in the adjustment process.

[49] We have made the small sample corrections for all our cointegration test statistics.

dence on how the structure of cointegration changes during the sample period. This is more evident if we look at the structure of cointegrating vectors.

The cointegrating vectors for all country pairs are presented in Tables 10-12 under the assumption that all models have one significant cointegrating vector. If we cannot detect cointegration for any sample period or if the cointegration structure changes during the sample period, we take that as a possible evidence of nonstability in the exchange rate and macrofundamentals relation (see De Grauwe and Grimaldi 2005). This is also in line with the findings of Goldberg and Frydman (2001), where they argue that different sets of macroeconomic fundamentals matter during different time periods, and hence it might be almost impossible to find a stable and statistically significant cointegrating vector for all the sample periods. However, if the traditional fundamentals affect exchange rates we should be able to present some empirical link between them. In this paper we assume that there is a linear long-run cointegration relation with nonlinear short-run dynamics as will be shown empirically later in the chapter.

The interpretation of our assumed cointegration vector in Tables 10-12 is an important task, and it is possible to state the restriction hypothesis based on the economic theory and try to test whether these restrictions are statistically relevant in the model. In the typical monetary model, a logical way to proceed is to put the theoretically consistent parameter estimates into the empirical model and try to identify the model based on these long-run parameter values. We proceed by estimating the most general linear long-run model with one cointegrating vector in all cases. Following Mark and Sul (2001) we then examine whether the theoretical model restrictions $\beta_1 = 1$, $\beta_2 = -1$ or $\beta_1 = -\beta_2 = 1$ are valid in our sample sets. As we can see from Tables 10-12 we can only state that for one case, namely the France-US set we may accept the theoretical null hypothesis for all sample periods. For the full sample period we could also accept the null hypothesis for the Germany-US case. However, this parameter estimate structure completely disappears when we examine only subperiods. For the cases of UK-US and Italy-US we can accept the null hypothesis of $\beta_1 = 1$ both for the full sample period and for the first subperiod. For most of the other cases we can see that the structure of parameter estimates changes dramatically between samples. There are several possible reasons why the parameter structure might change during the sample period. Perhaps the most important is that there is no reason to believe that foreign exchange rate market participants use the information in the same way over all policy regimes during the sample period (see more on this in De Grauwe and Grimaldi 2005).

Table 10. Cointegration tests for a full sample period (1974:1 –2001:3)

$e_t = \beta_0 + \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*)$

Country	Trace	Max	β_1	β_2	$\beta_1 = 1$	$\beta_1 = -\beta_2 = 1$
Germany	13.64	7.33	-0.04	0.38	1.01 (0.32)	1.60 (0.45)
UK	*33.51*	*19.30*	8.06	60.34	1.84 (0.18)	*14.61 (0.00)*
Japan	21.55	15.85	-21.98	-63.07	*9.75 (0.00)*	*10.25 (0.01)*
France	18.48	13.83	0.80	0.55	0.14 (0.70)	2.72 (0.26)
Canada	19.95	14.10	0.19	-4.85	*5.36 (0.02)*	*8.51 (0.01)*
Italy	29.73	*16.67*	0.74	0.98	0.24 (0.63)	*9.33 (0.01)*

Table 11. Cointegration tests for a sub-sample period (1974:1 –1990:2)

$e_t = \beta_0 + \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*)$

Country	Trace	Max	β_1	β_2	$\beta_1 = 1$	$\beta_1 = -\beta_2 = 1$
Germany	29.33	16.52	-2.09	3.85	*6.64 (0.01)*	*7.30 (0.03)*
UK	*38.13*	*23.23*	1.08	3.87	0.05 (0.82)	*11.39 (0.00)*
Japan	28.08	*18.78*	-3.31	-19.94	*7.70 (0.01)*	*10.09 (0.01)*
France	26.98	11.62	-2.47	-1.00	0.09 (0.70)	1.07 (0.58)
Canada	31.87	16.92	-1.20	1.33	*6.31 (0.01)*	*9.69 (0.01)*
Italy	*36.01*	*17.29*	0.18	2.15	1.71 (0.19)	*9.95 (0.01)*

Table 12. Cointegration tests for a sub-sample period (1990:3 – 2001:3)

$e_t = \beta_0 + \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*)$

Country	Trace	Max	β_1	β_2	$\beta_1 = 1$	$\beta_1 = -\beta_2 = 1$
Germany	*60.68*	*30.67*	0.56	-2.02	*4.88 (0.03)*	*14.05 (0.00)*
UK	*44.38*	*24.45*	-0.60	-4.56	*4.82 (0.03)*	*24.90 (0.00)*
Japan	26.40	17.92	1.05	2.75	0.08 (0.78)	*15.67 (0.00)*
France	30.07	19.34	4.38	9.97	1.27 (0.26)	4.40 (0.11)
Canada	25.30	17.87	0.29	2.78	*4.05 (0.04)*	*14.54 (0.00)*
Italy	37.02	*20.09*	-0.62	-3.74	*5.87 (0.02)*	*6.76 (0.03)*

Notes: We are testing for a single cointegrating vector in all cases. The italics indicates significance level at 10 per cent level for cointegration tests. The finite sample critical values for Johansen cointegration tests were obtained by adjusting the asymptotic critical values for the loss of degrees of freedom due to estimation of the parameters describing the short-run dynamics. The adjustment factor is given by *(T-p*m)/T*, where *T* is the sample size, *p* is the lag length and *m* is the number of variables. The number of lags were selected to 4 in all cases.

We should note that the results in Tables 10-12 do not indicate for which variable(s) the cointegration relationship might apply. By examining the long-run weak exogeneity in Tables 13-15 under the assumption of one significant cointegrating vector for every country pair, we can find out which variables can be considered as driving variables of the system. For the full sample period the exchange rate is found to be weakly exogenous variable in three cases, namely for Germany, UK, and Japan. Thus, for these countries the exchange rate against the US dollar is not adjusting in a linear cointegration models. This is an interesting result, since the above four currencies were most traded currencies during the sample period 1974-1998. The results of weak exogeneity are totally different for the cases of France, Canada, and Italy, where the exchange rates seem to be adjusting variables of the systems for the full sample periods. When analysing subperiods we can see that the property of the weak exogeneity in the exchange rates changes in many cases when we move from one sub-period to another. Thus, even if macroeconomic fundamentals and exchange rates might form a relatively stable linear long-run relationship, the way that this relationship matters in short-run dynamics might change considerably during the sample period. We argue that this is also significant evidence of possible nonlinear dynamics in the error correction structure. These findings are also consistent with Engel and West (2005), where they argue that it is not necessarily the exchange rates that are adjusting variables in the monetary models.

Table 13. Tests of long-run weak exogeneity for sample period 1974:1 – 2001:3

Country	Δe_t	$\Delta(m_t - m_t^*)$	$\Delta(y_t - y_t^*)$
Germany	1.50	2.24	0.02
UK	0.16	0.28	*10.17*
Japan	0.53	*4.91*	3.67
France	*9.07*	0.07	1.59
Canada	*4.76*	*4.21*	0.21
Italy	*3.94*	0.10	1.67

Table 14. Tests of long-run weak exogeneity for sample period 1974:1 – 1990:2

Country	Δe_t	$\Delta(m_t - m_t^*)$	$\Delta(y_t - y_t^*)$
Germany	*4.80*	*4.23*	0.01
UK	0.42	*8.72*	*7.03*
Japan	0.55	*7.63*	2.34
France	0.47	0.49	0.10
Canada	0.31	*4.89*	2.00
Italy	3.32	1.11	0.02

Table 15. Tests of long-run weak exogeneity for sample period 1990:3 – 2001:2

Country	Δe_t	$\Delta(m_t - m_t^*)$	$\Delta(y_t - y_t^*)$
Germany	0.00	*6.08*	*11.17*
UK	2.00	*4.07*	*6.41*
Japan	*8.66*	0.23	1.12
France	1.14	*5.89*	*4.98*
Canada	*3.87*	*7.75*	3.00
Italy	3.33	1.37	0.02

Notes: Test statistics are LR-tests for the hypotheses H_0^i: $\alpha_{ij} = 0$; $j = 1,\ldots,r$, where H_0^i is the hypothesis that a particular variable *i* does not adjust to the equilibrium errors and *j* is the number of long-run cointegrating vectors. The $\chi^2(1) = 3.84$ is 5 per cent critical value. The italics face indicates significant test statistics.

Unfortunately, the cointegration methods used in this chapter cannot suggest what type of nonlinearities may exist. One way to detect deviations from linearity is to relax the assumption of fixed-coefficient error correction model and examine whether the model's empirical specification is improved. The question then arises whether the nonlinear model can give more reliable results for the exchange rate behaviour. However, before we formulate a nonlinear error correction model it is important to obtain some idea about the parameter values of cointegrating vectors. We should point out also that the finding of cointegration between variables does not necessarily imply a linear error correction model. Before proceeding to the nonlinear estimation we should note that although the maximum likelihood estimates obtained from the Johansen procedure are unbiased, the estimates are also more dispersed than those provided by the alternative estimators, such as OLS (Stock and Watson 1993). Thus, we estimate OLS estimates for those parameters that we could not accept the parameter restrictions implied by the theory in Table 10. We will use these parameter estimates in subsequent analysis and they are presented in Table 16. In Table 16 we also tested, whether OLS formulation of parameter estimates might make the error correction terms stationary by using the ADF test for a regression with constant for the full sample periods. Testing for stationarity of the deviation series provides evidence of long-run comovements between fundamentals and exchange rate series. The results are very much in line with the Johansen cointegration tests in that there is no clear evidence of linear error correction structure. As with the Johansen tests, we could now reject the null of nonstationarity for the UK-US case. Furthermore, we could find possible cointegration also for the UK-US case that was rejected by the Johansen test statistics.

Table 16. The long-run parameter estimates for full sample period 1974:1 – 2001:3

$e_t = \beta_0 + \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*)$

Country	β_0	β_1	β_2	ADF(k)
Germany	0	1	-1	-2.43 (4)
UK	5.80	1	7.62	*-3.31 (11)*
Japan	5.14	0.02	1.68	-1.16 (11)
France	0	1	-1	*-3.76 (8)*
Canada	-0.23	-0.13	2.07	-2.04 (4)
Italy	13.56	1	5.88	-3.10 (4)

Notes: The lag length of ADF statistics is selected by the Akaike Information Criteria (AIC). The italics indicates significance at 5% level.

The results in section 4.4 indicate that the process of how exchange rates and monetary fundamentals affect each other's changes significantly during the sample period. For some sub-periods the macroeconomic fundamentals matter for exchange rate series and for another sub-period exchange rate series matter for behaviour of macrofundamentals. The assumption of weak exogeneity for all the explanatory variables with respect to the exchange rate is also made. We make this assumption, since our main focus is on the exchange rate dynamics. Furthermore, the empirical literature typically summarizes the various economic fundamentals underlying the monetary model to a single reduced form equation for the exchange rate (e.g. Mark and Sul 2001 and Rapach and Wohar 2004).

4.4 Estimated nonlinear error correction models

In this section we estimate a nonlinear extensions of the error correction models for the monetary model[50]. Given the assumption of weak exogeneity, the nonlinear analysis may be conducted for single equation exchange rate presented in equation (4.8)[51]. We specify six smooth transition threshold models, which use the same explanatory variables. We

[50] We are aware of the fact that before we begin to find the best fitting nonlinear model, it is important to test for possible presence of nonlinearity. However, since the linearity of long-run equilibrium is accepted by assumption we should only check the nonlinearity of error correction structure in our model, but this is done convincingly for similar type of data set in several previous studies (see e.g. Taylor and Peel 2000).

[51] It should be noted, however, that Engel and West (2005) show that exchange rates might affect to fundamentals. However, the existing monetary models of exchange rate cannot easily explain this result. Thus, we proceed by assuming that macrofundamentals affect to exchange rate.

have seen previously that the structural change in parameter estimates during the sample period and now we would like to know in which way the estimated long-run equilibrium relation matters for the short-run exchange rate movements. In formulating the nonlinear model specifications we will use the long-run error correction vectors that are taken directly from the Table 17. In moving from the long-run analysis to the short-run analysis, the emphasis shifts from the existence of long-run equilibrium to mean reversion to such equilibrium. The exchange rate deviation from monetary fundamentals can be presented as

$$z_t = e_t - \beta_0 - \beta_1(m_t - m_t^*) - \beta_2(y_t - y_t^*). \tag{4.12}$$

We estimate the parameters of the nonlinear version of the error correction model (4.8) by the maximum likelihood estimation methodology. Table 17 gives the results of time-varying error correction model after removing insignificant variables. We present only those parameter estimates that are significant at five percent critical level. As we can see from table 17, the nonlinear model specifications seem to be relevant in most cases. Only for Japan-USA we could not find any significant nonlinear error correction estimates during the sample period. The first conclusion from table 18 is that there are significant error correction parameters in most cases indicating that traditional macroeconomic fundamentals play a crucial role in the determination of exchange rate dynamics. The transition functions are significant, implying that different macroeconomic conditions (i.e. the change in inflation rate differentials) affect to exchange rate determinants. When different macroeconomic conditions occur, the same explanatory variables may play a very different role. This finding is consistent with the recent literature of the time-varying coefficient model of exchange rates (see Frydman and Goldberg 2001, De Grauwe and Vansteenkiste 2001, and Frömmel et al. 2003). We find, especially, that the error correction terms might be non-significant in one regime, but significant in another. We do not, however, argue that fundamentals do not affect to the exchange rates all the times, but that their effect depends on the inflationary environments of the countries.

We are also interested in the role of the interest rates in determining exchange rate dynamics. Frankel (1979) extended the monetary model of exchange rate by assuming that higher interest rates lead to lower money demand, which implies that higher domestic interest rates are related to an increase of the price of foreign currency. This is in contrast with the more traditional Mundell-Fleming-Dornbusch model, where higher domestic interest rates lead to an increase in capital inflows and, hence, to the appreciation of domestic currency. The estimation results of table 17 show the impact of the interest rate differential on exchange rate changes during the sample period. We find strong evidence of nonlinearity, with asymmetry relating to inflation differentials. We find that the relative interest rate parameter values have similar magnitudes but different signs between extreme regimes. This implies that interest rate variables seem to be relevant only to exchange rate changes when inflation differentials are either very large or very small. This makes sense for the large inflation differentials, since the monetary policy inventions usually occur when there is a risk of inflation instability. Hence, results support both views of interest rate effects to exchange rate dynamics. The interest rate effects to exchange rate changes occur primarily at lags of one to two quarters.

Table 17. Estimated parsimonious nonlinear error correction for the model

$$\Delta e_t = a_0 + \sum_{i=1}^{5} b_i \Delta e_{t-i} + \sum_{j=0}^{4} c_j \Delta f_t + \lambda_1 z_{t-1} + \sum_{k=0}^{4} \mu_k (i - i^*)_{t-k}$$
$$+ \left[a_0^* + \lambda_1^* z_{t-1} + \sum_{l=0}^{4} \mu_l^* (i - i^*)_{t-l} \right] G(s_t; \gamma, d)$$

Country	Germany	UK	France	Canada	Italy
Model	ESTAR	ESTAR	ESTAR	LSTAR	ESTAR
a_0	-0.03	0.00	0.01	0.02	-0.06
b_1	0.24	0.22	0.27		0.41
b_2	-0.11	-0.26	-0.14		-0.40
b_3	0.21	0.19	0.26	0.18	0.25
b_4	0.19		0.13		
b_5		-0.11		-0.11	
c_0		0.76	0.46		
c_1	-0.26	-0.55	-0.39	0.10	-0.61
c_2	-0.41	-0.30			
c_3	0.72			0.10	0.65
c_4	-0.67				-0.44
a_0^*	-0.04		0.23	-0.03	0.04
λ_1	-0.03 (0.05)	0.09 (0.05)	-0.01 (0.03)		-0.08 (0.03)
λ_1^*	-0.10 (0.06)	-0.19 (0.06)	-0.10 (0.04)	-0.68 (0.93)	-0.02 (0.05)
μ_0	4.65	1.69	1.87		-0.39
μ_1	-3.75	-1.21	-1.26	0.92	-0.34
μ_3					-0.41
μ_0^*	-4.44	-1.85	-1.81	-2.04	0.97
μ_1^*	3.72	1.35	1.05		
μ_3^*			-0.54		-0.54

Table 17 (Continued)

γ	0.15 (0.05)	0.16 (0.10)	0.84 (0.24)	0.40 (0.15)	0.06 (0.05)
d (%)	-1.73	-1.44	0.85	3.91	0
Statistics					
V_{NL}/V_L	0.82	0.84	0.63	0.75	0.78
JB	0.38 (0.83)	1.52 (0.47)	0.25 (0.88)	0.24 (0.89)	1.81 (0.40)
LR-test:	22.67	10.48	35.62	22.45	14.30
$\lambda_1 + \lambda_2 = 0$	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
ARCH(1)	1.03 (0.31)	1.47 (0.23)	0.07 (0.93)	0.02 (0.88)	1.27 (0.26)
ARCH(4)	1.26 (0.87)	1.93 (0.75)	0.76 (0.94)	0.75 (0.95)	6.18 (0.19)
Q(1)	<0.001 (0.99)	<0.001 (0.96)	<0.001 (0.93)	0.05 (0.83)	0.01 (0.92)
Q(4)	0.34 (0.98)	1.23 (0.87)	0.76 (0.94)	1.09 (0.90)	0.24 (0.99)

Notes: Estimation is by maximum likelihood. Figures in parentheses are estimated standard errors. V_{NL}/V_L is the ratio of estimated STR errors correcton model variance relative linear error correction model. *JB* is Jarque-Bera test for residual normality, LR-tests test the significance of error correction term, *ARCH* is a test for autoregressive heteroskedasticity, and *Q* is the Ljung-Box test of residual autocorrelation.

Table 17 also has diagnostic tests for our model. Tests are performed for normality, ARCH effects and residual autocorrelation. The errors appear to be normal and there seem not to be conditional heteroscedasticity or autocorrelation. For each estimated models we examined the restriction that $\lambda_1 + \lambda_1^* = 0$, which implies no significant error correction model. We could reject the null no cointegration for all cases in a nonlinear framework.

In sum, the nonlinear specification comprises two extreme regimes, in which the error-correction term is driven by two different error correction coefficients. Moving from the first regime ($G = 0$) to the second regime ($G = 1$) changes the error correction parameters to be significant for several cases. However, the transition speed depends on the transition variable. We can see for some cases very fast transition speeds and for some countries the transition speeds are much more slower. In Figure 15 we have plotted the time varying error correction terms. The error-correction parameters move sharply from large negative values to small negative or positive values. This finding might explain, at least partly, the difficulties of linear error correction models to identify error correction in certain periods.

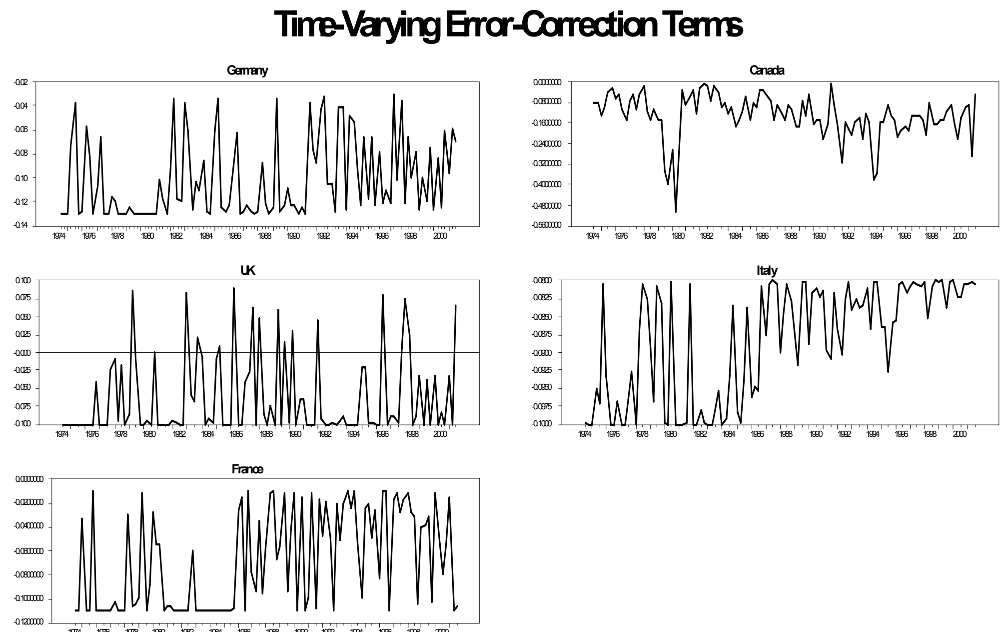

Fig. 15. Time-varying error correction parameters

The nonlinearity of the models are revealed by the plots of the estimated transition functions against the transition variables displayed in appendix 2 (Figures 16-20). In the upper right part of the plots we highlight the effect of the transition speed. We can see a slow speed of adjustment when the transition function, i.e., inflation differential, takes small values. Broadly speaking, the nonlinear error correction mechanism is only operating when the inflation differentials are sufficiently high. For very low inflation differentials, the error correction mechanism is only operating through the linear part of the model. As we can see also from figures, the STAR specification seems to be quite appropriate, since we have well-behaving transition functions for almost all cases. Furthermore, we can see the graphs of residual plots and model fits. In the upper left part of the figures we show fitted values of nonlinear error correction model and actual exchange rate series changes. As we can see, the models capture quite well the actual exchange rate series. The model does not only capture the changes in the actual data, but the nonlinear model also captures the long swings in the exchange rate changes. This is useful in examining the US dollar based exchange rate series, since there are long swings in the data during the 1980's. In lower left part we have presented the residual series. The lower right panels show how the transition variables evolve over time. The model dynamics change during the sample period and that nonlinear dynamics is very relevant explanation for the behaviour of exchange rate in 1980's.

An important question is, whether the nonlinear function form dominates linear error correction model. Statistical comparison across linear and nonlinear specification is not, however, a straightforward task. We address this question by examining both residual variances and significance of parameter estimates. The residual variances are much more lower for the nonlinear specifications. The results of the best fitting linear models give also a little support for the possible error correction mechanism for full sample period, since there is no cointegration in Tables 10-12 and, hence, no significant error correction mechanism for several cases. By contrast, the nonlinear error correction models are able to reject the null hypothesis of no cointegration at conventional risk levels in most of the cases.

4.5 Conclusions

Explaining the empirical behaviour of exchange rates has been a difficult task. In this chapter we have re-examined the flexible price monetary model by using a model, which includes a parametric nonlinear error correction presentation. We have built a time-varying error correction model, based on traditional fundamentals, short-term interest rates, and relative inflation rates in order to examine the G7 countries' exchange rates during the post Bretton Woods period. We assume the presence of linear long-run cointegration vector and nonlinear adjustment towards equilibrium level. We postulate that traditional macroeconomic fundamentals are important in driving the exchange rates. However, the underlying structure between fundamentals and exchange rates is nonlinear. The nonlinearity is presented as nonlinear exchange rate dynamics and the inflation differential is the threshold variable that governs this nonlinearity.

Our empirical evidence supports the hypothesis that economic fundamentals are significant driving forces of exchange rates. For several cases we can reject the random walk hypothesis. The nonlinear error correction model specification performs better than the linear error correction model. Particularly, the error correction parameters are almost always significant in nonlinear models. This result is totally opposite to results that we were able to find in analysing linear cointegration models, where we could only rarely find error correction mechanisms. The estimates also show that, when different macroeconomic environments occur, the same variable in the exchange rate model may have a different weight. Thus, in examining exchange rates it is crucial to recognise particular countries macroeconomic conditions.

An important finding is that by assuming nonlinearities in error-correction mechanism we are able to detect that macroeconomic fundamentals affect to exchange rates. This is in line with the existing monetary models of exchange rate suggestion for causality. However, in a recent paper by Engel and West (2005) it is suggested that exchange rate may also affect macrofundamentals. One possible area of future research is, thus, to examine nonlinearities in a multivariate cointegration context. In that way, it would perhaps be possible to find causality in both directions.

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Appendix 1 Table

Table 18. Tests of I(1) for series

Variable	ADF(k)	ADF(k)	ADF(k)
Germany	1974:1 – 1998:4	1974:1 – 1990:2	1990:3 – 1998:4
e	-2.21 (4)	-1.84 (4)	-1.04 (0)
m – m*	0.36 (1)	-0.99 (0)	0.09 (2)
y – y*	-2.18 (4)	-1.25 (4)	-0.98 (1)
i – i*	-2.29 (7)	-1.59 (2)	-1.54 (1)
$\pi_t - \pi_t^*$	-2.27 (3)	-1.73 (4)	-2.24 (4)
Δe	*-3.88 (3)*	*-6.79 (0)*	*-7.08 (0)*
Δ(m-m*)	*-5.88 (0)*	*-3.73 (1)*	*-3.11 (0)*
Δ(y-y*)	*-3.53 (3)*	*-9.62 (0)*	*-4.71 (1)*
Variable	ADF(k)	ADF(k)	ADF(k)
Japan	1974:1 – 2001:3	1974:1 – 1990:2	1990:3 – 2001:3
e	-1.64 (4)	-0.76 (0)	-2.55 (3)
m – m*	-1.65 (4)	-0.65 (1)	-1.79 (4)
y – y*	-1.64 (10)	-0.58 (10)	-0.75 (0)
i – i*	-2.39 (20)	*-3.30 (0)*	-0.87 (0)
$\pi_t - \pi_t^*$	*-4.59 (19)*	*-3.70 (10)*	-2.70 (10)
Δe	*-4.57 (4)*	*-6.57 (0)*	*-3.75 (0)*
Δ(m-m*)	*-4.00 (1)*	*-6.36 (1)*	*-2.91 (0)*
Δ(y-y*)	-1.88 (9)	-2.28 (11)	-1.53 (11)
Variable	ADF(k)	ADF(k)	ADF(k)
UK	1974:1 – 2001:3	1974:1 – 1990:2	1990:3 – 2001:3
e	-2.06 (11)	-1.99 (8)	-2.13 (0)
m – m*	-1.22 (3)	-0.73 (6)	-1.26 (2)
y – y*	-1.29 (19)	*-3.08 (16)*	-1.52 (7)
i – i*	*-4.19 (0)*	-2.00 (4)	-1.96 (0)
$\pi_t - \pi_t^*$	-2.48 (16)	-2.49 (4)	*-4.13 (10)*
Δe	*-5.01 (2)*	*-3.16 (2)*	*-6.46 (0)*
Δ(m-m*)	*-3.85 (1)*	*-4.04 (1)*	-2.81 (0)
Δ(y-y*)	*-5.36 (7)*	-2.46 (15)	*-7.13 (0)*

Table 18 (Continued)

Variable	ADF(k)	ADF(k)	ADF(k)
France	1974:1 – 1998:4	1974:1 – 1990:2	1990:3 – 1998:4
e	-2.51 (4)	-2.09 (4)	-1.06 (0)
m – m*	-2.11 (7)	-2.01 (0)	*-3.24 (4)*
y – y*	-0.37 (14)	-1.35 (3)	-1.15 (11)
i –i*	*-2.91 (5)*	-2.74 (2)	-1.14 (0)
π_t -π_t^*	-1.88 (12)	-1.31 (12)	-1.79 (7)
Δe	*-4.93 (2)*	*-6.33 (0)*	*-3.77 (2)*
Δ(m-m*)	*-3.89 (1)*	*-6.76 (1)*	*-3.17 (0)*
Δ(y-y*)	*-5.85 (1)*	*-4.70 (1)*	*-5.72 (0)*

Variable	ADF(k)	ADF(k)	ADF(k)
Canada	1974:1 – 2001:3	1974:1 – 1990:2	1990:3 – 2001:3
e	-1.24 (4)	-1.86 (3)	-0.71 (3)
m – m*	0.25 (2)	-0.68 (0)	-0.79 (7)
y – y*	-1.10 (4)	*-2.88 (1)*	*-3.10 (0)*
i – i*	-2.82 (3)	-2.06 (3)	-2.24 (0)
π_t -π_t^*	*-6.24 (0)*	*-4.34 (0)*	*-5.90 (0)*
Δe	*-4.68 (2)*	*-3.48 (2)*	*-4.84 (1)*
Δ(m-m*)	*-4.79 (1)*	*-8.29 (0)*	*-3.09 (2)*
Δ(y-y*)	*-11.24 (0)*	*-8.47 (0)*	*-7.51 (0)*

Variable	ADF(k)	ADF(k)	ADF(k)
Italy	1974:1 – 1998:4	1974:1 – 1990:2	1990:3 – 1998:4
e	-2.09 (4)	-2.16 (4)	-0.93 (0)
m – m*	*-2.88 (8)*	*-4.12 (0)*	-1.40 (3)
y – y*	0.30 (5)	-1.04 (0)	-0.32 (0)
i – i*	*-4.27 (0)*	-1.57 (6)	-1.12 (0)
π_t -π_t^*	-0.56 (19)	-0.88 (16)	-0.56 (7)
Δe	*-5.06 (2)*	-2.49 (5)	*-3.95 (2)*
Δ(m-m*)	*-7.28 (0)*	*-3.29 (1)*	-2.01 (2)
Δ(y-y*)	*-4.77 (4)*	*-5.36 (4)*	*-6.77 (0)*

Notes: The lag length of ADF statistics is selected by Akaike Information Criteria (AIC). The italics indicates significance at 5% level.

Appendix 2 Figures

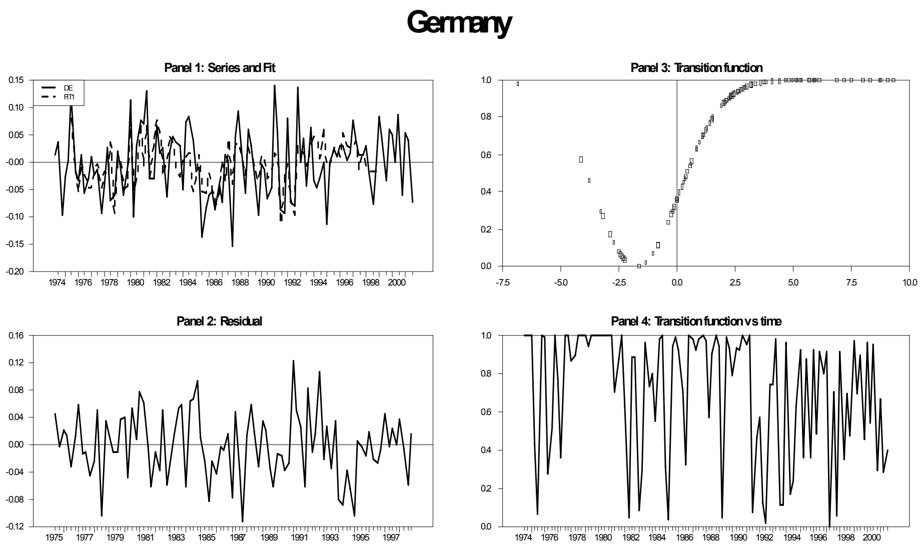

Fig. 16. The plots for Germany

UK

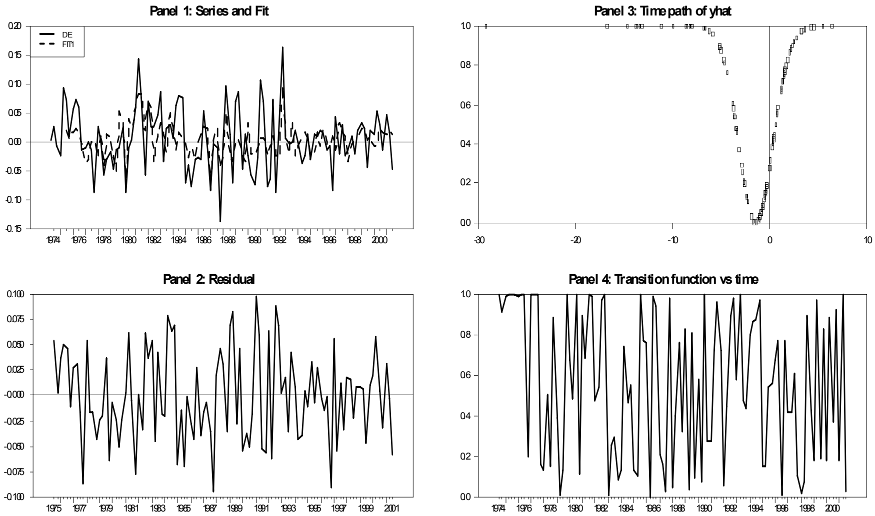

Fig. 17. The plots for the UK

France

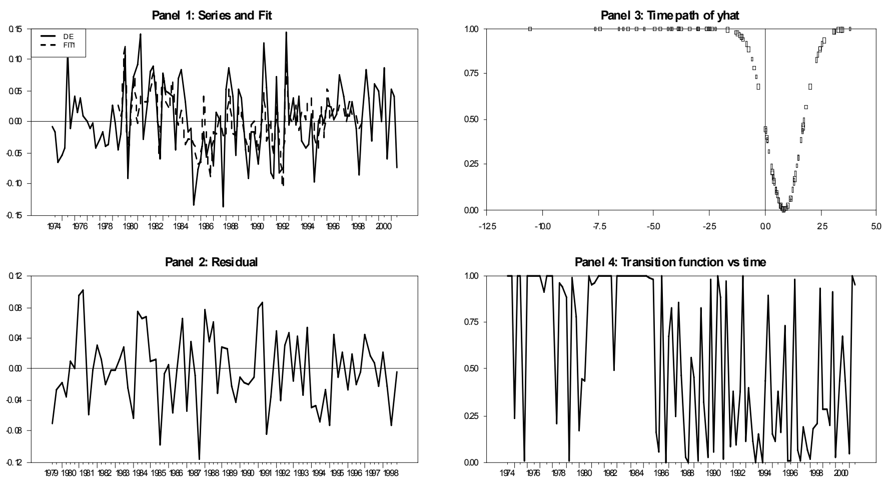

Fig. 18. The plots for France

Canada

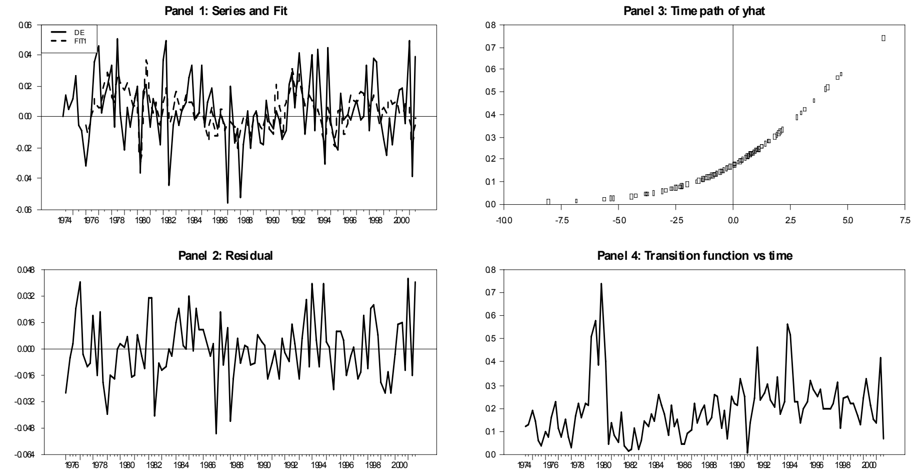

Fig. 19. The plots for Canada

Italy

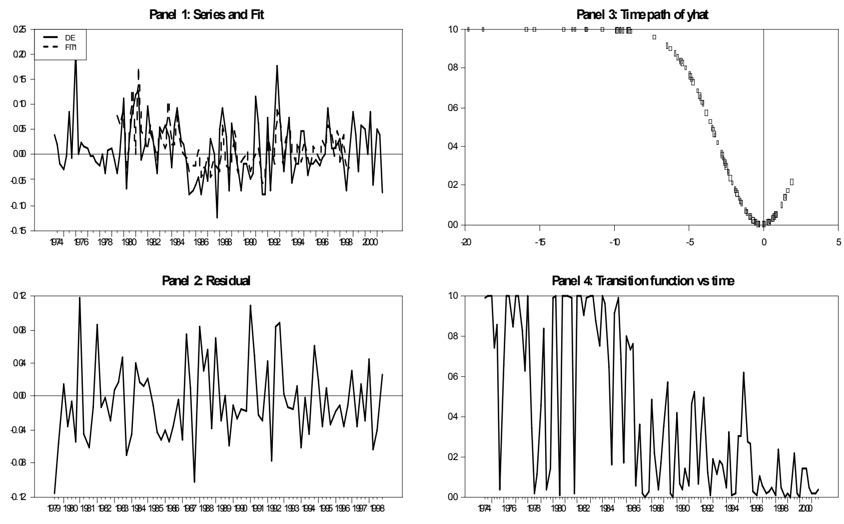

Fig. 20. The plots for Italy

5 Exchange rate pass-through into prices: Dependence on inflation regime

5.1 Introduction

One of the key issues in international economics is how exchange rate movements are transmitted into prices. The literature on this topic can be broadly categorized into two main strands. The first strand examines price discrimination or pricing to market (hereafter PTM)[52] across destination markets and analyzes the exchange rate change pass-through[53] to consumer prices (see e.g. McCarthy 2000), aggregate import prices (see e.g. Campa and Goldberg 2002) and import prices of a specific industry (see e.g. Goldberg and Knetter 1997)[54]. This strand has shown that exchange rate pass-through is not full and it varies both across and within industries. A stylized fact is that in many industrialized countries the pass-through of exchange rate changes is less than 100%. A median pass-through of currency depreciation to import prices is around 50% over a one-year horizon (Golberg and Knetter 1997)[55]. The second strand is closely related to the new open economy macroeconomics literature (see Obstfeld and Rogoff 1995, Lane 2001 and Sarno 2001, among others), which analyzes the situation of imperfect competition in the open economy macroeconomic models. It allows several forms of price and/or wage rigidity within fully specified dynamic general open economy macromodels. An incomplete exchange rate pass-through is usually interpreted as a combination of both pricing to market behaviour and sticky prices. The literature also divides firms' pricing behaviour as a producer currency pricing (hereafter PCP)[56] and local currency pricing (hereafter LCP) depending on in which currency firms set their export prices (see e.g. Betts and

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[52] Pricing to market is defined as the ability of firms to charge different prices for an identical good in different countries. For example, when the exchange rate changes, a firm may choose to pass the exchange rate completely into its selling prices or to keep the selling prices unchanged or use some combination of these both effects.

[53] Exchange rate pass-through is broadly defined as the degree to which exchange rate changes are reflected into the destination country local currency prices.

[54] See also Knetter (1989, 1993), among many others.

[55] See also Obstfeld (2002).

[56] For example, Obstfeld and Rogoff (1995, 2000) are good surveys for PCP models.

Devereux 1996, 2000)[57]. The PCP model has perfect exchange rate pass through, which implies that exchange rate fluctuations are the adjusting variables for imbalances in the economy and foreign good prices move one-for-one with changes in the nominal exchange rate. The LCP model has limited exchange rate pass through and there is only a limited expenditure switching effect between domestic and foreign goods. In the LCP models, the exchange rate changes have no short-run effects on consumer prices.

The recent literature of exchange rate pass-through suggests that the pass-through might be dependent on the monetary policy environment of importing country (see Taylor 2000, Choudhri and Hakura 2001, Campa and Goldberg 2002, and Devereux and Yetman 2003 among others). The pioneering paper is Taylor (2000), who argues that exchange rate pass through has declined in recent years as a result of a low inflation regime in industrialized countries. According to him a high inflation environment would increase exchange rate pass-through and a credible low inflation regime will decrease the pass-through. He shows that the pass-through is positively related to the rate of inflation, which suggests that low inflation itself might have caused the low pass-through. This decline in pricing power seems to be very closely related to the decline in consumer price inflations in many countries. Taylor explains that firms find it difficult to fully pass-through exchange rate changes to their export prices, because of intensified worldwide competitive pressure and low and stable inflation regimes that have been seen in industrialized countries in recent years. This is also an empirical fact, since for example in the US the inflation did not rise even if there were strong demand pressures in the late 1990’s. Furthermore, the US import prices did not raise much in the late 1980’s despite the fact that the US dollar depreciated drastically. Similar empirical evidence can also be found in papers that examine the pass-through for Japan (see Otani et al. 2003). Also, the formation of monetary union in Europe might have caused exchange rate pass through to decline, since it has caused a structural shift in the participating countries towards lower inflation rates. Moreover, as Devereux et al. (2003) have argued the single currency in euro zone will cause European consumer prices to be more insulated from exchange rate volatility than they have been before the common currency.

The recent empirical evidence has shown that exchange rate movements in the 1990’s and 2000’s seem to have less impact on import prices than they had in the 1970’s or 1980’s[58]. Gagnon and Ihrig (2001) and Choudhri and Hakura (2001) find that the decline in the pass-through is significantly related to the variability of inflation in several OECD countries. Moreover, Engel (2002) argues that the exchange rate pass-through is more related to the macroeconomic environment, such as inflation regime, rather than to microeconomic environment, such as imperfectly competitive firms. However, the evidence for pass-through and monetary policy regime relation is not conclusive. For example, Campa and Goldberg (2002) find that microeconomic factors, such as composition of import might be a more important factor than the inflation regime in determining pass-through. However, their results also show that exchange rate pass-through to import prices has declined significantly in the 1990’s. Furthermore, McCarthy (2000) finds that

[57] See also Chari et al. (2002)

[58] However, the empirical research has concentrated mostly on industrialized countries with floating exchange rates. Furthermore, several industrialized countries have used either explicit or implicit inflation targets after 1980s.

the pass-through tends not to be related to the inflationary experience in examining the pass-through to consumer prices for several countries. He estimates vector autoregreressive models for a group of industrialized countries and finds that the exchange rate pass-through to importing country inflation is larger in countries with larger import shares.

The exchange rate pass-through has been in a central role in debates over monetary policy and defining the optimal exchange rate regime. First, it is important to know how previous exchange rate changes affect on current consumer prices and how the exchange rate movements will affect central banks inflation forecasts. In this way the knowledge of the degree of the pass-through is crucial for conducting appropriate monetary and/or exchange rate policy. The question is important, since if pass-through is due to sticky prices, then pass-through is probably dependent on monetary policy as suggested by Taylor (2000). Second, there is also a direct effect to deciding how much central banks should tighten the monetary policy in response to exchange rate changes (Ball 1999). He argues that monetary policy rules are dependent on the degree of pass-through. An important question is, whether looser monetary policy (i.e. causing higher inflation), might cause also a higher rate of firms' price changes. There is now clear evidence that stabilizing role of flexible exchange rates might not be so strong as the traditional view assumes, when producers can set their prices in the consumer's currencies. If the monetary policy objective is price stabilisation, then it is very important to know the extent to which domestic consumer prices are influenced by exchange rate fluctuations. Third, there is clear evidence that exchange rate fluctuations are a result from the combination of monetary shocks and sticky prices (see e.g. Chari et al. 2002). The recent literature has shown that these effects cause much larger exchange rate fluctuations that the effects based on the more traditional view. Obstfeld and Rogoff (2000) and Obstfeld (2002) have shown that with the PCP it is not desirable to target exchange rates. The optimal policy can be achieved without any coordination between monetary authorities (Engel 2002). The situation is very different for the LCP case, since exchange rate fixing is the optimal monetary policy in the LCP environment as was shown in Devereux and Engel (2003). Thus, the effectiveness of monetary policy depends crucially on whether there is high or a low degree of exchange rate pass-through. The distinction between LCP and PCP is also an important question for the model structure, since most of the new open economy macroeconomic models assume only either one of these pricing types[59]. Hence, an important empirical question is which one of these two price assumptions should be followed in further studies.

There is some evidence that the pass-through from exchange rate changes to consumer prices might be nonlinear (Devereux and Yetman 2003). Devereux and Yetman examine several economies and find that there is a positive, but nonlinear relationship between exchange rate pass through and average inflation rate[60]. They find that when inflation raises then the exchange rate pass-through into import prices also increases in a nonlinear way. Their result is also supported by data. The nonlinearity in their paper comes from the finding, that when annual inflation rises above some threshold rate, there should be no further impact into domestic import prices, since all prices adjust continuously and hence, the exchange rate pass through is complete. Furthermore, some factors in the new open

[59] In the most of the new open economy models, there is no partial exchange rate pass-through to prices.

[60] They also find a positive relation between the pass-through and exchange rate volatility.

economy macroeconomic models support to nonlinear response to exchange rate shocks. A small change in exchange rates may leave the product prices unchanged in importing country currency terms due to, for example, menu costs or contract renegotiation costs. However, large and persistent shocks may trigger price adjustment for several firms and in the aggregate level this can be seen as a change in the pass-through elasticises.

We extend the recent pass-through literature in two main ways. First, we present a model that adds the importing country inflation regime as an explanation for the pass-through coefficient. Specifically, we assume that exporting firms use the present annual inflation rate as an indicator of future pass-through for their products. In this way, we will assume that the mark-up is time-varying rather than a constant. Thus, we can add time dependence also into the exchange rate pass-through elasticises. Our model is a highly conventional mark-up pricing model but nevertheless it includes features that are essential to implement nonlinear empirical analysis of exchange rate pass-through. Our proposed model suggests that a reduction in the average inflation rate alters the impact of exchange rate pass-through to prices. Specifically, the model shows how inflation regime affects to exchange rate pass through into import prices in a nonlinear way and the mark-up in our model will be dependent on the inflation regimes. Second, we will use nonlinear smooth transition estimation methodology (see Granger and Teräsvirta 1993 and Teräsvirta 1998 among others) to estimate the pass-through coefficients. To our knowledge, there is no study in the pass-through literature, which has used smooth transition models in examining the pass-through coefficients. Several previous studies have used a two-stage estimation procedure in estimating nonlinearity of pass-through elasticises[61]. In the first stage, these procedures are used to estimate the elasticity coefficients directly by using linear regression models. In the second stage, the possibility of nonlinearity is taken into account by using the previous stage coefficient as a dependent variable and inflation rates and/or inflation rate squares as an independent variables. We use one stage nonlinear estimation procedure that nests also a linear model specification for a full sample analysis. In this way we may obtain more robust parameter estimates for the pass-through coefficients.

We find that a positive but nonlinear relation between price changes and exchange rate pass-through for several OECD countries. Second, we find that consumer prices are rather insensitive to exchange rate changes. This implies that at a consumer level, the good prices are set in local currencies. For import prices, we find partial exchange rate pass-through for most of the sample countries. We also find that in both the consumer price and import price level, our pass-through coefficients do not vary as much across countries as many previous studies have shown.

The remainder of this chapter is organized as follows. In section 5.2 we give the theoretical and empirical model frameworks of exchange rate pass-through used in this paper. Section 5.3 gives the description of the data and basic statistics. In section 5.4 the main

[61] Several papers have used sub-sample estimation in examining pass-through changes after 1990. In these papers the break point is determined either exogenously or endogenously by some statistical procedure (see Campa and Goldberg 2002, among others). However, these sub-sample estimations have a serious small sample problem. Also, in sub-sample estimations one must assume that similar data generating process (DGP) is valid for both sub-periods.

estimation results are presented. Section 5.5 provides further discussion of results and examines the implications for monetary policy. Finally, section 5.6 concludes.

5.2 Theoretical framework

5.2.1 The mark-up model

In the model of perfect competition, a profit maximizing exporter with prices set in importing country currency will set its price equal to marginal cost,

$$P_j = EC_j^*, \tag{5.1}$$

where P_j is the local currency price of good j, Cj^* is exporter's marginal cost of good j production in its own currency, and E is the domestic currency price of foreign exchange. In this environment, the exchange rate movements pass-through is unity, or put it in other words, the elasticity of foreign currency price with respect to one unit exchange rate will be one (i.e. $\partial \ln P_j / \partial \ln E = 1$). Local currency import prices then fully respond to exchange rate changes. However, in the environment of internationally traded goods we must relax the assumption of perfect competition in many cases. In this case the exporter's profit maximization condition includes also a mark-up over marginal cost. With mark-up pricing the exporter firm will set its price in importing country currency as

$$P_j = \lambda EC_j^*, \tag{5.2}$$

where the mark-up λ is a function of the elasticity of demand. the equation (5.1) is a special case of equation (5.2), if the elasticity of demand is infinite. We should note that the mark-up coefficient in equation (5.2) would also depend on competition pressures and market conditions. The important point to note is that in equation (5.2) the exchange rate pass-through is not complete.

We consider a typical foreign exporting firm that has some degree of pricing power of its goods in importing country. We will assume that the firm sets its good export price in its own currency as a mark-up over its marginal costs of production ($C_{t,j}^*$) so that equation (5.2) holds. We will extend the model by assuming that the mark-up is time-varying in (5.2)[62]. Thus, we can write the importing price (5.2) at time t as

$$P_{t,j} = \lambda_{t,j} E_t C_{t,j}^* . \tag{5.3}$$

The mark-up (λ) is assumed to respond to competitive pressures in export market, demand pressures in importing country, and monetary policy conditions in the importing country. Specifically, we can write the mark-up in a functional form as

[62] See Kimball (1995) and Burstein et al. (2005).

$$\lambda_{t,j} = \lambda(\frac{P_{t,j}}{E_{t,j}C^{*}_{t,j}}, Y_t; S_t) . \tag{5.4}$$

Competitive pressure in importing country is measured by the gap between competitor's prices in the importing markets (P) and the production cost of exporting firm (EC^*). The demand pressures in importing country are measured by aggregate output (Y). We further allow a component that is dependent on the importing country monetary policy environment (S). We argue that the exchange rate pass-through is dependent on the importing country inflation regime. Thus, we set monetary policy variable as $S = \Pi$, where Π denotes the importing country annual inflation rate[63]. We view this monetary policy dependence as a firm strategic choice of how importing country monetary policy condition affects to exchange rate movements' pass through into export price. Now, the mark-up can be presented in a multiplicative form as

$$\lambda_{t,j} = \left(\frac{P_{t,j}}{E_{t,j}C^{*}_{t,j}} \right)^{\alpha} Y_t^{\beta} E^{\theta(\pi)} . \tag{5.5}$$

The value of α is expected to be between 0 and 1, and the value of β is expected to be positive. The function $\theta(\pi)$ denotes the pass-through elasticity to inflation rate, and it is expected to be between 0 and 1. The function $\theta(\pi)$ can be seen as a mark-up multiplier in the following way. As firms set prices for several periods in advance, them mark-ups responds more to exchange rate changes if the inflation regime is high. Thus, a high inflation rate regime would tend to increase the exchange rate pass through and the pass-through coefficient is dependent on inflation regime as suggested by Taylor (2000).

Substituting equation (5.5) into equation (5.3) we get

$$P^{im}_{t,j} = \frac{P^{\alpha}_{t,j}}{(E_{t,j}C^{*}_{t,j})^{\alpha}} Y^{\beta}_{t,j} E^{\theta(\pi)}_{t,j} E_{t,j} C^{*}_{t,j} . \tag{5.6}$$

Taking the logarithmic form of the equation (5.6) and dropping the good index j, we get

$$p^{im}_{t} = (1-\alpha)e_t + \theta(\pi)e_t + (1-\alpha)c^{*}_t + \alpha p_t + \beta y_t , \tag{5.7}$$

where the lower case letters denote the logarithmics of the variables.

[63] It is important to note that there are several other reasons that might have caused changes in mark-up coefficients. Mann (1986) lists some other factors than inflation variability that might have caused changes in firms mark-ups. First, exchange rate volatility is probably a crucial factor that makes importers more wary of changing prices and more willing to adjust profit margins. Second, if import price shocks are persistent, then the firms are more likely to change prices rather than adjust profit margins in response to exchange rate changes. Third, aggregate demand uncertainty will change importers profit margin in imperfectly competitive environment. Moreover, the degree of the openness of the economy and the credibility of monetary policy are important factors that might change mark-up coefficients.

In the above equation (5.7), exchange rate pass-through depends both on the direct effect to import price and the indirect effect via monetary policy environment to import price[64]. The direct pass-through coefficient *(1-α)* is expected to be between 0 and 1. If the foreign firm is a price taker in the domestic competitive markets, then $\alpha = 1$ and the direct pass-through effect is zero. In this case we can see that exchange rate changes have no direct effect in local currency prices and the firm is using LCP. If the firm passes exchange rate changes fully through to import prices, then $\alpha = 0$, and the firm is using PCP.

The equation (5.7) shows that the exchange rate changes have also an indirect effect, which is dependent on the inflationary environment of importing country. We assume next that there are only two extreme inflation regimes in the importing country, low inflation and high inflation. We will further assume that there is some threshold inflation rate (π^*), which divides these extreme inflation regimes. The low inflation regime ($\pi_t < \pi^*$) is defined as such a competitive environment that the firm cannot use pricing to market strategy. In this case the indirect exchange rate pass-through is zero and the firm will transmit all the exchange rate movements directly into export price. The high inflation regime ($\pi_t > \pi^*$) is such an environment that the firm can exercise full pricing to market strategy and the exchange rate pass-through is above zero. With these assumptions we can write the indirect exchange rate effect as

$$\theta(\pi) = \begin{cases} 0, if\ \pi \le \pi^* \\ \delta > 0, if\ \pi > \pi^*. \end{cases} \tag{5.8}$$

For these two extreme inflation regimes we find two different pass-through coefficients. If the importing country has a low inflation regime then the exchange rate pass-through to import price is $(1-\alpha)e$. If the importing country has a high inflation regime then the exchange rate pass-through to import price is $(1-\alpha+\delta)e$. We can see that the exchange rate pass-through is lower in low inflation regime than in a high inflation regime, since $(1-\alpha)<(1-\alpha+\delta)$. Intuitively, the low inflation regime is such a monetary policy environment that the firm faces competition in importing markets and it cannot pass-through all exchange rate changes into import prices. In a high inflation regime, the firm can pass-through all the exchange rate changes to import prices. Our model, thus, implies that a higher inflation would raise the pass-through coefficients in a nonlinear way. The model (5.7) can be written also in difference form as

$$\Delta p_t^{im} = (1-\alpha+\theta(\pi))\Delta e_t + (1-\alpha)\Delta c_t^* + \alpha \Delta p_t + \beta \Delta y_t \ , \tag{5.9}$$

where the monetary policy effect $\theta(\pi)$ is designed to account for any influence of inflation rates in exchange rate pass-through.

The above threshold effect might be plausible for a single firm, but for the case of aggregate variables we should smooth this nonlinear function in some way. A potential source of smoothness may be due to interaction of heterogeneous agents at the microeconomic level. First, there is probably a great diversity across firms when forming opinions

[64] Most of the previous empirical studies have examined only the direct effect of exchange rate changes.

of high and low inflation regimes. Second, when the annual inflation is only slightly below or above the predetermined threshold level, most of the firms take this as an evidence of only modest changes in importing country monetary policy and do not change their pricing strategies. However, when the annual inflation starts to grow above the threshold level, most of the firms are uncertain whether this is a signal of the changes in monetary policy. A growing number of the firms take this as an evidence of changes in importing country monetary policy and possible change their export prices in local currency terms. When inflation is high enough relative to threshold level, almost all the firms notice that importing country monetary policy has changed. The firms are supposed to take this environment change into account when deciding the pricing policies in importing country.

5.3 Empirical specification

We will examine the exchange rate pass-through in two empirical cases. In both cases we assume that the pass-through coefficients are nonlinearly related to inflation rates. By using nonlinear estimation we are able to examine in detail whether the exchange rate pass-through elasticises are dependent on importing country monetary policy regime. The first specification examines how the exchange rate changes translate to domestic countries inflation rates. The second specification examines how the exchange rate changes will translate to importing countries import prices. There is an important distinction between the pass-through to consumer prices and pass-through to import prices, since consumer prices usually also include the effect of a local distribution channel to the final consumers. Thus, we might expect much more higher pass-through coefficient to import prices in response to exchange rate shock.

5.3.1 Effects on consumer prices

We first analyze how the consumer prices respond to exchange rates at the aggregate level. This is probably the most relevant issue for monetary policy, since a large part of monetary policy literature is based on the fact that exchange rate movements will instantaneously be transmitted to domestic country consumer price inflation (see e.g. Svensson 2000 and Devereux 2001). Moreover, the examination of pass-through to consumer prices gives an appropriate test for the Taylor view of pass-through. Our model specification is a slight modification of model (5.1) and is closely related to the law of one price (LOP). Thus, at the aggregate level our consumer price model examines in some sense the familiar purchasing power parity (PPP) relation

$$P_t = E_t P_t^*, \tag{5.10}$$

where P_t and P_t^* are domestic and foreign currency price indices and E_t is nominal exchange rate between countries. If the LOP holds for all traded goods and preferences are identical across countries, then PPP relation holds in absolute terms. However, in reality, the different transportation costs, trade barriers and compositions of price indices lead to deviations from PPP. Thus, we concentrate on the relative PPP relation, which can be presented in a logarithmic form as

$$\Delta p_t = \Delta e_t + \Delta p_t^* . \tag{5.11}$$

The literature of PPP has shown very slow convergence rate[65] of parity reversion (see e.g. Rogoff 1996). Engel (1999) shows that even relative prices of tradables exhibit very slow convergence to the parity relation. One potential reason for the real exchange rate persistence might be the discriminatory pricing policy of importing firms. Froot and Rogoff (1995) consider the idea that partial pass-through of exchange rate may affect PPP. Feenstra and Kendall (1997) find that a substantial portion of observed PPP deviation is attributable to incomplete exchange rate pass through because of the importing firms pricing to market behaviour. Furthermore, Bergin and Feenstra (2001) finds that PTM and staggered price contracts cause persistently in real exchange rate dynamics in a general equilibrium setting. Another reason might be the nonlinear adjustment to parity level. For example, factors like transportation costs and trade barriers might drive a wedge between prices in separated markets. Furthermore, the interaction of heterogeneous agents and effects of interventions in the foreign exchange rate market might cause nonlinearity at microeconomic level (see Taylor 2003 among others)[66].

We use similar type of model as in Campa and Goldberg (2003), Devereux and Yetman (2003) and Choudri and Hakura (2001). However, we differ from these studies in an important way, since we assume that the exchange rate pass-through is a nonlinear function of importing country inflation rate series. Hence, we consider either nonlinear exponential or logistic smooth transition regression model. These functional forms are two possible specifications that take into account the pricing behaviour of exporting firms at the aggregate level as presented in previous section. The nonlinear function forms are bounded between zero and unity and assumed to be continuous. We argue that in model (5.7) the pass-through elasticity responds almost linearly to exchange rate movements near the threshold level, but that there is a smooth transition to full inflation dependent pass-through elasticities when the importing country inflation rate is increasing. The model is expressed in first-differences, with the addition of lagged effect of variables to allow gradual adjustment of consumer prices. The estimation equation can be expressed as

$$\begin{aligned} \Delta p_{t,j} = \beta_{0,j} + \sum_{i=1}^{4} \beta_{1,j}(i)\Delta e_{t-i,j} + \left[\beta_{0,j}^* + \sum_{i=1}^{4} \beta_{1,j}^*(i)\Delta e_{t-i,j} \right] \cdot G(s_t;\gamma,c) \\ + \sum_{i=1}^{4} \beta_{2,j}(i)\Delta p_{t-i,j} + \sum_{i=1}^{4} \beta_{3,j}(i)\Delta p_{t-i}^* + \varepsilon_{t,j} \end{aligned} , \tag{5.12}$$

[65] The consensus view is that the estimates of half-lives are from 3 to 5 years.

[66] The recent PPP literature has shown that the root of slow convergence might be in the different speeds of convergence for nominal exchange rates and prices (Engel and Morley 2001). According to this view the slow PPP reversion is probably due to slow nominal exchange rate adjustment rather than the slow price adjustment, hence, challenging the traditional view of price stickiness. (Cheung et al. 2004).

where Δp_t , Δe_t, and Δp_t^* are one-quarter domestic consumer price difference, exchange rate difference and foreign consumer price difference, respectively. The parameter i denotes the lag structure of the system and we have chosen to include up to four lags for all cases. The parameter j denotes importing country. We have chosen to use country j's bilateral US dollar exchange rates, since the US dollar is the most commonly used currency in international trade. We have also used the US inflation rates as a proxy of foreign country inflationary effects[67]. The parameter s_t is a stochastic transition variable, defined as $s_t = (\pi_t - \pi^*)$. The transition function $G(s_t;\gamma,c)$ is assumed to be either the logistic function

$$G(s_t;\gamma,c) = (1+\exp\{-\gamma(s_t - c)/\sigma_{s_t}\})^{-1} \tag{5.13}$$

or the exponential function

$$G(s_t;\gamma,c) = (1-\exp\{-\gamma(s_t - c)^2/\sigma_{s_t}\}), \tag{5.14}$$

where $\sigma_{s_t} = [\text{var}(s_t)]^{1/2}$ makes the parameter γ scale free.

The short-run exchange rate pass-through (1 quarter) to domestic prices of country j is defined by the coefficients

$$\textit{short-run pass-through} = \beta_{1,j}(1) + \beta_{1,j}^*(1)\cdot G(s_t;\gamma,c)\,. \tag{5.15}$$

The long-run pass-through is defined as a time-varying impulse response form. A problem is that the presentation (5.12) is self-exciting type of model, since the inflation rate is also the threshold variable. Thus, we should also take into account the endogeneity of transition variable when we are calculating pass-through ratios for longer than one period. We have calculated both medium-run exchange rate pass-through (1 year) and long-run exchange rate pass-through (5 years) by time varying impulse response form. The derivations of long-run pass-through ratios are presented in appendix 1. It should be noted that in calculating long-term pass-through ratios, we have to calculate the sum of all impulse response effects for the five years period, and then average all these effects to get long-run pass-through ratios.

It is important to note that the equation (5.12) should not be seen as a full specification of importing country's (j) inflation process. However, it reflects the aggregate effects of exchange rate pass-through to the consumer level (Devereux and Yetman 2003). We should also note, that we are not interested in the PPP hypothesis particularly, but the possibility of pass-through effects in consumer price changes. The parameter $\beta_{1,j}$ measures the direct exchange rate pass-through effect into importing country j consumer prices. As we can see from the above equation (5.12) the exchange rate effect depends also on the inflation rate regime (π_t) of the importing country.

[67] Devereux and Yetman (2003) have used similar data structure in their paper. The above model can also be seen as a relative PPP model against the US dollar, where the domestic inflation rate is a dependent variable.

There are now three extreme outcomes for exchange rate pass-through. First, the model is linear if parameter γ is not significant and the pass-through coefficients are determined solely by the linear part of the model. In this case the pass-through coefficient is $\beta_{1,j}(i)$. Thus, our model specification nests also the linear exchange rate pass-through alternative. Second, the inflation process might be near the threshold inflation rate (π^*) and the nonlinear part of the model collapses even if the parameter γ is significant. In this case the pass-through coefficient is also $\beta_{1,j}(i)$, while the model specification is nonlinear. Third, if the true inflation process differs significantly from the threshold level and the parameter γ is also significant then the nonlinear part affects to the exchange rate pass-through coefficients. In this case, the pass-through coefficient is a combination of direct and indirect effects and can be presented as $\beta_{1,j}(i) + \beta_{1,j}^*(i)$.

5.3.2 Effects on import prices

The empirical specification includes only the consumer prices and, hence, does not address the question of how the exchange rate changes are passed through into import prices. In this section we formulate the following empirical relation based on equation (5.7)[68]:

$$\begin{aligned}\Delta p_{t,j}^{IM} &= \alpha_{0,j} + \sum_{i=0}^{4} \alpha_{1,j}(i)\Delta ee_{t-i,j} + \left[\alpha_{0,j}^{*} + \sum_{i=0}^{4} \alpha_{1,j}^{*}(i)\Delta ee_{t-i}\right] \cdot G(s_t;\gamma,c) \\ &+ \sum_{i=1}^{4} \alpha_2(i)\Delta p_{t-i,j}^{IM} + \sum_{i=0}^{4} \alpha_{3,j}(i)\Delta mc_{t-1,j} + \sum_{i=0}^{4} \alpha_{4,j}(i)\Delta y_{t,j} + \eta_t^j\end{aligned}, \qquad (5.16)$$

where $\Delta p_{t,j}^{IM}$ and $\Delta ee_{t-i,j}$ are the one-quarter differences in the aggregate import prices and effective exchange rate for country j. The variables $mc_{t-1,j}$ and $y_{t,j}$ denote the exporter marginal costs and the importing country market demand conditions, respectively. Otherwise, the model is similar as the model for consumer prices in the previous subsection.

5.4 Data and sample countries inflation experiences

We use time series data for seven OECD countries: the United States, Japan, Germany, France, Italy, the United Kingdom and Canada. The source of data is the OECD Main Economic Indicators database. All the data are in logarithmic form. The data are quarterly and the sample period is 1974:1 – 2001:2[69]. The price series employed are consumer prices, aggregate import prices, nominal bilateral US dollar exchange rates, and nominal effective exchange rates. The consumer price index (CPI) provides the broadest measure of inflation in importing countries. Furthermore, we use aggregate import prices at the

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[68] For linear version see Hooper and Mann (1989), Yang (1997) and Campa and Goldberg (2002).

[69] The fourth quarter of 1998 is the end of the sample period for the eurozone countries, namely Germany, France and Italy.

country level. All the price data are in the form of price indices and in the local currency prices. Nominal exchange rate series are quarterly averages and defined as foreign currencies per unit of domestic currency. The real gross domestic products (GDP) are used as the proxy of demand pressure (y) in importing countries. The cost pressure measure is real unit labour costs, $mc^j = ulc^j/P^j$, where mc denotes marginal costs and P is the consumer price index of exporting country, and ulc is foreign exporters' unit labour costs for each country j.

Table 19 gives a summary of the behaviour of quarterly CPI inflations for our sample countries. The inflation rates are expressed in annual rates. We can see that the mean inflation varies between countries from a lowest 2.8 % for Germany to a highest 8.5 % for Italy. We can also see that for some countries there is much more volatility in inflation rates relative to others. Also, for some countries there are peak inflations, especially for the UK and Italy that are very high relative to the mean annual inflation rates. We note also that the inflation processes are very persistent, since the first order autocorrelations are high and for several countries we are not able to reject unit root in inflation rate at conventional 5 per cent significance level. It is noteworthy that several inflation series are *I(1)* processes, since many previous studies have considered inflation rates as a stationary processes. The finding that inflation series are nonstationary is, however, not an exception, since several papers (e.g. Juselius 1999 and Juselius and MacDonald 2004) suggest that several price indices are probably I(2) series[70].

Table 19. Sample countries inflation experience 1974:1 – 2001:3

Country	Mean	Max	Min	SD	ρ	ADF
Canada	5.0	13.3	-2.0	3.7	0.80	*-3.61 (0)*
Germany	2.8	10.3	-5.1	2.6	0.36	-2.55 (3)
UK	7.0	34.9	-3.0	6.7	0.58	*-3.66 (4)*
Japan	2.8	17.1	-3.8	4.3	0.43	*-2.98 (4)*
France	5.3	15.7	-1.1	4.4	0.90	-2.21 (1)
Italy	8.5	25.7	0.8	6.3	0.85	-0.99 (5)
USA	4.8	14.9	-1.1	3.3	0.78	-2.42 (2)

Notes: Max and Min denote maximum and minimum annual inflation rates, respectively. SD denotes standard deviation of inflation rate and ρ is the first order autocorrelation coefficient. The ADF denote augmented Dickey-Fuller unit root test and figure in parentheses is the lag length. The italics faces denote significance at the 5 per cent level.

[70] The problem might be, however, due to low power of unit root tests, structural changes, and/or few unusually high observations.

5.5 Empirical results

5.5.1 Estimates of the US dollar exchange rate pass-through into consumer prices

Table 20 reports the nonlinear estimation results for the US dollar nominal exchange rate pass-through into consumer prices when applying the conventional general-to-specific approach[71]. The estimation is based on equation (5.12). The parameter estimates are obtained by using nonlinear least squares, which provide consistent and asymptotically normal estimates (see Gallant 1987). We follow Teräsvirta (1994) and standardize the exponent by dividing it by the variance of the transition variable (π_t). Thus, the smoothness parameter (γ) is a scale free parameter. We have used linear model parameter estimates as the starting values for nonlinear estimation. We find that for all cases, the estimate of γ is significantly different from zero, which implies that there is a nonlinear relationship between exchange rate pass-through and annualized inflation rate from recent period. We find that the ESTAR model fits better for Canada, Germany and France, but the LSTAR model gives a better fit for the UK, Japan and Italy. We find also that the threshold inflation rate levels differ across countries. For countries Japan and France we find small threshold levels and for Italy we find relatively large threshold level of inflation. This might reflect countries' different inflationary policies during the sample period. The average threshold level of inflation is approximately 3 per cent.

The residual diagnostic statistics are also reported in Table 20. The test statistics show that error terms are free from autocorrelation and conditional heteroscedasticity. The residual variance ratio (V_{NL}/V_L) between the estimated nonlinear model and the alternative linear specification suggest for all countries a reduction in variances for the nonlinear case. Furthermore, both the Akaike (AIC) and Schwartz-Bayesian information criteria (SBC) are suggesting better model fit for the nonlinear regression models.

Table 20. Estimated parsimonious nonlinear regression model

$$\Delta p_{t,j} = \beta_{0,j} + \sum_{i=1}^{4} \beta_{1,j}(i)\Delta e_{t-i,j} + \left[\beta_{0,j}^{*} + \sum_{i=1}^{4} \beta_{1,j}^{*}(i)\Delta e_{t-i,j} \right] \cdot G(s_t;\gamma,c)$$
$$+ \sum_{i=1}^{4} \beta_{2,j}(i)\Delta p_{t-i,j} + \sum_{i=1}^{4} \beta_{3,j}(i)\Delta p_{t-i}^{*} + \varepsilon_{t,j}$$

Parameter	Canada	Germany	UK	Japan	France	Italy
Model	ESTAR	ESTAR	LSTAR	LSTAR	ESTAR	LSTAR
$\beta_1(1)$	0.03	0.03	-0.01	0.02	0.01	
$\beta_1(2)$	-0.09	0.01	-0.03	0.02	-0.02	
$\beta_1(3)$	-0.03		0.01	0.01		
$\beta_1(4)$	0.03		-0.01	0.05		

[71] We have only reported those parameter values that are significant at 5 per cent level.

Table 20 (Continued)

$\beta_1^*(1)$	0.03	-0.01	-0.01	-0.01		0.03
$\beta_1^*(2)$	0.20	-0.03	0.07	-0.01	0.03	-0.01
$\beta_1^*(3)$	0.07	0.06	-0.01		0.03	0.09
$\beta_1^*(4)$	-0.01	0.01	0.05	0.05	0.03	-0.03
$\beta_2(1)$	0.25	0.07	0.32	0.49	0.27	0.31
$\beta_2(2)$			0.14	-0.25		0.12
$\beta_2(3)$	0.39	0.30		0.24	0.23	0.24
$\beta_2(4)$		-0.13	0.11	0.27		-0.19
$\beta_3(1)$	0.28	0.18			-0.27	0.44
$\beta_3(2)$		-0.16	0.14	0.19	0.08	
$\beta_3(3)$			-0.18		-0.13	
$\beta_3(4)$	0.20	0.36	0.60	0.52	0.26	0.14
$\pi^*(\%)$	3.2	3.4	3.3	-0.7	0.8	6.2
γ	1.32	0.58	2.39	0.48	2.79	3.47
Statistics						
R^2	0.93	0.78	0.82	0.80	0.95	0.94
DW	2.03	1.99	1.42	1.75	2.02	1.95
AIC	-652.6	-626.9	-420.0	-621.3	-671.5	-566.3
AIC(linear)	-634.2	-619.4	-421.5	-601.9	-662.5	-554.9
SBIC	-617.6	-584.4	-377.7	-575.4	-631.6	-529.0
SBIC(linear)	-599,6	-584.8	-391.0	-567.3	-627.9	-520.3
V_{NL}/V_L	0.94	0.94	0.95	0.96	0.94	0.94
Q(1)	0.10	0.03	*6.05*	1.68	0.09	0.02
(p-value)	(0.76)	(0.85)	(0.01)	(0.19)	(0.75)	(0.87)
Q(4)	1.33	0.42	6.60	3.86	3.92	2.66
(p-value)	(0.85)	(0.98)	(0.16)	(0.42)	(0.42)	(0.62)
ARCH(1)	0.41	0.16	*6.26*	0.02	0.19	*5.27*
(p-value)	(0.85)	(0.69)	(0.01)	(0.89)	(0.66)	(0.02)
ARCH(4)	5.40	7.67	8.20	5.92	2.71	*13.85*
(p-value)	(0.25)	(0.10)	(0.08)	(0.21)	(0.61)	(0.00)

Notes: Estimation is by nonlinear least squares. R^2 denotes the coefficient determination, DW is Durbin-Watson statistics for error term correlation. *AIC* and *SBC* are Akaike information criteria and Schwartz-Bayesian information criteria, respectively. V_{NL}/V_L denotes variance ratio and is constructed as the ratio of the residual variance from the estimated nonlinear model with the residual variance from linear model. *Q(k)* denotes Ljung-Box residual autocorrelation statistics and *ARCH(k)* denotes residual heteroscedasticity. The figures in parentheses are the marginal significance level for test statistics. The italics faces denote significance at the 5 per cent level.

In Figure 2 we have plotted the estimated transition functions. The transition functions are presented as a function of the transition variable π_{t-1}. Overall, the transitions between extreme regimes seem to be rather smooth. We can see that for some countries, like the UK and Japan, the threshold functions are much more sharper than for other countries. However, the transition functions attain values from zero to unity for all countries over the sample periods. This is in line with Taylor's (2000) suggestion that the pass-through coefficients depend strongly on importing country's inflationary environment. However, for Germany we can see that the threshold function has only a few observations in the above regime. This is probably due to Germany inflation policy during the sample period. Furthermore, for inflation rates that are near the threshold inflation rates, the pass-through coefficients do not move with inflation rates. Broadly speaking, the nonlinear pass-through mechanism seems only to operate, when the previous annual inflation rate is sufficiently high relative to the threshold inflation rate. There is also some evidence that very low rates of inflation relative to threshold level gives to the model similar local dynamics as high inflation rates.

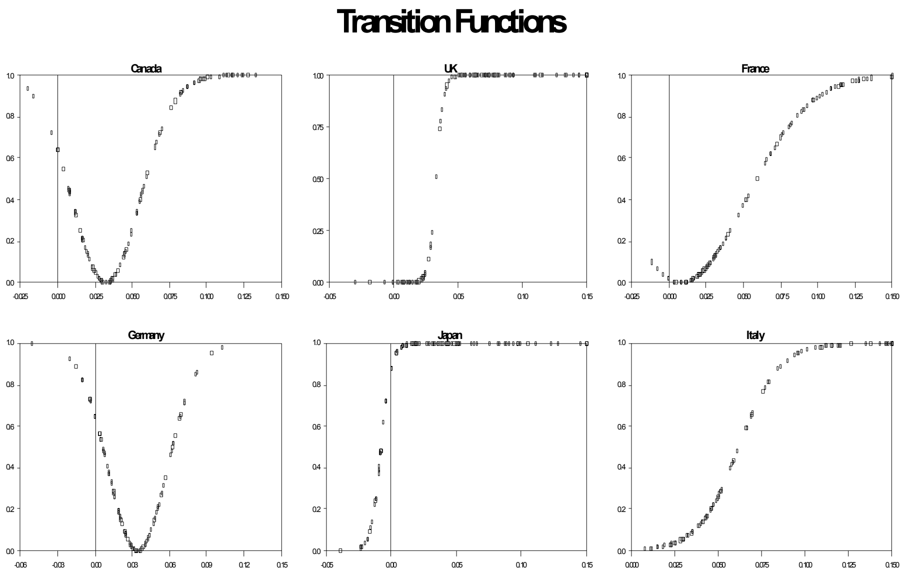

Fig. 21. The transition functions vs. inflation rates

Table 21 presents short-run (one-quarter), medium-run (one-year) and long-run (five-years) estimates of the US dollar exchange rate pass-through into consumer prices for our sample countries. For comparison, we also present the estimates in Choudhri and Hakura (2001) and Campa and Goldberg (2003). The average short-run, medium-run and long run elasticities are 0.02, 0.10 and 0.19, respectively. This means that on average a 1 percent depreciation of the domestic currency against the US dollar leads to about a 2, 10 and 19 percent rise in consumer prices in the short-run, medium-run and. For all countries we find that both short-, medium- and long-run exchange rate pass-through into consumer prices are clearly different from one. This implies that local currency pricing (LCP) characterize better exchange rate pass-through into consumer prices than producer currency pricing (PCP). We can see that short-term pass-throughs are very stable across countries. However, there are some differences in the estimates in the long run. The small open economies, such as Canada and Italy, have relatively higher pass-through estimates than other sample countries in the medium-run. Our pass-through estimates do not vary much across for large countries in the period of one to five years, which suggests that the way that foreign exporters set their prices is quite similar for the major OECD countries. However, for smaller countries we can again see that pass-through coefficients are much larger. Overall, our pass-through coefficients are very much in line with the previous studied at least in the short-run.

Table 21. Exchange rate pass-through into consumer prices

Country	T=1	T=4	T=20	CH Short-run	CH Long-run	CG Short-run	CG Long-run
Canada	0.06	0.16	0.41	0.00	0.11	-0.06	-0.06
Germany	0.02	0.08	0.13	0.05	0.13	-0.03	0.02
UK	-0.01	0.06	0.14	-0.01	0.02	-0.04	-0.12
Japan	0.01	0.06	0.12	n.a.	n.a.	0.00	0.09
France	0.01	0.04	0.08	-0.01	0.11	0.19	0.51
Italy	0.02	0.17	0.28	0.04	0.11	-0.06	0.09

Note: CH denotes Choudhri and Hakura (2001) estimates and CG denotes Campa and Goldberg (2003) estimates.

5.5.2 Estimates of the effective exchange rate pass-through into aggregate import prices

We have estimated the specification (5.16) for import prices in the UK, Japan and the USA[72]. The parameter estimates are presented in Appendix 2 (Table 23). Again, we have only presented those parameter estimates that are significant at five percent level. We can see that cost and output variables enter significantly in all regressions, suggesting the importance of these variables in import prices behaviour. We can also see that threshold levels differ significantly across countries. Especially, for the UK we find very large threshold level of the inflation. This is probably due to high inflation regime in the UK in 1970s. In Appendix 3 we have presented the plots of model fits, transition functions, and the relations between importing countries inflation rate and exchange rate pass-through. For Japan we find that the ESTAR model gives a better fit and for the UK and the USA we find that LSTAR model gives a better fit[73]. This reflects that the UK and the USA have very different exchange rate pass-through behaviour than Japan. This is not a new result, since similar results have also been find, for example, in the Campa and Goldberg (2002).

In Table 22 we present short-run and long run effective exchange rate pass-through estimates into importing country aggregate import prices. We have also presented Campa and Goldberg (2002) estimation results for comparison. The United States and the UK

[72] Since we have an effective exchange rate as our control variable instead of using single countries nominal exchange rate, we should take into account effective exchange rate movements' simultaneous effect to price changes. Thus, our lag structure goes from 0 to 4 in this part of the study. However, since we do not have additional control variables as our regressors, we will present only the long-run pass-through estimates.

[73] Actually, the threshold functions are very sharp so that they are very near of threshold models instead of the smooth transition models.

have the lowest pass-through, 32 and 35 per cent in the short-run. For Japan the pass-through elasticises to aggregate import prices are relatively higher. However, for all countries we can reject both local currency pricing and producer currency pricing in the short-run. Overall, we can see that import price pass-through estimates are much larger than consumer price estimates. We can reject the null hypotheses of LCP and PCP only for the USA in the long run. For Japan and the UK we could not reject the hypothesis of PCP in the long run. However, relative high pass-through estimate for Japan import prices is documented also in several other papers.

Table 22. Effective exchange rate pass-through into import prices

Country	short-run	long-run	CG short-run	CG long-run
UK	$0.35^{*,+}$	0.93^{*}	$0.39^{*,+}$	$0.47^{*,+}$
Japan	$0.59^{*,+}$	1.10^{*}	0.88^{*}	1.26^{*}
USA	$0.32^{*,+}$	$0.78^{*,+}$	$0.26^{*,+}$	$0.41^{*,+}$

Notes: *, + denotes significantly different from zero and one at 5 percent level, respectively. CG denotes Campa and Goldberg (2002) estimates.

5.6 Discussion

Recent development in open economy macromodels has assumed price rigidities and market imperfections in analysing monetary and/or exchange rate policy. The empirical results show that the rate of change in consumer prices does not follow very closely the changes in exchange rates. In the consumer price level, we can say that price setting is closer to local currency pricing than producer currency pricing. However, while consumer prices seem to be rather insensitive to exchange rate changes, there seems to be larger pass-through effect to import prices. Thus, our results suggest that there is a significant difference between price responds to exchange rate changes in consumer price and import price level. There are many reasons, however, why exchange rate changes will not pass through fully to consumer prices. First, the pass-through into domestic country consumer prices is affected by the way in which import and domestic prices interact. Second, the domestic policy, for example inflation targeting or wage settlement, may affect the exchange rate pass-through to consumer price level, while the import prices might change. An alternative explanation might be the domestic distribution sector. E.g. Corsetti and Dedola (2002) have explored models in which domestic distributor incorporate a substantial nontradable component so that the link between exchange rates and final consumer prices weakens further relative to exchange rate movements' effect to import prices.

The imperfect exchange rate pass-through changes the implication for optimal monetary policy. Devereux and Engel (2003) show that if prices are set in local currencies, then the optimal monetary policy is the fixing of nominal exchange rates. The reason is that a flexible exchange rate cannot achieve optimal relative price changes, since both foreign and domestic product prices are set in same currency. Thus, under the pure LCP the exchange rate has no role in adjusting goods prices. This might explain partly why the exchange rate changes are so volatile relative to price changes. In addition, the LCP might explain why there are short-run deviations from the law of one price also in the highly tradable goods sector. Furthermore, when prices are set in consumer's currency, an unexpected domestic currency depreciation has no traditional expenditure switching effect. This implies that the firms must allow their foreign mark-ups to adjust to keep their export prices fixed.

For many countries the objective of monetary policy is the price stabilisation. Thus, it is crucial to know the extent to which consumer prices are affected by the exchange rate fluctuations. In traditional models, a low exchange rate pass-through provides greater freedom for more independent monetary policy and, thus, it is easier to implement the low inflation policy. For example, if there is a monetary shock that increases inflation, government should take into account the pass-through effect when deciding to loosen its monetary policy. Furthermore, under inflation targeting policy, it is crucial to take the pass through endogeneity into account when deciding the target level of inflation rate. In Figure 22 we show the relationship between annual consumer price inflation and exchange rate pass-through in all the estimated countries. We have calculated the pass-through coefficients for every single period by using the formulation (5.12) and the plot them against the same period's annual inflation rates. The figure shows that exchange rate pass-through is positively related to inflation rate for many countries. Especially, we can see that very high inflation rates are associated with high pass-through elasticities for many countries.

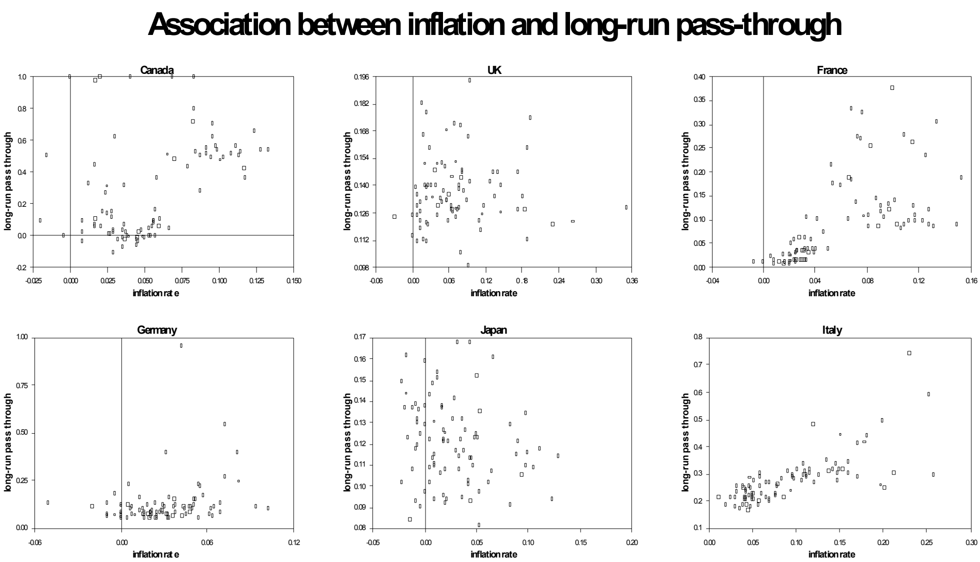

Fig. 22. The association between inflation and long-run pass-through

One of the major concerns in the euro area is whether a single monetary policy has different effects across different member countries. A major concern has been how shocks arising from outside of the euro area may affect the performance of euro area's consumer price inflation. The question is important, since a crucial part of ECB monetary policy is low consumer price inflation across the member countries. The serious inflationary pressures in some member countries might cause serious problems for conducting monetary policy in the whole area. Some authors have suggested that large euro depreciation may lead to serious problems to achieve low eurowide inflation, since depreciation may lead to differences in the structure of trade by the member countries. However, our results suggest that the US dollar exchange rate pass-through to consumer prices is quite similar for these countries[74]. The equality of pass-through elasticity suggest that a choice of single currency does not necessarily lead to large differences in member countries consumer price inflations even if there are large movements in the euro exchange rate. This finding is important for e.g. ECB monetary policy strategy, since ECB has not any specific target for the exchange rates. The exchange rate changes become relevant for monetary policy only if they influence inflation rates over all euro area.

5.7 Conclusions

In this chapter we have examined the exchange rate pass-through to consumer prices and aggregate import prices for several OECD countries. Our findings show that the exchange rate pass-through is dependent on the importing country inflation regime in a nonlinear way. The degree of exchange rate pass-through is determined endogenously by the importing country monetary policy. In other words, the degree of pass-through is affected by the inflationary environment that a firms face in importing country. In a low inflation regime we have low exchange rate pass-through elasticity and in a high inflation regime the exchange rate pass-through is more rapid. This result is quite similar for both larger and smaller industrialized countries. Contrary to some previous studies, we find also that there seem not to be so large differences across countries pass-through coefficients. However, import prices paid at importing country border display very different behaviour as the prices paid by final consumers.

The decline of exchange rate pass-through is in close relation with the low inflation regimes that have prevailed in industrialized countries after 1990s. In this way, our findings support very closely Taylor's (2000) conclusion that low inflation regime has caused lower exchange rate pass-through. Our results provide a strong support for the importance of some price stickiness. In particular, the firms take the inflationary environment into account when they adjust their prices. Overall, the importance of exchange rate pass-through and price stickiness should be taken into account when designing the optimal monetary policy for open economy. In this way, our results are also in line with Devereux and Engel (2003).

[74] The only exception is that in the medium- and long-run small open economy, Italy, has much higher pass-through coefficient than other EMU countries. This might suggest that the problem of price stability after exchange rate changes may concern more severely small open economies in the EMU.

A natural extension would be to examine the pass-through in a small open economy[75]. The question is very relevant, since exchange rate changes are likely to have a larger effect on small open economies' inflation than to larger economies. The reason is that small open economies have typically more concentrated markets and thus consumers of foreign imports can be seen as price takers. Furthermore, Calvo and Reinhart (2002) show evidence that the pass-through tends to be larger in small and high inflation countries compared to large and low inflation countries. It is thus important to examine whether the relation between inflation and pass-through is different across small and large countries, since it is possible that for small countries other factors than inflation regimes might be more important in determining the pass-through. Furthermore, one useful area for future research would be to examine cointegration relation between variables. Such a model might take into account the fact that there are long-run equilibrium relationships which drive the dynamics of exchange rate pass-through.

[75] Some studies (see e.g. Cunningham and Haldane 2000 and Garcia and Restrepo 2001) find that even relatively large changes in exchange rates in the 1990s had a very weak effect on inflation in some small open economies.

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Appendix 1 Time-varying impulse responses

For a STAR model

$$\Delta p_t = \beta_0 + \sum_{i=0}^{4} \beta_1(i)\Delta e_{t-i} + \left[\beta_0^* + \sum_{i=0}^{4} \beta_1(i)^* \Delta e_{t-i} \right] \cdot G(\gamma(\pi_{t-1} - \pi^*))$$
$$+ \sum_{i=1}^{4} \beta_2(i)\Delta p_{t-i} + \sum_{i=0}^{4} \beta_3(i)\Delta p_{t-i}^* + \varepsilon_t$$

we can calculate long-run pass-through ratios as follows. First, we can calculate time-varying impulse responses $(v_{j,t})$ for each time period j in a following way.

$$v_{0,t} = \frac{\partial \Delta p_t}{\partial \Delta e_t} = \beta_1(0) + \beta_1(0)^* \cdot G(\gamma(\Delta p_{t-1} + ...\Delta p_{t-4} - \pi^*)$$

$$v_{1,t} = \frac{\partial \Delta p_{t+1}}{\partial \Delta e_t} = \beta_2(1) \cdot v_{0,t} + \beta_1(1) + \beta_1(1)^* \cdot G(\gamma(\Delta p_t + ...\Delta p_{t-3} - \pi^*)$$
$$+ \left[\beta_0^* + \sum_{i=0}^{4} \beta_1(i)^* \Delta e_{t+1-i} \right] \cdot G'(\gamma(\Delta p_t + ... + \Delta p_{t-3} - \pi^*) \cdot \gamma \cdot v_{0,t}$$

$$v_{2,t} = \frac{\partial \Delta p_{t+2}}{\partial \Delta e_t} = \beta_2(1) \cdot v_{1,t} + \beta_2(2) \cdot v_{0,t} + \beta_1(2) + \beta_1(2)^* \cdot G(\gamma(\Delta p_{t+1} + ...\Delta p_{t-2} - \pi^*))$$
$$+ \left[\beta_0^* + \sum_{i=0}^{4} \beta_1(i)^* \Delta e_{t+2-i} \right] \cdot G'(\gamma(\Delta p_{t+1} + ... + \Delta p_{t-2} - \pi^*) \cdot \gamma \cdot (v_{1,t} + v_{0,t})$$

$$v_{3,t} = \frac{\partial \Delta p_{t+3}}{\partial \Delta e_t} = \beta_2(1)\cdot v_{2,t} + \beta_2(2)\cdot v_{1,t} + \beta_2(3)\cdot v_{0,t} + \beta_1(3) + \beta_1(3)^*$$

$$\cdot G(\gamma(\Delta p_{t+2} + ...\Delta p_{t-1} - \pi^*)) + \left[\beta_0^* + \sum_{i=0}^{4}\beta_1(i)^*\Delta e_{t+3-i}\right]$$

$$\cdot G'(\gamma(\Delta p_{t+2} + ... + \Delta p_{t-1} - \pi^*)\cdot\gamma\cdot(v_{2,t} + v_{1,t} + v_{0,t})$$

$$v_{4,t} = \frac{\partial \Delta p_{t+4}}{\partial \Delta e_t} = \beta_2(1)\cdot v_{3,t} + \beta_2(2)\cdot v_{2,t} + \beta_2(3)\cdot v_{1,t} + \beta_2(4)\cdot v_{0,t} + \beta_1(4) + \beta_1(4)^*\cdot$$

$$G(\gamma(\Delta p_{t+3} + ...\Delta p_t - \pi^*)) + \left[\beta_0^* + \sum_{i=0}^{4}\beta_1(i)^*\Delta e_{t+4-i}\right]\cdot$$

$$G'(\gamma(\Delta p_{t+3} + ... + \Delta p_t - \pi^*)\cdot\gamma\cdot(v_{3,t} + v_{2,t} + v_{1,t} + v_{0,t})$$

and

$$v_{j,t} = \frac{\partial \Delta p_{t+j}}{\partial \Delta e_t} = \sum_{i=1}^{4}\beta_2(i)\cdot v_{j-i,t} + \left[\beta_0^* + \sum_{i=0}^{4}\beta_1(i)^*\Delta e_{t+j-i}\right]\cdot G'(\gamma(\pi_{t+j-1} - \pi^*)\cdot\gamma\cdot(\sum_{i=1}^{4}v_{j-i,t})$$

for all $j \geq 5$.

Second, we calculate the sum of the impulse responses as $c_t = v_{0,t} + v_{1,t} + v_{2,t} + ... + v_{j,t}$ and finally we get the long-run pass-through ratios by taking the average of these impulse response sums.

Appendix 2 Table

Table 23. Estimated parsimonious nonlinear regression model for import prices

$$\Delta p_{t,j}^{IM} = \alpha_{0,j} + \sum_{i=0}^{4} \alpha_{1,j}(i)\Delta ee_{t-i,j} + \left[\alpha_{0,j}^{*} + \sum_{i=0}^{4} \alpha_{1,j}^{*}(i)\Delta ee_{t-i}\right] \cdot G(s_t;\gamma,c)$$
$$+ \sum_{i=1}^{4} \alpha_2(i)\Delta p_{t-i,j}^{IM} + \sum_{i=0}^{4} \alpha_{3,j}(i)\Delta mc_{t-1,j} + \sum_{i=0}^{4} \alpha_{4,j}(i)\Delta y_{t,j} + \eta_t^j$$

Parameter	UK	Japan	USA
Model	LSTAR	ESTAR	LSTAR
$\alpha_1(0)$	0.37	0.74	
$\alpha_1(1)$	0.16	-0.55	0.45
$\alpha_1(2)$		0.33	-0.24
$\alpha_1(3)$		-0.28	
$\alpha_1(4)$	0.05		0.28
$\alpha_1^*(0)$	-0.05	-0.46	0.37
$\alpha_1^*(1)$	0.21	1.09	-0.14
$\alpha_1^*(2)$	-0.32	-0.91	
$\alpha_1^*(3)$	-0.48	0.81	
$\alpha_1^*(4)$	0.28	-0.45	-0.14
$\alpha_2(1)$	0.21	0.49	0.44
$\alpha_2(2)$	0.11	0.17	
$\alpha_2(3)$		0.13	
$\alpha_2(4)$			
$\alpha_3(1)$		0.07	
$\alpha_3(2)$	0.06	0.16	-0.35
$\alpha_3(3)$	0.06		0.25

Table 23 (Continued)

$\alpha_3(4)$	-0.04		
$\alpha_4(1)$	-0.42	-0.02	-0.57
$\pi^*(\%)$	9.9	0.9	2.6
γ	3.31	0.85	4.47
Model statistics			
R^2	0.73	0.74	0.63
DW	1.92	2.02	1.92
AIC	-380.1	-210.7	-366.5
AIC(linear)	-306.5	-199.5	-338.2
SBIC	-331.7	-160.8	-335.0
SBIC(linear)	-265.6	-157.6	-297.4
V_{NL}/V_L	0.85	0.92	0.96
Q(1)	0.05	0.01	0.11
(p-value)	(0.82)	(0.89)	(0.73)
Q(4)	0.16	0.28	2.71
(p-value)	(0.99)	(0.99)	(0.61)
ARCH(1)	0.21	1.58	2.63
(p-value)	(0.65)	(0.21)	(0.10)
ARCH(4)	0.45	1.78	3.08
(p-value)	(0.97)	(0.78)	(0.54)

Notes: Estimation is by nonlinear least squares. R^2 denotes the coefficient determination, DW is Durbin-Watson statistics for error term correlation. *AIC* and *SBC* are Akaike information criteria and Schwartz-Bayesian information criteria, respectively. V_{NL}/V_L denotes variance ratio and is constructed as the ratio of the residual variance from the estimated nonlinear model with the residual variance from linear model. *Q(k)* denotes Ljung-Box residual autocorrelation statistics and *ARCH(k)* denotes residual heteroscedasticity. The figures in parentheses are the marginal significance level for test statistics.

Appendix 3 Figures

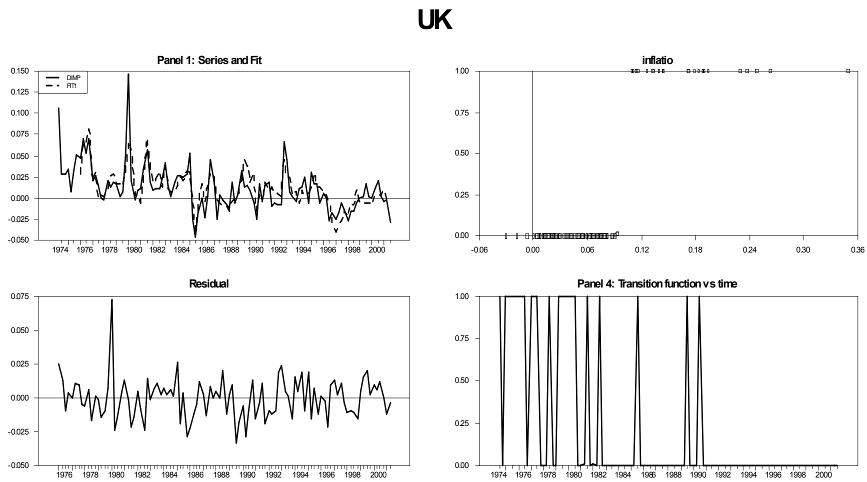

Fig. 23. The plots of the UK

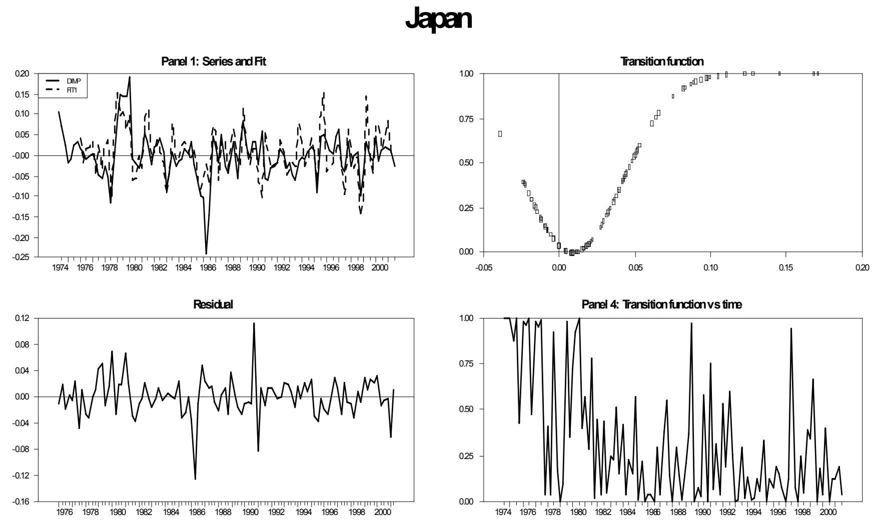

Fig. 24. The plots of Japan

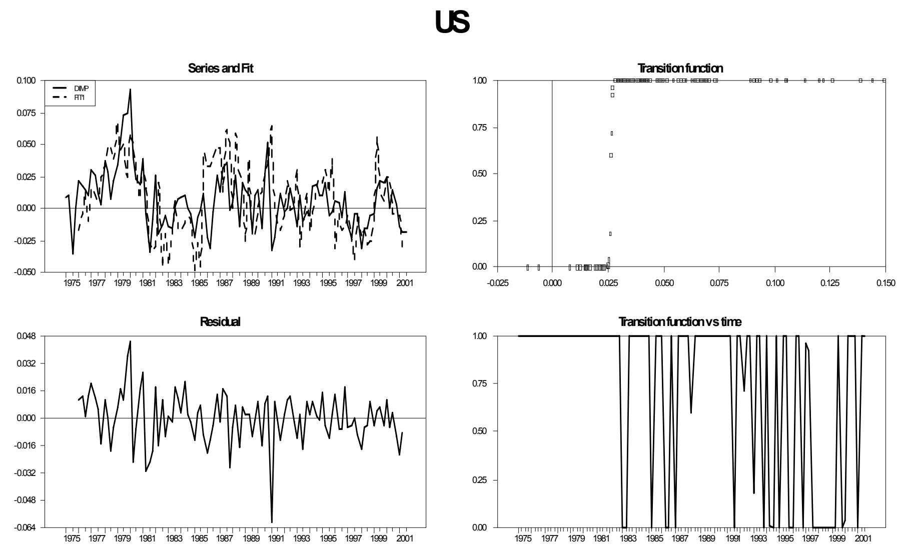

Fig. 25. The plots of the USA